

Labor Mobility and Racial Discrimination

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Keywords: Labor mobility, labor search, wage discrimination

JEL Classification: J15, J31, J6, J71

Mobilité du travail et discrimination raciale

Résumé: Dans cet article, nous étudions les effets des contraintes de mobilité du travail sur la discrimination raciale. Notre modèle de recherche d'emploi souligne que ces effets suivent une courbe en U inversé. En particulier, lorsque les contraintes de mobilité sont faibles, nous montrons que la discrimination salariale disparaît. Nous testons cette prédiction grâce à un choc exogène sur le marché du travail des footballeurs, appelé "arrêt Bosman". Cet arrêt, rendu en 1995 par la Cour de justice des Communautés européennes, a réduit les contraintes qui pesaient sur la mobilité des joueurs de football. À partir de données sur les clubs de première division anglaise, entre 1981 et 2008, nous comparons le marché avant et après l'arrêt Bosman. Si nos résultats empiriques montrent que la discrimination salariale existait avant Bosman, elle disparaît après Bosman, avec la réduction des contraintes sur la mobilité des travailleurs.

Labor Mobility and Racial Discrimination*

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Abstract

This paper assesses the effects of labor mobility constraints on racial discrimination. Our equilibrium search model shows that these effects follow an inverted U-shaped relationship. In particular, when mobility constraints are low, we find that discrimination disappears. We test 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. Consistent with a taste-based explanation, we find evidence that wage discrimination disappears 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 and evidence suggests that these differences are partly the consequence of racial discrimination.¹ “Why do people discriminate against one another, and how can we get them to stop?” (Gneezy and List, 2013, 5). Job-to-job mobility is a promising route to answer these questions. Constraints on mobility, such as quotas, work permits, or restrictive contracting rules on national labor, may limit the ability of workers to move from prejudiced firms to unprejudiced ones. When mobility is constrained, a firm is able to act on its prejudice because of the low cost of doing so.

In this paper, we assess the effects of job-to-job mobility constraints on racial discrimination. Our equilibrium search model reveals that these effects follow an inverted U-shaped relationship. In particular, when mobility constraints are low, we find that discrimination disappears. We test 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. Consistent with a taste-based explanation à la Becker (1957), we find evidence that wage discrimination disappears when constraints on worker mobility are lowered.

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

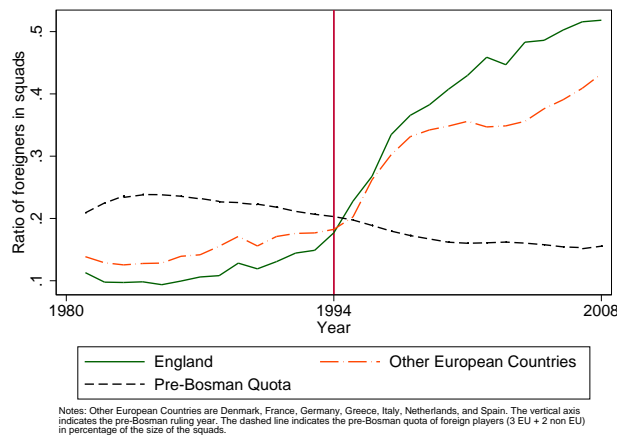
¹See Altonji and Blank (1999), Lang and Lehmann (2012) and Charles and Guryan (2011) for reviews of the literature.

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

(EU) nationals and the obligation to pay a fee for out-of-contract players.³ Figure (1) illustrates the intensified mobility of the soccer market in the wake of the Bosman ruling. Before 1995, the quota of foreign players in England, represented by the dashed grey line, was binding but not fulfilled. In 2008, the ratio of foreigners in squads exceeded 50% in the English first league. 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.

Second, extensive data on the career paths of professional soccer players can be gathered for most countries over long time periods.⁴ Third, we can match this extensive individual data with information on the skin-color of players.

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



Fourth, the soccer market offers a simple test for racial discrimination in salary setting (Szymanski, 2000). Assessing whether a group of workers is facing wage discrimination is typically difficult because we may not be able to

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

⁴Kleven, Landais, and Saez (2013) compiled data of all first-league soccer players for 14 Western European countries since the eighties.

capture productivity exactly.⁵ With this caveat in mind, Szymanski’s ‘market test’ is particularly elegant and parsimonious. Under the assumption that soccer is an efficient market,⁶ 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 clubs fielding an above-average proportion of black players systematically outperform clubs with a below-average proportion of black players. This implies that, for a given wage bill, a team can improve its performance by employing a higher share of black players and *in fine* that black players are being paid less than their talent would warrant. Szymanski finds evidence of discrimination while performing his test on a panel dataset of professional English clubs between 1978 and 1993, i.e., *before* the Bosman ruling. We extend the analysis one step further. First, we theoretically demonstrate why discrimination can survive in equilibrium with labor mobility constraints. Second, we exploit the Bosman ruling shock and provide empirical evidence that discrimination disappears with low mobility constraints. We briefly present these two contributions.

The first contribution consists of establishing a simple theoretical model to guide our empirical analysis. To derive equilibria in which group differentials persist, we merge ideas from search models of the labor market, à la Burdett and Mortensen (1998), with a Becker-style assumption of taste for discrimination.⁷ 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 job offer if the wage offered exceeds his

⁵Using a Mincer-type wage equation, for instance, with a dummy indicating whether a person belongs to a particular group is plagued by a common estimation problem: how do we account for unobserved variables? (See Charles and Guryan, 2011).

⁶Unlike 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.

⁷Altonji 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. There is thus little uncertainty over a player’s productivity.

reservation wage and searches on the job.⁸ Hence, the firm faces a job offer trade-off between diverting a share of the player's talent and seeing the player potentially poached by a rival, which induces a costly turnover.⁹

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

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

⁸In 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.

⁹The same mechanism can be found in Stiglitz (1974), where firms in urban areas increase wages to reduce costly turnovers.

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

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

aration rate. Prejudiced firms have a lower disutility in parting with disliked workers. Our results are consistent with Bowlus and Eckstein (2002) and race-based wage differentials. However, as in simpler models, such as Becker (1957), we predict that discrimination is eliminated through intensified labor mobility, i.e., increased competition to attract talent.

The second contribution consists of using the Bosman ruling on the English soccer market to estimate the causal effect of intensified labor mobility on discrimination. Remember that (1) players whose contracts have expired can now change clubs freely and (2) EU players can now move within the EU without taking up a valuable space in a foreign team’s quota. These two effects make it easier to “poach” employees and influence firms’ monopsony power. Interpreting the Bosman ruling as a shock on job-to-job mobility constraints, we find empirical evidence that wage discrimination disappeared in the post-Bosman period.¹²

The difference between the pre- and post-Bosman eras is robust to the use of various (1) estimators (OLS, within, IV and GMM), (2) samples (either league or match performance), and (3) groups of players (either black or white and either English or non-English). In particular, we find that black non-EU players still face some wage discrimination in the post-Bosman era due to higher mobility constraints relative to EU players. Moreover, the decrease in discrimination appears to be extremely fast: we find evidence of discrimination in the 5 years before the ruling, but no effect in the following 5 years.

Our results could be important for public policy. If we consider restrictive contracts to be an important component of the typically nebulous “labor

¹²In a fascinating book documenting how soccer can help economics, Palacios-Huerta (2014) reviews Szymanski’s paper. Concomitant to our work, Palacios-Huerta extends Szymanski’s sample from 1993 to 2008 and confirms the absence of discrimination in this period. His intuition is that the emergence of a market for corporate control of English professional clubs since the early 1980s has increased the competitiveness of English soccer. The added competition would have been able to drive discriminating firms out of the market. This effect may complement our mobility mechanism. A potential issue is, however, that a club experiencing very poor performance simply moves down from one division to another and is thus not driven out of the soccer market. Moreover, some firms operate under a soft budget constraint (Andreff, 2014). Our hypothesis is that the labor market became more competitive primarily because of relaxed constraints on mobility.

market frictions”, facilitating job changes and reducing frictions could lower discrimination.

A potential objection to our findings is that what we are measuring may not truly be the effect of the Bosman ruling but instead a change in attitudes toward racism and racial discrimination. Lang and Lehmann (2012) document a smooth decline in prejudiced views on school segregation, social interaction, and blacks in politics, as measured by national polls and surveys in the United States. The data reject, however, the idea of a dramatic reduction in discrimination over our period of analysis. Sport is not immune to racial bias, which exists on 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:¹³ in 2011, the English captain John Terry was accused of racial abuse, and in 2012, Marvin Sordell was racially abused by a Millwall fan. In August 2014, it came to light that the Cardiff City manager Malky Mackay shared racist e-mails and texts with the director of soccer in charge of transfers, Iain Moody.¹⁴ Frequent 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 of prejudiced employers in 1995.

Our result regarding intensified mobility is consistent with Becker’s argument that intensified product market competition can reduce race-based differentials caused by prejudice. Using regulatory reforms in U.S banking, Levine, Levkov, and Rubinstein (2008) show that the exogenous intensification of competition among non-financial firms has reduced the manifestation of racial prejudices in the demand for labor and has raised the wages for

¹³In the web appendix, we show strong evidence on racism in the English soccer in the post-Bosman era.

¹⁴For instance, the Daily Mail report 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.” In November 2014, the owner of Wigan Athletic FC, Dave Whelan stated while defending his decision to hire Malky Mackay that, “I think Jewish people do chase money more than everybody else.”

black workers to approach those of equally productive white workers.

Our paper is also related to more recent works examining the influence of labor market tightness on discrimination (Biddle and Hamermesh, 2013, Baert *et al*, 2014). These studies emphasize that employers discriminate less in labor markets with a small number of job seekers relative to vacancies. By offering more job opportunities, our mobility shock can be interpreted as a decrease in the ratio of employed job seekers to job offers, as there will be more competition between firms to attract workers. This situation should discourage employers from indulging in discriminatory tendencies.¹⁵

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 the identification strategy and the specifications of the market test for discrimination. Our empirical results on discrimination in the English soccer league are presented in section (5). The most important result is that discrimination disappears after the Bosman Ruling. Section (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 mobility and discrimination: (1) large observed variation in labor mobility constraints following the Bosman ruling; (2) an extensive collection of data on the career paths of professional soccer players, (3) data on the skin-color of individuals that we can match with our dataset, and (4) a simple test for racial discrimination in salary setting (Szymanski, 2000). Using a panel of all clubs in the English first division from 1981 to 2008, we now present three important features related to the competitiveness of the soccer market.

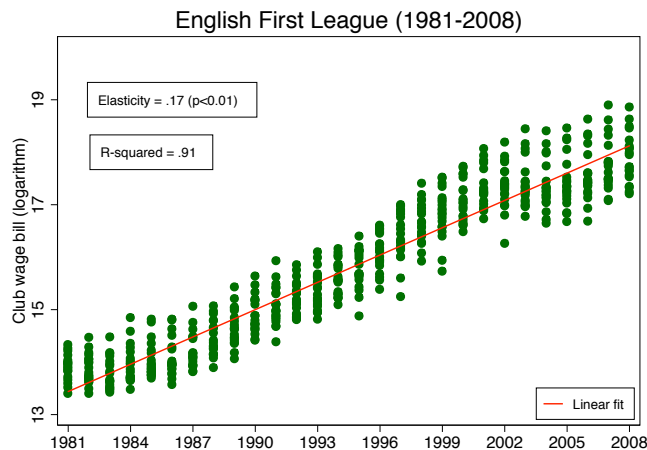
¹⁵The Bosman ruling shock may have also increased competition among workers by offering firms a larger pool of job seekers. However, to the extent that firms pay a turnover cost when workers quit, their monopsony power is reduced by the increased mobility of the labor market.

Our dataset is similar to Szymanski (2000) but differs in scope and time.¹⁶ In terms of scope, Szymanski uses a panel of 39 clubs from four divisions in the English soccer league over the 1978-1993 period. Thanks to Kleven *et al.* 2013, our dataset contains all professional soccer players, regardless of nationality, from the first league in England.¹⁷ In terms of time, we extend Szymanski’s sample to cover the post-Bosman era, from 1994 to 2008. Despite differences in scope and time, the most salient features remain.

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

Second, clubs are heterogeneous in wage bills, as shown in Figure (2). This figure reports the log of the clubs’ wage bills in the English first league from 1981 to 2008. It is worth mentioning that wage bills are linearly increasing over the years for all clubs, without being affected by the Bosman ruling.

Figure 2: Club’s Wage Bills

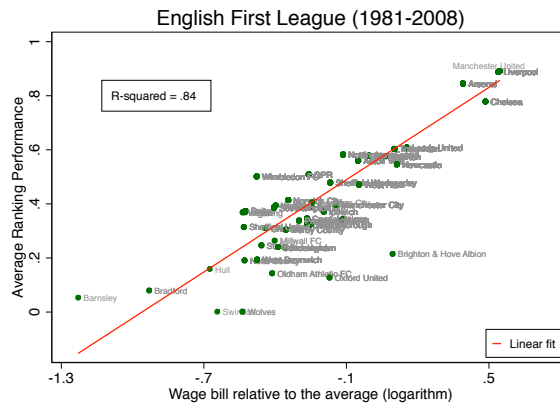


¹⁶See Appendix (A) for further information on variable definitions and data sources.

¹⁷We only have access to information on professional players from the top league, but our number of clubs is fairly similar to that in Szymanski’s sample: 41 teams from 1981 to 1993.

Third, soccer is a competitive market for talent. A common explanation for the increasing trend depicted in Figure (2) is the increasing price of talent. Players earn wages proportional to their talent. Thus, the higher the club’s sporting talent, the higher the wage bill,¹⁸ and the higher the ranking. Using clubs’ wage bill data and computing an index of performance based on league rankings (see 4.4 for detailed computation), we confirm that the English soccer market is fairly competitive. Wage expenditures and performance are heavily correlated in the English league between 1981 and 2008. This correlation is depicted in Figure (3) and holds despite the surge in player remuneration (see Figure 2).

Figure 3: Average Wage Bill and Performance



3 Theoretical framework

We establish an equilibrium labor market search model in the style of Burdett and Mortensen (1998), in which firms are homogeneous in size because the number of players per team is fairly rigid. Firms are instead heterogeneous in their wage bills (see section 2). We assume that a club’s budget is given and is spent searching for talent and paying wages. In this search setting, we introduce racial discrimination. We assume that some employers hold a ‘taste’ for

¹⁸In Appendix (B), we construct a crude a measure of the sporting quality of the team, as a proxy for its sporting talent. Figure (8) depicts a clear linear correlation between wage bill and team’s quality, with an r-squared of 0.81.

racial discrimination. 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) 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 mobility, as was the case before the Bosman ruling, may explain race-based wage differentials. This result is formalized below.

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

3.1 Workers

The mass L of workers is divided into two types according to their appearance, A and B . Type B could be discriminated against by employers. All workers are heterogeneous in talent independent of their type, and each type has the same distribution of talent. The decision problem faced by a worker in a traditional job search model is simple (see e.g. Burdett and Mortensen, 1998); he maximizes utility over an infinite horizon in continuous time by adopting a reservation wage strategy that is state dependent. At any moment in time, each worker is either unemployed (state 0) or employed (state 1). Firms are engaged in search, and, at random time intervals, workers receive information about new or alternative jobs. This information is encapsulated in the parameter of the Poisson arrival process, λ , which denotes the arrival rate of job offers. This parameter reflects the general state of the labor market, including contracting rules, institutional constraints on 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.¹⁹ 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 (see section 2), the worker searches on the job. However, the probability of receiving a perfect job offer, γ , depends on the state of the labor market and on firm behavior (described below).

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

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

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 em-

¹⁹Evidence of monopsony power in sport is documented in Kahn (2000).

ployment in an imperfect match, W_I , for a given value of k rather than a continuous distribution as follows:

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

and for a perfect match as

$$rW_P = t + \delta(U - W_P). \quad (3)$$

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.²⁰

3.2 Flow conditions

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

²⁰Given the high wages of soccer players, this assumption does not seem to be outlandish.

(ii) and perfect (iii) jobs:

i. Unemployment flows have to balance:

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

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

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

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

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

3.3 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. At any moment in time, we assume, for the sake of simplicity, that total resources are exogenously determined by sponsorship, past performances, and attendance in the stadium, among others. 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 $c(t)$. 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 :

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

$$J_P = \delta(V - J_P - c(t)) \Leftrightarrow J_P = \frac{\delta(V - c(t))}{r + \delta}. \quad (8)$$

We assume that when a worker leaves the firm, a vacancy is created, with V the value of a vacancy. However, 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 (7) 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, the firm will offer a perfect job offer if:²¹

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

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

²¹We 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: $V = 0$.

of turnover cost draws given by

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

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,²² such that $c(t) = ct^2$, yields a simple closed-form solution for γ . The probability of receiving a perfect offer is such that $J_P > J_I$:

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

As shown in equation (10), 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. For a high value of λ_1 , 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 decreases the number of employees that are stuck in imperfect matches (see equation 5), and reduces the probability of being in an imperfect match by increasing γ .

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 giving an imperfect offer to a worker is less attractive. Second, the higher is t , the more costly it is to replace talented workers, and the higher is the probability to offer a perfect job.

²²The 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 (Andreff (2014)).

3.4 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 (λ s) 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 λ s 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 . In our approach, firms do not segregate type B versus type A . 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.²³

The λ parameters reflect the general state of the labor market, and thus, the potential constraints on labor mobility. Holding everything else constant, lower constraints are associated with a higher arrival rate of job offers. Therefore, the end of both transfer fees for out-of-contract players and 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 states of the labor market.²⁴

In our approach, the prejudice applies to the separation rate. We assume that the two groups of workers, A and B , differ in their probability of drawing a perfect match, γ . They have different Pareto distributions governing the

²³Because talent and productivity are often Pareto distributed, adopting the same approach would give us a more plausible right-skewed distribution for wages.

²⁴We 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 .

turnover cost. In particular, we consider that $\tilde{c}_B = \tilde{c}_A - d$, where d represents the taste for discrimination and \tilde{c} represents the lower bound cost, such that the Pareto distribution of the turnover cost for B workers is shifted to the left. In our labor context, this shift to the left reflects the taste for discrimination d and the idea that a type B player who is unlikely to be a fan or manager favorite will be less costly to let go or easier to replace. This assumption implies a lower equilibrium probability of receiving a perfect job offer (γ) for workers in group B than in group A.

From this 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 an $i = A, B$ type player. Consider mobility constraints (i.e., low λ_1) and two teams with different shares μ of black players but with same resources. Because γ_B is lower, they will have a lower wage bill for the same talent. Thus, firms that hire more black players increase their performance compared with those that do not, which is exactly what Szymanski's market test aims to estimate.

3.5 Wage discrimination

Our perspective is that the Bosman ruling modifies the general state of the labor market. It lowers the constraints on labor mobility and thus affects the arrival rate of job offers for employed workers (λ_1). We thus consider the effect of different values of λ_1 on wage discrimination. We define wage discrimination as the difference in expected wages between individuals in groups A and B, such as:

$$\mathbb{E}_A(\omega|t) - \mathbb{E}_B(\omega|t) = bu_A + ktI_A + tP_A - bu_B - ktI_B - tP_B. \quad (11)$$

After some algebraic manipulation (reported in appendix C.1) and given

that both groups have the same likelihood of being out of work, equation (11) reduces to

$$\mathbb{E}_A(w|t) - \mathbb{E}_B(w|t) = t\lambda_0\delta(\gamma_A - \gamma_B) \left(\frac{(1-k)(\lambda_1 + \delta)}{(\delta + \lambda_1\gamma_A)(\delta + \lambda_1\gamma_B)(\delta + \lambda_0)} \right). \quad (12)$$

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

How does this 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 5), and increases the likelihood of receiving a perfect offer, γ (equation 10).

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 players cannot be poached and both A and B players receive imperfect wage offers. In this extreme case, there is no discrimination because A and B players with the same talent earn the same wage. As λ_1 increases, discrimination begins to *increase*. Because all firms endogenize the taste for discrimination (i.e., $\gamma_A \geq \gamma_B$), A players receive more perfect job offers than B players, and wage discrimination increases for a given level of talent. Then, as mobility constraints continues to decrease, the process is reversed. B players begin receiving a higher number of perfect job offers. As a result, the 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 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.2) 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 Interestingly, they are both increasing with the job offer arrival rate. 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).

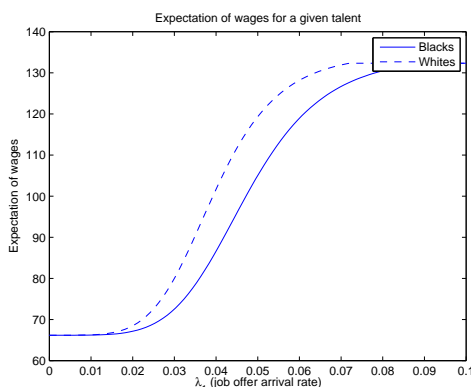


Figure 4: Wage Expectation

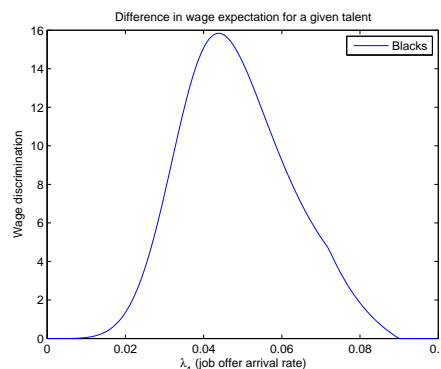


Figure 5: Wage discrimination

3.6 Job turnover

Our model also offers strong predictions regarding the turnover of type A and B workers, which explains the wage discrimination pattern. There exists an optimal job offer arrival rate (established in Appendix C.3), such that type A workers change firms more often than B workers when below this rate, and less often when above it. This optimal rate corresponds to the turning point of wage discrimination identified in Figure (5).

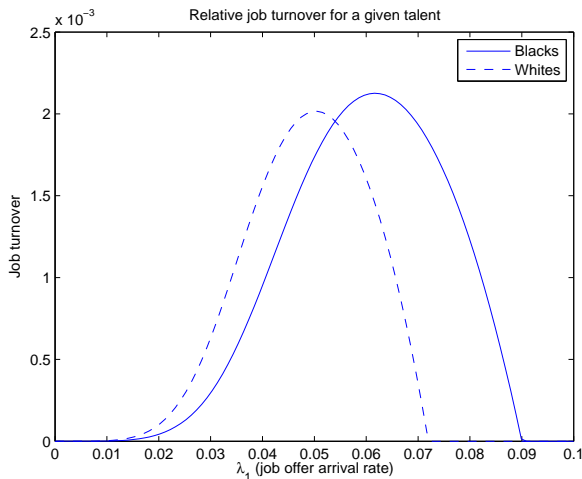
The expression for job turnover in a given period is $\lambda_1\gamma I$. For each group, A (whites) and B (blacks), this expression yields an inverse U-shaped curve

²⁵Although 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).

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

linking λ_1 and job turnover. These curves are simulated and depicted in Figure (6).

Figure 6: Job turnover



There are two reasons that job turnover is low at both extremes of these curves. Consider the case where λ_1 is low. First, workers already employed have a small likelihood of moving. Second, when λ_1 is low, the probability of receiving a perfect 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 new perfect job offers. However, as λ_1 increases, a “stock” effect comes into play: workers are less likely to move simply because those who are employed in an imperfect job become less numerous. Thus, two effects explain why job turnover follows an inverse U-shaped curve. As λ_1 increases, the pool of workers who want to change clubs (I) decreases, as does the number of moves ($\lambda I \gamma$). The differentiated turnover between A and B derives from the differences in γ s: the “stock” effect emerges earlier for A than for type B workers. 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.3) presents the proof. □

4 Empirical design

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

4.1 Model predictions

Our theory predicts an inverted U-shape relationship between race-based wage differentials 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. The pre- and post-Bosman period were characterized by high and low constraints on mobility, respectively. We expect the post-Bosman constraints to be sufficiently low to eliminate pre-Bosman wage discrimination. This is all the more possible given that European players may now receive job offers from all the EU clubs. The post-Bosman period is thus characterized by a much higher λ_1 , i.e., a higher arrival rate of job offers for employed workers.

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, wage discrimination decreases through two different channels. First, workers are more likely to receive “perfect” job offers. Second, the monopsony power of firms decreases because their employees are more likely to be poached.

Another interesting aspect of our model is that it predicts a racially differentiated effect on job turnover as constraints on job-to-job mobility decrease (see Figure 6).

4.2 The market test for discrimination

Our empirical analysis is based on 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). Szymanski's market test is perfectly compatible with our theoretical framework (see subsection 3.4). This test requires information on the skin-color of players that we can match with extensive individual data, as well as data on the wage bill of clubs. 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 top league from 1981 to 2008, and we explore the Bosman ruling as an exogenous mobility shock to the European soccer labor market. The Bosman ruling was decided on December 15, 1995, by the European Court of Justice. This important decision lifted restrictions on soccer player mobility based on the European Community Treaty of the free movement of labor (article 39). This decision had a profound effect on transfers in the European soccer market by banning restrictions on 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 because this case had been submitted to the Court on October 6, 1993. Thus, in December 1993, the European Union of Football Associations amended the regulations governing the *Status and Transfer of Football Players*. This amendment provided that a player may enter into a

contract with a new club when the contract between him and his club has expired, has been rescinded or will expire within six months. However, the two clubs were still forced to agree on a transfer fee with a specific action in case of disagreement. To prevent any contamination of the results caused by a possible anticipation, we omit the 1994-1995 and 1995-1996 seasons. We thus compare the pre-Bosman era (1981-1993) to the post-Bosman era (1996-2008) to identify the causal effect of intensified mobility on racial discrimination.

Exploiting this shock and using our extensive player data, we constructed some informative descriptive statistics (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; 70% of those players are English.²⁷ This number was higher before the Bosman ruling (93.2%) than after (52.6%) it. 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 1995 to 15% after the ruling. The consequence is that the number of white English players decreased from 83.3% before the ruling to 41.3% after the ruling.

Our strategy amounts to comparing black English to white English players 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 players did not play more matches than white players. By contrast, 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 is statistically insignificant and black players are now *less* tenured than white players. Interestingly, all the differences in differences are significant between the two periods. In the post-Bosman era, players on average played less matches, are qualitatively better, are more than 6 months older and have one-half semester more of tenure. Those statistics

²⁷We 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 together with English players makes sense in our context because these players were not considered foreigners in the English soccer market and were thus not subject to the foreign quotas.

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

Variable	Pre-Bosman (1981-1993)			Post-Bosman (1996-2008)			Overall D. in D.
	Black	White	Diff.	Black	White	Diff.	
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 budget, transfer fee records and average attendance are highly significant both economically and statistically. The only exception is the share of black English players, which is constant in both eras.

Table 2: Club differences in means

Variable	Pre-Bosman (1981-1993)	Post-Bosman (1996-2008)	Difference
Budget (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)
Average 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.

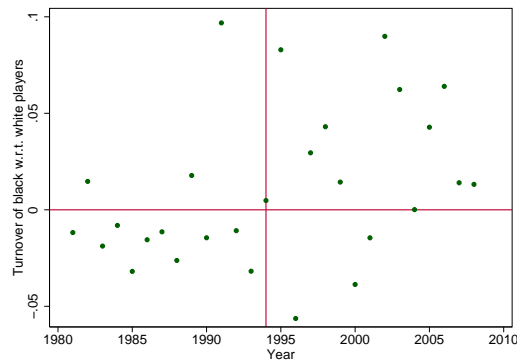
As mentioned above, we should also expect a racially differentiated change in job turnover to accompany a change in the job offer arrival rate. In our empirical context, we define job turnover as club transfers, i.e., moving from

one club to another during a given season. Figure (7) contrasts the turnover of black (B) and white (A) English players by comparing their share in the total number of transfers with respect to their share in the total population. The variable analyzed is the following:

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

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.

Figure 7: Relative turnover of black English players



As shown in Figure (7), before the Bosman ruling, white players tended to change clubs more often than black players, but this tendency was reversed after Bosman.²⁸ This reversed trend suggests that black players “voted with their feet” when the market was liberalized. 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.

²⁸It is unclear, however, when precisely this trend reversed. This may be because we imperfectly observe club transfers. We have data on the first division only, so what we are measuring are transfers within the first division. This may explain our outlying data point for the season 1991-1992, a season where four clubs were promoted and whose transfers were not perfectly captured.

4.4 Estimated equations

To apply Szymanski’s 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 Szymanski’s finding of apparent wage discrimination in the *pre*-Bosman era. We find, however, that wage discrimination has disappeared in the *post*-Bosman era. This is consistent with our theoretical predictions on the effect of relaxed mobility constraints.

The league performance specification

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

League Performance_{it}, is computed in relative terms as $\left(\frac{\text{Ranking}_{it} - \text{min}_t}{\text{max}_t - \text{min}_t}\right)$, 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 team *i* fixed effect, capturing permanent team-specific characteristics affecting performance, such as a location effect,²⁹ and ϵ_{it} is the usual error term. The team’s wage bill ($\text{WageBill}_{it} - \overline{\text{WageBill}}_t$) is measured as the log difference of the club wage bill relative to the annual average ($\overline{\text{WageBill}}_t$). The relative number of players ($\text{PlayerNb}_{it} - \overline{\text{PlayerNb}}_t$) is computed as the difference between the number of players used in a season *t* and a club *i* relative to the average. Controlling for a club fixed effect and time-varying wage bill, the relative number of players variable 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 in a given season *t* relative to

²⁹Some 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 annual average ($\overline{\text{Shareblack}_t}$). We compute this ratio based only on the share of black English players (including Irish, Scottish and Welsh players - see footnote 27). The share of black English players is relatively stable over time, but the share of black non-English players is constantly rising, which would lead us to estimate discrimination toward different individuals before and after Bosman. We later discuss results based on both black and white non-English players.

The coefficient of interest to us is β_3 , the effect of the share of black English players on performance. In case of race based wage discrimination, we expect β_3 to be positive.³⁰

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:

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

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

³⁰Compared with Szymanski, our transformation of the dependent variable implies, more intuitively, that in case of discrimination, an increase in the relative share of black players raises team's performance.

clubs, $\log(\text{WageBill}_{it}/\text{WageBill}_{jt})$. 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 The results are presented below.

5 Empirical results

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

5.1 Discrimination market test on league performance

We first use the ordinary least squares (OLS) estimator and the fixed-effects “within” estimator to eliminate the individual effect (α_i), with standard errors robust to 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 13) 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 effect. The relative number of players used exhibits a negative effect in line with “bad luck” because high turnover typically reflects a high level of injuries sus-

³¹Unfortunately, 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.

tained. 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 apparent discrimination. 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			
	<i>Pre-Bosman</i> (1981-1993)		<i>Post-Bosman</i> (1996-2008)	
	OLS	Within	OLS	Within
Sample:	(1)	(2)	(3)	(4)
Relative log wage bill	0.509 ^a (0.05)	0.396 ^a (0.12)	0.474 ^a (0.04)	0.157 ^b (0.08)
Share of black English players employed	0.577 ^a (0.18)	0.491 ^b (0.25)	-0.130 (0.15)	-0.000 (0.20)
Relative number of players used	-0.025 ^a (0.01)	-0.031 ^a (0.01)	-0.026 ^a (0.01)	-0.027 ^a (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 attendance, and relative record transfer fees.³² The key identifying assumption is that these variables do not affect league performance, except through their effect on the wage bill.

We briefly discuss the relevance of our three excluded instruments before presenting the results. First, good performance in cups in the previous season generates higher revenues in the contemporaneous season. Then, there is a strong link between the previous season attendance in the stadium and the current season resources and wage bill. Finally, we use record transfer fees to capture potential buyouts by rich owners. New owners often break the club's transfer fee record, but these purchases often have 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 the budget.

The instruments are constructed as follows. Relative lagged attendance is measured as the one-year lag of a club's attendance relative to the annual average. Performance in cups is measured as the one-year 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

³²Data are constructed from newspaper articles and online sources.

effect. Whereas the coefficient for the share of black players employed is positive and significant before the Bosman ruling, it is insignificant post-Bosman.

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 (0.25)	0.576 ^a (0.10)	0.200 ^c (0.12)	0.472 ^a (0.04)
Share of black English players employed	0.501 ^b (0.25)	1.364 ^b (0.54)	-0.111 (0.18)	-0.116 (0.48)
Number of players used	-0.030 ^a (0.01)	-0.036 ^c (0.02)	-0.027 ^a (0.01)	-0.010 (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.

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 restrictions for the excluded instruments is also passed and the Angrist-Pischke first-stage chi-squared statistics reject the null of underidentification (Angrist and Pischke, 2009).

In columns (2) and (4) of Table (4), we use the two-step generalized method of moments (GMM) approach of Arellano and Bond (1991). This estimator differences away time-invariant club specific effects. It relies on the

dynamic structure of the model for identification by using lagged levels of the independent variables as instruments for current differences. A problem with GMM estimators is that their validity is subject to the use of a relatively small or large number of instruments. A large number generates implausibly low values of Hansen tests of instruments exogeneity (Roodman, 2009), while using too few instruments is likely to generate a weak instruments problem and to deliver inaccurate estimates. Following Roodman's (2009) rule of thumb, the number of instruments is strictly lower than the number of clubs (groups) in the sample. This strikes a balance between estimate consistency and test validity. The diagnostic tests (Hansen and first and second order autocorrelation) presented at the bottom of the table reveal no evidence against the validity of the instruments used by the GMM estimator.

The GMM estimates of the share of black English players employed produce the same result as the other estimators: 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 should be relatively more constrained for non-

³³The 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.

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 coefficients for the share of non-EU black players are significant and positive after Bosman, even if its statistical significance is lower when we introduce club fixed effects (col. 3 and 4). Those coefficients imply that wage discrimination against black non-EU players could be present in the English first league post-Bosman.

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

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

Notes: the dependent variable is computed in relative terms as $\left(\frac{Ranking_{it} - \min_i}{\max_i - \min_i}\right)$, 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 discrimination

³⁴Non-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 few non-English black players playing in England (see section 4.3).

market test on matches (see equation 14). For the sake of simplification, we use our two preferred estimators for this purpose: ‘Within’ and ‘Within-IV’. This approach implies that for all estimations we include a dyadic fixed effect controlling for the pair of clubs involved in the match. We are thus exploiting the time series variation in our panel by measuring the effect of differences in the racial composition of the teams on the difference in performance. We first present results on long time spans 1981-1993 versus 1996-2008, and then, on short time spans 1989-1994 versus 1995-2000.

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 of the difference in the share of black players can be interpreted as follows. Consider two teams with the same wage bill and zero black players. The teams’ expected 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 coefficients of the difference in the share of black players, reported in columns (3) and (4), suggest that wage discrimination has disappeared post-Bosman.

As before, we instrument the wage bill in order to account for possible measurement errors and reverse causality. In columns (2) and (4), we

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

Dependent Variable:	Match Performance			
	<i>Pre</i> -Bosman (1981-1993)		<i>Post</i> -Bosman (1996-2008)	
	Within	Within-IV	Within	Within-IV
Estimator:	(1)	(2)	(3)	(4)
Difference in log wage bill	0.701 ^a (0.18)	1.546 ^a (0.47)	0.491 ^a (0.12)	0.330 (0.27)
Difference in black players employed	0.854 ^b (0.33)	1.079 ^b (0.43)	-0.148 (0.31)	0.039 (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).

instrument the difference in wage bills between the two clubs in a match 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 run our regressions on smaller time windows. In Table (7), we estimate equation (14) using the within estimator.³⁵ We check the extent of discrimination before and after the Bosman ruling by using 5-year time windows with two scenarios: (1) 1989-1993 vs. 1996-2000 and (2) 1990-1994 vs. 1995-1999. In the first scenario, we drop the years

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

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 the change in discrimination is following the Bosman ruling.

In the Bosman ruling period (col. 1 and 2), the estimates of the share of black English players are positive, significant and higher than in the 1981-1993 time frame (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 employed are statistically insignificant in the 5 years following the Bosman ruling (col. 3 and 4). This sharp contrast, 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

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

Dependent Variable:	Match Performance			
Sample:	Pre-Bosman		Post-Bosman	
	(1990-1994)	(1989-1993)	(1995-1999)	(1996-2000)
Estimator:	Within	Within	Within	Within
	(1)	(2)	(3)	(4)
Difference in log wage bill	0.517 ^c (0.29)	0.717 ^b (0.33)	0.432 ^c (0.26)	0.389 ^a (0.13)
Difference in black players employed	1.409 ^a (0.45)	1.203 ^b (0.55)	-0.820 (0.58)	-0.591 (0.70)
Difference in players used	-0.035 ^a (0.01)	-0.035 ^a (0.01)	-0.039 ^a (0.01)	-0.027 ^b (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).

³⁶This 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). An alternative to moving is, however, 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.

6 Conclusion

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

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

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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). 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. Precise 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 or more months of data, 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, this practice is unlikely to be a problem because the pay for most of these employees is

relatively small compared with the total wages of players and 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 these last measures to create an objective, albeit imperfect, measure of quality of the players (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

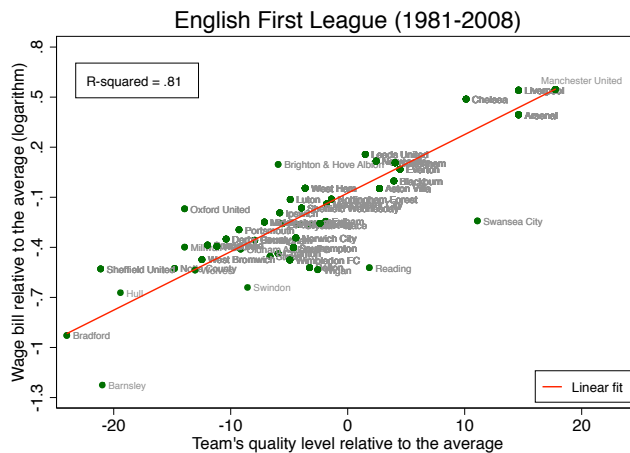
In order to check whether a team's talent is well measured by its wage bill, we create a team index that sums the talent of each team's player. This index is computed for each team based on the total number of matches played in the Premier League by their players, their age, and whether team's players 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,

³⁷Our full coding protocol is available upon request.

³⁸For an explanation of why this appearance-based method is appropriate, Palacios-Huerta (2014) considers the case of the legendary Manchester United player Ryan Giggs. He appears to be Caucasian, and it was unlikely that he faced discrimination as a professional player during his career because discriminators prejudge an individual based on appearances. However, after he became famous, he publicly revealed that he had been victim of racism as a child because of his father's skin color. This revelation came as a surprise to his fans.

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 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



C Proofs and derivations

C.1 A formula for wage discrimination

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

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

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

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

which amounts to the equation (12)

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

C.2 Proof of proposition 1

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

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

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

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

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

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

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, equation 17 is positive and equal to $\delta^3 \alpha$. For higher values of λ_1 (e.g. $\lambda_1 = 1$), this value is then negative. We can then establish that the function 17 is first positive, then negative at some value of λ_1 . It follows that there will be a single value for which the derivative

is equal to 0, and that below this value, equation 16 will be positive, and negative above it. Hence, the difference in expected wage will be a parabola with a single maximum. \square

C.3 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: $DM = \lambda_1\gamma_B I_B - \lambda_1\gamma_A I_A$

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

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

Table 8: Parameter values for the simulations

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

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 also presents some evidence of discrimination. Figures (9) and (10) depict the age density of black players in the squads of “discriminating” and “non-discriminating” teams before and after the Bosman ruling, respectively. We consider discriminating firms as the teams whose proportions of black players in the squad is lower than 75% of the other squads, and non-discriminating firms as those where the proportions of black players is higher than 25% of the other squads. Of course, this way of selecting discriminating and non-discriminating firms is far from perfect, but it fits in with our empirical strategy and our model. In our model, firms that discriminate are more likely to have their black players poached by rival firms. We should therefore expect these firms to have less black players.

In Figure (9), we observe that the age densities of black players pre-Bosman are quite similar in discriminating and non-discriminating clubs, but there is a huge change post-Bosman (Figure 10): the age density of black players is much more left-skewed in discriminating clubs. Why is this interesting? If we consider that mobility was constrained before Bosman, then we can propose the following explanation: when job-to-job mobility is high, players that were employed in discriminating clubs want to leave as soon as they get the chance. When players are young, they tend to play for their local clubs or simply for any club that wants them, whether these discriminate or not. Therefore, in discriminating clubs, the black players we find are mostly young: older players leave as soon as they get the chance. However, as job-to-job mobility was lower before Bosman, black players were less likely to gain from this strategy, hence the similar age profiles between discriminating and non-discriminating clubs.

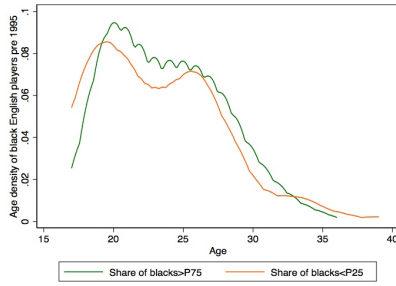


Figure 9: Pre-Bosman

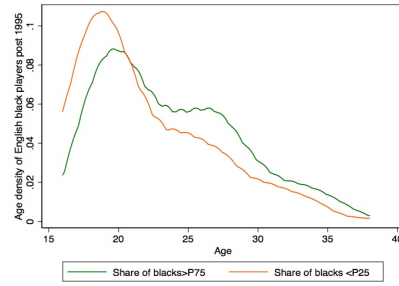


Figure 10: Post-Bosman

F First stage estimations

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

Dependent Variable:	Relative Log Wage Bill	
	Pre-Bosman (1981-1993)	Post-Bosman (1996-2008)
Sample:		
Share of black English players	0.009 (0.140)	-0.003 (0.140)
Number of players used	0.007 ^c (0.004)	-0.002 (0.003)
Lagged log attendances	0.248 ^a (0.008)	0.522 ^a (0.098)
Lagged cup performance	0.011 ^b (0.005)	0.004 (0.004)
Relative record transfer fee	0.125 ^a (0.036)	0.023 ^a (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: Sample:	Difference in Log Wage Bill	
	Pre-Bosman (1981-1993)	Post-Bosman (1996-2008)
Difference in share of black English players	0.035 (0.081)	0.116 (0.123)
Difference in number of players used	0.009 ^a (0.002)	-0.001 (0.002)
Difference in lagged attendances	0.011 ^a (0.003)	0.009 ^a (0.003)
Difference in lagged cup performance	0.009 ^a (0.003)	0.004 ^b (0.002)
Relative record transfer fee	0.088 ^a (0.032)	0.271 ^a (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.