Labor Mobility and Racial Discrimination*

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Abstract

We study the link between labor mobility constraints and racial wage discrimination. We show that when mobility constraints are low, the firms' monopsony power is reduced and discrimination disappears. We estimate this prediction 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. We find evidence that wage discrimination disappears abruptly when constraints on worker mobility are lowered. (JEL J15, J31, J6, J71)

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1 Introduction

Race differentials in labor market outcomes continue to persist. Evidence suggests that this is partly the consequence of racial discrimination.[1] In this paper, we explore the relationship between racial wage discrimination and job-to-job mobility constraints. In particular, we find that when mobility constraints are low, the monopsony power of prejudiced firms is reduced and discrimination disappears. We estimate this theoretical prediction 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 and find evidence that wage discrimination disappears abruptly when constraints on worker mobility are lowered.

Our results are of interest for several reasons. Since Becker (1957), two principal questions which the literature on the economics of discrimination has sought to address are 'how can discrimination survive in equilibrium?' and 'how can we get it to stop?' (Gneezy and List, 2013). The main mechanisms put forward to answer these questions are barriers to perfect competition in the product market[2] and frictions in the labor market. We contribute to the literature on labor market frictions by showing theoretically how monopsony power can allow firms to discriminate and empirically that discrimination disappears when frictions are low.

Even though there is a large theoretical literature on how frictions enable firms to discriminate against their employees, there is comparatively little empirical work on the effect of labor market competition on discrimination. Biddle and Hamermesh, 2013 and Baert *et al*, 2014 show that employers discriminate less in labor markets with a small number of job seekers relative to vacancies.[3] By offering more job opportunities, our mobility shock can be

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

[2]For empirical evidence of Becker's argument that intensified product market competition can reduce race-based differentials caused by prejudice, see e.g., Levine *et al.*, 2008.

[3]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

interpreted as a decrease in the ratio of employed job seekers to job offers. This situation should discourage employers from indulging in discriminatory tendencies.

Our paper also contributes to the literature on the effects of labor mobility constraints on wages. For instance, Naidu (2010) and Naidu, Nyarko, and Wang (2014) study the effect of lifting mobility constraints for a particular group of workers.[4] They confirm that wages rise due to a decrease in monopsony power, but without relating this trend to discrimination. We show that decreasing labor market frictions does not only reduce the monopsony power of firms, but also reduces discrimination even when employers are prejudiced.

Furthermore, our results could be important for public policy. Removing constraints on worker mobility, such as quotas, work permits, or restrictive contracting rules, may improve the capacity of workers to move from prejudiced to unprejudiced firms and reduce discrimination.[5] When mobility is constrained, a firm is able to act on its prejudice because of the low cost of doing so.

Using the European soccer market is a first step to study the link between discrimination and labor mobility constraints, and offers four important advantages. First, following the Bosman ruling, we observe large variation in labor mobility. 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.[6] The rul-

less conclusive, however. Baert *et al* (2014), using the method pioneered by Bertrand and Mullainathan, 2004, 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.

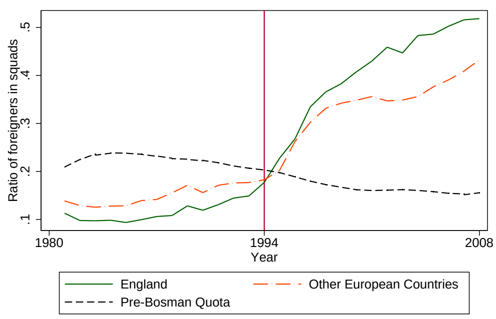
[4]Migrant workers in the United Arab Emirates in Naidu, Nyarko, and Wang (2014), and black Americans in the post-Bellum south in Naidu (2010).

[5]Examples of restrictive contracting rules are the use of non-compete rules or tacit agreements between firms to restrain labor competition. One recent example is the "Techtopus" case, where major Silicon Valley firms such as Google, Apple or Intel stand accused of passing tacit agreements not to hire each other's employees (see NY Times, April 24th 2014). Another worrying development is the rise of non-compete clauses in all types of jobs, such as summer camp counselors or hair stylists (see NY Times, June 8th 2014).

[6]In Europe, 5 foreigners were allowed to play per team, of which 2 only if they had

ing removed the quota barriers for European Union (EU) nationals and the obligation to pay a fee for out-of-contract players.[7] Figure (1) illustrates the intensified mobility of the soccer market after 1995 in the wake of the Bosman ruling. Before 1995, the quota of foreign players in England, represented by the dashed grey line, was binding but not fulfilled. After 1995, the ratio of foreigners in squads skyrocketed. 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 constraints on racial discrimination.

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 5 foreign players in percentage of the size of the squads.

Second, extensive data on the career paths of professional players can be gathered for most countries over long time periods.[8] Third, we can match this extensive individual data with information on the skin-color of players.

Fourth, the soccer market offers a simple test for racial discrimination in salary setting (Szymanski, 2000), which can be used to test the effect of our mobility shock on discrimination. Under the assumption that soccer is an

played in the host country for more than 5 years.

[7]Specifically, in June 1990, at the end of his contract, the former Belgian player Jean-Marc Bosman refused Liege's offer of a contract extension at only 25% of his old wage, and accepted a contract from the French club Dunkirk instead. 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.

[8]Kleven, Landais, and Saez (2013) compiled data of all first-league soccer players for 14 Western European countries since the eighties.

efficient market,[9] a team's wage bill should perfectly reward the talent of its players and explain the club's performance. Discrimination can then be said to exist if, for a given wage bill, clubs fielding an above-average proportion of black players systematically outperform clubs with a below-average proportion of black players. This implies 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. However, he does not explain why discrimination survives in equilibrium and which public policies can fight it. We first 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 low mobility constraints. We briefly present these two contributions.

We first develop a model to guide our empirical analysis. We merge ideas from search models à la Burdett and Mortensen (1998) with a Becker-style assumption of taste for discrimination to derive equilibria in which group differentials persist. Firms are heterogeneous in the talent of the workers whom they employ. For a given talent, firms offer workers two types of contracts: perfect and imperfect. A job offer is considered to be "perfect" when the worker receives a wage that perfectly rewards his talent, and "imperfect" when the worker earns only a fraction of what his talent is worth. The worker accepts the imperfect contract if the wage offered exceeds his reservation wage but searches on the job. Hence, the firm faces a trade-off when offering an imperfect contract: diverting a share of the player's talent but taking the risk of seeing the player poached by a rival, which induces a costly turnover. 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 them a perfect job offer is lower. This results in race-based wage differentials. We find, however, that prejudice does not translate into discrimination when labor mobility constraints are sufficiently low. Because

[9] 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.

monopsony power is reduced, firms should offer *perfect* contracts to avoid costly turnovers. Thus, for a given talent, preferred and disliked workers receive the same wage.

The second contribution consists of using the Bosman ruling on the English soccer market to estimate the causal effect of intensified labor mobility on discrimination. Interpreting this ruling as a shock on job-to-job mobility constraints, we find empirical evidence that wage discrimination disappeared in the post-Bosman period. The difference between the pre- and post-Bosman eras is robust to the use of various estimators (OLS, within, IV and GMM), dependent variables (either league or match performance), and groups of players (either black or white and either English or non-English). Moreover, the decrease in discrimination appears to be extremely fast: we find evidence of discrimination in the 5 years before the ruling, but not in the following 5 years. Interestingly, due to higher mobility constraints, we find that *black non-EU* players still face some wage discrimination in the post-Bosman era.

Potential objections to our empirical findings are related to events contemporaneous to the Bosman ruling. First, concomitant to our work, Palacios-Huerta (2014) confirms the absence of discrimination in the English soccer market from 1993 to 2008 using a different sample. His insightful intuition is that the emergence of a market for corporate control of English professional clubs has increased the competitiveness of English soccer, which would have driven discriminating firms out of the market. However, the market for corporate control started since the early 1980s, i.e., far before the ruling. Moreover, a club experiencing very poor performance simply moves down from one division to another and is not necessarily driven out of the market.[10] Our hypothesis is that the labor market became more competitive primarily because of relaxed constraints on mobility.

A second potential objection is that what we are measuring may not truly be the effect of the Bosman ruling but instead a general change in attitudes toward racism and racial discrimination. The data reject, however, the idea of a general and dramatic reduction in prejudiced views since 1995. Sport is

[10]Note also that some firms may operate under a soft budget constraint (see Andreff, 2014) relaxing the impact of a very poor performance.

not immune, as shown by recent episodes of racial bias in basketball courts (Price and Wolfers, 2010) and soccer fields. Since the Bosman ruling, racist incidents in soccer, whether from fellow soccer players, owners, managers or supporters continue to make the headlines of English newspapers.[11] Frequent racist 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 caused by a decrease in job-to-job mobility constraints, in the wake of the Bosman ruling, rather than a dramatic change in attitudes or prejudiced views in 1995.

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 constraints affect discrimination. In section (4), we present our 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 when constraints on worker mobility are lowered. Section (6) concludes.

2 The competitive market for soccer players

We have already discussed the four important advantages offered by the soccer market to the study of labor mobility and discrimination: (1) large observed variation in mobility constraints following the Bosman ruling, (2) an extensive collection of data [12] on the career paths of professional soccer players, (3) data on the skin-color of individuals that can be matched with our dataset, and (4) a simple test for racial discrimination in salary setting

[11] In a web appendix, we show strong evidence of racism in English soccer in the post-Bosman era. As a recent example, 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. The Daily Mail reports that 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 that "Not many white faces amongst that lot but worth considering."

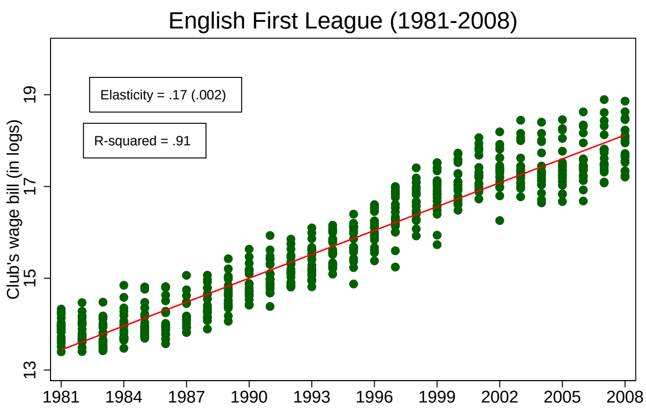
[12] See Appendix (A) for further information on variable definitions and data sources.

(Szymanski, 2000). We document four facts about the soccer market that will guide the development of the model in Section 3.

Fact 1: league competition is hierarchical. Competition is 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.

Fact 2: clubs are heterogeneous in wage bills. Figure (2) reports the log of the clubs' wage bills in the English first league from 1981 to 2008. Each year, we observe a dispersion of wage bills supporting club heterogeneity. It is also worth mentioning that wage bills are linearly increasing over the years for all clubs, without being affected by the Bosman ruling shock.

Figure 2: Club's Wage Bills

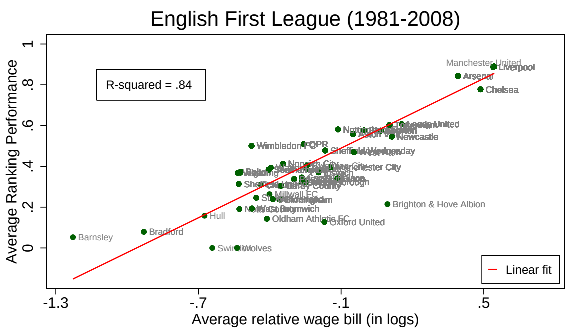


Notes: Each dot stands for one club. The regression line is depicted. Elasticity coefficient from the OLS regression of the log-club's wage bill on year dummies is reported with standard error in parenthesis.

Fact 3: soccer is a competitive market for talent. A common explanation for the increasing trend in wage bills depicted in Figure (2) is the increasing price of talent. Players earn wages proportional to their talent and demand for talent is increasing, which raises the price of talent.

Fact 4: talent explains performance. A team's sporting talent is highly correlated with its success and performance. Thus, the higher the club's sporting talent, the higher its wage bill, and the higher its ranking.[13] In Figure (3), we observe that wage expenditures and performance are heavily correlated in the English league between 1981 and 2008.

Figure 3: Average Wage Bill and Performance



Notes: Average ranking performance is the average club's performance from 1981 to 2008. Performance of club *i* is computed each year *t* as $[(\text{Ranking}_{it} - \max_t)/ (\min_t - \max_t)]$, where Ranking_{it} min and max are the lowest and highest possible rankings in *t*.
Average relative wage bill is the average club's relative wage bill from 1981 to 2008. The relative wage bill of a club *i* in year *t* is the log difference of the club's wage bill. relative to the annual average.

3 Theoretical framework

We develop a 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 (as motivated by the above fact 2). We assume that a club's budget is given and is spent searching for talent and paying wages. In this search setting, we assume that some employers hold a 'taste' for racial discrimination.[14] If this assumption holds, then disliked workers 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. Becker (1957)

[13]In Appendix (B), we construct a proxy for a team's sporting talent, based on a crude measure of team quality. Figure (8) depicts a clear linear correlation between wage bill and team quality, with an r-squared of 0.81.

[14]Altonji and Blank (1999) reviews the pioneering works introducing search into taste-based theories of discrimination. Note that statistical discrimination, based on stereotypes and made possible by imperfect information, is not an issue in our setting since employers easily observe the performance of soccer players.

argues, however, that prejudice against disliked workers does not necessarily result in economic discrimination and race-based wage differentials. In other words, an employer's 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 labor mobility, as in the pre-Bosman period, may explain race-based wage differentials. This result is formalized below.

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

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; 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, including contracting rules, institutional constraints and barriers to mobility. This market parameter also depends on the worker's current situation (employed or not). 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.[15] 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 worker 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 (motivated by fact 3), 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 (described below).[16]

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., see 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.

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

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

[16] 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.

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). A bad or imperfect job offer is rejected by an employed worker and accepted by an unemployed one. A worker with an imperfect contract 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.[17]

3.2 Firms

Firms maximize team performance, which is measured by the percentage of wins, championship success, or, as is usually the case in league soccer, league position (Szymanski, 2000). Team performance depends on the quantity of sporting talent hired by the club. Remember that firms are homogeneous in size but heterogeneous in wage bill (or total sporting talent). For a given wage bill, we are agnostic about the talent distribution across workers within firm and we abstract from analyzing diversity and spillover effects to focus on the role of prejudice and the type of job offers (perfect or imperfect).

Firms maximize team performance simply by minimizing the total cost of talent given total resources. We assume, for the sake of simplicity, that total

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

resources are exogenously given when the decision is made. In the steady state, each firm equalizes its wage bill to its total resources.

In this model, we do not detail how workers and firms are matched. Once a random match is made, we endogenize the probability of offering a perfect job offer, γ, to a worker based on the firm's behavior. In general there will be many causes leading particular firms to prefer one worker to another, such as the role of appearance that will be considered below. For the moment, the ensemble of hiring considerations is symbolized by 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 positive turnover cost $c(t)$[18]. This tension can be simply modeled by comparing the expected costs of offering a perfect (J_P) versus an imperfect (J_I) contract to a given worker with talent t:

$$\begin{aligned} J_I &= (1-k)t + (\delta+\lambda_1\gamma)(V - J_I - c(t)) \\ &\Leftrightarrow \; J_I = \frac{(1-k)t + (\delta+\lambda_1\gamma)(V - c(t))}{r+\delta+\lambda_1\gamma}, \end{aligned} \tag{4}$$

$$J_P = \delta(V - J_P - c(t)) \Leftrightarrow J_P = \frac{\delta(V - c(t))}{r+\delta}. \tag{5}$$

We assume that when a worker leaves the firm, a vacancy is created, with V the value of a vacancy. Because the turnover cost $c(t)$ is positive, the firm incurs a loss whenever a worker leaves. This happens either through the exogenous separation process, which occurs at rate δ, or because a rival firm has poached the worker, which occurs at rate $\lambda_1\gamma$. Equation (4) states the tension between the loss that occurs from the player leaving or being poached $(V - J_I - c(t))$ and the gain from diverting part of the talent $(1-k)t$. So,

[18] The same mechanism can be found in Stiglitz (1974), where firms in urban areas increase wages to reduce costly turnovers.

the firm will offer a perfect job offer if:[19]

$$
\begin{aligned}
J_P > J_I \quad &\Leftrightarrow \quad \frac{\delta(V - c(t))}{r + \delta} > \frac{(1-k)t + (\delta + \lambda_1\gamma)(V - c(t))}{r + \delta + \lambda_1\gamma} \\
&\Leftrightarrow \quad c(t) > \frac{(1-k)t(r+\delta)}{\lambda_1\gamma r}. \qquad (6)
\end{aligned}
$$

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

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

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 becomes more concentrated at these higher cost levels. Assuming that c is distributed Pareto and convex in talent,[20] such that $c(t) = ct^2$, yields a simple closed-form solution for γ. The probability of receiving a perfect offer is such that $J_P > J_I$:

$$\gamma(t) = \left(\frac{\tilde{c}\lambda_1\gamma r t}{(r+\delta)(1-k)}\right)^{\alpha} = \left(\frac{\tilde{c}\lambda_1 r t}{(r+\delta)(1-k)}\right)^{\frac{\alpha}{1-\alpha}}. \qquad (7)$$

As shown in equation (7), the probability γ depends primarily on three important variables, λ_1, k, and t, which govern differences in wages. Everything else being equal, λ_1 determines the strength of the firm's monopsony power. A high value of λ_1 implies a low monopsony power, and consequently firms are less likely to offer low wages to their employees because they anticipate that other firms can poach them. Thus, a higher λ_1 reduces the probability of being in an imperfect match by increasing γ and decreases the number of employees that are stuck in imperfect matches (see equation 13).

[19] We use 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: $V = 0$.

[20] The convexity of the turnover cost is empirically supported by the fact that star players are more talented than journeymen and face a higher demand for their talents (see Andreff, 2014).

The effect of the other two parameters is straightforward. First, as k increases, firms divert a smaller share of the monetary value of a worker's talent; therefore offering an imperfect contract to a worker is less attractive. Second, the higher is t, the more costly it is to replace talented workers, and the higher is γ the probability to offer a perfect job.

3.3 Employer prejudice

Considering that some firms hold a 'taste' for discrimination against type B workers, Bowlus and Eckstein (2002) and Bowlus *et al.* (2001) assume that the arrival rates (λs) and the job destruction rate (δ) vary according to the type of worker (A and B). Accordingly, job offer rates from prejudiced firms are lower for B than 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 worker's type, 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 between types A and B. The second reason is that models based on Mortensen (1988) have unrealistic, left-skewed wage distributions. This problem does not occur in our model, as the wage distribution depends on the distribution of talent.[21]

In our approach, the prejudice applies to the separation rate. We assume that A and B have different Pareto distributions governing the turnover cost. In particular, we consider that $\tilde{c}_B = \tilde{c}_A - d$, where d represents the employer's taste for discrimination and $\tilde{c}$ is the lower bound cost, such that the Pareto distribution of the turnover cost for B workers is shifted to the left. In our labor context, this shift to the left reflects 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 implies a lower equilibrium probability of

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

receiving a perfect job offer (γ) for workers in group B than in group A.

From the assumption on γ, we can ground Szymanski's market test. In the steady-state, resources (R) are equal to the club's wage bill (Ω), plus the sum of realized turnover costs (C). We can write this down as:

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

where μ is the share of black players in the squad, and t_i is the talent of player $i = A, B$. Consider high mobility constraints (i.e., low λ_1) and two teams with the same resources (R) but different shares μ of black players. Because $\gamma_B < \gamma_A$ when λ_1 is low, the team with the higher share of black players will have a lower wage bill for the same level of sporting talent. This implies that when the two teams saturate their resource constraints, they will have the same wage bill but different sporting talent. The team with the higher share of black players will have a higher performance, which is exactly what Szymanski's market test aims to estimate.

3.4 Wage discrimination

Our perspective is that the Bosman ruling modifies the general state of the labor market. It lowers the constraints on labor mobility and thus affects the arrival rate of job offers for employed workers (λ_1).[22] Holding everything else constant, we assume that lower constraints are associated with a higher arrival rate of job offers. Indeed, the end of transfer fees for out-of-contract players and of quotas within the EU imply that players, independent of their race, 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

[22] We are not the first to consider that changes in λ may reflect changes in the state of the labor market. As stated in Manning (2003) "in the model of Burdett and Mortensen (1998), the extent of employer market power is determined by the rate at which job opportunities arrive relative to the job destruction rate: the lower the arrival rate of job offers, the more market power employers will have."

states of the labor market.[23]

We 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:

$$\mathbb{E}_A(\omega|t) - \mathbb{E}_B(\omega|t) = bu_A + ktI_A + tP_A - bu_B - ktI_B - tP_B.^{24} \tag{8}$$

After some algebraic manipulation (reported in appendix C.2) and given that both groups have the same likelihood of being out of work, equation (8) reduces to

$$\mathbb{E}_A(w|t) - \mathbb{E}_B(w|t) = t\lambda_0\delta(\gamma_A - \gamma_B)\left(\frac{(1-k)(\lambda_1+\delta)}{(\delta+\lambda_1\gamma_A)(\delta+\lambda_1\gamma_B)(\delta+\lambda_0)}\right). \tag{9}$$

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 offer does not depend on the type of worker.

How does wage discrimination vary when we modify the job offer arrival rate (λ_1)? Everything else held equal, the effect of a higher λ_1 on discrimination is twofold: it decreases the steady-state share of individuals in bad or imperfect matches I (see equation 13), and increases the likelihood of receiving a perfect offer, γ (equation 7).

Overall, through γ, the effect of λ_1 on wage discrimination is non-linear. For values of λ_1 close to 0, firms anticipate that their workers cannot be poached and both A and B receive imperfect wage offers. In this extreme case, there is no discrimination because workers A and B earn the same (imperfect) wage for the same talent. As λ_1 increases, discrimination begins to *increase*. Because all firms endogenize the consequence of the taste for discrimination (i.e., $\gamma_A \geq \gamma_B$), A-type workers receive an increasing number of perfect job offers compared to B workers, and wage discrimination increases for a given level of talent. Then, as λ_1 continues to increase, mobility

[23] 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.

[24] See C.1 for the equilibrium flow conditions. These conditions enable us to compute I and P, the number of workers with imperfect and perfect contracts.

constraints continue to decrease, and the discrimination process is reversed. B-type workers begin receiving a higher number of perfect job offers. As a result, the wage gap narrows and even disappears when constraints are sufficiently low. Thus, the relationship between λ_1 and wage discrimination follows an inverted-U shape pattern. Formally we state this in Proposition (1).

Proposition 1. *The race-based difference in expected wage is a parabola that has a single maximum. Wage discrimination first increases and then decreases with constraints on job-to-job mobility.*

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

To gauge the effect of mobility constraints on wage discrimination, we simulate the effect of an increase in λ_1 on $\mathbb{E}_A(\omega|t)$ and $\mathbb{E}_B(\omega|t)$ using plausible parameter values.[25] The simulated expected wages for workers of type A (whites) and B (blacks) are reported in Figure (4).[26] They are both increasing with the job offer arrival rate.[27] Then, we report the difference in expected wages for type B in Figure (5), which depicts the inverted-U shape pattern proved in Proposition (1).

3.5 Job turnover

Our model also offers interesting predictions regarding the turnover of type A and B workers, which explains the wage discrimination pattern. The expression for job turnover in a given period is $\lambda_1 \gamma I$. We can show that there exists a job offer arrival rate, λ_1^*, such that the turnover of A workers is

[25] Although our labor market differs from the one analyzed by Bowlus and Eckstein (2002), we use their estimated values for λ_0 and δ based on the National Longitudinal Survey of Youth. The parameter values for the simulations are tabulated in Appendix (D).

[26] Values of λ_1 are low but consistent with the estimated values in Bowlus and Eckstein (2002) and Bowlus, Kiefer, and Neumann (2001).

[27] Naidu (2010) and Naidu, Nyarko, and Wang (2014) document empirically that wages are increasing with higher labor mobility.

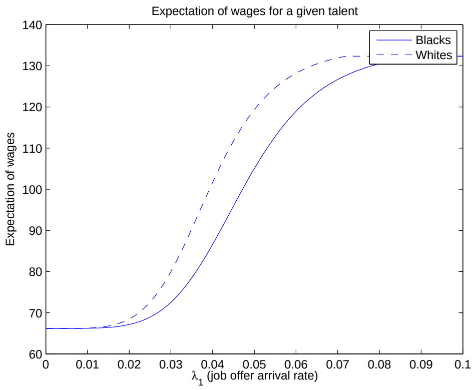


Figure 4: Wage Expectation

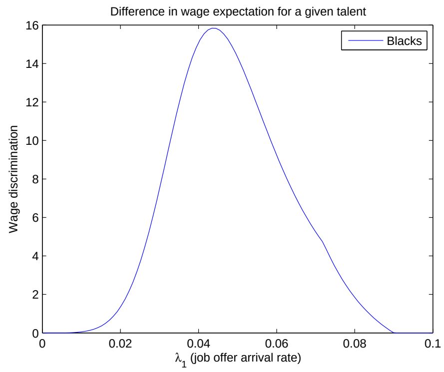


Figure 5: Wage discrimination

higher than that of B workers when below this rate, and lower when above it. This pattern is stated formally in Proposition (2).

Proposition 2. *A decrease in constraints on job-to-job mobility causes a racially differentiated change in job turnover.*

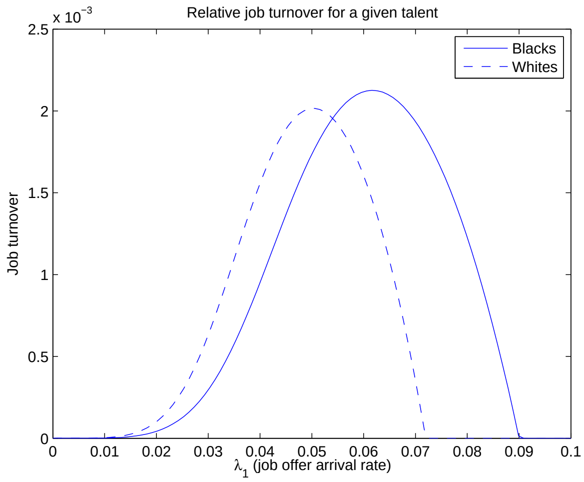
Proof. Appendix (C.4) presents the proof. □

As in the preceding section, we simulate for each group, A (whites) and B (blacks), the relationship between λ_1 and the job turnover using the same plausible parameter values. The simulated inverse U-shaped curves are depicted in Figure (6).

There are two reasons that job turnover is low at both extremes of the curves in Figure (6). As an illustration, consider one extreme case: a low λ_1 (i.e., high mobility constraints). First, in this case, employed workers have a small likelihood of moving. Second, the probability of receiving a perfect job offer (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 perfect job offers. However, as λ_1 increases, a 'stock' effect comes into play: workers are less likely to move simply because they are more and more numerous to be employed in a perfect job. Thus, as λ_1 increases, the pool of workers

Figure 6: Job turnover



who want to change clubs (I) decreases, and so the number of moves ($\lambda I \gamma$) decreases as well. The differentiated turnover patterns of A and B come from the differences in γs: the "stock" effect emerges earlier for A than for type B workers.

4 Empirical design

We first discuss how we move from our theory to empirics (4.1). Then, in subsection (4.2), we present the market test to detect wage discrimination and discuss our data on racial information, as well as its sources and limitations. In subsection (4.3), we explain why we use the Bosman ruling as a mobility shock, and present some descriptive statistics. Finally, we detail the equations that we estimate to evaluate the extent of race-based wage discrimination in subsection (4.4).

4.1 Model predictions

Our theory predicts an inverted U-shaped relationship between wage discrimination and constraints on job-to-job mobility (see Proposition 1). In moving from theory to estimation we face, however, one major issue. The levels of mobility constraints in the European soccer market are unfortunately not readily quantifiable. In lieu of quantifiable measures, we use the Bosman ruling shock to infer information on the level of constraints. We assume that the pre-Bosman period was characterized by high constraints on mobility, while constraints are low in the post-Bosman era. Recall that European players may now receive job offers from all the EU clubs without quota limitations and out-of-contract restrictions. We expect the post-Bosman constraints to be sufficiently low to eliminate the pre-Bosman wage discrimination.

Our theory provides a plausible explanation of why wage discrimination could have decreased post-Bosman despite prejudice against black players: as the arrival rate of job offers increases, the monopsony power of firms decreases and workers are more likely to receive "perfect" job offers.

Another interesting aspect of our model is Proposition (2). It predicts a racially differentiated effect on job turnover as constraints on job-to-job mobility decrease (see Figure 6), which is what we will observe in the data.

4.2 The market test for discrimination

Our empirical analysis uses Szymanski's market test. The intuition behind this test is simple: if all individual talent is perfectly rewarded, the team's performance (a function of talent) should be perfectly explained by the team's wage bill. 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). This test is perfectly compatible with our theoretical framework (see subsection 3.3).

The market test for discrimination requires information on the skin-color of players that we can match with extensive individual data, as well as data on clubs' wage bills. We explain how we collected and coded the data in Appendix A.

4.3 The Bosman shock as a source of identification

We apply the market test for discrimination to a panel of all English clubs in the first 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 EU players in the EU's national leagues and by allowing players in the EU to move to another club at the end of their contract without a transfer fee being paid.

Though this decision came into force in December 1995, it could have been anticipated. Indeed, the Bosman case had been submitted to the Court two years before, 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 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 can construct some informative descriptive statistics (see Appendices A and B for data details). We have information on 3,788 players who participated in the first English league during our two periods (1981-1993 vs. 1996-2008); 70% of those players are English.[28] This number was higher before the Bosman

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

ruling (93.2%) than after it (52.6%). In total, 10.7% of players are English and black; this number is fairly stable before (9.9%) and after (11.2%) Bosman. However, the number of black non-English players has skyrocketed from 0.8% before the ruling to 15% after it. The consequence is that the number of white English players decreased from 83.3% before the ruling to 41.3% after the ruling.

In the descriptive statistics, we keep only black English and white English players in our sample to avoid comparing players of different nationalities. Table (1) reports average characteristics for black and white English players in the first league, both before and after the Bosman ruling. We observe in the pre-Bosman period that black and white players played a fairly similar number of matches. However, black players were slightly more qualified, one year younger and a bit more tenured. In the post-Bosman era, we observe notable changes: the quality difference becomes statistically not significant and black players are now *less* tenured than white players.

Interestingly, all the differences in differences are significant between the two periods (reported in the last column of Table 1). In the post-Bosman era, players on average played less matches, are qualitatively better, more than 6 months older and more tenured. These statistics clearly show that there was a Bosman effect on players and that this change differed along racial lines.

Table 1: Individual differences in means: black vs white English players

	Pre-Bosman (1981-1993)			Post-Bosman (1996-2008)			Overall
Variable	Black	White	Diff.	Black	White	Diff.	D. in D.
Player's number of matches	20.79	20.46	.33	18.29	17.70	$.59^c$	2.68^a
	[9.3]	[9.6]	(.37)	[8.3]	[8.6]	(.30)	(.18)
Player's quality level	2.42	2.24	$.18^b$	2.88	2.78	.10	$-.54^a$
	[1.9]	[2.1]	(.08)	[1.9]	[1.9]	(.07)	(.04)
Age of players	23.79	24.87	-1.08^a	24.92	25.59	-0.67^a	$-.71^a$
	[4.0]	[4.4]	(.16)	[4.7]	[5.0]	(.17)	(.09)
Tenure (in years)	2.50	2.38	$.12^a$	2.4	2.72	$-.32^a$	$-.25^a$
	[1.2]	[1.2]	(.05)	[1.1]	[1.8]	(.07)	(.03)
Observations	737	5152		980	3531		

Notes: This table reports average characteristics for black and white English players in the first league. 'Diff.' means difference in means between blacks and whites; 'D. in D.' means difference in difference between pre- and post-Bosman period. Standard deviations are in brackets; standard errors are reported in parentheses, with ^{a}, ^{b} and ^{c} denoting significance at the 1%, 5% and 10% level, respectively.

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

Table (2) using club statistics. All differences in terms of wage bill, transfer fee records and stadium attendance are highly significant both economically and statistically. The only exception is the share of black English players, which is constant in both eras. These variables are used as regressors or instruments and will be discussed later on.

Table 2: Club differences in means

Variable	Pre-Bosman (1981-1993)	Post-Bosman (1996-2008)	Differences in Difference
Wage bill (in millions of pounds)	2,9	35,5	32,6[a]
	[2,4]	[25,5]	(1,5)
Transfer fee record (in millions)	1,31	8,96	7,65[a]
	[1,3]	[6,5]	(0,38)
Stadium attendance (in thousands)	21,5	33,2	11,7[a]
	[8,9]	[12,0]	(0,9)
Share of black English players	0.13	0.14	0.01
	[.1]	[.1]	(0.01)
Observations	304	258	

Notes: This table reports average characteristics for English first league clubs. 'Difference' means difference in means between pre- and post-Bosman period. Standard deviations are in brackets; standard errors are reported in parentheses, with [a] denoting significance at the 1% level. Wage bills and transfer fee records are in nominal values.

As emphasized in Proposition (2), we should 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:

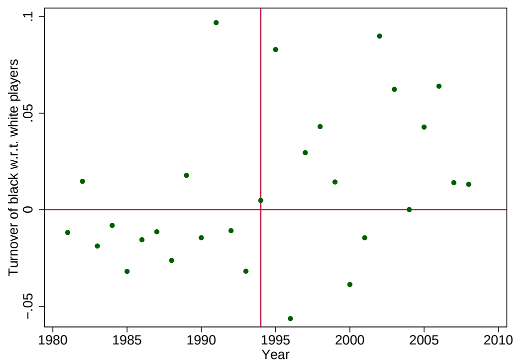
$$\text{Turnover} = \frac{\text{Share in transfers}_B}{\text{Share in population}_B} - \frac{\text{Share in transfers}_A}{\text{Share in population}_A}.$$

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.

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.[29] This reversed trend suggests that black players "voted with

[29]It is unclear, however, when precisely this trend reversed. This may be because we

Figure 7: Relative turnover of black English players



their feet" when the market was liberalized. This is in line with the reduction in job tenure observed for black players in Table (2) after the Bosman ruling. We also find evidence, presented in Appendix (E), 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 wage discrimination is present in the soccer labor market.

4.4 Estimated equations

To apply the market test, relating a club's performance to its wage bill and its share of black English players, we use two different measures 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 the finding of apparent wage discrimination in the *pre*-Bosman era. However, we find that wage discrimination has disappeared in the *post*-Bosman era. This is consistent with our theoretical predictions on the effect of relaxed mobility constraints.

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 were not perfectly captured.

The league performance specification

$$\begin{aligned}\text{League Performance}_{it} = \alpha_i &+ \beta_1(\text{WageBill}_{it} - \overline{\text{WageBill}}_t) \\ &+ \beta_2(\text{PlayersNb}_{it} - \overline{\text{PlayersNb}}_t) \\ &+ \beta_3(\text{Shareblack}_{it} - \overline{\text{Shareblack}}_t) + \epsilon_{it}. \qquad (10)\end{aligned}$$

League Performance$_{it}$, is computed in relative terms as $\left(\frac{Ranking_{it} - min_t}{max_t - min_t}\right)$, where $Ranking_{it}$ is the final ranking of team i at the end of season t; and min and max are the lowest and highest possible rankings each season. α_i is a club i fixed effect, capturing permanent team-specific characteristics affecting performance, such as a location effect,[30] 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 club i in a season t relative to the annual average. Controlling for club fixed effects and wage bills, the relative number of players captures "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 i 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 (including Irish, Scottish and Welsh black players - see footnote 28). The reason is that the share of black English players is relatively stable over time, while the share of black non-English players is constantly rising. Thus, we avoid estimating discrimination toward different individuals before and after Bosman. We later discuss results based on both black English and non-English players.

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

[30]Some players could prefer to live in London rather than Liverpool for instance, which would allow London-based clubs to have better performances for lower wage bills.

The match performance specification

One drawback of the team performance measure, based on final rankings, is the relatively limited number of observations. We thus consider a new dependent variable based on match results that offers many more observations:

$$\begin{aligned}\text{Match Performance}_{ijt} = \xi_{ij} &+ \beta_1 \log(\text{WageBill}_{it}/\text{WageBill}_{jt}) \\ &+ \beta_2(\text{PlayersNb}_{it} - \text{PlayersNb}_{jt}) \\ &+ \beta_3(\text{Shareblack}_{it} - \text{Shareblack}_{jt}) + e_{ijt}, \qquad (11)\end{aligned}$$

Match Performance_{ijt} between the home team i and the away team j in year t is simply the goal difference in their match. 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 a large difference in wage bills, which explains a large difference in sporting talent, to lead to a large difference in performance (i.e., in goals scored). The match performance is conditioned on a match fixed effect (ξ_{ij}) and e_{ijt} is the usual error term. Our wage bill variable is the log difference between the wage bills of the two clubs, $\log\left(\text{WageBill}_{it}/\text{WageBill}_{jt}\right)$. We also add the difference in the number of players used ($\text{PlayersNb}_{it} - \text{PlayersNb}_{jt}$). Our variable of interest is the difference between the two teams in the share of matches played by black English players ($\text{Shareblack}_{it} - \text{Shareblack}_{jt}$).[31]

5 Empirical results

We estimate models (10) and (11) 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.

[31] 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.

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 club clustering (to allow for a correlation of the error within clubs 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 because it eliminates the club fixed effects (through first-differencing) and 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 10) without instrumenting for the wage bill. The relative wage bill variable has a positive effect on performance, which is economically and statistically significant. Unsurprisingly, this effect is larger if we do not control for the club fixed effects. 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 share of black English players. In the pre-Bosman era (columns 1 and 2), we find that performance depends significantly on the team's racial composition, suggesting a race-based wage discrimination: if talent of black players is underpaid compared to talent of white players, controlling for wage bill will not fully explain performance. In other words, part of talent is not fully accounted by the club's wage bill. Therefore, teams fielding an above-average proportion of black players outperform clubs with a below-average proportion of black players. After Bosman (columns 3 and 4), the apparent wage discrimination disappears. Performance is now independent of the racial composition of teams. This result is consistent with our theoretical predictions on the effect of relaxed mobility constraints on wage discrimination.

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

Dependent Variable:	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 employed	0.577^a	0.491^b	-0.130	-0.000
	(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 is computed in relative terms as $\left(\frac{Ranking_{it}-min_t}{max_t-min_t}\right)$, where $Ranking_{it}$ is the final ranking of team i at the end of season t; and min and max are the lowest and highest possible rankings each season. Robust standard errors are 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 in the pre-Bosman period? Let us compare the 1993 situation of two clubs that are identical except in their share of black English players: (1) a club that does not employ black players, and (2) a club that employs 3.7 black players (the 1993 average number). Based on our estimates, we find that to obtain the same performance with equally talented white players, the club that does not employ black players should pay 800,000 pounds more than the other 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 issues? 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 stadium attendance, and relative record transfer fees.[32] The key identifying assumption is that these variables do not affect league performance, except

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

through their effect on the wage bill.

We briefly discuss the relevance of our three excluded instruments before presenting the results. Good performance in cups and higher stadium attendance in the previous season both generate higher revenues in the contemporaneous season. Then, we use record transfer fees to capture potential buyouts by rich owners. New owners often break the club's transfer fee record, but these purchases have been documented to have on average little effect on final performance. Many record purchases prove to be poor value for money, and the effect of transfer fees on performance are insignificant when controlling for wage bill.

The instruments are constructed as follows. Relative lagged stadium 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 lag 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 of the within-IV results are reported in Table 9 (see Appendix F) and show that these instruments have a significant effect on the relative log wage bill, except for the lagged cup performance in the post-Bosman era. The second-stage of the within-IV results are presented in columns (1) and (3) of Table (4). These results confirm the post-Bosman effect on wage discrimination. Whereas the coefficient for the share of black players employed is positive and significant before the Bosman ruling, it is insignificant post-Bosman.

The IV results could be affected by a weak instrument problem. If the instruments correlate 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 F). 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 re-

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

Dependent Variable:	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 is computed in relative terms as $\left(\frac{Ranking_{it}-min_t}{max_t-min_t}\right)$, where $Ranking_{it}$ is the final ranking of team i at the end of season t; and min and max are the lowest and highest possible rankings each season. AB means Arellano-Bond. Robust standard errors are in parentheses, clustered by club, with ^{a}, ^{b}, and ^{c} denoting significance at the 1%, 5%, and 10% percent level, respectively.

strictions 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: wage discrimination appears to be significant before the Bosman ruling but not after the ruling. Again, these results are consistent with our theoretical predictions on the effect of relaxed mobility constraints.

Before using match data, which allows for more statistical power to run the discrimination market test, we present a further test that supports our claim that the Bosman shock is driving the decrease in discrimination. This test is on the differences between black English and black foreign players.

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.[33] As a consequence, despite a general decrease in labor mobility constraints, mobility is relatively more constrained for non-EU players, even after the Bosman ruling. We expect black non-EU players, since they are more constrained, to still face wage discrimination.

We find evidence of wage discrimination against black non-EU players by performing the market test after Bosman on different shares of players: black English, EU black, non-EU black, and non-EU white.[34] Results are reported in Table (5). We find that the estimates of 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 estimates imply that wage discrimination against black non-EU players could be present in the English first league post-Bosman.

[33] 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 must 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.

[34] Non-EU players are non-member players of the common market or the European free trade association zone. We do not report results for the pre-Bosman period, as there were only few non-English black players playing in England (see section 4.3).

Table 5: *Post*-Bosman: are non-EU black players discriminated? (1996-2008)

Dependent Variable:	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 black EU players employed	0.056	0.029	0.125	0.174
	(0.20)	(0.20)	(0.31)	(0.32)
Share of *black non-EU* players employed	0.566^{a}	0.580^{a}	0.587^{c}	0.613^{c}
	(0.20)	(0.20)	(0.34)	(0.34)
Share of white non-EU 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	258	257	251	249
Club fixed effect	No	Yes	No	Yes

Notes: the dependent variable is computed in relative terms as $\left(\frac{Ranking_{it}-min_t}{max_t-min_t}\right)$, where $Ranking_{it}$ is the final ranking of team i at the end of season t; and min and max are the lowest and highest possible rankings each season. 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 (F) and available upon request with the corresponding usual statistics.

5.2 Discrimination Market-test on Match Performance

To test the robustness of our results, we employ a new methodology to detect discrimination based on a different performance variable: the market test on matches (see equation 11). For the sake of simplification, we use our two preferred estimators: 'Within' and 'Within-IV'. This choice implies that for all estimations we control for a match fixed effect (ξ_{ij}). 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 performance. We first present results on long time spans, and then, on short time spans (using 5-year time windows).

Pre- versus post-Bosman periods: long time spans

The results on the pre-Bosman (1981-1993) and the post-Bosman (1996-2008) periods are reported in Table (6). In all regressions, standard errors are two-way clustered at the match and the receiving club level, to allow for a correlation of the error within matches across years, and within receiving clubs across years (e.g., any specific trend at home). We consider that the match$_{ij}$ between team i and team j is in the same cluster as ji because a common trend may influence the score of some club pairings independently of which club is receiving or visiting. Note that the results hold for different ways of clustering the standard errors.

Using match level data, we confirm the evidence that discrimination is present before Bosman and disappears after the ruling. The estimates on 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 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 match by one goal. In other words, because of wage discrimination, the club with only black players can hire more talented players. A similar effect could be achieved by doubling the wage budget. By contrast, the non-significant estimates on the share of black players, reported in columns (3) and (4), suggest that wage discrimination has disappeared post-Bosman.

Table 6: Market-test: Match Performance and Discrimination - Within and Within-IV

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	1.546^a	0.491^a	0.330
	(0.18)	(0.47)	(0.12)	(0.27)
Difference in share of black players employed	0.854^b	1.079^b	-0.148	0.039
	(0.33)	(0.43)	(0.31)	(0.40)
Difference in players used	-0.048^a	-0.054^a	-0.030^a	-0.031^a
Observations	4494	3129	4402	3264
Hansen p-value		0.79		0.22
Club-pair fixed effect	Yes	Yes	Yes	Yes

Notes: the dependent variable is the goal difference in the match. Robust standard errors in parentheses with ^{a} denoting significance at the 1% percent level. Standard errors are two-way clustered at the visiting and receiving club level (see text for details).

In columns (2) and (4), we used an instrumental variable approach to account for possible measurement errors and reverse causality. The difference in wage bills is instrumented with the corresponding difference in the instrument variables used in subsection (5.1). The result of the first-stage estimates are reported in Table (10) in Appendix F. 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 is 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).

Pre- versus post-Bosman periods: short time spans

One of the advantages of this new dependent variable is that we have many more observations, and so we can estimate equation (11) on smaller time windows. We check the extent of discrimination before and after the Bosman ruling by using 5-year time windows and two scenarios: (1) 1989-1993 vs. 1996-2000 and (2) 1990-1994 vs. 1995-1999. In the first scenario, we drop the years 1994 and 1995 to avoid any anticipation effect (see section 4.3), while in the second case we include these two years to stress how dramatic and quick the change in discrimination is following the Bosman ruling. The results are reported in Table (7) using the within estimator.[35]

Table 7: Market-test: Match Performance and Discrimination - Within

Dependent Variable:	Match Performance			
Sample:	*Pre*-Bosman		*Post*-Bosman	
	(1989-1993)	(1990-1994)	(1995-1999)	(1996-2000)
Estimator:	Within	Within	Within	Within
	(1)	(2)	(3)	(4)
Difference in log wage bill	0.517^c	0.717^b	0.432^c	0.389^a
	(0.29)	(0.33)	(0.26)	(0.13)
Difference in black players employed	1.409^a	1.203^b	-0.820	-0.591
	(0.45)	(0.55)	(0.58)	(0.70)
Difference in players used	-0.035^a	-0.035^a	-0.039^a	-0.027^b
	(0.01)	(0.01)	(0.01)	(0.01)
Observations	1898	1726	1558	1571
Club-pair fixed effect	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 at the match and the receiving club level (see text for details).

In the Bosman ruling period (col. 1 and 2), the point estimates of the

[35]Results using the Within-IV are qualitatively similar and available upon request.

share of black English players are positive, significant and higher than in the longer time frame (1981-1993; see col. 1 and 2 of Table 6). This seems to indicate that discrimination was increasing in the years leading up to the Bosman ruling. As expected by our theory, we find that the estimates of the share of black English players are statistically insignificant in the 5 years following the Bosman ruling (col. 3 and 4). This sharp difference before and after the ruling, when using smaller time windows, supports the idea that only a profound change in labor market conditions could have brought down discrimination this quickly.[36]

6 Conclusion

In this paper, we find strong evidence that wage discrimination has been eliminated following a decrease in labor market frictions. As shown in our model, a reduction in labor market frictions can erode the monopsony power of firms, leading to an eradication of the black-white wage gap. Our model appears to fit the empirical facts quite well.

A heartening interpretation of our results is that the labor market conditions can cause wage differentials between white and black employees to disappear even if racist attitudes remain.

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[36]This change is remarkable when we consider that contracts for soccer players last around 3 to 4 years. Some players were thus not immediately able to move from one club to another after the Bosman ruling. We may therefore expect the discrimination coefficient to decrease more slowly. We tested for this possibility by interacting the share of black English players with a year trend in our Post-Bosman regression, and found a negative trend that was only significant at the 10% level (results are available upon request). Note that an alternative to changing club is to renegotiate the contract to avoid being poached by other clubs. This renegotiation is quite frequent post-Bosman and typically leads to increases in wage and contract length.

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A Data appendix

Our dataset contains all professional soccer players, regardless of nationality, from the first English league from 1981 to 2008 and comes from Kleven, Landais, and Saez (2013). Our dataset is similar to Szymanski (2000) but differs in scope and time. Szymanski uses a panel of 39 clubs from four divisions in the English soccer league over the 1978-1993 period, i.e., pre-Bosman. In terms of scope, we have access to information on all players from the first division in England. Despite our focus on elite players, the number of clubs in our sample is fairly similar to Szymanski's pre-Bosman sample: 41 teams from 1981 to 1993 against 39. In terms of time, we cover the pre- and post-Bosman eras, from 1981 to 2008. To this dataset we add the following variables.

Wage bills. We use wage bills from the Companies House website, a British government agency that collects annual reports from registered companies. Wage data are provided for almost all the English clubs in our sample. 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.

These wage data are considered reliable because they are obtained from audited annual accounts. There are some issues however. First, the reports are not homogeneous over the 30-year period. Some clubs changed the ending date on their company accounts and reports annual results over thirteen months or more, in which case the data were adjusted on a pro rata basis. Then, 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. Finally, the wage bill is given for all staff, including salaries for scouts, statisticians, physiotherapists, and coaches. However, there are two reasons why it is unlikely to be a significant problem for our analysis. One is that the pay for most of the non-player employees is relatively small compared with the total wages of players; and the other reason is that the share of the pay for non-player employees likely accounts for a similar share of the wage bills in all clubs.

Racial information. 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).[37] 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).[38] These pictures were obtained primarily from the reputable website `transfermarkt.de`, and when pictures from that site were not available, we 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.

Control variables. In addition to information on race and wage bill, we have 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 part of this information to create an objective, albeit imperfect, measure of player's sporting quality (see Appendix B). Moreover, we added information on team's ranking and attendance from the European soccer statistics website.

B Sporting talent and wage bill

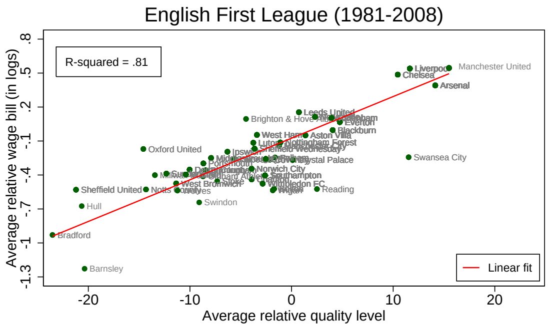
In order to check whether sporting talent is well approximated by wage bill, we created a team specific talent index that sums the talent of each team's player. This index is computed for each club based on the experience of their players (i.e., their total number of matches played in the Premier League), the age of their players, and whether they are selected by their national team, and at what level (youth team, B team, A team). This variable is a crude measure for the team's sporting talent, but gives a rather sensible proxy for the quality of the team. As an example, the best performers, identified in Figure (3), which are Arsenal, Chelsea, Liverpool and Manchester United, are also the best teams in terms of quality with the highest wage bills. A

[37]Our full coding protocol is available upon request.

[38]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 appearance. 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.

regression of the average team's quality on average wage bill gives a R-squared of 0.81. The wage bill does control quite accurately for the team's sporting talent.

Figure 8: Average Wage Bill and Team's Quality



Notes: Average relative wage bill is the average club's relative wage bill from 1981 to 2008. The relative wage bill of a club *i* in year *t* is the log difference of the club's wage bill relative to the annual average. Average relative quality level is the average club's quality from 1981 to 2008. The club's relative quality level is computed as the difference of the club's quality level relative to the annual average.

C Proofs and derivations

C.1 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 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:

$$\lambda_0 u = \delta(1-u) \Rightarrow u = \frac{\delta}{\lambda_0 + \delta}. \tag{12}$$

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

$$I(\delta + \lambda_1\gamma) = \lambda_0(1-\gamma)u \Rightarrow I = \frac{\lambda_0(1-\gamma)\delta}{(\delta+\lambda_1\gamma)(\delta+\lambda_0)}. \tag{13}$$

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

$$\delta P = \lambda_1\gamma I + \lambda_0\gamma u \Rightarrow P = \frac{\lambda_1\gamma\lambda_0(1-\gamma)}{(\delta+\lambda_1\gamma)(\delta+\lambda_0)} + \frac{\lambda_0\gamma}{(\lambda_0+\delta)}. \tag{14}$$

C.2 A formula for wage discrimination

In this appendix we detail how we obtain the formula for wage discrimination (9). Let define the wage discrimination as the difference in expected wages between individuals in groups A and B, such as:

$$D = \mathbb{E}_A(\omega|t) - \mathbb{E}_B(\omega|t) = bu_A + ktI_A + tP_A - bu_B - ktI_B - tP_B. \tag{15}$$

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 (9)

$$D = \mathbb{E}_A(\omega|t) - \mathbb{E}_B(\omega|t) = t\lambda_0\delta(\gamma_A - \gamma_B)\left(\frac{(1-k)(\lambda_1+\delta)}{(\delta+\lambda_1\gamma_A)(\delta+\lambda_1\gamma_B)(\delta+\lambda_0)}\right).$$

C.3 Proof of proposition 1

Proposition 1: *The race-based difference in expected wage is a parabola that has a single maximum. Wage discrimination first increases and then decreases with constraints on job-to-job mobility.*

Proof. We start from equation (9), $D = \mathbb{E}_A(\omega|t) - \mathbb{E}_B(\omega|t)$, derived in appendix (C.2) and establishing the difference in expected wages between A and B. We look at the maximum and minimum of D by taking its derivative with respect to λ_1:

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

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

Since, $\gamma_A \geq \gamma_B$, the sign of 16, depends on the sign of the numerator of the expression in parentheses:

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

It follows that to understand the sign of the derivative of (9), we need to study the variation of:

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

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

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

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, 17 is positive and equal to $\delta^3\alpha$. Then, for higher values of λ_1 (e.g. $\lambda_1 = 1$), 17 is negative. We can thus 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. □

C.4 Proof of proposition 2

Proposition 2: *A decrease in constraints on job-to-job mobility causes a racially differentiated change in job turnover.*

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 $\frac{\gamma_A}{\gamma_B} > \frac{1-\gamma_B}{1-\gamma_A}$. $\square$

D Parameters for the simulations

Even though our labor market is very different from the one in Bowlus and Eckstein (2002), we use the same values for the common parameters λ_0 and δ (see Table 8). Our specification imposes $0 < \alpha < 1$. This implies that the expectation of c is not defined. However, 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	Parameters	Values
k	0.5	α_W	0.8
λ_0	0.04	α_B	0.8
δ	0.004	$\tilde{c}_W$	0.05
t	150	$\tilde{c}_B$	0.04
r	0.05		

Note: λ_0 and δ are estimates from Bowlus and Eckstein (2002).

E Further evidence of discrimination

The age profile of black players before and after the Bosman ruling presents also some evidence of discrimination. Figures (9) and (10) depict the age density of black players in "discriminating" and "non-discriminating" teams before and after the Bosman ruling, respectively. We consider discriminating teams as the ones whose proportions of black players in the squad is lower than 25% of the other squads, and non-discriminating teams as those where the proportions of black players is higher than 75% of the other squads. Of course, this way of selecting discriminating and non-discriminating teams is far from perfect, but it fis consistent with our empirical strategy and model.

In Figure (9), we observe that the age densities of black players in pre-Bosman are quite similar in discriminating and non-discriminating clubs, but there is a huge change post-Bosman (Figure 10): the age density of black players is more right-skewed in discriminating clubs. These differences could be explained by job-to-job mobility constraints. When players are young, they tend to play for their local clubs or for any club that wants them, whether these discriminate or not. Then, when labor mobility is low, as in pre-Bosman, black players employed in discriminating clubs can not

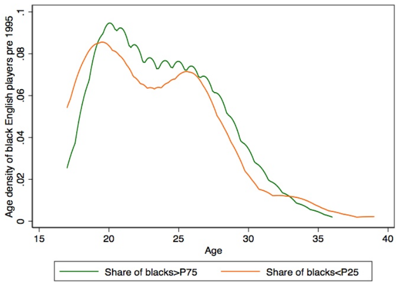


Figure 9: Pre-Bosman

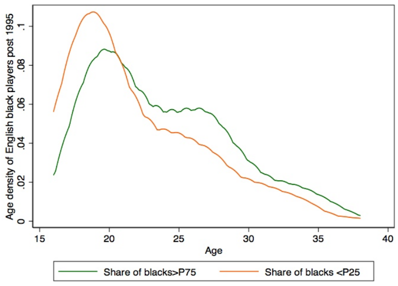


Figure 10: Post-Bosman

move. This explains the similar age profiles depicted in Figure (9) between discriminating and non-discriminating clubs. By contrast, when job-to-job mobility is high, as in post-Bosman, black players in discriminating clubs are mostly youngsters since older players leave as soon as they can. This explains why the age density of black players is more right-skewed in discriminating clubs in the post-Bosman era (Figure 10).

F 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)
Observations	259	268
Club fixed effects	Yes	Yes
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 at the club level, with ^{a}, ^{b} and ^{c} denoting significance at the 1%, 5% and 10% level respectively.

Table 10: Market-test: Match Performance and Discrimination, long span - 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^a	-0.001
	(0.002)	(0.002)
Difference in lagged stadium attendance	0.011^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	0.271^a
	(0.032)	(0.044)
Observations	3129	3264
Pair of clubs fixed effects	Yes	Yes
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	0.461	3.034
$\chi^2(2)$ p-value	0.79	0.22

Notes: robust standard errors in parentheses are two-way clustered at the visiting and the receiving club level, with [a] and [b] denoting significance at the 1% and 5% level respectively.