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PAY AND PERFORMANCE IN BASEBALL: MODELING REGULARS, RESERVES AND EXPANSION

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ABSTRACT

Although the relationship between pay and performance in baseball has been convincingly demonstrated by Scully, a number of unresolved questions remain. Using a large sample of player salaries from contracts on file at the American League office, new estimates of this relationship are reported. The primary findings are as follows. First, while Scully's basic results are qualitatively robust, the salary elasticities for various performance and experience variables are substantially lower for our sample and specification. Second, for most variables, recent performance, as well as career average, contributes to the explanation of salary differences. Third, expansion has a significant effect on salary structure, and, in our model, makes it statistically invalid to estimate a single salary equation from pooled time-series data that includes an expansion year.

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I. INTRODUCTION

The labor market for professional baseball players has been extensively studied and the connection between the performance of athletes and their salaries convincingly demonstrated. The theoretical argument relating pay and performance can be summarized as follows. If teams maximize profits, choosing all inputs except player talent competitively, then players will earn their marginal revenue products net of monopsony rents accruing to the teams.[1] Marginal revenue products are then represented in empirical specifications by summary performance measures such as batting average or earned run average. The derived demand for player inputs is a function of player performance.

Gerald Scully's (1974) definitive study demonstrates that most of the variation in observed player salaries can be explained by a handful of player productivity variables. While impeccable in both theoretical treatment and empirical technique, Scully's study (like all other work in the area) relies upon salary data of questionable accuracy that are gleaned from popular publications. Moreover, the players reported in such sources are not representative of the population of major-leaguers. Popular sources are much more likely to publish information about well-known players than about the journeymen who comprise the bulk of any major-league roster. This sampling bias raises questions about whether the pay and performance relationship previously discovered holds true for players of lesser abilities and lower salaries.

Another issue pertains to the effect of changes in industry structure on salaries. For example, Scully's data were from 1968 and 1969. The latter was an expansion year, when two new teams were admitted to the American League. Two more teams joined the National League in 1970. For reasons discussed herein, expansion should cause a change in the relationship between salary and performance, raising the possibility that pooling observations over a period of substantial structural change is not an acceptable procedure.

In this paper, we report the results of an analysis of salaries from a player sample that is dramatically different from those used in past studies. The data were collected by agents of the Internal Revenue Service and employees of the American League office under our general direction, in connection with <u>Selig v. U.S.</u>, no. 81-C-334 (U.S. District Court for the Eastern District of Wisconsin), and cover 461 observations directly from player contracts for the years 1968, 1969, and 1970. These data were painstakingly verified by the Internal Revenue Service since they were used in litigation. The salary observations are accurate to the level of the thoroughness of the

Internal Revenue Service.

The major findings are as follows. First, our data are substantially different than those in Scully's sample, and are probably more representative of the population of interest. Second, when Scully's model is estimated using our data, substantially less of the total variation in salaries is explained and some important differences emerge in the estimations. Third, our new specification of the salary model increases its explanatory power for our sample to roughly the level achieved by Scully on his sample. Finally, we find statistical evidence that a change in industry structure does have a significant effect on player salaries. Therefore, observations on marginal players that span an expansion cannot be pooled.

II. COMPARING THE DATA SETS

The data made available to us were based on trades or sales where at least one of the parties to the transaction was an American League team. Associated with each transaction are the name(s) of the player(s) subject to the transaction, the teams involved, the transaction date, and the terms of the transaction. Salaries were reported in the following way. At the very least, the salary was reported for every player during the season in which the transaction occurred. Usually the salaries for one season prior and one season after the transaction were also included, and occasionally additional years were reported.

Table 1 shows the comparison between Scully's and our data.

TABLE 1

Comparison of Scully and Fort-Noll Data

Characteristic	Scully[a]	Fort-Noll
Number of Observations: Total	148	461
1968	87	116
1969	67	166
1970	-	179
Multiple Observations on Same Player		
1968 and 1969, not 1970	41	63
1969 and 1970, not 1968	-	57
1968 and 1970, not 1969	-	6
1968, 1969, and 1970	-	22
Annual Observations involving Sale or Trade (%)		
1968	-	41
1969	-	55
1970	-	42
Annual Observations in American League (%)		
1968	-	73
1969	-	80
1970	-	76
Salary Range ($ thousands)	10-125	4-90
Mean Salary ($)	48,100	22,390
Salary Observations under $25,000 (%)	23	66
Mean Years of Experience		
Hitters	8	6
Pitchers	6	6
Mean Playing Time per Season		
Hitters (At Bats)	385	261
Pitchers (Innings Pitched)	207	106

[a] Reported by Scully (1974, pp. 926-7).

Scully's sample clearly is comprised primarily of veteran "regulars"--players who are usually in the starting lineup. For example, for pitchers the sample mean number of innings pitched is approximately the amount of work expected from a starter who is in the regular rotation, pitching every four or five days, and is far greater than even the best relief pitchers would be expected to pitch. For hitters, the average number of at bats in a season for all major-league players (other than pitchers) is approximately half the mean in Scully's sample. The mean salary in Scully's sample is about $6,000 higher than the mean for veteran ballplayers during the period.

By contrast, our sample of players contains a much larger proportion of mediocre players, judged on experience and playing time. Further, although the two data sets exhibit the same range of salaries, our players were paid substantially less on average. Although precise information is not available on the issue, our sample mean is closer to the average for all players on the forty-man major-league roster, from which the sample was taken.[2]

All-in-all, our sample appears generally representative.[3] But even if it is not (i.e. neither Scully's nor our sample is judged generally representative), the two samples are substantially different. The importance of this difference is revealed when Scully's salary model is estimated on each of the two data sets.[4] The results are reported in Table 2. The 1968 and 1969 data from our sample are used in one estimation to facilitate comparison with Scully's original findings and the results for the entire 1968-1970 sample are also

TABLE 2

Estimation of the Scully Salary Model For both Samples[a]

Variable	Scully Sample (1968-9)	Fort-Noll Sample (1968-9)	Fort-Noll Sample (1968-70)
I. Hitters			
At Bat Share[b]	.275 (3.10)	.273 (5.38)	.274 (7.19)
Slugging Ave.[c]	1.072 (4.76)	.481 (2.44)	.373 (2.75)
Banjo Hitter[d]	.058 (1.61)	.006 (0.06)	-.063 (0.85)
Marginal Attendance[e]	.265 (2.81)	.101 (1.82)	.125 (3.07)
Population 1970[f]	-.062 (0.78)	.103 (2.22)	.068 (1.91)
Experience[g]	.522 (7.53)	.297 (5.61)	.300 (7.54)
National League[h]	.019 (0.62)	.055 (0.72)	.006 (0.11)
Constant	.670 (0.82)	9.206 (13.07)	9.26 (17.37)
R^2	.81	.68	.67
II. Pitchers			
Innings Pitched Share[i]	.970 (4.22)	.276 (4.88)	.385 (7.96)
Power Pitching Ratio[j]	.808 (4.00)	.235 (2.26)	.108 (1.32)
Marginal Attendance[e]	-.062 (0.61)	.070 (1.34)	.087 (2.02)
Population 1970[f]	-.053 (0.53)	.057 (1.09)	.004 (0.09)
Experience[g]	.502 (7.70)	.277 (6.10)	.289 (7.78)
National League[h]	.007 (0.20)	-.044 (0.52)	-.082 (1.32)
Constant	3.385 (4.16)	9.181 (13.02)	9.876 (16.55)
R^2	.78	.59	.64

a. The dependent variable is the natural log of player salary. All independent variables are in logs except, of course, for dummy variables. The Scully results are from the 1974 paper, pp. 926-7. All reports are for the indicated pooled sample. In the Fort-Noll sample, player histories were compiled from the Official Baseball Register and The Baseball Encyclopedia. A number of Fort-Noll observations are not included in the applications of Scully's model shown in this Table. The reasons are that (1) the marginal attendance variables for expansion teams (1970) could not be calculated and (2) players we distinguish later as Hopefuls are not included. The t-statistics are reported in parentheses.

TABLE 2 (cont'd)

b. The fraction of his team's total at bats accounted for by the player, on average, for all his years in the major leagues.

c. Lifetime total bases from safe hits divided by official lifetime at bats.

d. A dummy variable equal to one if the player is an above-league-average hitter and below-league-average slugger, zero otherwise.

e. The estimated effect of win percentage on annual attendance, from a regression for each major-league team. Because Scully's actual numbers were unavailable at the time this project was undertaken, the effect was estimated independently here as explained (but not reported) by Scully (1974). The coefficient estimates in Table 2 for this variable may not be strictly comparable.

f. Reported by SMSA, Bureau of the Census, used in millions.

g. Number of years that the player's name appears on a major-league roster.

h. A dummy variable equal to one if the salary observation pertains to a year in which the player was on a National League roster, zero otherwise.

i. The fraction of his team's total innings pitched accounted for by a pitcher, on average, during his major-league career.

j. The pitcher's lifetime ratio of stikeouts to walks.

reported. A number of results deserve mention: (1) R^2 is much higher for Scully's sample, for both the hitter and pitcher equations; (2) all variables that are statistically significant for Scully's original sample are also significant for our sample, except that the dummy variable distinguishing hitters with high batting average but low slugging average (banjo hitters) is marginally significant in Scully's sample but insignificant in ours; (3) every variable that Scully found statistically important is closer to zero in absolute magnitude in the regression for our data, and in most cases, these differences are large; (4) the variables related to market size (population and Scully's estimated marginal effect of winning on attendance) appear to contribute to the salary explanation from our sample but not Scully's.[5]

From these observations, we conclude that Scully's model is qualitatively robust; the important variables in his model retain their significance for a completely different sample. Nevertheless, the differences in the regressions suggest that (1) Scully's model may be misspecified in terms of functional form, and (2) the process by which salaries are determined for mediocre players, who are represented only in our sample, may be substantially different than the process for the more established veterans that comprise Scully's sample.

III. ALTERNATIVE SPECIFICATIONS

Given the preceding regression comparison, a number of plausible respecifications suggest themselves. All previous attempts at widening the repertoire of performance variables have reached the same conclusion: the variables are so highly correlated that, once a few are included, more variables provide little by way of added explanation.[6] Nonetheless, one potentially important independent variable is missing from the pitching equation: whether the pitcher is a starter or fulfills a relief role. The ratio of starting appearances to total appearances is included in an attempt to capture the difference between types of pitchers. Because innings pitched gives undue weight to starters, correcting for starting appearances should lower the value of starters relative to relievers. Hence, the starter ratio should have a negative sign.

Our sample contains several young, marginal players with little or no major-league experience. Career major-league performance records are presumably less meaningful ability indicators for these players than for more experienced veterans. Two dummy variables were used to account for the lower end of the experience and performance range: "Rookie" indicates that an observation was for a player in his first major-league season, and "Hopeful" indicates that the observation pertained to a player with less than five years experience and fewer than forty at bats (thirty innings pitched) in each season spent on a major-league roster if a hitter (pitcher). Hence, a player in his first year, who did not meet the at bats or innings pitched cutoff in that year, is treated as a Hopeful.

Another refinement is to account for the possibility that recent performance receives different weight than career performance in the salary process. For variables that measure playing time (at bats and

innings pitched), a substantial change from historical averages can occur because of injuries that actually have no bearing on a player's long run abilities. Poor players may get more playing time in a given season due to an injury to a regular in the starting lineup, or a starting regular may be out of the lineup temporarily due to an injury that is not permanently disabling. Concerning other performance measures, a player may be detectably improving or declining so that more recent data better indicate his productivity to a team. Hence, we include variables that measure the extent to which performance indicators in the year preceding the salary observation depart from career averages for the player.

The techniques just described pose some problems of estimation. First, having defined a major-league season as greater than forty at bats for hitters or thirty innings pitched for pitchers, players defined as "Hopefuls" are treated as having no performance data. Hence, natural logarithms of these variables are undefined. To cope with this problem, all performance variables were measured as one plus the actual value.[7] As explained below, this transformation substantially affects the estimates of the salary elasticities for the transformed independent variables. Second, rookies have only one year of history, so that calculating departures from average yields meaningless zeroes. Departure variables for established players who were absent from the majors in the year prior to a salary observation (say, due to injury or a stint in the minor leagues), likewise are misleading. To handle cases of established players absent from the major leagues in the prior season, we constructed the dummy variable "No Last Year." Then, for Hopeful, Rookie, and No Last Year observations, the logarithm of departure variables was simply set equal to zero.

Finally, more refined measures of major-league experience and "banjo hitting" were devised. For the former, the effects of age and experience were separated. We expect that age does not enter the salary equation linearly in logarithms. Instead, age has a more pronounced effect in the early and late years than in the middle of a player's career. Accordingly, age was entered as a deviation from the mean, and separated for players of above average (Old Age) and below average (Young Age) maturity. The square of these deviations from the mean age was also included to detect nonlinear effects. Furthermore, because age and experience are obviously highly correlated, one more adjustment was made. First, years of experience was regressed on age, and an expected amount of experience was calculated for each player. The difference between expected and actual experience was entered for each player as the independent effect of experience that was not already accounted for by age. We expect this experience residual to be inversely related to salary because experience is thought to improve performance, especially early in a player's career; for two players of equal age, the player with more experience can be expected to perform better. This should result in a negative sign for the Positive Experience Residual variable while the opposite is expected for the Negative Experience Residual variable.

To deal with banjo hitters, we regressed slugging average on

batting average, finding the expected strong correlation. As with experience, the difference between predicted and actual slugging average (Positive or Negative Slugger Residual) was used to measure the extent to which a player was truly a slugger.

Estimation results of this respecification for the sample period 1968-1970 are shown in Table 3. Model I is the initial attempt to include all of the ideas just described while Model II drops those notions which yielded no added explanation. Note that hitters and pitchers are reported in a single equation with a dummy variable distinguishing the two (Pitcher). The common variables related to experience showed no statistically detectable difference for hitters and pitchers in any of the respecified regressions, or for that matter in Scully's model applied to our data. Also, measures of market effects (population and marginal attendance) have been dropped. Indeed, the market effects variables were always insignificant in the respecified regressions (including some unreported results) which was cause for their subsequent omission. This is consistent with the profit-maximizing theory of the operation of sports teams.

In Table 3, the respecification produced significant and plausible results. For hitters and pitchers, the performance variables proposed by Scully (At Bats and Innings Pitched Shares, Slugging Average and Power Pitching Ratio) all retain their significance. All of these measures have higher t-values in our Model II (Table 3) than in Scully's original results (Table 2).

TABLE 3

Estimation of Respecified Salary Equation

Using 1968-1970 Data[a]

Variable	Model I		Model II	
At Bats Share + 1[b]	11.04	(11.67)	10.83	(11.87)
Slugging Average + 1[b]	.932	(2.25)	1.156	(3.29)
Positive Slugger Residual[c]	.424	(.047)	-	
Negative Slugger Residual[c]	.987	(1.18)	-	
Positive At Bats Departure[d]	.043	(3.06)	.041	(3.51)
Negative At Bats Departure[d]	.031	(1.98)	.027	(2.05)
Positive Slugging Average Departure[e]	.957	(1.11)	.855	(1.01)
Negative Slugging Average Departure[e]	-1.204	(2.53)	-1.126	(2.42)
Innings Pitched Share + 1[b]	9.236	(9.54)	9.243	(9.96)
Power Pitching Ratio + 1[b]	.357	(3.29)	.386	(3.90)
Starter Ratio + 1[f]	-.601	(4.79)	-.588	(4.83)
Positive Innings Pitched Departure[g]	.007	(0.37)	-	
Negative Innings Pitched Departure[g]	.010	(0.53)	-	
Positive Power Pitching Departure[h]	.030	(0.32)	.032	(0.35)
Negative Power Pitching Departure[h]	.191	(1.50)	.204	(1.66)
Old Age[i]	-.014	(0.21)	-	
Old Age Squared[i]	.052	(1.69)	.045	(4.26)
Young Age[i]	.219	(2.66)	.228	(3.19)
Young Age Squared[i]	-.218	(4.96)	-.221	(5.42)
Positive Experience Residual[j]	-.059	(1.76)	-.061	(1.83)
Negative Experience Residual[j]	.237	(8.19)	.236	(8.26)
Hopeful[k]	.292	(2.69)	.309	(3.33)
Rookie[l]	.090	(1.21)	.074	(1.16)
No Last Year[m]	.000	(0.00)	-	
Pitcher[n]	.112	(1.71)	.124	(1.93)
Constant	8.944	(75.76)	8.918	(89.29)
R^2	.80		.80	
Degrees of Freedom	435		441	
Sum of Squared Residuals	29.74		29.87	

TABLE 3 (cont'd)

a. The dependent variable is the logarithm of player salary. All independent variables that are not dummies are also in logarithmic form. Logarithms of pitching variables are zero for hitters and vice versa for pitchers. Data sources are the same as in Table 2.

b. These variables are the same as defined in Table 2, except that one has been added to insure that all observations have strictly positive values, as noted in the text.

c. The estimated relationship between slugging average and batting average was:

$$SA = \underset{(2.35)}{.042} + \underset{(17.5)}{1.25}\ BA, \quad R^2 = .570.$$

The variable Slugger Residual is set equal to the predicted SA from the above equation minus the actual SA for hitters. If positive, then Positive Slugger Residual equals the residual plus one, and the logarithm of Negative Slugger Residual equals zero. If the residual is negative, then Negative Slugger Residual equals one plus the absolute value of the residual and the logarithm of Positive Slugger Residual equals zero.

d. Let At Bats Departure equal the difference between the hitter's share of team at bats last season and his career average share of team at bats. If positive, Positive At Bats Departure equals At Bats Departure plus one and the logarithm of Negative At Bats Departure equals zero. If negative, then Negative At Bats Departure equals one plus the absolute value of At Bats Departure and the logarithm of Positive At Bats Departure equals zero.

e. Let Slugging Average Departure equal the difference between the hitter's slugging average last season and his career slugging average. Positive and Negative Slugging Average Departures are calculated in an identical manner as Positive and Negative At Bats Departure (see note d).

f. The ratio of lifetime games started to total lifetime pitching appearances.

g. Let Innings Pitched Departure equal the difference between the pitcher's share of team innings pitched last season and his career average share of team innings pitched. Positive and Negative Innings Pitched Departures are calculated in an identical manner as Positive and Negative At Bats Departures (see note d).

TABLE 3 (cont'd)

h. Let Power Pitching Departure equal the difference between the pitcher's power pitching ratio last season and his career average power pitching ratio. Again, for the calculation of Positive and Negative Departures, see note d.

i. Age is represented as the deviation from the sample mean, which was 28 years. If this deviation is positive, Old Age equals one plus the deviation while its logarithm is zero if the deviation is negative. Young Age equals one plus the absolute value of the deviation if the deviation is negative and the logarithm of Young Age is zero if the deviation is positive. Squared variables are the square of the logarithms of the age variables.

j. The estimated relationship between experience and age was:

$$\text{EXPER} = -16.8 + \underset{(20.0)}{} \; .771 \text{ AGE}, \quad R^2 = .611.$$
$$(20.0) \qquad (27.0)$$

Let Experience Residual equal the predicted EXPER from the above equation minus the player's actual EXPER. Positive and Negative Experience Residuals are calculated in a manner identical to Positive and Negative Slugger Residuals (see note c).

k. If the player has less than five years of major-league experience and fewer than forty at bats in each season (hitters) or fewer than thirty innings pitched in each season (pitchers), then Hopeful equals one. Otherwise, Hopeful equals zero.

l. If a player has one year in the major leagues with playing time in excess of the cutoffs in the preceding note in that season, Rookie equals one, and is zero otherwise.

m. According to the cutoffs in note k for experience below five years, and for those with experience exceeding or equal to five years, No Last Year equals one if the player was not on a major-league roster the previous season. Otherwise, No Last Year equals zero.

n. Pitcher equals one for pitchers, zero otherwise.

Many of the new variables, are also significant. Apparently when sluggers have a bad year it is regarded as a negative signal in the salary determination process (estimated Negative Slugging Average Departure has a negative sign). Both Positive and Negative At Bats Departures are positively related to salaries. While the interpretation for players whose at bats share was increasing is clear (an improving player, on average), for players with recent at bats below their career averages, the dominant factor is most likely an injury which need not detract from their long-run performance. For hitters, the failure to distinguish sluggers from banjo hitters is disappointing (Positive and Negative Slugger Residuals are both insignificant). Regarding pitchers, the respecified model distinguishes relief pitchers from starters (Starter Ratio significantly negative). However, Innings Pitched Departure adds essentially nothing to the explanation of salaries.

Our treatment of age and experience is generally successful. Age has the predicted effect, especially for the very young and old: youthfulness is detrimental to the salaries of the very young, in middle years age has relatively little effect, and age again becomes an important contributor to increased salary for older players. Players with long careers presumably have more lucrative outside earnings potential (broadcasting, acting, advertising, public relations) because of their fame, and accordingly, if a team wants to keep them, their salaries must exceed these opportunity costs. The experience residual also is significant, especially for players with more experience than their age predicts (Negative Experience Residual significantly positive). Players who have less experience than their age indicates earn lower salaries (Positive Experience Residual significantly negative).

Finally, the various dummies that were intended to represent players with discontinuous playing histories proved to explain little. Only the "Hopeful" variable is statistically significant; however, its magnitude is quite small. The same can be said for the Pitcher variable.

The overall quality of the model is about the same as for Scully's original formulation, despite the omission of the market measures. Moreover, nothing is lost in choosing Model II over Model I (about .009 on R^2); we reject the hypothesis that Model I provides a better fit. Comparisons between Model II and Scully's model (Table 2) are somewhat obscured by the data transformations used to produce the former. Coefficients of the continuous variables in Model II are not elasticities, as is usually the case for log-linear estimation. Moreover, Model II does not produce constant elasticities across the range of the independent variables. The salary elasticities of the independent variables are increasing in the magnitude of the variables, asymptotically approaching the estimated coefficient as the value of the independent variable approaches infinity. But this is rather academic, for the values of the primary performance variables--playing time shares, slugging average, power pitching ratio--can never become very large and, for all but the last, are definitionally less than one.

Rising elasticities over the relevant range of the independent variables seems to us to be a desirable characteristic, for it is consistent with the results of the constant-elasticity model reported in Table 2. There, Scully's sample of regulars produced higher elasticity estimates than did our sample. This suggests that elasticities rise with increases in skill. Comparisons of elasticities between Table 2 and Model II in Table 3 can be made by calculating point elasticities in Model II for values of the independent variables that are likely to be observed. Table 4 shows a set of such calculations and comparisons. Each postulated value of the independent variable of interest, from which point elasticities are calculated, lies within the normal range for players who typically spend an entire season on a major-league roster.[8]

The elasticity estimates for the two playing time variables exhibit the broadest variability in Model II. Elasticity ranges for both playing time variables (At Bats and Innings Pitched Shares) include the estimates obtained from the regressions in Table 2. In addition, the elasticity estimates for both variables obtained from the Scully model using the Fort-Noll sample are quite close to the Model II estimates at the low end of the performance range, as would be expected from the nature of the sample. Innings Pitched Share has the further property that the original elasticity estimate by Scully is essentially the same as the elasticity estimate from Model II that arises when the variable takes the value of .15, which corresponds to the mean of that variable in Scully's sample (.14).

TABLE 4

Salary Elasticity Comparisons[a]

	Salary Elasticity				
	Scully Model[b]			Model II[c]	
Variable	Scully Data 1968-9	Fort-Noll Data 1968-9	Fort-Noll Data 1968-70	Chosen Value	1968-70
At Bats Share	.28	.27	.27	.02	.21
				.05	.52
				.08	.80
Slugging Average	1.07	.48	.37	.30	.27
				.40	.33
				.50	.39
Innings Pitched Share	.97	.28	.38	.03	.27
				.09	.76
				.15	1.21
Power Pitching Ratio	.81	.24	.11	.75	.17
				1.25	.21
				1.75	.24

a. Entries are estimated elasticity of salary with respect to the corresponding independent variable. Variables are defined in Table 2.

b. Repeated from Table 2, these are the estimated coefficients from the Scully model applied to both sets of data.

c. Calculated from Model II of Table 3. The elasticity, in general, for variables transformed by adding one is:

$$\frac{x}{y}\frac{\delta y}{\delta x} = \frac{xb(x+1)^{b-1}}{y} = \frac{xb(x+1)^{b-1}}{(x+1)^b} = b(\frac{x}{x+1}),$$

where x is the independent variable, b is the regression coefficient, and y is the dependent variable. The elasticity is calculated by taking the exponent of the log-linear relationship.

The elasticity estimates from Model II for the other two performance variables, slugging average and power pitching ratio, are lower than those in Table 2 throughout the entire range of chosen alternative values. For the estimates of Scully's model using the Fort-Noll sample (Table 2), the elasticity estimates are at the upper extreme of the range of feasible elasticities calculated from Model II. Slugging averages and power pitching ratios that in Model II would generate elasticities roughly comparable to these estimates from Scully's model using our data would pertain only to league leaders in those categories. The estimated elasticities from the Scully sample in Table 2 are so far above the range of feasible choices that in Model II they pertain only to record holders.

In Model II, playing time elasticities are low for reserves (.2 to .3) and near unity for regulars while elasticities remain low for the other performance measures over the entire range of player classification. These results hold for both hitters and pitchers. While salaries are essentially proportional to playing time for regulars, with some additional payment to sluggers and strikeout pitchers, playing time is less important for players at the low end of the skill range. A twofold explanation is plausible. At the lower end of the salary scale, a dominant factor is simply payment of the minimum salary necessary to keep the player in professional baseball (or to meet the minimum salary limit established through collective bargaining). Many players in this salary range are on major-league rosters only part-time and their key to staying in the majors is in performing a specialized role (e.g. pinch hit effectively against a right-handed fastball pitcher, or succeed as a relief pitcher against power-hitting right-handed batters). At the upper salary ranges are the regulars (stars) who draw fans beyond the importance of their performance to the outcome of the game, an element in the salary determination process that generally becomes more important as performance measures increase.

Based on the analysis in this section, we reach the following conclusions. First, with a different sample than previously available, we find that the general qualitative predictions of the economic theory of wage determination in professional sports and the value of the original specification by Scully are confirmed. Second, with respect to our data, plausible respecifications enable us to achieve explanatory power on a par with previous work while omitting the theoretically suspicious variables that measure market size. Third, the difference in the quantitative effects of playing time and direct performance variables, including those present in our model but not used elsewhere, mark our specification as both a plausible and important improvement.

IV. THE EFFECTS OF INDUSTRY STRUCTURE

During the period covered by our data, both major leagues expanded. The American League added the Seattle Pilots (now the Milwaukee Brewers) and Kansas City Royals in 1969 while the Montreal Expos and San Diego Padres began National League play in 1970. There are a number of reasons why expansion should affect the distribution of

player salaries. Assuming that expansion perceptively dilutes average player quality, major-league veterans of a given ability prior to expansion become more valuable. Even though veterans' absolute talents might remain constant, their talents improve relative to the new diluted level and, hence, so do contributions to the chances that their teams will fare well. Further, the label of "star" is in some respects a relative, not absolute, characteristic.[9] For example, even though they may not be considered stars in the league at large, home fans grant star status to the one or two best players on their team. Expansion, then, increases the number of players who will be regarded as stars by increasing the number of home-team partisans. On another front, the presence of minimum salary makes expansion place upward pressure on the entire salary structure. Although the minimum ability required to obtain a position on some major-league roster is lowered, the minimum salary requirement negotiated through collective bargaining does not decrease when expansion occurs. Analytically, the effect is similar to the consequences of raising the minimum wage while holding skills fixed, as examined by Scoville [1974]. Finally, the supply of players with sufficient talent to play major-league baseball is unlikely to be perfectly elastic. Even if it is technically possible to undertake an expansion that does not dilute average and minimum playing skills, such is likely to be achieved only at some increase in salaries. For all of these reasons, expansion should not be expected to leave the salary determination process unaffected; for players of all abilities, expansion should lead to higher salaries.

The regressions reported in the previous two sections of the paper are based upon pooled samples of salary observations taken from years in which expansion occurred. In this section, we consider two issues: whether pooling data from these years is a statistically acceptable procedure and whether expansion had identifiable effects on salary determination. It turns out that the two issues are not independent.

To examine the first issue, separate regressions for Model II were run for each of the sample years 1968, 1969, and 1970. Table 5 reports the results of these and two other estimations, one for pooled 1968-69 observations and one for pooled 1969-70 observations. The procedure used to test pooling is the standard method using F-tests (for example, see Maddala [1977]). Table 6 shows calculated and critical F-values relevant to all the possible pooling questions. Pooling is rejected in all cases. The result that the three sample years do not conform to a single model lends empirical support to the theoretical argument that expansion significantly affects the salary determination process. The pooling of observations to estimate the models reported in Tables 2 and 3 is not supported statistically. Consequently, we regard the 1968 regression in Table 5 as the pre-expansion estimation while the 1969 and 1970 regressions in that table are taken to represent different states of adjustment for the post-expansion industry.

Inspection of Table 5 provides some insight into the elements of Model II that were and were not stable during the expansion period. Career performance variables generally are quite stable: the coefficients for at bats share, slugging average, power pitching ratio,

TABLE 5

Estimation of Model II with Unpooled Data[a]

Variable	1968	1969	1970	1968-69	1969-70
At Bats Share +1	8.862(4.29)	11.93 (8.15)	10.64 (7.81)	10.79 (8.93)	11.55 (11.72)
Slugging Average +1	1.230(1.58)	.830(1.36)	2.010(3.89)	.884(1.82)	1.263 (3.32)
Pos. At Bats Departure	.049(1.92)	.067(3.45)	.016(0.86)	.053(3.42)	.044 (3.45)
Neg. At Bats Departure	.041(1.38)	.052(2.61)	.002(0.07)	.040(2.34)	.030 (2.08)
Pos. Slug. Ave. Depart.	.540(0.25)	3.010(1.61)	.042(0.04)	1.850(1.35)	.505 (0.56)
Neg. Slug. Ave. Depart.	-.092(0.11)	-2.340(3.53)	.455(0.38)	-1.130(2.15)	-1.773 (3.18)
Innings Pitched Share +1	9.295(4.00)	9.517(6.20)	9.875(7.24)	9.295(7.26)	9.642 (9.52)
Power Pitching Ratio +1	.280(1.15)	.519(3.09)	.504(3.29)	.427(3.11)	.468 (4.42)
Starter Ratio +1	-.531(1.85)	-.700(3.80)	-.600(3.07)	-.624(4.03)	-.643 (4.78)
Pos. Power Pitch. Depart.	-.145(0.73)	.039(0.32)	.457(1.88)	-.016(0.15)	.102 (0.98)
Neg. Power Pitch. Depart.	-.338(0.98)	.079(0.38)	.317(1.93)	-.052(0.28)	.253 (2.00)
Old Age Squared	.069(3.31)	.053(3.37)	.021(1.06)	.056(4.47)	.036 (2.94)
Young Age	.344(2.08)	.174(1.54)	.204(1.86)	.239(2.53)	.209 (2.71)
Young Age Squared	-.302(3.16)	-.195(2.97)	-.190(3.10)	-.241(4.41)	-.204 (4.64)
Pos. Experience Residual	-.097(1.33)	-.065(1.20)	-.029(0.57)	-.063(1.46)	-.057 (1.57)
Neg. Experience Residual	.253(3.99)	.236(5.31)	.207(4.78)	.251(6.80)	.227 (7.33)
Hopeful	.218(1.00)	.469(2.91)	.363(2.70)	.367(2.83)	.387 (3.92)
Rookie	.248(1.59)	.182(1.90)	-.202(1.96)	.223(2.69)	.009 (0.13)
Pitcher	.314(2.15)	-.002(0.02)	.102(0.97)	.085(1.03)	.065 (0.96)
Constant	8.770(37.7)	8.911(53.8)	8.841(60.5)	8.889(64.8)	8.912 (83.3)
R^2	.81	.84	.83	.80	.82
Degrees of Freedom	96	146	159	262	325
Residual Sum of Squares	7.79	7.92	9.30	17.94	19.06

a. Estimates of the model in Table 3 for indicated subsets of the sample. Variables are as defined in Table 3 and t-statistics appear in parentheses.

TABLE 6

F-Tests for Pooling Validity[a]

Years Pooled[b]	Critical F[c]	Calculated F[d]
1968 and 1969	1.62	1.72
1969 and 1970	1.61	1.64
1968-69 and 1970	1.60	2.03
1968 and 1969-70	1.60	2.74
1968 through 1970	1.42	1.97

a. Critical and calculated values of the F-statislic for testing whether the given subset can be pooled. The null hypothesis is that the coefficients in the stated years are equal.

b. The years refer to observation subsets. The estimated equations used for the tests are those in Table 5, plus the equation applied to the entire sample, 1968-70, reported as Model II in Table 3.

c. The values of F (n,d) from tables of the F-distribution, n = degrees of freedom in the numerator and d = degrees of freedom in the denominator. If the calculated F exceeds the critical value, pooling is rejected at the .05 level.

d. Calculated in the standard fashion:

$$F = \frac{(RSS_p - RSS_s) \;/\; (DF_p - DF_s)}{RSS_s \;/\; DF_s}$$

where RSS indicates the residual sum of squares, DF indicates degrees of freedom, p indexes the pooled regression, and s indexes the separate regressions.

and starter ratio are stable and usually significant by conventional standards. Stable as well are the age and experience variables, including the previously discussed inverse relationships between salary and experience residual although the effect is much stronger for players who are more experienced than their age would predict (Negative Experience Residual).

Not surprising is that the unstable elements in Model II include the variables measuring the departures from usual, or average, performance. This result is intuitively plausible. The predictive power of the previous year's performance, relative to a player's historical average, must surely be affected by a change in the league-average skill level owing to expansion. In addition, except for the Hopeful variable during post-expansion, the various dummy variables are unstable. Most notable among these is the Rookie variable which is approximately equally significant between 1969 and 1970, but changes sign between the two years.

In an attempt to eludicate expansion effects statistically, we took the following approach. Most players on expansion teams are selected from among major-league reserve players and minor-league rosters. Hence, the performance variables of expansion players--especially playing time variables--are likely to have a different meaning in the salary determination process than the same variables will have for players on established teams. After all, even a team of dregs must elevate some players to the status of regulars. Hence, in the Fort-Noll sample, we separated players who finished a given season playing for an expansion team from those players that did not, i.e. we did not count players who were "just passing through" an expansion team as expansion players. We then reestimated Model II allowing both the intercept and slope parameters to vary according to whether or not the player was an expansion player as follows:

$$\ln\text{ (salary)} = \alpha_0 + D\alpha_1 + \beta' X + \gamma' DX + u.$$

The αs are constant terms, X is the vector of independent variables, βs and γs are coefficients, u is the error term, and D equals one if the player is an expansion player (zero otherwise). The test is simply to determine whether the set of γ coefficients, the difference between expansion players and others, add significantly to the explanatory power of Model II and, individually, are significantly different from zero.

The results appear in Tables 7 and 8. There were 55 expansion players in the 1969 sample and 35 in 1970. The first table is the full expansion characterization of Model II in Table 3 while the latter reports a trimmed version. Based on F-tests (using an unreported pooled expansion characterization for the entire sample, 1969-70), the following two results were found using the first two columns of Table 7: (1) the expansion characterization of Model II revealed that the expansion variables were statistically significant in 1969 but not in 1970 and (2) even with the expansion characterization, pooling of the 1969 and 1970 is invalid.

The rationale for the trimmed version of the expansion

TABLE 7

Expansion Characterization of Model II[a]

Variable	1969 Expansion Coefficients	1969 Other Coefficients	1970 Expansion Coefficients	1970 Other Coefficients
At Bats Share +1	-6.880(1.97)	13.140(7.57)	3.131(0.74)	10.230(6.49)
Slugging Ave. +1	.405(0.30)	.816(1.16)	-2.070(1.43)	2.158(4.04)
Pos. At Bats Depart.	-.065(1.71)	.095(3.72)	.119(1.48)	.010(0.49)
Neg. At Bats Depart.	-.045(0.86)	.073(3.08)	.052(0.56)	.012(0.50)
Pos. Slug. Ave. Depart.	-3.650(0.16)	3.068(1.68)	3.188(0.63)	-.487(0.42)
Neg. Slug. Ave. Depart.	3.379(1.67)	-2.847(4.15)	-1.127(0.33)	.856(0.63)
Innings Pitched Share +1	-11.110(3.02)	12.380(7.10)	2.851(0.56)	9.327(6.19)
Power Pitching Ratio +1	.151(0.44)	.506(2.40)	-.329(0.29)	.616(3.86)
Starter Ratio +1	.762(1.89)	-.914(4.28)	.147(0.23)	-.558(2.54)
Pos. Power Pitch. Depart.	-.034(0.13)	-.027(0.17)	-.813(1.12)	.637(2.29)
Neg. Power Pitch. Depart.	.579(1.33)	-.060(0.20)	-.950(1.31)	.485(2.84)
Old Age Squared	.066(1.41)	.043(2.64)	.061(1.14)	.000(0.02)
Young Age	.125(0.54)	.141(1.03)	.027(0.08)	.195(1.60)
Young Age Squared	-.016(0.12)	-.205(2.53)	.004(0.02)	-.176(2.57)
Pos. Experience Residual	.016(0.14)	-.042(0.69)	-.140(0.90)	.018(0.31)
Neg. Experience Residual	.063(0.67)	.235(4.69)	-.247(1.66)	.242(5.07)
Hopeful	-.219(0.65)	.588(2.91)	-.175(0.17)	.428(3.03)
Rookie	-.572(2.94)	.534(3.87)	.599(1.52)	-.239(2.16)
Pitcher	-.006(0.03)	.037(0.31)	.393(0.35)	.087(0.80)
Constant	.181(0.49)	8.771(44.1)	.028(0.05)	8.703(53.8)
R^2	.88		.86	
Degrees of Freedom	126		139	
Residual Sum of Squares	5.94		7.73	

a. As described in the text. Variables are defined in Table 3 and t-statistics are in parentheses.

b. Coefficient estimates pertaining to players who are characterized as expansion players, as described in the text.

c. Coefficient estimates for the sample as a whole. Again, the technique is described in the text.

TABLE 8

Trimmed Expansion Characterization[a]

Variable	1969 Expansion Coefficients[b]	1969 Other Coefficients[c]	1970 Expansion Coefficients	1970 Other Coefficients
At Bats Share +1	-7.705(2.42)	13.670(8.76)	2.287(0.69)	10.140(6.59)
Slugging Ave. +1	.658(1.03)	.718(1.20)	-1.816(2.12)	2.537(4.28)
Pos. At Bats Depart.	-.053(1.52)	.086(3.54)	.081(1.66)	.011(0.58)
Neg. At Bats Depart.	-.046(1.04)	.067(2.93)	.026(0.43)	.013(0.55)
Pos. Slug. Ave. Depart.	-	2.978(1.67)	-	-.371(0.34)
Neg. Slug. Ave. Depart.	3.817(2.41)	-2.850(4.26)	-1.574(0.62)	.903(0.68)
Innings Pitched Share +1	-7.946(2.93)	11.680(7.23)	2.480(0.77)	9.403(6.41)
Power Pitching Ratio +1	-	.570(3.49)	-	.602(3.97)
Starter Ratio +1	.698(1.88)	-.936(4.46)	.249(0.47)	-.522(2.56)
Pos. Power Pitch. Depart.	.039(0.16)	-.041(0.27)	-1.143(1.80)	.633(2.33)
Neg. Power Pitch. Depart.	.640(1.50)	-.126(0.44)	-1.410(2.77)	.492(2.95)
Old Age Squared	.041(1.10)	.046(2.97)	.063(1.64)	.002(0.07)
Young Age	-	.194(1.80)	-	.194(1.76)
Young Age Squared	-	-.214(3.42)	-	-.176(2.87)
Pos. Experience Residual	-	-.044(0.89)	-	-.001(0.01)
Neg. Experience Residual	.059(0.76)	.239(5.15)	-.132(1.30)	.232(5.08)
Hopeful	-	.498(3.15)	-	.446(3.36)
Rookie	-.427(2.72)	.450(3.63)	.491(1.68)	-.221(2.07)
Pitcher	-	.012(0.14)	-	.099(0.95)
Constant	.182(2.32)	8.789(53.7)	.090(0.83)	8.705(57.9)
R^2	.88		.86	
Degrees of Freedom	133		146	
Residual Sum of Squares	6.13		7.84	

a. Trimmed according to the omitted variables, due to multicollinearity problems described in the text.

b. See note b, Table 7.

c. See note c, Table 8.

characterization for Model II is as follows. In the 1970 estimation of the full expansion characterization, the presence of individually significant expansion variables that, taken as a group, are no longer significant is a distinct indication of acute multicollinearity. This problem may be especially serious for precisely those players that the expansion characterization is designed to handle—major-leaguers with marginal skills. Lacking more data, the standard remedy for such problems is to drop variables. The result is shown in the last two columns of Table 7. While pooling is still disallowed under the trimmed version of the expansion characterization, we do find that expansion variables are statistically significant for both 1969 and 1970.

These results, gleaned from distinguishing expansion players, provide evidence that our respecified model (Model II) has captured important aspects of the salary process. We believe that the following additional evidence is the clincher. When an expansion characterization of the original Scully model (as just discussed for Model II) is applied to the Fort-Noll sample (1969, 1970, and pooled 1969-70), one finds that: (1) expansion effects are not statistically significant in either 1969 or 1970, and (2) pooling remains valid for the Scully model, even under the expansion characterization. These results hold for both hitters and pitchers. Thus, the respecified model, but not the Scully model, enable us to detect the effect of expansion on salary structure: players on expansion teams are paid according to a different pay-performance relationship than our players on established teams. Unfortunately, we cannot be very specific about exactly how these structures differ, except to point out that most of the expansion coefficients in Table 8 are negative, implying the obvious. Expansion players are generally paid less than players of like histories on established teams. This is hardly surprising, of course, since the fact that a player was made available in an expansion draft is indicative of his value to the team that gives him up.

V. CONCLUSIONS

Using data that are more representative of the population of baseball players on major-league rosters and a respecification of salary determination that is both theoretically and empirically plausible, the following general results were obtained. First, the empirical models of the past, culminating in the work by Scully, are qualitatively but not quantitatively robust when applied to a different data set. Substantially less of the total salary variation in the Fort-Noll sample is explained by Scully's model than was explained in his original sample, and some important differences in the explanation itself emerge. Most notable among these differences is the overall reduction in the magnitude of the estimated coefficients (elasticities) and the change in the importance of the market size variables between the samples. We conclude that the salary determination process for the more general player population represented in our sample is different than the process for the more established players which dominate Scully's sample.

Second, our respecification of the salary relationship provides interesting insights. Roughly the same degree of explanatory power is achieved even though the theoretically suspicious market size variables are omitted from our model. Performance and playing time are just as important as they have been in past studies, and a more detailed analysis of recent departures from averages of these variables yields additional explanation of the salary outcome. Relief pitchers are distinguished from starters, and the separate effects of age and experience are untangled in a detailed fashion. Further, while salary elasticities with respect to playing time variables are close to those from Scully's model, the elasticities for the other performance variables are much lower than Scully's model would indicate from the sample at hand; only all-time great sluggers and strikeout pitchers would conform to the elasticities estimated from the Scully model. Finally, under an expansion characterization of our salary model, we find statistical evidence that changes in industry structure have a significant effect on player salaries. First, pooling observations across sample years is not statisticaly supported; the observations in separate years do not conform to a single salary model. Second, expansion players are statistically distinguishable from other players in both post-expansion years. While pooling these two years, even under the expansion characterization, remains invalid, the evidence strongly supports the hypothesis that expansion is the major cause of the inapplicability of a single salary model over the entire sample. The most important conclusion from the foregoing analysis is that one must keep one's eye on the ball if one expects to hit it; studies of wage payments to players must take account of changes in industry structure if the endeavor is to provide a complete explanation of the pay-performance relationship.

FOOTNOTES

1. The evidence generally supports the proposition that most teams maximize profits but that a few (most notably for our sample, the Boston Red Sox) are more aptly described as maximizing victories subject to a break-even revenue constraint (Noll, 1974).

2. Several salary observations in our sample fall below the minimum major-league salary. The reason is that some transacted players were on minor-league rosters for all or part of the sample period. Although major-league teams have a 25-man roster limit (until September 1), the teams actually have access to forty players. Owing to injuries, demotions to the minor-leagues, and late-season additions, nearly every player on the forty-man roster will play on the major-league team at some point during the season. The mean salary for the forty-man roster can be estimated by dividing the average payroll for players of major-league teams by forty. The result is an average salary between $25,000 and $30,000 during the sample period, based on data reported in Noll (1974), Chapter 1.

3. Two possible sources of sample bias in our data are: (1) salary observations are for players who were traded or sold and (2) most of our players are from the American League. We believe that neither problem is important. First, nearly all major-league

players are transacted during their careers and, in any case, because we were supplied with player salaries for seasons in which the player was not subject to trade or sale, about half the salary observations are for players who were not transacted in the year of observation (Table 1). In addition, the neoclassical theory of the firm based on profit-maximization predicts that the transfer of players from one team to another should not be correlated with their salaries. Players are subject to transaction when their marginal value products are higher in the new location than at the old, a circumstance unaffected by salaries paid by former teams (see El Hodiri and Quirk [1971], [1974]). Salaries should then determine the terms of the transaction, but not whether it takes place. If teams do not maximize profits, but instead maximize victories subject to a lower-bound constraint on net revenue, then players who, _ex post_, are overpaid relative to their productivity will be more likely to be transferred from teams in relatively poor markets to teams in relatively better ones. Regarding the preponderance of American League players, we report below statistical analysis that finds no difference between American League and National League salaries.

4. Logically, salary at time t depends upon performance up to time t-1 (i.e. last season); performance summary variables are calculated up to 1967 for 1968 salary observations and so on.

In our definition of a baseball season, we set the beginning of a fiscal year at October 1, for that is approximately the end of the season.

5. Population is a major factor influencing attendance, as shown in Noll (1974), Chapter 4, and it is very likely to be highly correlated with Scully's marginal attendance estimate. Consequently, the t-values reported in _both_ our sample and Scully's for these variables are unreliable. Nevertheless, the estimates are more plausible in the regressions on our data for both variables. First, the estimated coefficients have the predicted positive sign. Second, if market characteristics have an effect on salaries, there is no theoretical reason to expect that the magnitude of the effect should differ substantially between pitchers and hitters. Indeed, in our sample, the magnitudes are close for the two player categories.

6. Indeed, one single measure--playing time--seems to capture a large amount of the performance explanation. A rational team will naturally give more playing time to players contributing most to the team.

7. This transformation is not as innocuous as it might first appear. Some of the performance variables are ratios that take values near zero for all players (e.g. at bat share). Near zero,

the logarithmic function has much more curvature than it does near one, so that the data transformation is tantamount to a change of the assumption that the salary function is highly nonlinear to an assumption of near linearity.

8. Players of star caliber would normally have statistics a little above the top of the range chosen, and marginal players who are rarely used should have statistics lower than the bottom of the range. But the range is representative of the vast majority of players.

9. For evidence on the independent effect of star status, see Noll (1974), Chapter 4.

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