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Abstract: We use data from the National Basketball Association (NBA) to analyze the impact of minimum salaries on an employee's career length. The NBA has a salary structure in which the minimum salary a player can receive increases with the player's years of experience. Salary schedules similar to the NBA's exist in public education, federal government agencies, the Episcopalian church, and unionized industries. Even though the magnitude of the salaries in the NBA differs from other industries, this study provides insight to the impact of this type of salary structure on career length. Using duration analysis, we find statistically significant evidence that minimum salaries shorten career length.

Introduction

One of the tenets of the U.S. private sector industrial relations system is that union and management use collective bargaining to reach a mutually agreeable contract. Compromise is inevitable and neither side is ever totally satisfied with the outcome. Confounding the whole process however is the law of unintended consequences. On occasion both sides agree to a mutually beneficial position only to have profit maximizing market behavior thwart their efforts. Such may have been the case when the National Basketball Players Association (NBPA) and the NBA owners reached an accord in 1998 to end a lockout that included a provision to pay veteran players a minimum wage scale based on years of experience. This egalitarian wage policy was agreed to because it was thought to provide a fairness to the players in the lower part of the wage distribution.

Unions have been documented to shift rents disproportionately to the lower skilled workers. (Freeman and Medoff 1984; Black and Parker 1985; and Parsons 1991.) Hill and Groothuis (2001) suggest that union rents can be redistributed from superstar's salaries to the average player's salaries because of the median voter rules in collective bargaining. Freeman and Medoff (1984) state: "In a simple voting model of union behavior in which union polices are set by the median voter a pattern of lower median than mean wages is likely to result in a policy of greater gains for the lower paid." Parsons (1992) finds that union wage distributions are more egalitarian than the median voter model predicts. One explanation he offers is that egalitarian wage polices may be favored by the majority as a means of union solidarity during strikes. These union voice median voter tendencies may have been the catalyst to the implementation of a minimum wage scale based on experience.

The minimum wage scale, however, may provide market incentives to shorten careers of low skilled experienced players by replacing the high paid experienced players with cheaper less experienced players. Kahn (2000) suggests that sports business decisions provide labor market laboratories to tests implications of the overall labor market. We suggest that egalitarian minimum wage increases in sports markets may provides insights where salary schedules are similar to the NBA's such as in public education, federal government agencies, the Episcopalian church, and other unionized industries.

Institutional Background on NBA Minimum Wage Scale

The dataset for this analysis begins with the 1990-91 NBA season. At this time the NBA had a single minimum wage scale for players.[1] Actual figures are listed in Table 1. Beginning with the 1995-96 season, the collective bargaining agreement (CBA) between the owners and players' union called for a rookie minimum scale and a somewhat higher minimum scale for veteran players. For instance, as shown in Table 1, the rookie minimum was $200,000 in 1995-96 and the veteran minimum was $225,000. In the compromise CBA that concluded the lockout at the beginning of the 1998-99 season a minimum salary scale was set for each separate experience level from rookies through ten year and up veterans. Each additional year of experience garnered a higher guaranteed minimum salary (see Table 2 for details).

With the introduction of higher salaries for veteran players the CBA also created an economic incentive for teams to jettison older players in favor of younger ones. For instance, in

[1] Actually beginning with the introduction of a salary cap in the NBA in 1983 teams that had limited or no room under the salary cap for their first round draft picks were required to pay them a minimum scale salary that was higher than the overall minimum used in this study. These first round picks were under guaranteed contracts and received higher salaries when cap room became available. Therefore we did not use these as an actual minimum wage scale constraint in this analysis.

the 2008-09 season a ten year and up veteran player making the minimum salary earned $820,161 more than a second round draft pick earning the league minimum. Anticipating this dilemma, the negotiators of the CBA added requirements that only the portion of any veteran's salary at or below the four year level counts against the salary cap and any additional salary above the four year level is reimbursed to the team from a league-wide fund following the completion of the season. In the 2005 CBA the experience level for the "counts against cap" and payment for overage from a league-wide benefit fund was lowered from four to two years. Under these terms the difference between a rookie minimum salary in the 2008-09 season and a two year minimum salary was only $355,467.

The lowering of the number of years for the "counts against cap" and payment from the league-wide benefit fund makes one suspicious that teams were eliminating veterans in favor of rookies. This study will investigate this issue using duration analysis.

Theory and Model

A number of previous studies of professional sports leagues have made the assumption that teams maximize profits (Hamlen Jr., 2007; Fort and Quirk, 1995; Scully, 1974; Vrooman, 1995). We also assume that NBA teams maximize profits. The revenue generated by player *i* is equal to the marginal revenue product from employing player *i*, MRPi. MRPi is also the maximum amount a team is willing to pay player *i*. In a perfectly competitive industry an employee's salary is equal to their marginal revenue product. In the NBA, employers may have monopsony power because players chosen in the NBA draft can sign contracts only with the team that drafts them.

Even though teams have some monopsony power, evidence reveals that player salaries are close to their marginal revenue products (MacDonald and Reynolds, 1994; Rosen and Sanderson, 2000). The salary paid to player *i*, Si, is an amount that is negotiated between the player and the team (Conlin and Emerson, 2003). In the seasons following the 1997-98 season, a player faces a different mandated minimum salary, MMSi, at each level of experience.

Consider the following model based on Ducking and Bollinger (2012) where a retained player's negotiated salary is greater than or equal to his mandated minimum salary and less than or equal to the marginal revenue product plus LS_i, where LS_i is the subsidy paid by the league out of a league wide fund to the team for player i's salary, $MMSi \leq Si \leq MRPi + LSi$. Profits per player are the difference between the player's MRP and the total amount of compensation paid to the player by the team. The total compensation paid to the player includes the salary plus any bonuses. If a player earns the minimum salary for his experience level without any bonuses, $S_i = TS_i + LS_i$, where TS_i is the amount of player i's salary paid by the team. For simplicity, we define profits per player as the difference between the value of a player's marginal product and the salary paid to the player by the team, $MRP_i - TS_i$. Even though profits per player are MRP minus the player's total compensation, MRP – TS captures the impact of minimum salaries. Teams employ the N players for which the sum of profits per player is the largest. Therefore, each team's total profits are a function of profits per player,

$$\text{profits} = f(MRP_1 - TS_1, MRP_2 - TS_2, \ldots, MRP_N - TS_N) \quad (1)$$

where

$$MRP_1 - TS_1 > MRP_{N+1} - TS_{N+1},\ MRP_2 - TS_2 > MRP_{N+1} - S_{N+1},\ \ldots, MRP_N - TS_N > MRP_{N+1} - TS_{12}.$$

$MRP_{N+1} - TS_{N+1}$ represent the profits per player for the player with the highest profits per player who is currently not on the team. Even though the player's marginal revenue product may depend on the productivities of the other players employed by the team, team management has an idea of a player's marginal revenue product given various combinations of players. Management ultimately chooses the combination of players in a manner that maximizes profits.

The minimum salary schedule has the ability to shorten the career length of players through two mechanisms. The first mechanism operates when the minimum salary schedule causes the salary paid by the team to exceed the player's marginal revenue product. When the salary paid by the team exceeds player *i*'s marginal revenue product, the team is going to dismiss player *i*. Player *i*'s career length is shortened if no other team employs player *i*. Therefore, if

$$MRP_i - TS_i < 0, \quad (2)$$

Player *i*'s career length may be shortened. This mechanism can occur when a team is forced to give a player a raise that makes the salary paid by the team to player *i* greater than his marginal revenue product. The NBA's salary structure along with the salary structure of any other industry that employs a minimum salary schedule has the ability to artificially increase the minimum cost of an input for teams whether there is a corresponding increase in the value of the player or not. The second mechanism operates when the minimum salary schedule causes the profits per player from hiring player *i* to fall below the profits per player from hiring player *j*. Player *j* is the individual with the largest profits per player not previously employed by the team who is capable of replacing player *i*'s duties on the team. If this occurs, the team is going to dismiss player *i* and hire player *j*. Player *i*'s career length is shortened if no other team employs player *i*. Therefore, when

$$MRPi - TSi < MRPj - TSj, \quad (3)$$

player *i*'s career length may be shortened. This mechanism can occur when a team is forced to increase the part of a player's salary that is paid by the team that makes the profits per player from employing player i less than the profits per player from employing player *j*.

Duration Analysis

Berger and Black (1998) developed a semi-parametric hazard function technique to examine the duration of periods in which recipients received Medicaid payments. Groothuis and Hill (2004) adapted the methodology to examine issues regarding racially motivated retention of white players in the NBA brought to light by Hoang and Rascher (1999); Groothuis and Hill (2008) used the same methodology to revisit claims by Jiobu (1988) of "so-called" racial exit discrimination in Major League Baseball. Ducking and Bollinger (2012) used a Weibull proportional hazard model to estimate the effect of mandatory minimum salaries in the NFL on career length. To examine the effect of mandatory minimum salary scales in the NBA on player duration we have chosen to use the technique employed previously by Groothuis and Hill (2004 and 2008). While the Ducking and Bollinger (2012) approach allows for direct estimates to be made concerning the length of a player's career it suffers from restrictions on the sample. All players must start their careers during the 2000 through 2008 NFL seasons therefore veteran players who are already in the league prior to the 2000 NFL Season cannot be considered. This may bias the estimates. The semi-parametric hazard function technique used in this study will allow for estimates of the impact of mandatory minimum salary scales on the probability that a player is retained from one season to the next. It does not require any left censoring of the dataset. Since our data is at the season level we calculate our hazard model as a discrete random variable. As with Berger, Black, and Scott (2004), we model the durations of a single spell and

assume a homogeneous environment so that the length of the spell is uncorrelated with the calendar time in which the spell begins. This assumption lets us treat all the players' tenure as the same regardless of when it occurred in the panel study. For instance, all fourth year players are considered to have the same base line hazard regardless of calendar time so a fourth year player in 1990 has the same baseline hazard as a fourth year player in 1997.

To understand how stock data influences a likelihood function we follow the notation of Berger, Black, and Scott (2004). Suppose the probability mass function (pmf) of durations is defined as $f(t,x,\beta)$, where t is the duration of the career, x is a vector of performance and personal characteristics, and β is a vector of parameters. Now denote F(t,x,β) as the cumulative distribution function; then the probability that a career lasts at least t° years is simply 1 - F(t°,x,β). If we define the hazard function as h(t,x,β) ≡ f(t,x,β) / S(t,x,β) where S is the survivor function, $S(t,x,\beta)=\prod_{i=1}^{t-1}[1-h(i,x,\beta)]$ and apply the definition of conditional probabilities, we may express the pmf as

$$f(t_i,x_i,\beta)=\prod_{j=0}^{t_i-1}[1-h(j,x_i,\beta)]h(t_i,x_i,\beta). \qquad (4)$$

If we have a sample of n observations, $\{t_1, t_2, \ldots, t_n\}$, the likelihood function of the sample is

$$L(\beta)=\prod_{i=1}^{n} f\ (t_i,x_i,\beta)=\prod_{i=1}^{n}\left(\prod_{j=1}^{t_i-1}[1-h(j,x_i,\beta)]\,h(t_i,x_i,\beta)\right). \qquad (5)$$

Often it is not possible to observe all careers until they end, hence careers are often right-censored. Let the set A be the set of all observations where the players' careers are completed and the set B be the set of all observations where the careers are right censored. For the set of right-censored observations, all we know is that the actual length of the career is greater than t_i, the observed length of the career up through the last year. Because we know that the actual length of the career is longer than we observe then the contribution of these observations to the likelihood function is just the survivor function (S).

To introduce stock sampling, let the set C be the set of careers that were in progress when data collection began. For these observations, we know that the career i has lasted for r years before the panel begins so the likelihood must be adjusted by the conditional probability of the career having length r. Of course, some stock-sampled observations may be right-hand censored. Let the set D be the set of all stock-sampled observations that are also right-hand censored. An example of a career that is both right and left censored would be a player that starts his career prior to 1989 and ends his career after 2008, an unlikely event. Taking into account all four sets: A, B, C, and D the likelihood function becomes

$$L(\beta) = \prod_{i \in A}\left(\prod_{j=1}^{t_i-1}[1-h(j,x_i,\beta)]\,h(t_i,x_i,\beta)\right) \times \prod_{i \in B}\left(\prod_{j=1}^{t_i-1}[1-h(j,x_i,\beta)]\right) \times \prod_{i \in C}\left(\prod_{j=r_i}^{t_i-1}[1-h(j,x_i,\beta)]\right) h(t_i,x_i,\beta) \times \prod_{i \in D}\left(\prod_{j=r_i}^{t_i-1}[1-h(j,x_i,\beta)]\right) \quad (6)$$

In equation (3) the contribution of censored, stock-sampled observations to the likelihood function is strictly from the last two terms; such observations simply provide information about the survivor function between (r,t).

Thus we, as Berger, Black and Scott (2004), have expressed the likelihood function as a function of the hazard functions. All that remains is to specify the form of a hazard function and estimate by means of maximum likelihood estimation. As the hazard function is the conditional probability of exiting NBA given that the NBA career lasted until the previous season, the hazard function must have a range from zero to one. In principle, any mapping with a range from zero to one will work. For our purposes we choose the logit model.

The intuition behind the logit model for the hazard function is relatively simple. For each year during the survey in which the player is in NBA, the player either comes back for another season or ends his career. If the career ends, the dependent variable takes on a value of one; otherwise, the dependent variable is zero. The player remains in the panel until the player exits NBA or the panel ends. If the panel ends, we say the worker's spell is right-hand censored. Thus a player who begins his NBA career during the panel and plays for 6 years will enter the data set 6 times: the value of his dependent variable will be zero for the first 5 years (tenure one through five) and be equal to one for the sixth year.

To illustrate a stock sample consider another player who enters the panel with 7 years of NBA job tenure prior to 1989 the first year of the panel, then plays for an additional 3 years for a 10 year career. For this player we ignore his first 7 years of tenure because he is left-hand censored. As the equation of the likelihood function with stock data indicates, the duration of a NBA career prior to the beginning of the panel makes no contribution to the value of the likelihood function. Therefore only years 8 through 10 will enter the data set with the dependent variable taking on the value zero for years 8 and 9 and in the 10^{th} year it takes on a value of one with this player appearing in the data set a total of 3 times. Note for all players who are right-

hand censored, we do not know when their career ends so their dependent variables are always coded as zero.

Because the players in the panel have varying degrees of job tenure prior to the beginning of the panel, we identify the hazard function for both long and short careers. The disadvantage to this approach is that the vector γ_t of equation (3) can be very large. In our study it would require 21 dummy variables. We also run into problems with the dummy variable technique because we have too few players who have long careers. To simplify the computation of the likelihood function and be able to keep the long careers, we simply approximate the γ_t vector with a 4th order polynomial of the players' tenure in NBA, which reduces the number of parameters to be estimated from 21 to 4. Thus, the hazard function becomes

$$\Pr(t, x\beta) = \Pr(\phi(t) + x\beta), \tag{7}$$

where $\phi(t)$ is a 4th order polynomial of the player's tenure in NBA. The 4th order polynomial therefore includes tenure to the first, second, third, fourth and fifth powers. Once again, we choose the Taylor series approximation technique over using tenure dummies due to the small number of observations for high tenures.[2]

In addition to the experience variable and the related first, second, third, and fourth order polynomial of experience we include several player performance measures. Independent performance variables included in the model are: points scored per minute of playing time; total

[2] When higher order polynomials of the fifth and sixth power are included results do not change suggesting that a fourth order polynomial is flexible enough to capture the influence of the base line hazard.

rebounds per minute; assists per minute; steal per minute; blocks per minute; and turnovers per minute.[3] For positive performance measures such as points, rebounds, assists, steals, and blocks it is expected that apriori higher values of any of these variables should increase the probability of duration. Likewise, an increase in turnovers should decrease the probability of duration. The only biographical data included in the study is height; it is expected that taller players last longer in the league. The draft number of each player is included as a proxy for player potential following college or a foreign league career; the draft number ranged in value from one for the first player selected in the draft to 60 for the last player selected in the second round.[4] It is expected apriori that a lower draft number will improve duration; therefore the coefficient of draft number should be negative.

The number of teams in the NBA rose from 27 teams at the beginning of this study to 29 teams at the beginning of the 1995-96 season to 30 teams beginning with the 2004-05 season. Since the number of teams obviously affects a players probability of retention on a roster from one season to the next we created dummy variables for the time periods 1995-2003 and 2004-2008, leaving 1990-1994 as the control group.

To examine the effect of the collectively bargained minimum wage scale salaries on the duration we utilize three separate measures. First, following the approach of Ducking and Bollinger (2012) we use the mandatory raise in base salary that a player must receive from year t

[3] Annual statistics were obtained from Doug's MLB and NBA Stat Home: http://www.dougstats.com/. Biographical information on each player came from a variety of sources including the third edition of the NBA Encyclopedia, various editions of The Sporting News Official NBA Player Register, Wikipedia, and/or NBA.com.
[4]Since 2005 there are 30 teams in the league. Beginning with the 1989 draft there were only two rounds. There were three rounds in the 1988 draft. There were 7 rounds from 1985-1987; there were 10 rounds from 1974-1984. Draft numbers in these seasons had much higher ranges in value. For undrafted players in seasons with only two rounds players were assigned a draft number of 65.

to year t+1 in order for the player to earn the minimum salary for year t+1. All salaries are adjusted to 2008 dollars to alleviate inflation impacted results. Second, to examine the dilemma a team general manager faces in trying to decide whether or not to keep a veteran player earning a league minimum or a rookie earning a lower minimum we subtract the rookie minimum from the veteran minimum, again using real dollar figures. Lastly, to take in to consideration the collectively bargained institutional requirements designed to encourage teams to retain veteran players we subtract the rookie minimum from the counts against the cap figure for veteran players.

Empirical Results

Descriptive statistics for the variables used in the logit regression are found in Table 3. Of particular interest are the values of the three measures of pay differentials used in this study. All of the values of these variables are measured in 2008 dollars. The figures shown in Table 3 are the means and standard deviations of the variables. The real mandatory raise is significantly lower than the real difference in a veteran's minimum salary versus the rookie minimum or the real difference between a veteran's minimum salary versus the counts against cap amount. Unlike the NFL analyzed by Ducking and Bollinger (2012), the NBA does not have as many veteran players that earn the league minimum from one year to the next. The NFL has a larger roster, more injuries, and special teams. All of these factors lead to situations where journeyman players earn minimum salaries while biding their time for a chance to start. The NFL also has a hard salary that puts pressure on teams to keep the salaries of these players at minimum levels. The NBA has a soft salary cap with several exceptions some of which are designed to increase the salaries of veteran non-star players (Hill and Groothuis 2001).

The results of the logit analysis utilizing real mandatory raise as an explanatory variable are shown in Table 4. In all of the time periods shown, draft number, points per minute, rebounds per minute, assists per minute, blocks per minute, and turnovers per minute are correctly signed and significant at the 1% level. Steals per minute are significant at the 5% level in the overall time period and the 1998-2008 period. Overall these results suggest that the draft is very efficient at determining player potential and that performance is the prime consideration determining player longevity. The experience variables are not individually significant; this is not surprising given the collinearity involved and the number of different experience measures employed. The only result that is an apparent anomaly is the sign and significance of the coefficients for the time dummy variables. The time dummy variables were included to capture the effect of the expansion of the league from 27 teams in the 1990-1994 period to 29 teams in the 1995-2003 period and 30 teams in the 2004-2008 period. A priori, one would expect the sign of the coefficients for the two included dummy variables to be positive. Perhaps these variables are capturing some other league changes during the time period. Exclusion of the variables did not change the sign or significance of any of the other explanatory variables.[5]

The result of main interest in Table 4 is that the coefficient for the real mandatory raise is not significant in any of the time periods. This stands in stark contrast to the results of Ducking and Bollinger (2012) for the NFL but is not unexpected given the differences in the NFL and NBA outlined above.

Turning attention to Table 5 the logit analysis results are shown in which pay differential is measured as the real difference in a veteran's minimum salary versus the rookie minimum. All of the results are very similar to the results in Table 4 concerning the effects of draft number and

[5] Results of these logit regressions are omitted to conserve space but are available upon request.

performance on duration. Again the key focus is the coefficient for the real difference in a veteran's minimum salary versus the rookie minimum. In all time periods the coefficient of this variable is highly significant, 1% level, and properly signed in all periods tested, even in the brief 1995-1997 time period in which there was only a rookie and a veteran minimum that differed by only $25,000 to $30,500.

When the NBA Players Association negotiated the veteran minimum pay scale by experience shown in Table 2 they obviously feared that cost-minimizing owners would favor cheaper rookies over veteran players. As previously mentioned, in the CBA in operation from the 1998-1999 season to the 2004-2005 season, only the portion of any veteran's minimum salary at or below the four year level counts against the salary cap and any salary above this level is reimbursed to the team from a league-wide fund following the completion of the season. In the subsequent seasons this was changed from a four year minimum to a two year minimum. Therefore, in Table 6 the time periods shown are the standard overall period, 1990-2008, the period 1998-2004, and the 2005-2008 period. A comparison of results from the latter two periods will allow us to see if the changes in the 2005 CBA eliminated the incentive to replace veterans with rookies.

Logit results in Table 6 show similar findings for draft number and performance variables in the overall time period, 1990-2008, and in the 1998-2004 period. However draft number and some performance variables are only significant at the 5% level in the 2005-2008 period and other performance variables are not significant at all. Obviously the lower number of observations is hurting the explanatory power of some of the determinants.

The coefficient of the real difference between the rookie minimum and the portion that counts against the cap for veteran players earning a minimum salary is highly significant in all cases, including the 2005-2008 period. It would appear that the efforts by the union in the 2005 CBA to alleviate economic incentives for teams to retain cheaper rookie players in place of veteran players with comparable statistics was not successful.[6]

Summary and Conclusions

The conclusions from the empirical work of this paper are in some sense a simple affirmation that basic microeconomic principles of profit maximization will always prevail in the marketplace. Profit-maximizing owners will hire lower cost workers if the marginal revenue product of the workers is the same. The NBA is no exception to this rule.

Efforts by the NBA Players Association to negotiate higher minimum salaries for veteran players and limit economic incentives not retain such veterans were not totally successful. We find strong evidence that these veteran minimum salaries have a negative impact on career length. Obviously the NBA, NBA Players union, and/or players' agents were monitoring the consequences of the original format of the veteran minimum pay plan negotiated in the 1998 CBA and attempted to remedy the disincentives to retain veterans in the 2005 CBA. Our results indicate that these changes still were not enough.

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[6] In the interest of fairness we must point out that any attempts at rent-sharing in collective bargaining agreements can only be evaluated in hindsight. Since the players in question are not stars their roster spot on an NBA team is tenuous at best. Even if all else is equal teams may decide to keep a rookie over a veteran player hoping the rookie has some upside potential. The economic welfare gain by the few veterans who retain a roster spot and the higher minimum salary may offset the economic welfare loss of the few veterans who are not retained.

These results have important implications for other industries because whenever minimum salaries force a worker's pay to be higher than their value to the firm, the worker's career length is expected to be shortened regardless of industrial structure, gender, or magnitude of minimum salaries. We argue that career lengths will be shortened any time the employer has the ability to dismiss or fire workers and the salary structure forces the firm to pay the worker more than the value of the worker to the firm. Therefore, any industry that has a minimum salary schedule in place should be aware that this salary structure has the ability to shorten the career length of an employee when it forces the employer to pay the employee more than the employee's value to the firm.

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Table 1.

Minimum Annual Scale by Season

Season	Minimum	Rookie Minimum	Veteran Minimum
1990-1991	120,000		
1991-1992	130,000		
1992-1993	140,000		
1993-1994	150,000		
1994-1995	150,000		
1995-1996		200,000	225,000
1996-1997		220,000	247,500
1997-1998		242,000	272,500

Figures were obtained from Patricia Bender's website entitled: Patricia Various Basketball Stuff: http://www.eskimo.com/~pbender/index.html

Table 2.

Minimum Annual Scale by Season and Experience

Exp.	1998-1999	1999-2000	2000-2001	2001-2002	2002-2003	2003-2004	2004-2005	2005-2006	2006-2007	2007-2008	2008-2009
0	287,500	301,875	316,969	332,817	349,458	366,931	385,277	398,762	412,718	427,163	442,114
1	350,000	385,000	423,500	465,850	512,435	563,679	620,046	641,748	664,209	687,456	711,517
2	425,000	460,000	498,500	540,850	587,435	638,679	695,046	719,373	744,551	770,610	797,581
3	450,000	485,000	523,5000	565,850	612,435	663,679	720,046	745,248	771,331	798,328	826,269
4	475,000	510,000	548,500	590,850	637,435	688,679	745,046	771,123	798,112	826,046	854,957
5	537,500	572,500	611,000	653,350	699,935	751,179	807,546	835,810	865,063	895,341	926,678
6	600,000	635,000	673,500	715,850	762,435	813,679	870,046	900,498	932,015	964,636	998,398
7	662,500	697,500	736,000	778,350	824,935	876,179	932,546	965,185	998,967	1,033,930	1,070,118
8	725,000	760,000	798,500	840,850	887,435	938,679	995,046	1,029,873	1,065,918	1,103,225	1,141,838
9	850,000	885,000	923,500	965,850	1,000,000	1,000,000	1,000,000	1,035,000	1,071,225	1,108,718	1,147,523
10+	1,000,000	1,000,000	1,000,000	1,000,000	1,030,000	1,070,000	1,100,000	1,138,500	1,178,348	1,219,590	1,262,275

These figures were taken from the 1999 and 2005 NBA Collective Bargaining Agreements.

Table 3.
Descriptive Statistics

Variable/ Mean Standard Deviation	Overall Time Period: 1990-2008	1995-1997 Time Period	1998-2008 Time Period	2005-2008 Time period
Draft Number	27.311 22.851	27.885	26.755 21.085	27.528 20.98
Height	79.180 3.782	79.198 3.783	79.195 3.721	79.063 3.644
Experience	4.668 3.908	4.654 3.882	4.844 3.983	4.639 3.815
Points per Minute	.3714 .124	.371 .119	.359 .121	.368 .124
Rebounds per Minute	.176 .081	.173 .084	.176 .079	.176 .080
Assists per Minute	.082 .059	.083 .057	.078 .056	.076 .055
Steals per Minute	.032 .017	.034 .020	.031 .016	.029 .016
Blocks per Minute	.022 .023	.021 .023	.021 .023	.021 .024
Turnovers per Minute	.060 .030	.065 .032	.0595 .031	.057 .033
Dummy for 1995-2003	.481 .50	---------	.533 .499	---------
Dummy for 2004-2008	.282 .450	--------	-------	----------
Real Mandatory Raise	7680.287 41550.82	1898.962 11191.78	12206.23 52674.22	10453.15 46035.17
Real Difference vs. Rookie Minimum	41968.43 148607.3	4215.431 11795	68406.05 186489.3	62198.1 174751.6
Real Difference vs. Counts against Cap	27579.4 88003.89	--------	44573.03 109820.8	43573.65 112764.4
Duration	.126 .332	.107 .309	.133 .340	.137 .344
Number of Obs.	7533	1195	4548	1694

Table 4.
Logit Results for Real Mandatory Raise

Variable	Overall Time Period: 1990-2008	1995-1997 Time period	1998-2008 Time Period
Intercept	1.353 (1.378)	-6.749 (4.075)	2.034 (1.717)
Draft Number	.013** (.002)	.0126** (.004)	.013** (.002)
Height	-.013 (.017)	.085 (.051)	-.021 (.021)
Experience	-.079 (.105)	.280 (.30)	-.096 (.136)
Experience Squared	-.034 (.036)	-.153 (.109)	-.003 (.046)
Experience Cubed	.008 (.005)	.023 (.015)	.002 (.006)
Experience to Fourth Power	-4.655e-04 (2.857e-04)	-.001 (9.38e-04)	-4.12e-05 (3.715e-04)
Experience to Fifth Power	7.54e-06 (7.54e-06)	2.65e-05 (2.03e-05)	-2.17e-06 (8.78e-06)
Points per Minute	-4.125** (.355)	-5.70** (1.03)	-4.799** (.472)
Rebounds per Minute	-2.804** (.649)	-5.497** (1.85)	-2.585** (.865)
Assists per Minute	-7.038** (1.023)	-8.337** (3.013)	-7.123** (1.326)
Steals per Minute	-5.335* (2.334)	-4.040 (5.450)	-8.067* (3.284)
Blocks per Minute	-8.872** (2.225)	-16.316** (6.169)	-7.388** (2.826)
Turnovers per Minute	6.627** (1.414)	21.298** (3.719)	5.782** (1.986)
Dummy for 1995-2003	-.376** (.097)	-------	-.231** (.095)
Dummy for 2004-2008	-.128 (.104)	--------	(omitted)
Real Mandatory Raise	1.11e-06 (7.09e-07)	1.05e-05 (6.26e-06)	8.27e-07 (7.39e-07)
Number of Obs. Log Likelihood	7533 -2540.041	1195 -343.649	4548 -1553.872

Standard errors are shown in parentheses.
* indicates significance at the 5% level.
** indicates significance at the 1% level.

Table 5.
Logit Results for Real Difference in Pay versus Rookie Minimum

Variable	Overall Time Period: 1990-2008	1995-1997 Time Period	1998-2008 Time Period
Intercept	.592 (1.387)	-6.947 (4.222)	.877 (1.735)
Draft Number	.012** (.002)	.006 (.005)	.011** (.002)
Height	-.003 (.017)	.086 (.053)	-.007 (.022)
Experience	-.158 (.106)	.062 (.309)	-.236 (.138)
Experience Squared	-.012 (.037)	-.080 (.111)	.037 (.047)
Experience Cubed	.006 (.005)	.014 (.016)	-.002 (.006)
Experience to Fourth Power	-3.721e-04 (2.947e-04)	-.001 (.001)	1.86e-04 (3.852e-04)
Experience to Fifth Power	6.68e-06 (6.48e-06)	1.83e-05 (2.02e-05)	-5.66e-06 (9.08e-06)
Points per Minute	-3.813** (.359)	-5.437** (1.056)	-4.377** (.477)
Rebounds per Minute	-2.848** (.652)	-5.376** (1.920)	-2.697** (.873)
Assists per Minute	-6.721** (1.025)	-6.675* (3.059)	-6.543** (1.332)
Steals per Minute	-5.060* (2.341)	-6.632 (4.784)	-7.647* (3.287)
Blocks per Minute	-8.915** (2.232)	-14.005* (6.280)	-7.438** (2.851)
Turnovers per Minute	6.386** (1.410)	20.355** (3.627)	5.267** (1.97)
Dummy for 1995-2003	-.478** (.097)	-------	-.232* (.096)
Dummy for 2004-2008	-.277** (.106)	--------	(omitted)
Real Difference vs. Rookie Minimum	1.80e-06** (1.93e-07)	4.34e05** (6.74e-06)	1.76e-06** (2.04e-07)
Number of Obs. Log Likelihood	7533 -2498.882	1195 -325.189	4548 -1518.244

Standard errors are shown in parentheses.
* indicates significance at the 5% level.
** indicates significance at the 1% level.

Table 6.
Logit Results for Real Difference in Pay versus Counts against Cap

Variable	Overall Time Period: 1990-2008	1998-2004 Time Period	2005-2008 Period
Intercept	.603 (1.388	.898 (2.258)	1.40 (2.827)
Draft Number	.011** (.002)	.010** (.003)	.008* (.004)
Height	-.003 (.0172)	-.003 (.028)	-.018 (.035)
Experience	-.244* (.108)	-.45 (.182)	-.117 (.241)
Experience Squared	.010 (.037)	.112 (.059)	-.063 (.094)
Experience Cubed	.004 (.005)	-.011 (.008)	.0185 (.016)
Experience to Fourth Power	-3.06e-04 (2.905e-04)	.001 (4.367e-04)	-.001 (.001)
Experience to Fifth Power	5.85e-06 (6.37e-06)	-.1.43e-05 (9.85e-06)	3.86e-05 (3.31e-05)
Points per Minute	-3.789** (.358)	-4.667** (.620)	-4.002** (.760)
Rebounds per Minute	-2.880** (.652)	-3.913** (1.127)	-1.000 (1.387)
Assists per Minute	-6.642* (1.027)	-8.109** (1.742)	-4.307* (2.067)
Steals per Minute	-5.172* (2.343)	-7.295* (4.016)	-9.952 (5.834)
Blocks per Minute	-8.750** (2.228)	-6.029* (3.687)	-9.159* (4.539)
Turnovers per Minute	6.340** (1.411)	6.578** (2.518)	3.169 (2.469)
Dummy for 1995-2003	-.493** (.098)	-.265* (.161)	(omitted)
Dummy for 2004-2008	-.3480** (.108)	(omitted)	(omitted)
Real Difference vs. Counts Against Cap	3.32e-06** (3.42e-07)	2.97e-06** (4.81e-07	3.60e-06** (5.66e-07)
Number of Obs. Log Likelihood	7533 -2496.021	2854 -920.949	1694 -588.289

Standard errors are shown in parentheses.
* indicates significance at the 5% level.
** indicates significance at the 1% level.