Racial Differences in Professional Basketball Players' Compensation

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This article investigates racial differences in 1985–86 salaries of individual professional basketball players. White and black players earn similar mean compensation; however, controlling for a variety of productivity and market-related variables and for the endogeneity of player draft position, we find a significant ceteris paribus black compensation shortfall of about 20%. Further, we find that all else equal, including team performance and market factors, replacing one black player with an identical white player raises home attendance by 8,000 to 13,000 fans per season. The compensation and attendance results together are consistent with the idea of customer discrimination.

I. Introduction

The National Basketball Association (NBA) appears to be an example of racial progress. Blacks comprise roughly 75% of the players and about 80% of starting players. There are more black coaches in the NBA than in other professional sports (Berry, Gould, and Staudohar 1986). Many of the most highly paid players are black: the top three NBA salaries for the

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1985–86 season went to Earvin Johnson ($2.5 million), Moses Malone ($2.145 million), and Kareem Abdul-Jabbar ($2.03 million), all black players. Several other black players earned over $1 million: Ralph Sampson ($1.165 million), Marques Johnson ($1.1 million), Otis Birdsong ($1.1 million), Julius Erving ($1.485 million), and Patrick Ewing ($1.25 million).[1] Blacks have traditionally held leadership positions in the National Basketball Players Association (NBPA), although its director is white.[2]

Despite this evidence of black economic reward, blacks remain a small minority of coaches (two out of 23) and executives (one out of 23 general managers) in the NBA (see Berry et al. 1986). Bradley (1976) and Halberstam (1981) document the existence of racial prejudice in the NBA for the 1960s and 1970s. In particular, basketball fans often suggest that there is a premium for being a white player. Indeed, this last outcome would be predicted in a model of customer discrimination (Becker 1975). Since basketball players are more highly visible (both on the court and on the bench) to fans than players in other sports, customer discrimination may well be more prevalent in basketball than elsewhere, especially in view of the relative scarcity of white players.

Economic theory predicts that customer discrimination can persist under competition, while employer or co-worker discrimination is more likely to diminish (Becker 1975). In customer discrimination, the consumer is willing to pay a premium for white workers. In basketball, the white fan acts as if the white player is producing more entertainment value than a comparable black player. If there is customer prejudice in general, then black workers may be able to escape discrimination by working in sectors without producer-consumer contact. However, if black service workers have a large enough comparative advantage in services, or if, compared to the size of the service sector, the relative labor intensity of service production or the relative size of the black labor force is large enough, then customer prejudice will result in discrimination even under general equilibrium conditions (Kahn 1987). On the other hand, under employer or co-worker discrimination, buy outs by nondiscriminators or segregation by firm may tend to reduce wage discrimination (Becker 1975).

This article investigates NBA salary determination and particularly racial salary differentials in the NBA. We are able to provide considerably more detailed controls for performance than is done in typical studies of wage determination using data such as the Current Population Survey or the National Longitudinal Surveys. While blacks and whites earn comparable

[1] These salary figures are from our data base, to be described below. The highest-paid white players include Larry Bird ($1.8 million), Jack Sikma ($1.6 million), Mitch Kupchak ($1.15 million), and Kevin McHale ($1 million).

[2] Among the NBPA leaders have been Oscar Robertson, Paul Silas, and Bob Lanier, all black players (Berry et al. 1986).

overall salaries, we find that, all else equal (i.e., performance and market-related variables), blacks are paid less than whites by about 20%, an effect that is highly significant statistically. These results are robust to a variety of independent variable specifications and estimation techniques. Our findings suggest that although blacks have made major gains in professional basketball, unexplained racial salary differentials are of a similar order of magnitude as for the rest of the economy.[3] In addition, comparison of uncorrected mean salaries with regression results underscores the importance of controlling for other influences on salaries in studies of wage determination. Further, we find that, all else equal, white representation on a team contributes to home attendance, providing evidence consistent with the idea of customer discrimination. On the other hand, we do not find evidence of discrimination in the draft selection process.

II. Labor Relations and Salary Determination in the NBA

The NBA players have had union representation since the 1960s.[4] However, as in other professional sports, collective-bargaining agreements in basketball set the rules for salary determination rather than calling for a specific salary scale.[5] The era of free agency (i.e., a player selling his services in an open market, once the initial contract has been completed) in the NBA began with the 1976 collective-bargaining agreement. This contract provided for two distinct periods regarding free agency: 1976–81 and 1981–87. During the 1976–81 interval, a player could become a free agent when his existing contract expired. If he were bid away by another team, the two clubs involved needed to agree on compensation; the NBA commissioner acted as arbitrator in the event of no agreement. In 1981, this system was replaced by one in which the teams were given the right of first refusal: a team could keep a player's contract by matching the bidding team's offer.

By greatly reducing employer monopsony power, the free-agency system appears to have led to a rapid escalation of salaries. According to Berry et al. (1986), the average player salary was $109,000 in 1976, while our data for 1985–86 show an average of about $380,000.[6] At the same time, perhaps

[3] The 20% range is consistent with findings for the male labor force as a whole. For example, Blau and Beller (1984) found that, all else equal, black males faced a 14.5% weekly earnings shortfall in 1981 compared with white males.

[4] Much of this section is drawn from Berry et al. (1986).

[5] Basketball, as well as football and baseball, has a minimum salary ($70,000 for 1985–86). In addition, in 1983, a team salary cap was instituted in the NBA (see below).

[6] Similar increases have been observed in baseball following the institution of free agency (see Hill and Spellman 1983). The NBA compounded average annual rate of salary increase from 1976 to 1986 was about 13%, a substantially higher figure than the 7% annual hourly wage increase from 1976 to 1984 for the private sector (see *Monthly Labor Review* [January 1986]).

as a result of this salary escalation, the owners began reporting severe financial difficulties. In 1982, it was claimed that the average team lost \$700,000 and that only six of 23 were making positive profits (Berry et al. 1986, p. 181). As a result, in 1983 a cap on total team salaries was agreed upon by the owners and the NBPA. While the operation of the salary cap is very complicated, involving multiple qualifications and clauses, its intent is clear—to slow down the spiral in player salaries.[7]

It is probably too early to tell whether the cap has had any impact on overall salaries.[8] However, for the purposes of our study, the existence of the cap poses no obvious difficulty in comparing players' salaries. As of 1985–86, salary determination in the NBA, at least for experienced players, is basically market determined, as players use the free-agent system to find the best offer. If there is customer discrimination and if owners are profit maximizers, the free-agency system will establish a white salary premium. This premium will not be eroded by competition as long as white fans prefer to watch white basketball players.

III. Estimated Methods and Data

Our basic model for the salary-determination process is of the following form:

$$S = V^a, \tag{1}$$

$$\ln S = a \ln V, \tag{2}$$

where

S = individual player's compensation for 1985–86, including an imputation for deferred payments;[9]
V = a measure of the player's market value (marginal revenue product).

If there is customer discrimination, then V will depend on a player's race as well as his actual performance:

$$\ln V = cP + dR, \tag{3}$$

where

P = the player's performance level;
R = 1 for white players, 0 for black players.

[7] The basic salary cap for 1985–86 was the maximum of \$3.8 million/team, or 53% of gross revenues. On the other hand, there are enough exceptions to the cap to put 20 out of 23 teams over the \$3.8 million figure. According to Berry et al. (1986), the cap has, at the margin, caused some teams to alter their lineups to accommodate the cap.

[8] The increase from 1984–85 to 1985–86 in average salaries was about 11%, similar to the 13% average from 1976 to 1986 reported earlier (see Berry et al. 1986).

[9] Data are described more fully below.

If we had data on V and P, then we could directly test the model of customer discrimination as described by (1)–(3). In particular, the model would predict that once we controlled for V, race would not affect compensation (i.e., race would not enter eq. [1]) if customers are the only source of discrimination. However, given performance (P), race would affect compensation through its impact on the player's market value. The racial discrimination coefficient would equal ad. If teams or white players as well as customers discriminated, then, even controlling for V, race would still affect salaries (i.e., R would enter into eq. [1] as well as [3]). Finally, if customers did not discriminate but teams/players did, then $d = 0$ and race would affect compensation given V.

While data on V, P, and S would allow us to distinguish the differing models, we do not observe V or P. Instead, we must proxy P by a vector X of player performance indicators and team and local metropolitan area characteristics.[10] Rather than (2), we have:

$$\ln S = B_4'X + \delta R + \varepsilon_4, \tag{4}$$

where B_4 is a coefficient vector, δ is the race effect, and ε_4 is an error term.

Equation (4), while estimable, does not give us the same information that estimating (1)–(3) would. In particular, as noted, by adding a race term to (1), we could separately identify team/player and customer discrimination. However, in (4), lack of data on P and V forces us to combine these forms of discrimination into one coefficient.

To estimate the compensation model, we merge three data bases: (1) player salary information, (2) individual performance data, and (3) team and metropolitan area data. Player salary data for 1985–86 are taken from a survey of the entire NBA by a team of newspapers including *The Sporting News, Houston Chronicle, Phoenix Gazette, Detroit Free Press,* and *New York Daily News.*[11] As noted, the salary data include an imputation for the value of deferred compensation (as well as bonuses). More specifically, the player-compensation figure includes "base salary, signing bonuses, 'reasonably attainable' performance bonuses, and deferred payments" (*New York Daily News,* December 26, 1985, p. 100). Thus, a serious attempt was made in the newspapers' survey to go beyond a base-salary measure. However, our compensation variable does not include future deferred payments, and it is impossible to break down current compensation into salary and nonsalary components. If black players came from poorer families than whites, liquidity problems might lead the former to receive more of their

[10] The possible omitted variable problem caused by use of X is discussed below.

[11] Actual player compensation is published in the December 26, 1985, edition of the *New York Daily News* (p. 100). This survey was described verbally to us by Mike Douchant, basketball editor of *The Sporting News.*

compensation as salary rather than as deferred payments. Lack of data on future deferred payments, then, may understate white compensation by more than black compensation. Finally, the bonus component (which obviously is not separable in our sample from deferred payments) is not an actual "piece rate" but an estimate of what the surveyers thought was reasonably attainable by any player. It should, therefore, be thought of as a deferred payment.

Player performance and team data are taken from the 1985–86 editions of *The Sporting News NBA Guide* and *NBA Register*. Finally, metropolitan area data for each player's team are taken from the *Statistical Abstract of the United States 1985 and 1986.*

The vector X of explanatory variables includes the following (measured for the start of the 1985–86 season):

SEASONS = total seasons played,[12]
MINS = average minutes played per game,
GAMES = average games played per season,
FTPCT = career free throw percentage (fraction made),
FGPCT = career field goal percentage (fraction made),
POINTS = career points scored,
CENTER = dummy variable for centers,[13]
FORWARD = dummy variable for forwards,
OFFREB = career per game offensive rebounds,[14]
DEFREB = career per game defensive rebounds,
ASSISTS = career per game assists,
PFOULS = career per game personal fouls,
STEALS = career per game steals,
BLOCKS = career per game shots blocked,
HOMEATT = 1984–85 home attendance of one's 1985–86 team,
RACEMSA = 1980 percent of Standard Metropolitan Statistical Area (SMSA) population that was black in the SMSA where one's team was located,
POPMSA = 1980 population of one's team's SMSA,

[12] A small number of players (e.g., Moses Malone, Julius Erving, George Gervin, Artis Gilmore) had ABA experience before the NBA-ABA merger in 1976. For these players, we computed the entire pro career, including ABA statistics.

[13] In the case of those who play more than one position (e.g., Kevin McHale plays forward and center), we code his primary position as defined by the *NBA Guide*.

[14] Rebounds were not broken down into offensive and defensive categories before 1973. For those whose careers started before 1973, we imputed offensive and defensive rebounds by prorating the total rebounds for these years by the post-1973 reported percentages of total rebounds in these two categories. The number of players affected was small.

INCMSA = 1983 real per capita income in the SMSA,[15]
DRAFTNO = number a player was picked in the draft (e.g., the top player was picked number 1),
WINPCT = 1984–85 winning percentage of one's 1985–86 team,
WHITE = 1 if player is white, 0 if player is black.

The shooting, scoring, rebounding, assists, fouls, and defense statistics are all obvious measures of on-court player performance. Longevity (SEASONS) combines a pure seniority effect and the value of a sustained level of per game and per season performance. The position variables (CENTER, FORWARD) are included to control for the possible differential value of these positions, especially for centers. It is often alleged that a good center is of special importance, and that the best teams usually have a dominant center. In addition, good forwards may be more scarce than good guards. By including CENTER and FORWARD, we test for the notion that the measured performance variable may not completely control for this position phenomenon. Minutes per game may measure a player's importance but may also be a negative productivity indicator since we control for per game statistics (e.g., more minutes for the same points per game). Games per season measures freedom from injury or willingness to play injured.

The team-related variables HOMEATT, RACEMSA, INCMSA, and POPMSA all capture market factors. The customer-discrimination argument would suggest interacting RACEMSA with WHITE: we would expect the premium paid to white players to be a negative function of the proportion of the local population that is black, and this test was attempted (see below). WINPCT is included to test for the idea that the team "settles up" with its players on the basis of the past year's performance.

Finally, DRAFTNO is included as a further indicator of player quality, potential, or fan appeal. It is an inverse ordinal measure of the league's desire for a given player as he entered the NBA. While one should control as completely as possible for player quality, DRAFTNO may confound player quality and racial aspects of fan appeal. The customer- (as well as team/player-) discrimination model would suggest that, all else equal, white players would be drafted before black players. The endogeneity of DRAFTNO suggests using a two-stage least-squares (2SLS) procedure for the salary equation. We therefore also estimate with 2SLS the following equation for DRAFTNO:

$$\text{DRAFTNO} = \alpha' Z + B\,\hat{\ln}\, S + e_S, \tag{5}$$

[15] Money income was taken from *Statistical Abstract of the United States 1985 and 1986*. This was then deflated by the relative cost of living in the area. Cost of living was calculated by inflating the 1976 budget for an intermediate living standard (for a family of four) by the area's increase in the CPI from 1976 to 1983. Cost-of-living figures were taken from U.S. Department of Labor, Bureau of Labor Statistics, *Handbook of Labor Statistics,* 1978 and 1985.

where Z includes

WHITE = (defined earlier),
EARLY = dummy variable for those who left college as underclassmen,
SEASONS = seasons as a pro (defined earlier),
COLLSEA = college seasons,
CGAMES = total college games played per season,
CPTS = total college points scored per season,
CFGPCT = college field goal percentage,
CFTPCT = college free throw percentage,
CREB = total college rebounds per game (data on offensive and defensive rebounds were not available),
CAWARDS = number of times named to *The Sporting News* first or second All-America team or won its College Player of the Year award,
$\hat{\ln} S$ = predicted log compensation from a reduced-form model.

In the DRAFT equation, we have included measures of college performance, plus SEASONS (as a pro) in order to control for cohort effects in the player draft. Further, if there is discrimination, then at the same price (i.e., compensation) whites should be drafted before blacks. Therefore, the DRAFT equation was estimated by including predicted log compensation as a measure of anticipated cost of a player at the time of the draft.

While the compensation data are for only 1 year, it is possible that multiyear considerations affect compensation. Indeed, the notion of deferred payments is clearly a multiyear phenomenon. Further insight into compensation determination can be gained by interacting seniority with performance variables. Such a test was also done.

The salary regression can shed light on racial differences in compensation controlling for qualifications. However, these techniques do not provide direct tests of discrimination, although their results can be suggestive. To shed light on fan preferences (and, implicitly, the idea of customer discrimination), we estimate a model of home attendance. We have obtained data on home attendance for the six-season period 1980–81 through 1985–86 (collected from *The Sporting News NBA Guide* for each of these years). We then pose the following equation (units of observation are a team in a given year):

$$\begin{aligned}\text{ATTEND} = m_0 &+ m_1\,\text{YEAR} + m_2\,\text{WINPCT} + m_3\,\text{STARS} \\ &+ m_4\,\text{ARENA} + m_5\,\text{RACEMSA} + m_6\,\text{POPMSA} \\ &+ m_7\,\text{TEAMS} + m_8\,\text{PRICE} + m_9\,\text{INCMSA} \\ &+ m_{10}\,\text{PCTWHITE} + \varepsilon_6,\end{aligned} \tag{6}$$

where

ATT = team's home attendance for a given year,
YEAR = time trend (i.e., 81 for 1980–81, . . . , 86 for 1985–86),
WINPCT = team's winning percentage for a given year,
STARS = number of team members on first or second all-NBA team (chosen by *The Sporting News*),
ARENA = arena capacity,
TEAMS = number of other major league sports franchises in the area,
PRICE = minimum ticket price in 1967 dollars,
PCTWHITE = fraction of team members who were white

(RACEMSA, INCMSA, and POPMSA have been defined earlier).

Estimation of (6) is a direct test of the idea of customer discrimination. The PCTWHITE coefficient tells us the impact on attendance of adding white players of equal ability to the black players replaced. WINPCT and STARS are further indicators of team fan appeal, ARENA controls for the upper limit on attendance, and RACEMSA, INCMSA, POPMSA, and TEAMS are demand-related variables. Minimum ticket price is used since average ticket price data were unavailable.[16] Since teams choose ticket prices, PRICE is likely to be endogenous. Therefore, (6) was also estimated with PRICE omitted—that is, a reduced form. However, even in this context, PCTWHITE may yield interesting information about fan preferences.

IV. Basic Empirical Results

Table 1 provides mean values for the sample used in the wage and draft regressions (cases with missing values were deleted). Black players make up 74.3% of the sample, a figure very close to Berry et al.'s (1986) estimate of 75%. Blacks on average have a \$10,620 compensation advantage (2.7%) over whites, while the difference in mean log compensation would imply a 10.5% black advantage ($e^{.1} \approx 1.105$). However, blacks on average have more experience (both seasons and minutes per game), better shooting statistics, more points per game, more points per minute, and are drafted earlier than whites. Blacks are more likely to leave college before using up their eligibility, score and rebound more as collegians than whites, and win more college awards. On the other hand, whites play for teams with better records, a possible Celtic effect.[17] The other team-related variables suggest that whites play for teams with higher home attendance, located

[16] For some cases, even minimum ticket price was missing. In these instances, we interpolated to approximate this variable.

[17] With Celtics excluded, average WINPCT is .490 for whites and .487 for blacks, but the race coefficient in the log wage regression is unchanged.

Table 1
Mean Values of the Wage-Regression Sample

Variable	Whites	Blacks
SEASONS	4.414	4.8571
FTPCT	.729	.737
FGPCT	.486	.490
MINS	20.862	25.446
POINTS	8.69	12.11
GAMES	66.17	71.09
CENTER	.345	.173
FORWARD	.448	.399
OFFREB	1.372	1.634
DEFREB	3.052	3.142
ASSISTS	1.858	2.665
PFOULS	2.457	2.699
STEALS	.626	.971
BLOCKS	.648	.638
HOMEATT	473,540	454,080
RACEMSA	11.0%	13.6%
POPMSA	271.5×10^4	314.8×10^4
DRAFTNO	29.1	24.8
WINPCT	.524	.491
INCMSA	$9,410.9	$9,640.7
EARLY	.1207	.1905
COLLSEA	3.8103	3.5774
CGAMES	27.292	27.993
CPTS	13.271	16.155
CFGPCT	.5393	.5243
CFTPCT	.7201	.7051
CREB	6.6063	7.3616
CAWARDS	.3793	.4167
S	$396,570	$407,190
ln*S*	12.57	12.67
Sample size	58	168

in smaller metropolitan areas, and with a greater percentage of whites in the area. These team-variable means are consistent with the customer-discrimination idea, as teams with more white players draw more fans and are located in relatively white areas. The effect of white players on home attendance is tested more formally below.

A. Compensation Regression Results

While table 1 shows a slight black salary advantage over whites, tables 2 and 3 present regression results suggesting that, all else equal, blacks are paid significantly less than whites in the NBA.[18] First, the ordinary least

[18] The existence of a $70,000 minimum salary suggests the possibility of a truncated dependent variable. However, in our sample, those earning $70,000 and having complete data on their explanatory variables comprised less than 1% of the regression sample. Thus the truncation problem is not likely to be severe in this case. On the other hand, the mean salary figures in table 1 are 2%–5% higher than the $380,000 figure reported earlier, indicating that low-paid players were slightly less likely to have complete data than highly paid players.

Table 2
OLS and 2SLS Results for ln (Compensation)

Explanatory Variables	OLS	OLS	OLS	2SLS
CONSTANT	10.378	10.660	10.574	10.643
	(.4265)	(.4203)	(.4512)	(.4887)
SEASONS	.0300	.0352	.0316	.0286
	(.0105)	(.0102)	(.0101)	(.0111)
FTPCT	−.2191	−.3434	−.2735	−.2761
	(.4225)	(.4080)	(.3977)	(.4189)
FGPCT	1.2761	1.4045	1.1678	.9052
	(.6799)	(.6597)	(.6451)	(.6751)
MINS	.0095	.0128	.0142	.0124
	(.0129)	(.0125)	(.0122)	(.0128)
POINTS	.0558	.0470	.0444	.0471
	(.0114)	(.0111)	(.0108)	(.0125)
GAMES	.0081	.0057	.0066	.0097
	(.0027)	(.0027)	(.0027)	(.0032)
CENTER	.0808	.0599	.0844	.1105
	(.1318)	(.1283)	(.1252)	(.1308)
FORWARD	−.0297	−.0175	−.0132	−.0197
	(.0915)	(.0889)	(.0865)	(.0900)
OFFREB	.1022	.0660	.0854	.1166
	(.0758)	(.0736)	(.0720)	(.0780)
DEFREB	.0491	.0464	.0352	.0288
	(.0370)	(.0357)	(.0352)	(.0366)
ASSISTS	.0396	.0296	.0281	.0250
	(.0325)	(.0313)	(.0304)	(.0322)
PFOULS	−.0732	−.0418	−.0315	−.0498
	(.0530)	(.0515)	(.0509)	(.0569)
STEALS	−.0612	−.0630	−.0552	−.0488
	(.1019)	(.0982)	(.0956)	(.0993)
BLOCKS	.0184	.0342	.0389	.0426
	(.0620)	(.0604)	(.0591)	(.0625)
HOMEATT (%10^3)	. . .	. . .	.00009	.00008
			(.0003)	(.0003)
RACEMSA	. . .	. . .	.0003	−.0012
			(.0040)	(.0042)
POPMSA	. . .	. . .	.0004	.0004
			(.0001)	(.0001)
INCMSA (%10^3)	. . .	. . .	−.0092	−.0144
			(.0177)	(.0187)
DRAFTNO	. . .	−.0048	−.0050	−.0063
		(.0011)	(.0011)	(.0033)
WINPCT	. . .	.0730	.0199	.0086
		(.1955)	(.2449)	(.2559)
RACE	.2065	.1914	.2130	.2262
	(.0679)	(.0660)	(.0647)	(.0674)
Sample size	226	226	226	226
R^2	.7224	.7462	.7660	. . .
S.E.E.	.4002	.3845	.3728	.3870

NOTE.—(Asymptotic) standard errors in parentheses.

squares (OLS) and 2SLS race effects for ln (compensation) in table 2 are highly significant and amount to a white premium of about 20%. The first column of table 2 includes only individual performance measures and race,

Table 3
Selected OLS Log Compensation Results from a Seniority-Interaction Model

		$\left[\frac{\partial \ln(\text{Compensation})}{\partial \text{Race}}\right]$ by Seniority	
Explanatory Variables	Coefficients	Seasons	$\frac{\partial \ln(\text{Compensation})}{\partial \text{Race}}$
RACE	.0968	1	.1251
	(.1164)		(.0975)
DRAFTNO	−.0047	2	.1534
	(.0017)		(.0811)
RACE * SEASONS	.0283	3	.1817
	(.0232)		(.0688)
DRAFT * SEASONS	.0001	4	.2100
	(.0003)		(.0631)
FTPCT * SEASONS	−.4413	5	.2383
	(.1267)		(.0656)
FGPCT * SEASONS	.4901	6	.2666
	(.2283)		(.0755)
POINTS * SEASONS	.0041	7	.2949
	(.0023)		(.0904)
OFFREB * SEASONS	.0399	8	.3232
	(.0217)		(.1084)
DEFREB * SEASONS	−.0373	...	...
	(.0102)		
ASSISTS * SEASONS	.0021	...	...
	(.0094)		
FOULS * SEASONS	.0055	...	...
	(.0176)		
STEALS * SEASONS	.0149	...	...
	(.0294)		
BLOCKS * SEASONS	.0004	...	...
	(.0178)		
S.E.E.	.3479	...	...

NOTE.—See table 2, col. 3, for a list of the other explanatory variables.

while the second column adds DRAFTNO and WINPCT, and the third column adds these two and the market variables. Although the raw means of log compensation show a 10% advantage for blacks (table 1), as soon as we control for individual performance, we get a significant 20% black compensation shortfall. This effect holds with the inclusion of DRAFTNO, WINPCT, and the market-related variables.

Among other results from table 2, scoring appears to strongly affect salary; for example, a 10 point per game scoring differential contributes to an all else equal 40%–50% compensation premium. Longevity (SEASONS) and durability within a season (GAMES) both contribute to salaries, although these variables may be indications of playing ability: the coach (or general manager) chooses to keep the better players on the court. Players in large markets (POPMSA) make a clear salary premium. For example, moving from Washington, D.C., to Los Angeles raises salaries about 18%.

Such a finding may indicate that in more lucrative markets, the marginal revenue product of, say, the fifth star of a team is greater than in less lucrative markets. Finally, the negative effect of DRAFTNO suggests that this variable is a further indicator of player quality or fan appeal, beyond what is measured by the performance statistics.[19]

1. *Life-Cycle Considerations*

While our data are for only 1 year, it is possible that multiyear considerations affect salary determination. The table 2 findings control for seniority, so whites and blacks are in the regressions roughly at the same point in their life-cycle. However, additional insight into compensation can be gained by investigating the interaction of performance variables and seniority. Table 3 shows the results of such a test. As a group, the interactions are significant at better than 1%, $F(11,204) = 3.3$, with a 1% critical value of about 2.3. Important wage-influencing variables such as scoring or field goal percentage have significant positive interaction effects with seniority. Evidently, sustained performance in these areas is particularly rewarded. On the other hand, free throw percentage and defensive rebounding, both less important determinants of salary, have significant negative seniority interactions.

The race-seniority interaction is positive but insignificant. However, the derivative of log wages with respect to race is insignificant at 0 or 1 year seniority but becomes significant with 2 or more years in the league. While this derivative rises with seniority, the insignificance of the interaction suggests caution in interpreting such a rise.

Finally, the effect of draft does not decay with seniority—the point estimate for DRAFT∗SEASONS is insignificant and very small compared to the main DRAFTNO effect. Evidently, DRAFTNO is close to a "fixed" effect, controlling for the other variables. Comparing two hypothetical players with the same performance but different draft numbers, we conclude that the league drafted the earlier player because of some fixed, unmeasured playing ability or fan appeal. For example, the Celtics recognized something in Bill Russell, Larry Bird, and Kevin McHale that could not be measured with the usual player performance statistics.

[19] Heteroscedasticity tests suggested by Dutta (1975) did not indicate that this problem was present in our data. In addition, regressing the squared residuals from the full OLS ln (compensation) model on the predicted value of the dependent variable (and a constant) yielded a coefficient of .0154 (.0181). Further, reverse regression tests indicated black overqualification to obtain a given compensation level. That is, when the predicted value of ln (compensation) for RACE = 0 was regressed on RACE and actual ln (compensation), we obtained a race effect of −.2359 (.0476) and a coefficient on ln (compensation) of .7652 (.0283). Therefore, this test did not indicate an omitted-variable problem.

2. *Alternative Specifications*

The race effect on compensation was robust to alternative specifications. For example, inclusion of height, weight, all-star games, number of team changes, or team ability to pay (as measured by team salary minus the player's salary) did not affect the race results. In particular, the insignificant effects for "ability to pay" are consistent with team profit maximization: teams are only willing to pay players what they are worth. Further, tests for structural differences between black and white log salary regressions accepted the hypothesis of a common structure—except for the race dummy variable. In particular, interactions between RACE and RACEMSA were insignificant.

3. *Extreme Values*

Having established a finding of a significant black shortfall in compensation, all else equal, we now investigate the extent to which such a finding reflects extreme values. For example, in our sample of 58 white players, a raise for one player from $100,000 to $2,000,000 salary raises the average white salary by about $33,000, or over 8%. Figures 1 and 2 shed light on the extreme value issue: the figures plot, respectively, white and black residuals from a pooled OLS log compensation regression from table 2, column 3. There are several highly paid white players with very large residuals, although, as noted, heteroscedasticity tests did not indicate that this was a problem.[20] A large positive residual for such players could indicate fan appeal due to their race or unmeasured playing ability.

To assess the impact of such extreme values on racial salary differentials, we estimated the basic log salary regression but excluded players making over $800,000/year, that is, we excluded a total of 20 players (five whites and 15 blacks). The race coefficient became: .1560 (.0646). When we restricted the sample to those earning less than $600,000 (an exclusion of 12 whites and 25 blacks), the RACE effect became: .1293 (.0693). Finally, when we restricted the model to those earning at least $100,000 (due to large negative black residuals at low salary levels), the RACE effect was .2327 (.0665).

The findings for the samples excluding highly paid players indicate that extreme values contribute to some but not all of the ceteris paribus racial compensation differential. Even among low to moderately paid players (earnings less than $600,000), there is a race effect of about 13% that is significant at 6.1% (two-tailed test).

B. Draft Position Results

Table 4 contains OLS and 2SLS results for the determination of a player's draft position. In all specifications, white players are drafted later than

[20] Ibid.

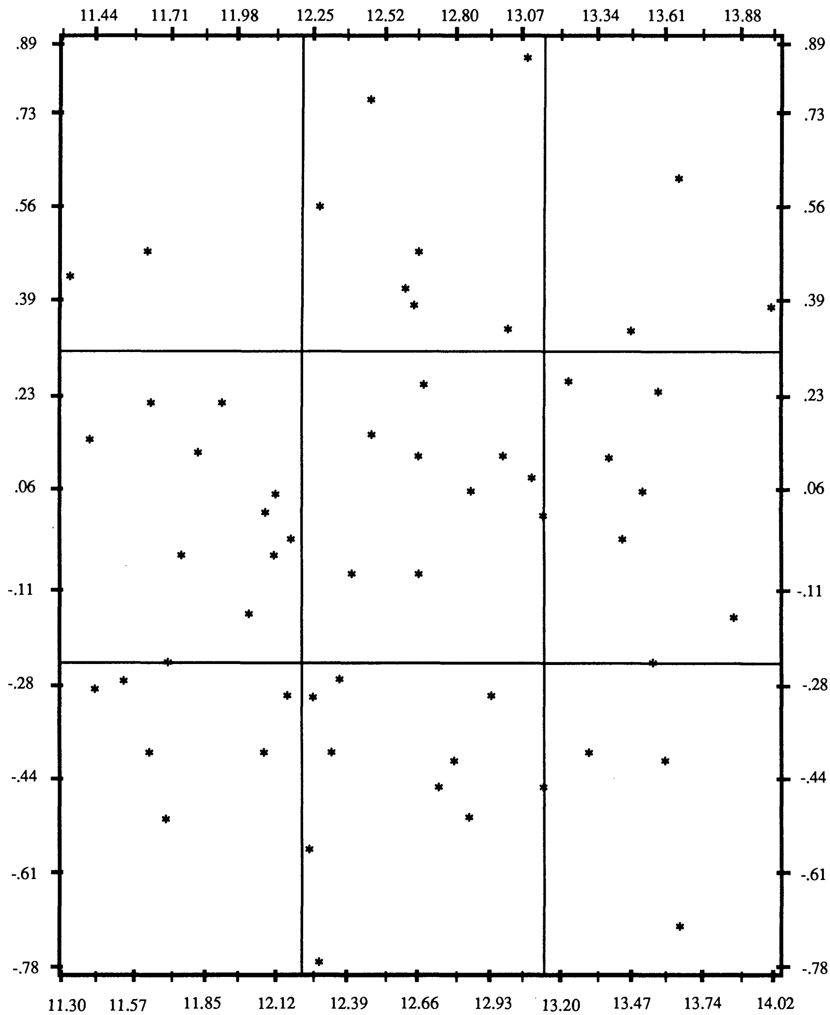

FIG. 1.—White residuals from basic OLS log salary model by predicted level of log salary

black players, other things equal, although the RACE coefficient is never significant. Such a finding does not support the idea of discrimination. Further, the compensation effect is negative and significant in all cases: in the 2SLS models, this means that predicted compensation is related to being selected earlier in the draft. Clearly, the compensation variable is not acting as a price term, which should have a positive coefficient, controlling for ability. It is likely that compensation (or its predicted value) in table 4 is really a measure of player quality.

The overall results for the draft equations are disappointing. Other than compensation, the only significant coefficients are EARLY, COLLSEA,

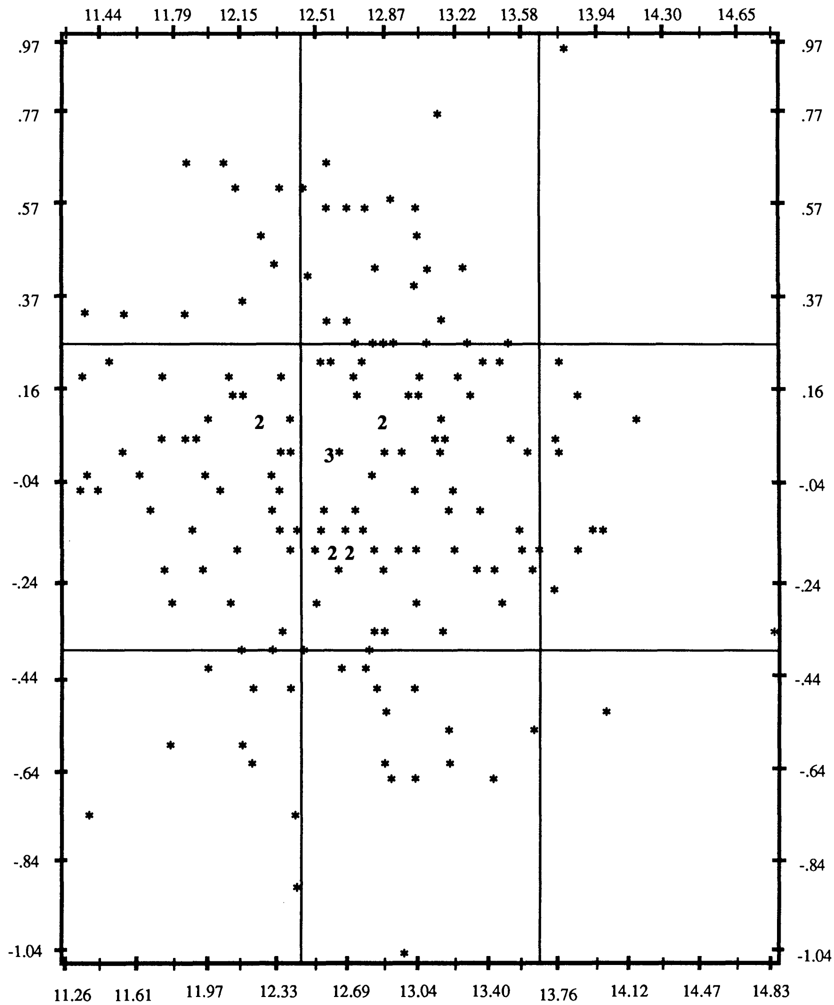

FIG. 2.—Black residuals from basic OLS log salary model by predicted level of log salary

CGAMES, and CAWARDS. To make a full study of the draft process, we would need college statistics and draft numbers for all players drafted and for those not drafted. Our sample includes only those who were drafted *and* who made it into the NBA. Unfortunately, complete data on the draftable population are unavailable.

C. Home Attendance Results

Table 5 contains regression results for the determinants of home attendance. We report both OLS and generalized least squares (GLS) coeffi-

Table 4
OLS and 2SLS Results for Position in Draft

Explanatory Variables	Dependent Variable: DRAFTNO OLS	DRAFTNO 2SLS	ln(DRAFTNO) OLS	ln(DRAFTNO) 2SLS
CONSTANT	320.07	325.71	17.225	17.638
	(39.335)	(45.91)	(2.1234)	(2.3019)
SEASONS	.5322	.6143	.0528	.0604
	(.7589)	(.8433)	(.0268)	(.0302)
EARLY	−8.4788	−8.2754	−.4023	−.3832
	(4.6858)	(4.902)	(.1714)	(.1812)
COLLSEA*	−6.0410	−6.0382	−.1727	−.1733
	(2.7404)	(2.8307)	(.0994)	(.1038)
CGAMES	−1.4035	−1.3759	−.9635	−.9054
	(.6033)	(.6319)	(.5684)	(.5998)
CPTS*	−.1737	−.1648	−.0574	−.0542
	(.4881)	(.5055)	(.2350)	(.2455)
CFGPCT	6.3856	8.8789	−.9577	−.7437
	(41.845)	(44.255)	(1.5240)	(1.6234)
CFTPCT	−34.488	−34.710	−.8758	−.8752
	(24.464)	(25.291)	(.9126)	(.9530)
CREB*	.8441	.8445	.0611	.0684
	(.5827)	(.6019)	(.1383)	(.1448)
CAWARDS	−3.5230	−3.4365	−.4573	−.4480
	(1.9739)	(2.0656)	(.0708)	(.0752)
RACE	3.1735	3.1728	.1997	.1987
	(3.8542)	(3.9823)	(.1439)	(.1502)
ln(Compensation)	−17.085	−17.731	−.7573	−.8192
	(2.7646)	(3.7618)	(.1007)	(.1401)
S.E.E.	22.754	23.511	.8301	.8668
R^2	.3244	...	.5109	...

* These variables are logged when the dependent variable is ln(DRAFT).

cients—the latter estimates correct for serial correlation in our time series using a method proposed by Pindyck and Rubinfeld (1976). We interpret the GLS results with caution due to the short time series used. In both OLS specifications, PCTWHITE has positive significant (at 2% in col. 1 and 11% in col. 2 on two-tailed tests) effects on home attendance. These OLS results imply that going from an all-black to an all-white team increases attendance by 137,885 fans (col. 3's coefficient applied to the mean attendance level) to 157,040 fans (col. 1). At $10 per head (a conservative estimate of average ticket price plus concession revenue) and 12 players per team, these estimates imply an arena revenue effect of $114,904 to $130,887 per white player. This estimate is somewhat higher than the 20% ceteris paribus racial salary differential (about $80,000) in table 2, suggesting that teams and white players share in the gains from serving fans' desires to see white players. The GLS results show positive but smaller and less significant effects of PCTWHITE on attendance (or its log). The effect on

Table 5
OLS and GLS, Annual Home Attendance (1980–81 to 1985–86)

	Dependent Variable			
	ATTEND (%10³)		ln (ATTEND)	
	(Mean = 440 × 10³)		(Mean = 12.952)	
Explanatory Variables	OLS	GLS	OLS	GLS
CONSTANT	−1254.1	−1247.4	3.6227	−2.8322
	(340.48)	(349.76)	(1.5406)	(1.8165)
YEAR	14.010	14.814	.0344	.0359
	(3.9825)	(4.0703)	(.0100)	(.0104)
WINPCT	473.670	448.25	1.2532	1.1682
	(66.154)	(60.913)	(.1665)	(.1572)
STARS	18.377	13.479	.0230	.0134
	(16.569)	(14.450)	(.0418)	(.0371)
ARENA*	5.2067	.0416	.2220	.1385
	(1.5808)	(.0176)	(.0893)	(.1021)
RACEMSA*	−2.1441	−2.8483	−.0067	−.0084
	(1.1652)	(1.3887)	(.0033)	(.0040)
POPMSA*	.0140	.0170	.1011	.1330
	(.0053)	(.0062)	(.0497)	(.0592)
INCMSA*	.0116	.0126	.3116	.3313
	(.0049)	(.0059)	(.1137)	(.1403)
TEAMS	.8327	.0971	−.0005	−.0004
	(4.7144)	(5.6476)	(.0121)	(.0146)
PRICE	13.533	−3.8426	.0260	−.0259
	(18.132)	(17.926)	(.0448)	(.0458)
PCTWHITE	157.040	96.089	.2726	.1619
	(68.117)	(60.888)	(.1708)	(.1558)
S.E.E.	79.55	65.30	.1992	.1671
R^2	.6304	...	.6079	...
n	138	138	138	138

NOTE.—Units of observation are team years for 23 teams and 6 years.
* These variables are logged when the dependent variable is ln(ATTEND). When the dependent variable is ATTEND, these variables are in thousands.

ATTEND is still significant at 11% (two-tailed test), however, and suggests a per player revenue effect of $80,000.

Table 5 also indicates that teams playing in large arenas, with high winning percentages, and in areas with small numbers of blacks and high incomes, have higher home attendance, other things equal. The effect of price is small, and much smaller than its standard error. Price is obviously an endogenous variable, and it may be picking up demand effects, although excluding price did not affect the results. Finally, when PCTWHITE was interacted with other variables, there was no evidence that the impact of white players on attendance varied with the racial composition of the market. Table 6 contains the results of such interactions. The only significant interaction is a negative one with ARENA: whites add more fans in small arenas.[21] Whites may be especially good draws in cities with small

[21] The other interactions were insignificant as a group as well as individually.

Table 6
Selected OLS Coefficients from Attendance
(Model with PCTWHITE Interactions)

	ATTEND (%10³)	ln(ATTEND)
PCTWHITE	347.88	5.0104
	(746.60)	(13.920)
PCTWHITE * RACEMSA	15.606	.0289
	(12.934)	(.0364)
PCTWHITE * POPMSA*	.0701	.5369
	(.0642)	(.6295)
PCTWHITE * INCMSA*	.0449	.8656
	(.0514)	(1.2736)
PCTWHITE * WINPCT	45.410	−.2098
	(586.86)	(1.5796)
PCTWHITE * STARS	66.944	−.0639
	(139.89)	(.3641)
PCTWHITE * PRICE	−180.79	.5262
	(156.16)	(.4101)
PCTWHITE * ARENA*	−41.052	−1.6329
	(13.809)	(.8130)
PCTWHITE * TEAMS	−8.303	−.0986
	(56.023)	(.1582)
S.E.E.	76.535	.1979
R^2	.6794	.6375

NOTE.—For other variables, see table 5.
* These variables are logged when the dependent variable is ln(ATTEND). When the dependent variable is ATTEND, these variables are in thousands.

arenas such as Portland or Sacramento, as compared with those with large arenas like Detroit.[22]

V. Today's NBA Salaries in Historical Perspective

Since the 1950s, many factors have combined to raise NBA players' salaries: unionization, free agency, the increasing popularity of professional sports, television, and so on (Berry et al. 1986). To provide an indication of how far players' salaries have come, we have, based on our salary model, computed predicted 1985–86 salary levels for selected all-time great players. The results are shown in table 7. Of this group of players, only Wilt Chamberlain at $2.317 million would have been paid at the top levels of the salary structure, as he was when he played. All of these players actually

[22] With the exception of PRICE, our results for attendance are similar to those of Noll (1974, chap. 4, pp. 115–57), who did not include racial makeup of the team or arena size. He suggests including each variable other than RACEMSA or POPMSA as an interaction with POPMSA but with no main term. When this specification was attempted, the results were similar to those in table 5. In particular, the PCTWHITE effect in the Noll-type specification was positive and significant. Finally, when we omit Celtics (due to the uniqueness of Larry Bird, Kevin McHale, et al.), the attendance results get stronger: in OLS regressions the PCTWHITE effect becomes 188.57 (81.38) (ATTEND) and .4239 (.2048) (ln ATTEND).

Table 7
Actual and Predicted Salaries of Selected All-Time Greats

Player	Calendar Year	Year of Career	Actual Salary at the Time ($)	Salary in 1985 Dollars	Predicted Salary ($)*
Elgin Baylor	1962–63	5th	30,000	106,656	1,712,357
Wilt Chamberlain	1972–73	14th	450,000	1,157,143	2,317,489
Bob Cousy	1962–63	13th	50,000	177,760	956,233
Dave DeBusschere	1971–72	10th	100,000	265,622	1,021,142
Walt Frazier	1973–74	7th	300,000	726,221	946,418
John Havlicek	1969–70	8th	140,000	410,820	797,346
Elvin Hayes	1971–72	4th	87,500	232,420	1,349,803
Earl Monroe	1972–73	6th	200,000	514,286	743,270
Bob Pettit	1962–63	9th	30,000	106,656	1,506,648
Oscar Robertson	1971–72	12th	233,333	619,785	1,312,598
Bill Russell†	1966–67	11th	125,001	414,149	1,558,021
Jerry West	1973–74	14th	300,000	726,221	1,331,624

SOURCES.—Koppett (1968)—Hayes and Russell; Koppett (1973)—Monroe; Hirschberg (1963)—Baylor, Pettit, and Cousy; Bradley (1976)—Frazier; Libby (1977)—Chamberlain and West; U.S. Senate Subcommittee on Antitrust and Monopoly (1972)—DeBusschere and Robertson; Halberstam (1981)—Havlicek.

NOTE.—The results are based on wage-level regressions, not log wage regressions. For superstars, it may be unrealistic to say that, for example, each increase in points per game raises salary by the same percentage (as a log wage regression requires). The salary level regression produced a larger race effect ($111,590) (34,743) than did the log wage regression.

* Based on SALARY level OLS regression and statistics as of the indicated year of career.

† Player-coach.

earned in 1985 dollars considerably less money than they would have if they were playing today. In particular, players such as Elgin Baylor, Wilt Chamberlain, Bob Pettit, Bob Cousy, and Bill Russell paid a high price by being born too soon.

VI. Conclusion

In this article, we have investigated racial salary differentials in the NBA. A variety of specifications and statistical techniques indicate that, ceteris paribus, black NBA players earn significantly less than white players by about 20%. In addition, ceteris paribus, home attendance is a positive function of white representation on the team. Such findings are consistent with the notion of customer discrimination. On the other hand, our results for draft position do not indicate discrimination in hiring.[23]

[23] As this article goes to press, we have become aware of two recently completed papers by sociologists on the NBA. First, Wallace (in press) analyzed NBA salary determination. While race was not his major concern, he did find a significant salary advantage for whites, other things equal. However, he did not control for market-related variables other than population or team winning percentage, did not consider the endogeneity of draft position, and did not perform reverse regression tests or seniority interactions. Further, he did not deal with the issue of the source of discrimination against blacks and did not provide an economic framework (profit-maximizing or otherwise) to analyze the consequences of prejudice. Second, Schollaert and Smith (1987) analyzed the impact of team racial composition on

While some of our findings are consistent with the existence of discrimination against blacks in the NBA, two recent analyses of professional baseball during the free-agency era fail to find racial discrimination there (Raimondo 1983; Hill and Spellman 1984). What accounts for the differences between our findings and theirs? We believe that the relative scarcity of white NBA players and the greater visibility of basketball players to fans (compared with baseball) account for our findings. For example, Hill and Spellman (1984) report that about 30.8% of major league baseball players in 1976 were black, a far smaller figure than we find in the NBA. In addition, while the ceteris paribus racial compensation differential we found is large (20%), it exists in the context of roughly equal overall black and white salaries. Social pressures to eliminate discrimination in this situation may well be less than if there were a highly visible black shortfall in mean salaries. The fact that several black players are among the most highly paid in the league may deflect any suspicion that blacks face discrimination.

It is noteworthy that such estimates of blacks' ceteris paribus salary shortfalls are found 10 years after the advent of free agency in the NBA. In addition, the high losses claimed by owners (Berry et al. 1986) suggest that they are under some pressure to maximize profits. If discrimination exists in such a competitive market, then it is either profitable, or owners are willing to take negative profits in order to indulge their preferences for white players. Our results suggest that customer (fan) discrimination may be the ultimate cause of the black shortfall. As long as fans prefer to see white players, profit-oriented teams will make discriminatory salary offers.

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attendance. Using a similar specification to ours, but for a smaller sample of teams and earlier time period (1969–82 or 1977–82), they did not find a race effect. Our attendance results for 1980–86 are probably due to the fact that the league has become three-quarters black in these years. In the earlier periods, blacks comprised 60%–70% of the league. As white players have become more scarce, an increase in their presence now has a greater marginal effect on attendance. The results of our article, compared to earlier work, are thus consistent with the idea of customer discrimination.

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