Pay Discrimination in the NBA Revisited

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An investigation of pay discrimination in the National Basketball Association is conducted using a panel data set spanning the 1990 to 2000 time period. Studies using 1984-1986 NBA season data suggested that salaries for black players were approximately 14 percent to 20 percent lower than those of comparable white players. Evidence suggested this shortfall might be the result of fan discrimination. Studies using data from some seasons in the 1990s found either no evidence of pay discrimination or evidence of pay discrimination only at the highest player performance levels. The results of the current study suggest a lack of pay discrimination in the NBA during this time period. False positive conclusions are possible because of a correlation between height and race.

Introduction

The use of the residual method to measure wage discrimination is fraught with problems. Oaxaca (1973) stated:

> If it were possible to control for virtually all sources of variation in wages, one could pretty well eliminate labor market discrimination as a significant factor in determining wage differentials by sex (or race)...The other extreme is to control for virtually nothing and thereby minimize the role of productivity differences, ΔZ. This is tantamount to declaring at the outset that the two labor inputs are near perfect substitutes and thereby attributing virtually all of the observed wage differential to labor market discrimination. (p. 699)

Between these two extremes the researcher must find adequate and appropriate proxy variables to measure the productivity of workers as well other market factors that could impact wages. Accomplishment of this task still can lead to an understatement or overstatement of wage discrimination using wage equation residuals. For example, if female or minority workers do not invest as heavily in education or training as a rational market response to a lower rate or return on investment due to

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market discrimination, then the resulting estimates using the residual methodology understate the overall effect of discrimination on wages. On the other hand, if race or sex is correlated with an omitted variable that helps to explain productivity, then the resulting estimates of wage discrimination overstate wage discrimination or can lead to a false positive conclusion of the existence of wage discrimination.

Labor market studies of wage discrimination in professional sports do not suffer from data shortages. According to Kahn (2000), the plethora of statistics designed to measure the performance of our professional athletes provides a virtual labor market laboratory for analysts. Problems of correlation between race and productivity measures may still present problems.

This study uses panel data over the 1990-2000 seasons to examine for evidence of pay discrimination.[1] Models are estimated using the standard OLS technique and the random effects formulation.

Previous Studies

Kahn and Sherer (1988), using salary data from the 1985-86 season, found evidence that black players, who represented almost 75 percent of team rosters, received approximately 20 percent less in pay, ceteris paribus, compared to white players. The source of this discrimination was suggested to be fans. An article by Brown, Spiro, and Keenan (1991) using 1984-1985 season salary data found similar results, about a 15 percent pay premium for white players; another article by Koch and Vander Hill (1988) using the same salary data as Brown, Spiro, and Keenan (1991) but more dependent variables estimated about a 9 percent pay premium for whites. Both studies also cite fan discrimination as the driving force behind pay discrimination.

Several articles (Dey, 1997; Gius and Johnson, 1998; and Bodvarsson and Brastow, 1999), using salary data from the late 1980s and 1990s failed to find pay discrimination in this more recent time period. Hamilton (1997) found evidence of racial pay differences only at the upper end of the 1994-1995 season's salary distribution. Dey (1997) replicated the regression models of Brown, Spiro, and Keenan (1991) for each year of the period 1987-1993 as well as a pooled sample. In two of the years and in the pooled sample there was a premium for white players, but this premium disappeared when a dummy variable for centers was added to the model. Dey (1997) concludes that the salary premium for whites reported by Brown, Spiro, and Keenan (1991) may actually be a pay premium for centers because they were more likely to be white. Kahn and Sherer (1988) did have position dummy variables for center and forward, however, and still found pay discrimination.

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[1] The salaries for the study begin with the 1990-1991 season and end with the 2000-2001 season. Because statistics for the three previous seasons are averaged on a per game basis, player stats actually range from the 1987 season to the 1999 season.

Bodvarsson and Brastow (1999) found evidence to suggest that the disappearance of the pay discrimination was the result of a decrease in owner monopsony due to the negotiation of a new NBA collective bargaining agreement between the owners and players in 1988 combined with the addition of four new teams. Further, they provide empirical evidence to support the hypothesis that at least part of the racial salary gap was the result of owner and manager discrimination and not just fan discrimination.

Except for the article by Dey (1997), all the research in this area used single season salary data for model estimation. Dey (1997) used pooled data from five seasons in some of his empirical work; however, he used only OLS regressions.

Model and Data

The model used to estimate player salaries is a standard log-linear format:

$$\mathrm{Ln}(S_t) = \alpha + \beta_1 X_1 + \beta_2 X_{2t-1} + \gamma_t + \varepsilon_t \qquad (1)$$

where:

S_t = The salary of the player in current dollars in year t;
X_1 = A vector of time-invariant player characteristics;
$X_{2(t-1)}$ = A vector of past performance statistics and experience measures;
γ_t = A vector of time trend variables; and
ε_t = A vector of disturbances.

The vector of time-invariant explanatory variables in the model includes:

White = 1 if the player is white, 0 otherwise;
Drnumber = The position the player was taken in the NBA draft;
Height = The player's height in inches.

As discussed previously, the race of a player may impact salary if general managers pay white players a premium based on a prejudice by predominantly white fans. The draft number of each player indicates the position they were selected in the draft. Values range from 1 to 60.[2] Salary should be inversely related to draft number. While other articles had dummy variables for player position, e.g., center and forward, the inclusion of a continuous measure of height would appear to be more appropriate. Height is designed to capture any pay premium over and above performance statistics such as rebounds and blocked shots.[3] Taller players may

[2] In the most recent years there are 29 teams in the NBA, so the first round draft picks ranged from 1 to 29. The second round picks ranged from 30 to 58. In other seasons there were fewer teams in the league. Also there were more than two rounds in the draft prior to the 1989 season. I entered a value of 60 for any player drafted lower than the second round in any season.

[3] Height is not significant as a determinant of team win percentage when blocks and rebounds are included in a regression model as independent variables; it is if they are excluded. Team data on win percentage from the 1995-1996 seasons was regressed on team performance statistics including field goal percentage, free throw percentage, three-point field goal percentage, rebounds, steals, assists, blocks, turnovers, and average team height.

provide an intimidating presence on the court causing opponents to alter their shots or passes; this may decrease opponents' shooting percentages and increase their turnovers. Given the inelasticity of supply in regards to height the coefficient of height is expected to be positive. [4]

The vector of past performance statistics and experience measures includes:

Yrs = The player's years of NBA experience;
Yrsq = The square of the player's years of NBA experience;
Pts = Points scored per game for the three previous seasons;
Reb = Total rebounds per game for the three previous seasons;
Ast = Assists per game for the three previous seasons; and
Blk = Blocks per game for the three previous seasons.

All of the performance measures should have a positive influence on salary. Salary should also rise as players gain more experience, so the coefficient of years is expected to be positive. Player skills deteriorate over time; the inclusion of years squared is designed to capture this effect. Its sign should be negative.

The vector of time trend variables includes:

YR90 = 1 if the year is 1990, 0 otherwise, through …
YR99 = 1 if the year is 1999, 0 otherwise.

No dummy variable is established for the most recent year of salary data.

The model is estimated first using the standard OLS regression technique assuming each observation for a player is separate and independent. Obviously however, this technique may yield biased estimators for the model because the performance measures may display a high degree of correlation between years for the same player. In this case the model needs to be rewritten as:

$$\mathrm{Ln}(S_{it}) = \alpha + \beta_1 X_{1it} + \beta_2 X_{2it-1} + \gamma_t + \varepsilon_{it} \tag{2}$$

where i refers to the individual player. Two options for estimating this model are the fixed effects approach and the random effects approach. In the fixed effects formulation of the model differences across individuals are captured in differences in the constant term; thus, any time-invariant personal characteristics are dropped from the regression. In this formulation of the model it is impossible to determine if differences exist between white and black players due to race or some other time-invariant variable. Therefore, the fixed effects model will not be used.

In the random effects formulation the differences between individuals is modeled as parametric shifts of the regression function. This technique of estimating panel data allows for estimates of all of the time-invariant personal characteristics as well as the performance statistics. Breusch and Pagan (1980) developed a Lagrange

[4] It is interesting to note that players drafted in the first round were, on average, a half inch taller than those drafted in second round, who were, in turn, a half inch taller on average than those who were undrafted but made rosters.

multiplier test (LM Test) for the appropriateness of the random effects model compared to the OLS format.[5]

Salaries for the players were obtained from a website maintained by Patricia Bender; she gathered the data from articles in the *Dallas Morning News*, *USA Today,* and other sources. Player statistics and characteristics were obtained from the *Official NBA Encyclopedia* and various issues of the *Sporting News Official NBA Register*.

Empirical Results

OLS Results

In Table 1 OLS regression results are reported. In column I the standard model shown as equation 1 was estimated excluding the variable height. The results indicate a pay premium of approximately 7.4 percent for white players; the coefficient of the race variable is statistically significant. An examination of the mean values of the independent variables, shown in Table 2, indicates that white players in the sample are about two inches taller on average than black players. Adding height to the regression model causes the coefficient of race to become insignificant while the coefficient of height is statistically significant; these results are shown in column II of Table 1.

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[5] The Baltagi and Li modification allows for unbalanced data and reduces to the standard formula:

$$\lambda_{LM} = \frac{nT}{n(T-1)}\left(\frac{\sum_i\left(\sum_t \varepsilon_{it}\right)^2}{\sum_i\sum_t \varepsilon_{it}^2} - 1\right).$$

Under the null hypothesis, λ_{LM} is distributed as chi-square with one degree of freedom. See Stata Release 6, Reference SU-Z p. 439 for details or Greene (2000, pp. 572-573). Unfortunately, a rejection of the null hypothesis that a single constant term is inappropriate for the data does not preclude the appropriateness of the fixed effects model. The Hausman test is designed to test for fixed and random effects regressions. The test measures whether there is a statistically significant difference between the coefficients estimated by the fixed effect versus the random effects model.

The test statistic for the Hausman test is:

$$W = (\beta_f - \beta_r)'(V_f - V_r)^{-1}(\beta_f - \beta_r)$$

where all vectors and matrices have the row and column corresponding to the intercept removed along any row and column corresponding to a parameter that cannot be estimated by the fixed-effects estimator. Under the null hypothesis, W is distributed as chi-square with k degrees of freedom, where k is the number of estimated coefficients in □ excluding the intercept and time-invariant regressors. See Stata Release 6, Reference SU-Z, p. 439 for details. If there is a significant difference, then either the model is mis-specified or the assumption that the random effects are uncorrelated with the independent variables is incorrect. If the model is assumed correctly specified, then rejection of the null hypothesis suggests that the random effects model is inappropriate.

Table 1—Combined OLS Regression Results: (1989-2000)

Variable	I	II
Constant	13.5067	11.3520
	(286.413)	(32.508)
Race	.0739	.0253
	(2.844)	(.939)
Draft Number	-.0069	-.0065
	(-14.504)	(-13.625)
Years	.0955	.0969
	(10.044)	(10.239)
Years Squared	-.0055	-.0057
	(-8.631)	(-9.00)
Points	.0530	.0533
	(18.027)	(18.245)
Rebounds	.0801	.0675
	(12.283)	(9.933)
Assists	.0732	.1008
	(10.652)	(12.369)
Blocks	.2117	.1635
	(8.534)	(6.318)
Height		.0274
		(6.227)
Yr90	-1.3856	-1.3930
	(-27.672)	(-27.950)
Yr91	-1.2148	-1.2262
	(-25.274)	(-25.619)
Yr92	-1.133	-1.1408
	(-23.700)	(-23.975)
Yr93	-.9576	-.9609
	(-19.896)	(-20.065)
Yr94	-.8374	-.8454
	(-17.577)	(-17.829)
Yr95	-.6723	-.6820
	(-14.450)	(-14.725)
Yr96	-.6441	-.6514
	(-13.641)	(-13.861)
Yr97	-.4916	-.4981
	(-10.591)	(-10.781)
Yr98	-.3184	-.3225
	(-6.793)	(-6.914)
Yr99	-.1141	-.1200
	(-2.411)	(-2.547)
Adjusted R-Square	.62	.63
Number of Observations	3781	3781

T-statistics are in parentheses below the coefficients

The coefficients of all the other explanatory variables in the model are properly signed and statistically significant. The time trend dummy variable coefficients are negative and get progressively smaller in absolute value as the years progress. This is to be expected because the omitted time trend dummy variable is for the most recent salary year in the study.

Table 2—Mean Values of Variables in Model (1989-2000)

Variable	Black	White
Race (number of obs.)	2986	795
Draft Number	25.56	30.16
Years	5.48	5.42
Years Squared	44.38	42.68
Points (per game)	10.25	7.77
Rebounds (per game)	4.23	3.69
Assists (per game)	2.34	1.77
Blocks (per game)	.52	.48
Height (in inches)	78.79	80.77
Ln (S)	14.10	13.87
Salary	$2,188,708	$1,739,788
Salary inYr90	$920,271 (n = 214)	$830,013 (n = 72)
Salary inYr91	$1,038,540 (n = 248)	$976,691 (n = 81)
Salary inYr92	$1,213,008 (n = 259)	$1,022,600 (n = 77)
Salary inYr93	$1,476,928 (n = 249)	$1,069,325 (n = 78)
Salary inYr94	$1,644,364 (n = 265)	$1,408,572 (n = 72)
Salary inYr95	$1,892,293 (n = 300)	$1,598,463 (n = 67)
Salary inYr96	$2,322,487 (n = 275)	$1,692,146 (n = 72)
Salary inYr97	$2,596,779 (n = 299)	$1,834,608 (n = 70)
Salary inYr98	$2,843,519 (n = 284)	$2,553,585 (n = 71)
Salary inYr99	$3,536,160 (n = 278)	$3,272,555 (n = 64)
Salary in Yr2000	$3,777,413 (n = 315)	$3,275,943 (n = 71)

The simple OLS regression results indicate that possible pay discrimination is simply the result of model misspecification. Correlation between height and race generates a false positive result for pay discrimination. While other studies have included dummy variables for position in their regressions, e.g., center =1 if the player is a center, zero otherwise, none have included a continuous measurement of height. It should be noted also that Hoang and Rascher (1999) found evidence of "exit discrimination"—white players had longer career lengths, ceteris paribus, compared to black players. Groothuis and Hill (2004), however, have found no such evidence when height is included in their duration model.

Random Effects Results and LM Test Results

The Lagrange Multiplier test statistic is 488.89, which greatly exceeds the 95 percent chi-squared with one degree of freedom, 3.84. Thus, the simple OLS regression model with a single constant term is inappropriate. Table 3 reports results of the random effects regression formulation. The overall results (column I) again suggest that race is a statistically significant determinant of the natural log of salary. The inclusion of height (column II) causes the coefficient of race to again become insignificant while the coefficient of height is significant, the same as in the OLS results.

The coefficients of the experience and performance measures are again all properly signed and significant. The coefficients of the time-invariant personal characteristics, except for race in column II, are also all properly signed and significant.[6]

Table 3—Random Effects GLS Combined Regression Results (1989-2000)

Variable	I	II
Constant	13.5164	11.3154
	(243.428)	(22.767)
Race	.0855	.03418
	(2.060)	(0.798)
Draft Number	-.0083	-.0079
	(-11.113)	(-10.561)
Years	.0874	.0882
	(8.604)	(8.711)
Years Squared	-.0047	-.0049
	(-6.849)	(-7.102)
Points	.0532	.05379
	(12.342)	(12.528)
Rebounds	.0793	.0677
	(8.175)	(6.778)
Assists	.0723	.0965
	(6.917)	(8.229)
Blocks	.1849	.1401
	(4.865)	(3.577)
Height		.0280
		(4.456)
Yr90	-1.3262	-1.3333
	(-26.577)	(-26.805)
Yr91	-1.1902	-1.1992
	(-25.174)	(-25.432)
Yr92	-1.1300	-1.1358
	(-24.430)	(-24.624)
Yr93	-.9666	-.9713
	(-21.238)	(-21.397)
Yr94	-.8614	-.8686
	(-19.541)	(-19.741)
Yr95	-.6930	-.7010
	(-16.358)	(-16.568)
Yr96	-.6787	-.6852
	(-16.113)	-(16.292)
Yr97	-.5288	-.5341
	(-12.984)	(-13.134)
Yr98	-.3617	-.3651
	(-8.898)	(-8.995)
Yr99	-.1445	-.1484
	(-3.594)	(-3.695)
R-Sq: within	.3941	.3954
between	.7117	.7148
overall	.6247	.6285

Z-statistics are in parentheses below the coefficients

[6] The Hausman Test statistic is 83.63; the critical chi-square value is 27.59. I cannot therefore conclude that the individual effects are uncorrelated with the other regressors in the model.

OLS Results for Individual Years

The simple log linear format of the model, without a time trend variable, was run for each separate year. The results are shown in Table 4. In a couple of the yearly results, 1990 at the 95 percent level and 1991 at the 99 percent level, the coefficient of race is significant. When height was added to the model (these results are not shown in the interest of brevity), the coefficient of race became insignificant.

Table 4—Yearly OLS Regressions Results (1990-2000)

Variable/ Year	1990	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000
Constant	12.28	12.56	12.55	12.73	12.89	12.90	12.89	13.07	12.95	13.07	12.94
	(104.92)	(121.38)	(91.99)	(109.37)	(105.43)	(122.70)	(99.62)	(107.64)	(111.40)	(90.95)	(105.27)
Race	0.15	0.19	0.130	-0.05	-0.002	0.01	-0.01	-0.03	0.09	0.16	0.13
	(1.88)	(2.61)	(1.36)	(-.592)	(-.022)	(0.08)	(-.160)	(-.364)	(1.16)	(1.59)	(1.61)
Draft	-0.004	-0.01	-0.01	-0.01	-0.01	-0.01	-0.01	-0.01	-0.01	-0.01	-0.01
Number	(-.3.42)	(-5.33)	(-4.52)	(-3.89)	(-5.01)	(-5.76)	(-5.64)	(-6.25)	(-4.02)	(-2.70)	(-3.30)
Years	0.07	0.05	0.08	0.08	0.05	0.07	0.11	0.10	0.17	0.12	0.19
	(2.14)	(1.69)	(2.07)	(2.54)	(1.52)	(2.49)	(3.33)	(3.45)	(5.82)	(3.45)	(6.66)
Years	-0.005	-0.004	-0.005	-0.005	-0.003	-0.004	-0.01	-0.01	-0.01	-0.01	-0.01
Squared	(-1.84)	(-1.86)	(-1.61)	(-2.18)	(-1.27)	(-2.04)	(-3.54)	(-3.29)	(-4.98)	(-2.85)	(-5.23)
Points	0.04	0.04	0.06	0.05	0.05	0.06	0.05	0.05	0.05	0.07	0.06
	(5.26)	(5.56)	(5.45)	(4.95)	(5.46)	(6.52)	(5.01)	(4.83)	(5.37)	(5.94)	(5.47)
Rebounds	0.08	0.09	0.05	0.09	0.09	0.08	0.09	0.08	0.06	0.06	0.09
	(4.04)	(4.71)	(1.98)	(4.42)	(4.44)	(4.49)	(4.15)	(3.83)	(2.97)	(2.52)	(3.77)
Assists	0.04	0.04	0.04	0.07	0.07	0.08	0.10	0.09	0.10	0.09	0.10
	(2.16)	(2.57)	(1.65)	(3.09)	(3.07)	(3.90)	(3.86)	(3.78)	(4.14)	(3.45)	(4.26)
Blocks	0.22	0.18	0.23	0.10	0.06	0.14	0.19	0.29	0.33	0.27	0.32
	(3.17)	(2.64)	(2.56)	(1.35)	(0.79)	(1.74)	(2.12)	(3.38)	(4.33)	(2.81)	(3.64)
Adjusted R-Square	0.56	0.56	0.47	0.54	0.53	0.62	0.58	0.61	0.61	0.54	0.63
Number of Obs.	286	329	336	327	337	367	347	369	355	342	386

T-statistics are in parentheses below the coefficients

Application of Model to 1984-1985 Data

To determine if the findings of pay discrimination in the Kahn/Sherer article could be replicated with the current model, the basic OLS format was applied to the salary data from the 1984-1985 season. The results are reported in Table 5, column I. The coefficient of race in the model suggests about a 16 percent pay premium for white players, ceteris paribus. These results are comparable to the findings of Kahn and Sherer. Inclusion of the height variable in the model (column II) causes the coefficient of race in the model to become insignificant at the 95 percent level but still significant at the 90 percent level. The coefficient of height is also not significant at commonly accepted levels.

Table 5—OLS Regressions Results (1985-1986 Salary Data)

Variable	I	II
Constant	11.3634	9.6275
	(112.802)	(8.886)
Race	.1574	.1215
	(2.356)	(1.731)
Draft Number	-.0054	-.0054
	(-5.348)	(-5.382)
Years	.0805	.0854
	(2.566)	(2.717)
Years Squared	-.0031	-.0037
	(-1.277)	(-1.486)
Points	.0515	.0530
	(7.417)	(7.591)
Rebounds	.0773	.0631
	(4.778)	(3.428)
Assists	.0370	.0521
	(2.124)	(2.639)
Blocks	.0926	.06951
	(1.913)	(1.380)
Height		.0222
		(1.609)
Adjusted R-Square	.69	.69
Number of Observations	244	244

T-statistics are in parentheses below the coefficients

The coefficients of all the performance measures are again all properly signed and significant, except for the coefficient of blocks in column II. The coefficients of the experience measures do not provide as much explanatory power in a one-year time frame. The coefficients of years squared are insignificant in both versions of the model. The coefficients of draft number are significant in both models.

It is interesting to note that the adjusted R-square of the regressions using the 1985-1986 salary data (the same as in the Kahn/Sherer research) are substantially higher than any of the individual yearly regression results reported in Table 4 from the 1990-2000 period. As salaries have skyrocketed and the dispersion of wages has exploded, the ability of regression analysis to predict salaries has been weakened.

Conclusions

The results of the empirical research indicate there does not appear to be any pay discrimination present in the 1990s in the NBA. While simple OLS regression results from the pooled data or individual years could yield a false positive, in all cases these are the result of correlation between race and height. White players in the NBA are two inches taller on average than black players. The apparent pay premium for white players is a height premium. Results from the application of the current model to data from a previous time period in which other authors reported pay discrimination is not as instructive. Previous conclusions of pay discrimination are not as robust when a measure of height is included.

The results of this study have implications in a much broader context than just the market for professional basketball players. First, findings of wage discrimination based on observations from a brief time period should be viewed cautiously. Second, alternatives specifications of the model should be explored to examine the robustness of the results in order to avoid false positive conclusions.

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