Race, Compensation and Contract Length in the NBA: 2001–2002*

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We study race and pay in the NBA for 2001–2002. For players who were neither free agents nor on rookie scale contracts, there were large, statistically significant *ceteris paribus* nonwhite shortfalls in salary, total compensation, and contract duration. But for players under the rookie salary scale (first-round draft picks) and free agents, race effects were small and insignificant. These results suggest discrimination against marginal nonwhite players.

The National Basketball Association (NBA) is seen by many as an oasis of economic opportunity for highly skilled African American athletes. Roughly 80% of the league's players are black. And of the 42 players who had signed contracts as of the 2001–2002 season with annual salaries of more than $10 million (the top decile of the league's 415 players), 37 (88%) were black. Despite this clear evidence of black success in the NBA, the question of discrimination against African Americans remains a salient one. It has been reported, for example, that black players look with suspicion at specific, marginal white players' generous contracts as possible evidence of discrimination (Platt 2000). The 1998 NBA lockout, involving a league of white team owners and an 80% black union, was seen by many players as a racial confrontation (Shropshire 2000); and sports league punishments imposed on players who commit acts of violence are seen by some as biased against black players (Shropshire 2000). Whether these perceptions are accurate in the individual cases involved, and whether, assuming they are accurate, they represent the pattern and practice of the NBA, the fact that they are held at all is reason enough to explore the issue of race and compensation in the NBA.

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This paper examines the impact of race on compensation in the NBA among the players under contract for the 2001–2002 season. Our principal contributions are as follows. First, unlike prior research, we explicitly account for the differing negotiation rights that different players enjoy. As detailed in succeeding discussion, there are three categories of players with respect to negotiating rights in the NBA: (1) first-round draft picks, who have a specific salary scale laying out salary for each of their first three years in the NBA; (2) players who were not drafted on the first round and who do not have free agency rights; and (3) free agents. Category 2 is the only one over which individual teams have monopsony power, and we explicitly account for these different rights in our empirical analysis. We suggest that discriminatory pay differentials are more likely to be observed for players over which individual teams have market power; if so, then it is essential to examine these groups separately.

Second, unlike other research on race and the NBA, we use data on total contract duration and compensation, as opposed to specific year salary, giving a more complete picture of player compensation than single-year salary does. Moreover, knowing when a player signed his current contract allows us to collect productivity and seniority information that was known at the time the contract was signed. Because free agency rights depend on seniority, if we want to know whether a player was a free agent when his contract was signed, we need to know when in his career the negotiation took place. Our data provide us with this information, again unlike other work on the NBA and salaries. Third, our data cover a relatively recent season, 2001–2002, allowing us to update the results of prior research.

Previous Research on Racial Pay Gaps in the NBA

Using data from the mid-1980s, several authors found apparently discriminatory wage differentials favoring white players in the NBA. Controlling for a variety of performance and market-related variables, there were statistically significant black salary shortfalls of 11 to 25 percent, depending on sample and specification.[1] On the quantity side, two studies obtain mixed findings on discrimination in employment in the NBA. On the one hand, Kahn and Sherer (1988) found no significant racial differences in draft order in the mid 1980s, conditional on college performance. On the other hand, Hoang and Rascher (1999) found that, other things equal, during the 1980–1991

[1] See, for example, Kahn and Sherer (1988), Koch and Vander Hill (1988), Wallace (1988), and Brown, Spiro, and Keenan (1991).

period, white players' careers were about two years longer than blacks' (a 36% effect), suggesting the possibility of retention discrimination.

By the mid-1990s there was evidence of unexplained black salary shortfalls only among the elite players in the NBA. In particular, Hamilton (1997) found for the 1994–1995 season that, all else equal, there were no overall significant racial salary differentials in the NBA.[2] However, to examine discrimination across the distribution of playing talent, he used quantile regressions, and found no significant racial salary differentials at the 10th, 25th, and 50th percentiles, but positive and significant or marginally significant white pay premia at the 75th and 90th percentiles amounting to 0.18 to 0.19 log points. Groothuis and Hill (2004) found no evidence of retention discrimination in the 1989–1999 period, in contrast to Hoang and Rascher's (1999) study of the 1980–1991 years.

Consistent with the salary evidence mentioned previously, there is a variety of evidence from the 1980s suggesting that fans preferred white NBA players. For example, Kahn and Sherer (1988) found that, all else equal, during the 1980–1986 period, each white player generated 5700 to 13,000 additional fans per year. The dollar value of this extra attendance more than made up for the white salary premium. Hoang and Rascher (1999) also examined NBA attendance during the 1980–1991 period, and found that, other things equal, larger values of (a variable defined as percent white on the team/percent white in the city) were significantly positively associated with attendance. This finding is consistent with the idea of customer preferences for white players. And other researchers found a close match between the racial makeup of NBA teams in the 1980s, and of the areas where they were located, again suggesting the importance of customer preferences (Brown, Spiro, and Keenan 1991; Burdekin and Idson 1991; Hoang and Rascher 1999; Bodvarsson and Partridge 2001), although player preferences could also have produced a similar result. While the evidence is not unanimous (see for example, McCormick and Tollison 2001), overall, most studies based on 1980s data find evidence consistent with customer discrimination in the NBA.

In contrast to those in the 1980s, most NBA studies based on data from the 1990s do not find evidence of customer discrimination, although the number of studies is admittedly smaller than those for the 1980s. Dey (1997), for example, found that all else equal, white players added a statistically insignificant and economically relatively unimportant 60 fans apiece per season during the 1987–1993 period. Moreover, Stone and Warren (1999) studied

[2] Other analyses of basketball salaries in the 1990s also failed to find *ceteris paribus* racial salary differentials. See Dey (1997) and Bodvarsson and Brastow (1998).

1993 basketball trading card prices, an indicator of fan preferences, and found no racial differences in prices, controlling for player performance. On the other hand, Kanazawa and Funk (2001) found that, other things equal, more white players had a significantly positive effect on Nielsen ratings of televised NBA games during the 1996–1997 season. But overall, the evidence for customer discrimination in the 1990s seems weaker than for the 1980s. This comparison of the two decades is consistent with the decline in the NBA's overall unexplained white salary premium from the 1980s to the 1990s.

This review of empirical work on discrimination in the NBA has suggested that evidence for discrimination appears weaker for the 1990s than for the 1980s, and our paper will update this information for both compensation and customers into the post-2000 period. Moreover, two recent papers on discrimination in the NBA have added some theoretical insights that we will take into account in our work. First, Bodvarsson and Partridge (2001) build a model in which customer, employer, and coworker discrimination can occur at the same time. Their models imply that market-related variables such as racial composition of the area where a team is located should be considered potentially endogenous when studying individual white and nonwhite players' salaries. We take this factor into account in some analyses by excluding these variables, in effect estimating reduced forms.

Second, McCormick and Tollison (2001) pose an alternative explanation for the *ceteris paribus* racial salary differentials we observed in the 1980s. Specifically, if black players have a less elastic labor supply curve to their team than white players, and if teams have monopsony power over the players, then a simple discriminating monopsony argument implies a black salary shortfall. If white players have closer income-generating substitutes to an NBA career than black players, then this monopsony outcome could result. We should also point out that in a Nash bargaining framework, whites are likely to have higher status quo (i.e. during negotiations) incomes than black players. Thus, a Nash bargaining model between one team and individual players is likely to have the same qualitative predictions as the discriminating monopsony model.

The monopsony model can only apply to players who are not free agents and to those who actually negotiate with their team over salary, and our empirical work pays careful attention to the institutional rules for free agency set out by the collective bargaining agreements in the NBA. Moreover, unlike earlier work, our data source has information on the duration and total compensation in each player's contract, as well as when it was signed. We are also able to match productivity data and free agency rights that pertain to the time the player's contract was signed. In contrast, earlier work has matched productivity data to the current year.

Institutional Setting

In our data, which as noted cover players under contract for the 2001–2002 season, 413 of the 415 NBA players signed in 1995–1996 or later. Thus, virtually everyone was subject to either the 1995 or the 1999 collective bargaining agreement (NBA Collective Bargaining Agreement, September 1995, January 1999). Key features of both of these contracts are free agency rules and the rookie salary scale. Starting with the 1995 contract, all first-round draft picks (29 each year) had their salaries delineated in the agreement for each of their first three seasons. For example, the tenth pick in the 1997 draft was to receive \$1.0762 million in 1997–1998, \$1.2376 million in 1998–1999, and \$1.399 million in 1999–2000.

The 1999 agreement (which began in the 1998–1999 season, ending the lockout) augmented the three-year schedule of salary by giving the team an option to hold the player for a fourth year if a raise of a specified amount was given. And in the fifth year, the player becomes a restricted free agent who can be reserved on his original team as long as the team offers at least a stipulated raise, and matches any outside offer—the so-called right of first refusal. Under the 1999 agreement, for nondrafted rookies, or for rookies drafted in the second of the two draft rounds, there is a right of first refusal in the fourth year. Moreover, in addition to the league minimum, the 1999 agreement established seniority-specific minima for each year under the contract. Thus, first-round draft picks under rookie contracts as of 2001–2002 had salaries set according to the 1999 agreement. Controlling for where one was picked in the draft (an issue we explore in succeeding discussions), there is no room for racial differentiation in pay for this group during their first three years, and little room in the fourth year.

Free agency under either the 1995 or the 1999 agreement generally comes after three years of service in the NBA. Exceptions to this include rookies drafted in the first round of 1998 or later in their fourth year, which as noted is a team option year. All salary offers are potentially influenced by complicated team salary cap arrangements that were in place in the 1995 agreement, and tightened in the 1999 agreement. Further, in both contracts, there was a maximum individual player contract duration (seven years in 1995 and six years in 1999 for players changing teams, and seven years for players staying with their team).

This review of the collective bargaining contracts implies that the salary determination process is likely to differ, depending on one's free agency rights and rookie contract status. First, for rookies drafted on the first round, there is no negotiation over their salary. Second, those with less than three years of service who were not drafted in the first round, or who entered the league before 1995 are only allowed to bargain with their current team. They are the most likely group to be subject to monopsony power. Third, the free agents can negotiate with many teams, and may therefore end up with salaries approaching their marginal productivity. As discussed further below, we take these differences in bargaining rights into account.

Data and Estimating Equations

For information on pay, we use the *USA Today* 2001–2002 NBA salary database, which has information for every NBA player on contract duration, total cash compensation, 2001–2002 salary, and ending year of the contract (see http://asp.usatoday.com/sports/basketball/nba/salaries/default.aspx). From this information, we were able to determine when the contract was signed. We then collected career performance data on each player as of the date the contract was signed, using *The Sporting News NBA Register* (various volumes). Local area data including 1994 per capita income (in U.S. dollars), 1997 metropolitan area population, and the white percentage of 1996 area population were collected for U.S. cities from the U.S. Department of Commerce (1998). For the two Canadian NBA cities (Toronto and Vancouver) that were eligible for players to sign with during the pre-2001–2002 period, we collected 1994 provincial per capita income data from the Department of Finance, Canada (1995), converted into U.S. dollars using the Organization for Economic Cooperation and Development's (OECD) purchasing power parity figures. Data on 1996 metropolitan area population and percent visible minority population for Toronto and Vancouver were taken from the Statistics Canada Web site (http://www.statcan.ca). Since the Vancouver Grizzlies moved to Memphis for the 2001–2002 season, any person signing a contract for this team that begins with these two seasons is assigned the Memphis values for these metropolitan area variables.

Our basic salary analyses can be expressed in the following functions $F(-)$, $G(-)$ and $H(-)$, whose form will be discussed shortly:

$$\ln \text{salary} = F(X, \text{white}, u_a) \tag{1a}$$

$$\ln \text{compensation} = G(X, \text{white}, u_b) \tag{1b}$$

$$\text{duration} = H(X, \text{white}, u_c), \tag{1c}$$

where salary is the average annual compensation (including bonuses) over the duration of the player's contract, compensation is total cash compensation for the contract, duration is years' duration of the contract, X is a vector of explanatory variables to be described in succeeding discussions, white is an indicator for white players, and u_a, u_b, and u_c are disturbance terms. Because the salary distribution is truncated at the league minimum ($332,817 for 2001–2002), we use Tobit analyses for estimating 1a and 1b, and because duration is a discrete variable, we estimate 1c using ordered probit.

The explanatory variables in *X* include: a draft indicator equaling the inverse of a player's draft number for drafted players, and zero for undrafted players; individual years of experience dummies for one through seventeen years, with zero years being the omitted category (i.e. rookies); points, rebounds, assists, and blocked shots per game in the NBA, with zeroes for rookies; number of NBA all-star teams to which one was named; the number of games per season a player was out of the lineup in his NBA career, again with zeroes for rookies; position dummies (center and forward); and the previously mentioned local market variables—log of 1997 population, log of 1994 per capita income, and percentage of the metropolitan area's population that in 1996 was white.

The key explanatory variable of course is the white dummy variable. The relatively small number of whites in the NBA makes running separate regressions by race impractical. In addition, there is a small number of Hispanic or Asian players who are given values of zero for the white variable. Results were similar when these players were omitted from the analysis. There were no significant differences between North American whites and whites from Europe.

The draft variable is a measure of anticipated player quality when he entered the league. It also places a person in the rookie salary scale. It is defined as 1 divided by draft number for those drafted, and 0 for those not drafted. It therefore ranges from 0 to 1, and assumes that players drafted very late (e.g. with draft number 58) were only marginally more desired than those not drafted at all. We experimented with the functional form of the draft measure with no change in the results. In addition, in some analyses, we included interaction forms between the draft variable and our experience measures to test the idea that the impact of being drafted at a given level decays over time as teams observe one's actual performance.[3]

Experience is measured by individual season dummies so as to impose the least structure possible on the experience-earnings profile. Results were similar with the more familiar quadratic or quartic experience specifications. Scoring, rebounding, making assists, and blocking shots are the major

[3] As there may be discrimination in hiring, we also investigated whether there were racial differentials in the draft variable, controlling for college performance, and whether one skipped college and came directly to the NBA. We found no such differentials. And acknowledging the possibility that draft position is potentially endogenous, we also estimated linear versions of 1a–1c with the draft variable instrumented by these college-related variables, with no change in the basic findings.

performance indicators for NBA players. They are expressed on a per-game basis, and equal zero for rookies. Thus, they are in effect interactions between veteran status and performance. The "main effect" of being a rookie is the predicted dependent variable when the experience and performance variables are all zero. Finally, number of all-star teams made is an indicator of playing quality, and number of games per season missed is largely a measure of health, although it is also affected by the coach's lineup decisions. Again, these variables equal zero for rookies, making them interactions with nonrookie status.

The position variables may represent relative scarcity of particular skills. Recognizing that teams may practice positional segregation, we also estimated the models with the position variables excluded. The final explanatory variables are the market measures: population, per capita income, and racial composition of the area. Because teams may select players' race based on market characteristics, we also estimated models with these variables excluded.[4]

While equations 1a–1c were estimated for the entire NBA sample, our discussion of free agency suggested that we should expect different results depending both on free agency status and where one was drafted. We therefore also estimate the model in three different subsamples: (1) nonrookies with less than three years of experience, plus rookies who were not drafted on the first round, or who entered the league before 1995 (the first year in which there was a set rookie salary scale for first-round draft picks); (2) rookies who were drafted on the first round 1995 or later; and (3) players with at least three years of experience. It is important to note that seniority, and therefore free agency eligibility status, is measured as of the signing of the player's contract.

Note that we include nondrafted rookies in group 1, even though they are technically free agents. We follow this procedure because the NBA had the chance to exercise monopsony power over these players by drafting them, but chose not to. However, our results are virtually identical if we call these undrafted rookies free agents, and move them to group 3. Finally, if, controlling for productivity, whites in general are wealthier than nonwhites, and if risk aversion is similar across players, then contract duration should be longer for white players in all groups than nonwhite players.[5]

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[4] In light of the possibility of customer discrimination, in some models we interacted a player's race with the racial composition of the area, and we briefly discuss these results in footnote 11.

[5] We do not control for date the contract was signed, as this is likely to be an inverse measure of contract duration, given our focus on the 2001–2002 season, and duration is endogenous (with no obvious identifying instruments that would not also belong in the salary and compensation equations). However, when we included year of signing as a control, our qualitative conclusions were unchanged.

Because a major possible source of NBA racial compensation differentials is customer discrimination, we also analyze the determinants of home attendance pooling the 1996–1997 through 2000–2001 seasons:

$$\ln \text{attend}_{it} = B'X_{it} + C'Z_i + a_t + e_{it}, \quad (2)$$

where for team i and year t, attend is home attendance; B and C are coefficient vectors; X is a vector of time varying explanatory variables including percent of the team's players who were white, log of average ticket price, log of arena capacity, team winning percentage, number of major league sports franchises in the metropolitan area, and number of players on that year's team named to the NBA all-star game; Z is a vector of the same time invariant market variables used in the salary analysis, including log of the 1997 area population, percentage of the area's 1996 population that was white, and log of the area's 1994 per capita income; a is a time effect; and e is a disturbance term.

The attendance and ticket price data for equation 2 come from Professor Rodney Fort's Web site (http://users.pullman.com/rodfort/PHSportsEcon/Common/OtherData/DataDirectory.html); the team racial composition and arena size data come from *The Sporting News NBA Guide* (various issues) and *The Sporting News NBA Register* (various issues); the market variables have already been discussed (see previous discussion); and the number of franchises was determined from http://www.espn.com. We correct the standard errors for the correlation across years for a given team. In addition, in some models we also control for team effects, although in these equations, we are not able to enter the time-invariant market characteristics. Equation 2 is designed to test whether adding white players brings in fans, all else equal. The results of this analysis will help us understand the race effects in the analysis of compensation.

Results

Table 1 contains mean values for white and nonwhite players. Exactly 20 percent of the league consists of white players, a figure consistent with what others have found in the NBA since at the least the 1980s (see the papers referred to in the discussion of the literature). Like previous research on the NBA, our data show that nonwhite players generally outperform white players, and are more highly paid as well. Average annual compensation under current contracts is 0.174 log points, or roughly 19 percent, higher for nonwhite than white players ($e^{0.174} \approx 1.19$), but the total value of nonwhite players' contracts averages only 0.093 log points (9.7%) more than that of white players. In an accounting sense, this difference between annual and

TABLE 1

MEAN VALUES FOR WHITE AND NONWHITE PLAYERS

	White players	Nonwhite players	Difference (white–nonwhite)
log(annual salary in thousands of dollars)	7.535	7.709	−0.174
log(total compensation in thousands of dollars)	8.797	8.889	−0.093
Contract duration (years)	4.036	3.931	0.105
Draft	0.079	0.126	−0.047*
Experience (years)	3.012	3.708	−0.696
Points per game	4.404	7.082	−2.678***
Rebounds per game	2.195	2.931	−0.736**
Assists per game	0.993	1.543	−0.550**
Blocks per game	0.308	0.391	−0.083
Points per game (nonrookies)	7.778	10.268	−2.490***
Rebounds per game (nonrookies)	3.876	4.249	−0.373
Assists per game (nonrookies)	1.753	2.236	−0.483
Blocks per game (nonrookies)	0.544	0.567	−0.023
Number of all-star teams	0.181	0.476	−0.295
Games missed per season	14.195	15.696	−1.501
Number of all-star teams (nonrookies)	0.319	0.690	−0.371
Games missed per season (nonrookies)	25.067	22.756	2.311
Center	0.301	0.123	0.178***
Forward	0.494	0.425	0.069
Guard (omitted category)	0.205	0.452	−0.247***
log area population	15.003	15.224	−0.221**
log area per capita income	10.036	10.055	−0.018
Percentage white population	81.480	80.637	0.843
Sample size	83	332	

*, **, *** denote, respectively, significant racial differences at the 10, 5, and 1 percent levels on two-tailed tests based on a linear regression of each variable on a white dummy variable.

contract-wide compensation is because of the 0.105 years' longer contract duration for white players.[6] Recall that the draft variable equals 1/draft number for those drafted, and 0 for those not drafted into the league. The higher value for nonwhites implies that they were in general taken earlier than whites. Nonwhites have more experience, better performance statistics, and more all-star team appearances than white players. Recall that the basic performance variables (scoring, rebounds, assists, and blocked shots) are coded 0 for rookies. Our conclusion about nonwhites' better performance can be seen by examining data in Table 1 on performance among those with NBA experience, although only the scoring difference is statistically significant among veterans. Looking at those with NBA experience,

[6] This is a true difference in contract duration, because according to the *USA Today* Web site (http://www.usatoday.com), almost all long-term contracts in the NBA are guaranteed.

nonwhite players miss fewer games per season, although the racial difference is not significant.

White players are more likely to be forwards or centers, and less likely to be guards than nonwhite players, with statistically significant differences in the incidence of guard or center. It has been argued in the sociology of sports literature that historically, black players were kept out of leadership positions such as quarterback, catcher, or middle infielder (Kahn 1991). While such arguments imply that guard would be a white position, the NBA data suggest otherwise. Finally, nonwhite players are located in areas with larger populations, slightly higher per capita incomes, and a smaller white population percentage, although only the total population difference is significant. But data in Table 1 do suggest some slight sorting by race across teams, a phenomenon consistent with either customer discrimination or player preferences.

Table 2 shows the racial differences in salary, contract-wide compensation, and duration among each of the three groups that face different rules for contract negotiation. In panel A, we see that among the players subject to individual teams' monopsony power, whites on average have annual salaries

TABLE 2

SELECTED MEANS BY RACE FOR SUBSAMPLES

A. Nonrookies with less than three years' experience, plus rookies who were not drafted on the first round, or who entered before 1995

	White players	Nonwhite players	Difference (white–nonwhite)
log(annual salary in thousands of dollars)	6.864	6.719	0.145
log(total compensation in thousands of dollars)	7.912	7.421	0.492
Contract duration	3.571	2.653	0.918
Sample size	21	75	

B. Rookies drafted on the first round in or after 1995

	White players	Nonwhite players	Difference (white–nonwhite)
log(annual salary in thousands of dollars)	7.271	7.435	−0.164
log(total compensation in thousands of dollars)	8.657	8.820	−0.163
Contract duration	4.000	4.000	0.000
Sample size	24	70	

C. Players with three or more years' experience

	White players	Nonwhite players	Difference (white–nonwhite)
log(annual salary in thousands of dollars)	8.072	8.208	−0.136
log(total compensation in thousands of dollars)	9.374	9.504	−0.131
Contract duration	4.316	4.417	−0.101
Sample size	38	187	

TABLE 3

SELECTED MARGINAL EFFECTS FOR BEING WHITE FOR THE FULL SAMPLE

	White coefficient for:					
	log(annual salary × 10^{-3})		log(total compensation × 10^{-3})		Contract duration	
Specification	coef	asy se	coef	asy se	coef	asy se
Basic	0.0443	0.0872	0.1769	0.1411	0.2678*	0.1403
Basic, excluding position and market variables	0.0807	0.0841	0.2313*	0.1359	0.2929**	0.1336
Basic, with draft–experience interactions	0.1165	0.0810	0.2759**	0.1351	0.3099**	0.1438
Basic, with draft–experience interactions, Excluding position and market variables	0.1365	0.0780	0.3070**	0.1297	0.3160**	0.1369
Estimation method	Tobit		Tobit		Ordered probit	
Sample size	415	415	415	415	415	415

Entries for the salary and compensation equations are the marginal effects of each explanatory variable, evaluated at the sample incidence of uncensored observations (0.954). The marginal effects for duration are the impact of each variable on the index b'x, which is already in contract duration units (years).

Basic specification includes controls for: draft, individual years of experience dummies, points per game, rebounds per game, assists per game, blocked shots per game, number of all-star teams, games missed per season, position dummies for center and forward, and market variables (log population, log per capita income, and percentage white population). *, **, *** denote, respectively, significance at the 10, 5, and 1 percent levels on two-tailed tests.

that are 0.145 log points (15.6%) higher than nonwhites, have contract-wide compensation that is 0.492 log points (63.6%) higher than nonwhites' compensation, and sign contracts for 0.918 years longer duration than nonwhites do. In contrast, for the other groups, nonwhites have higher salary, higher compensation, and contract duration that is either the same as or longer than that of whites.[7] None of these differences are statistically significant, possibly because of the small within-group sample sizes. However, the pattern of racial differences does suggest that whites do better relative to nonwhites among players facing the most monopsony power.

Table 3 shows basic Tobit and ordered probit results for the impact of being white on salary, contract-wide compensation, and duration for the full sample of NBA players.[8] Results from several specifications are shown; in particular, based on our earlier discussion, we estimated models including or excluding position and market variables, and also models with interactions between the draft variable and the experience dummies. Unconditional effects of race on log salary or log compensation are shown (they were

[7] *USA Today* reports a four-year contract duration of 92 of the 94 first-round draft picks under rookie contracts. Presumably, this reflects the very likely event that teams will pick up the fourth year option of these players.

[8] Results for the other variables are available from the authors upon request.

obtained by multiplying the raw Tobit coefficients by the sample incidence of uncensored observations, 0.954—only 19 of 415 players earned the league minimum); and the results for duration are also partial derivatives of the dependent variable.

Table 3 shows that the white effect on log of annual salary is positive in each case, although relatively small to moderate in magnitude: it ranges from 0.0443 to 0.1365 log points, and is significant only once at the 10 percent level on a two-tailed test (the draft–experience interaction model with the position and market variables excluded). However, the white effect on contract-wide compensation is larger, ranging from 0.1769 to 0.3070 log points, and is statistically significant three of four times. The positive effects of being white on contract duration, which range from 0.2678 to 0.3160 years and are 1.91 to 2.31 times their asymptotic standard errors, help account for the compensation results in comparison to the salary findings. Overall, other things equal, there is some weak evidence that white players make higher annual salaries than comparable nonwhites; and there is stronger evidence that whites have longer contracts and larger amounts of guaranteed money than comparable nonwhites.

As discussed previously, discrimination may vary according to whether the employer has monopsony power. Tables 4 and 5 show the effect of being white among the three subsamples described previously. As in Table 3, we show several kinds of models: some with and some without the position and market variables; and some with and some without draft–experience interactions. We note that for panel B in Table 4, the first-round draft picks in their rookie contract signed since 1995, experience is zero for everyone, implying that we cannot estimate a draft–experience interaction model for them (see Table 5).

In Tables 4 and 5, the only group for which there are large or significant *ceteris paribus* racial salary, compensation, or contract duration differentials consists of players without set salary scales, and without free agency rights: nonrookies with less than three years' experience, plus rookies who were not drafted on the first round or who entered the league before 1995. For this group under all of the specifications shown in Tables 4 and 5 (panel A in each case), there are large, statistically significant, positive white salary, compensation, and contract duration effects, all else equal. After controlling for other factors, there are white–nonwhite differentials in salary of 0.3117 to 0.4628 log points (36.6 to 58.9 percent); compensation differences of 0.7080 to 0.9221 log points (103.0 to 151.5 percent); and differentials in contract duration of 1.31 to 1.49 years. For first-round draft picks under rookie scale contracts (all of whom signed after 1995), there are very small and statistically insignificant salary and compensation effects, and virtually

TABLE 4

SELECTED MARGINAL EFFECTS OF BEING WHITE FOR SUBSAMPLES (MODEL DOES NOT INCLUDE EXPERIENCE–DRAFT INTERACTIONS)

	Dependent variable					
Specification	log(annual salary × 10^{-3})		log(total compensation × 10^{-3})		Contract duration	
Estimation method	Tobit		Tobit		Ordered probit	
	coef	asy se	coef	asy se	coef	asy se
A. Nonrookies with less than three years' experience, plus rookies who were not drafted on the first round, or who entered before 1995						
Position and market vars in	0.3144**	0.1329	0.7113***	0.2243	1.3106***	0.3360
Position and market vars out	0.4391***	0.1219	0.8872***	0.2070	1.4148***	0.3010
Sample size	96	96	96	96	96	96
B. Rookies drafted on the first round 1995 or after 1995						
Position and market vars in	−0.0218	0.0714	−0.0199	0.0724	—	—
Position and market vars out	−0.0423	0.0663	−0.0396	0.0671	—	—
Sample size	94	94	94	94	94	94
C. Players with three or more years' experience						
Position and market vars in	0.0113	0.1188	0.0928	0.1908	0.0629	0.2027
Position and market vars out	0.0139	0.1175	0.0889	0.1891	0.0796	0.1983
Sample size	225	225	225	225	225	225

Entries for salary and compensation equations are the marginal effects of each explanatory variable, evaluated at the sample incidence of uncensored observations. Only in panel A were there any censored observations (19), implying that the incidence for this group was 0.802. The marginal effects for duration are the impact of each variable on the index b'x, which is already in contract duration units.

For sample B, 92 of 94 players had contract duration of exactly four years; thus, duration models were not estimated for this group. Controls include: draft, individual experience dummies, points per game, rebounds per game, assists per game, blocked shots per game, all star teams, and games missed per season. Position dummies are for centers and for forwards; market variables include log population, log per capita income, and percent white population. *, **, *** denote, respectively, significance at the 10, 5, and, percent levels on two-tailed tests.

everyone has a four-year contract (Table 4, panel B). And for veterans with some free agency rights (Table 4, panel C and Table 5, panel B), again there are small and statistically insignificant racial differentials in salary, compensation, and contract length.[9]

In Tables 4 and 5, we have defined the group subject to individual team monopsony power to be those players without free agency rights, who also

[9] The fact that we observe positive white effects on duration only for the group subject to monopsony power argues against a pure insurance explanation as we don't observe any effect for white free agents. Rather, the longer contracts for the group A white players should be seen as part of their total compensation level. In addition, in light of Hamilton's (1997) results, using quantile regressions for the free agents, we found no evidence of discrimination at any quantile from 0.1 to 0.9 with 0.1 increments.

TABLE 5

Marginal Effects of Being White: Specification Includes Draft × Experience Dummy Interactions, Subsamples

		Position and market variables included		Position and market variables excluded	
Dependent variable	Estimation method	Effect	Asy SE	Effect	Asy SE
A. Nonrookies with less than 3 years' experience, plus rookies who were not drafted on the first round or who entered before 1995					
Ln salary	Tobit	0.3117***	0.1062	0.4628***	0.0990
Ln compensation	Tobit	0.7080***	0.1800	0.9221***	0.1676
Duration	Ordered probit	1.3664***	0.3448	1.4946***	0.3078
B. Players with 3 or more years' experience					
Ln salary	Tobit	0.0245	0.1223	0.0269	0.1208
Ln compensation	Tobit	0.1018	0.1971	0.1006	0.1951
Duration	Ordered probit	−0.0189	0.2130	0.0094	0.2075

Note: a model is not estimated for the sample of rookies drafted on the first round in or after 1995 because experience is zero for everyone in this group.

Controls include: draft, individual experience dummies, interactions between each experience dummy and draft, points per game, rebounds per game, assists per game, blocked shots per game, all star teams, and games missed per season. Position dummies are for centers and for forwards; market variables include log population, log per capita income, and percent white population.

*, **, *** denote, respectively, significance at the 10, 5, and 1 percent levels on two-tailed tests.

were not under the rookie salary scale (panel A in each case). Earlier, we noted for players not drafted in the first round after 1994 (i.e. not subject to the rookie salary scale), that as of the 1999 contract, they became restricted free agents in their fourth year, subject to the right of first refusal. But for such players signing their contracts under the 1995 agreement, there was no right of first refusal. Our basic model treats all players in their fourth year, but not under the rookie salary scale as free agents. However, recognizing that players under the right of first refusal may be somewhat restricted in their job search, we have also re-estimated our models by including all such fourth-year players who were under the right of first of refusal in the group subject to monopsony power. Our results were very similar to those reported in Tables 4 and 5.

The results in Tables 4 and 5 are consistent with the idea first mentioned in the NBA context by McCormick and Tollison (2001) that NBA teams are discriminating monopsonists.[10] Players subject to the rookie salary scale for first-round draft picks, as well as free agents, are not directly affected by individual team monopsony, although overall league market power may have

[10] Madden (1973) suggested discriminating monopsony based on different labor supply elasticities by gender to the firm as a cause of the gender pay gap.

an effect on the rookie salary scale itself. But given where one is drafted, there is no scope for racial differences in treatment among first-round picks. And Table 4, panels B and C, and Table 5, panel B indeed show no racial differentials among these two groups. However, for players negotiating subject to their individual team's monopsony power, we indeed see large differentials favoring white players in salary, compensation, and contract duration. As noted earlier, an alternative, but equivalent view to the discriminating monopsony model, is a Nash bargaining framework in which white players have higher status quo income than nonwhite players. Basic Nash bargaining theory would then imply better pay for white players.[11]

Whether racial differences in non-NBA alternatives or in status quo income are large enough to matter is an empirical question. If, for example, playing in Europe, coaching or broadcasting are sufficiently lucrative, and if there are large racial differences in these opportunities (or in family wealth), then the monopsony or Nash bargaining argument may be valid. The median free agent in our sample earned about $4.3 million per year, suggesting that alternative jobs or status quo wealth may not be empirically relevant for such a player's negotiations. However, for players subject to team monopsony (i.e. rookies not drafted on the first round or nonrookies who are not free agents), median compensation was $492,000. Non-NBA alternatives or family financial circumstances might be relevant for some players in this latter group, especially if they are disproportionately marginal players, and thus anticipate a short NBA career.

An alternative explanation for our basic results is that it may be less costly for teams to discriminate against marginal black players than against black players who have lasted in the league long enough to become free agents or in choosing top draft picks. That is, discrimination hurts the team's chances of winning, but this effect will be smallest for the marginal players. Therefore, teams may choose to discriminate more against the marginal players.

Finally, Table 6 shows very small and insignificant team composition effects on fan attendance. This analysis suggests that customer discrimination is

[11] In both the full sample and for each of the subsamples in Tables 4 and 5, we also estimated models where the white dummy variable was interacted with the percentage white population. In the full sample, contract duration was the only equation to have a significant white player–percent white population interaction, which was positive. In the subsamples, these interaction effects were insignificant in every case, except for a significantly positive interaction for contract duration for the group subject to team monopsony. In the subsamples, the fact that the only significant interaction effect is observed in the monopsony sample is evidence against a general sorting of risk averse white players into heavily white areas. Perhaps these junior white players without rookie scale contracts are willing to settle down in white areas to a greater extent than in less white areas. The effect might not be present for veteran free agent white players, who may have the wealth to live where they want, and therefore value a long-term contract in white areas no more highly than anywhere else.

TABLE 6

REGRESSION RESULTS FOR LOG HOME ATTENDANCE (1996–1997 THROUGH 2000–2001)

	Coef	SE	Coef	SE
Percentage white players	0.0009	0.0016	0.0001	0.0015
Percentage white population	0.0009	0.0018	—	—
Log ticket price	0.2810***	0.0637	0.1640*	0.0854
Arena size	0.00005***	0.00001	0.00003**	0.00002
Winning pct	0.0003***	0.0001	0.0003***	0.0001
Number of allstars	0.0124	0.0137	0.0136	0.0151
Log population	−0.0245	0.0251	—	—
Log per capita income	−0.1362	0.1110	—	—
Year dummies	Yes		Yes	
Team dummies	No		Yes	
R^2	0.8697		0.9323	
Sample size	145		145	

NOTE: standard errors are corrected for correlation across observations for the same team.
*, **, *** denote, respectively, significance at the 10%, 5% and 1% levels on two-tailed tests.

not behind the results in Tables 4 and 5, although we pointed out that recent research on customer discrimination in the NBA was mixed on the question of the existence of customer discrimination. But our wage and contract duration results in Tables 4 and 5 argue against a customer discrimination-based explanation, because if customer discrimination were a powerful force in this context, we would have expected to see race effects for free agents as well. While earlier research suggests that customer discrimination may well have affected NBA salaries in the 1980s, it appears to be a less powerful force in post-2000 period.

Conclusions

In this paper, we have examined racial salary, contract-wide compensation, and contract duration differentials in the NBA for those under contract for the 2001–2002 season. For the league as a whole, there were small to moderate and generally statistically insignificant annual salary differentials favoring white players, all else equal. However, there were larger and usually statistically significant racial differentials in total compensation favoring white players, and whites also had statistically significantly longer contract duration. Upon disaggregation, however, we found such racial effects only among players subject to individual teams' monopsony power—rookies who were not on the rookie scale plus players who were not free agents. Among this subgroup, there were indeed large, statistically significant *ceteris* *paribus* nonwhite shortfalls in salary, total compensation, and contract duration. But for the other groups—players under the rookie salary scale, as well as veteran free agents, there were only very small and statistically insignificant racial effects for each of these dependent variables. Finally, we found no evidence of customer discrimination when we examined home attendance and the racial composition of teams. Overall, there appears to be less evidence of racial compensation discrimination in the 1990s and early 2000s than the 1980s, and the apparent decline in customer discrimination since the 1980s is consistent with these changes.

Our results are consistent with models of discriminating monopsony, on the assumption that white players have a more elastic labor supply curve than nonwhites (McCormick and Tollison 2001). Additionally, the results may also be explained by a Nash bargaining model, in which white players have higher status quo incomes than black players. An alternative explanation is that it may be less costly for teams to discriminate against marginal black players than against black players who have lasted long enough in the NBA to qualify for free agency or in choosing top draft picks. While there may be omitted variables such as one's ability to work with teammates, the pattern of our results across negotiation status groups suggests that something more than omitted variables is at work here.

It is interesting to note that in our data, there were two situations in which racial pay differentials for similarly qualified players were absent—players under the rookie salary scale and free agents. The rookie salary schedule is a union pay scale, and the absence of racial salary differences once we control for draft position is analogous to the small racial pay differences one would expect to observe under union standard rate schemes, controlling for occupation and seniority (Ashenfelter 1972; Freeman 1982). The nonwhite free agents are protected by employer competition, at least in the seeming absence of customer discrimination, as predicted by Becker's (1957) model of discrimination. Thus, union pay scales and competition are both potential mechanisms that can reduce racial pay differences.

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