PAY, PERFORMANCE, AND COMPETITIVE BALANCE IN THE NATIONAL HOCKEY LEAGUE

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INTRODUCTION

This paper uses data from the National Hockey League (NHL) to test several propositions that arise from the literature on the economics of professional team sports. One of the first problems studied in this literature was the effect of the player reservation system on players' salaries. In a pioneering work in this field, Scully [1974] showed that prior to the introduction of free agency the marginal revenue products of major league baseball players were considerably in excess of the players' salaries. Since professional hockey is the only professional sport that still has a strong player reservation system, it is of interest to measure the extent of monopsony power in this market. To accomplish this, we estimate the marginal revenue products for NHL players using data in 1993/94 and compare them with players' salaries. These comparisons form the basis of tests of two competing hypotheses on the relationships between pay, productivity, and experience. The data are also used to test a hypothesis suggested by Telser [1995] that outcomes in a monopsonist labor market can be viewed as the result of playing the *ultimatum game.*

Perhaps the most fundamental issue in the sports literature is the extent to which competitive balance among the teams is affected by institutional arrangements. The NHL differs from other professional sports leagues in that there is no gate sharing, very little sharing of television revenues, and no salary cap. The institutions considered in this paper are the NHL system of limited free agency and its reverse-order-of-finish entry draft. The standard model of a sports league views teams as participating in a non-cooperative game in which profit-maximizing behavior by the teams leads to an equilibrium outcome with unequal playing strength among the teams. If sales and trades of player contracts and draft choices are allowed, the model generates an interesting invariance proposition whereby the movement from a player reservation system to free agency or the introduction of a rookie draft does not alter the competitive balance in the league.[1]

The major empirical tests of the invariance theorem have considered the introduction of free agency in major league baseball and rookie drafts in professional football and baseball. In each of these cases, the institutional change occurred in a given year and thus afforded a natural experiment for testing the invariance hypothesis using time-series data. Testing the invariance proposition using hockey data is more difficult because the changes in the player reservation system have been minor and

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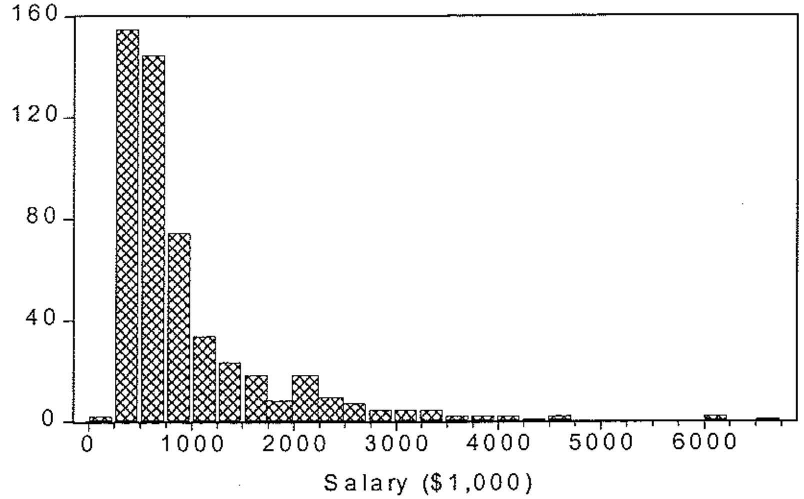
FIGURE 1
Histogram of Players' Salaries, 1995-96



have been introduced gradually rather than abruptly. The problem with the entry draft in hockey is that it has been operating since 1963, so any test of its impact using time-series data would be complicated by the expansion of the NHL from the original 6 teams to the current 27 teams, as well as the presence of the rival league, the World Hockey Association, from 1972 to 1979. In spite of these difficulties, this paper will attempt to shed some light on the validity of the invariance proposition with respect to free agency and the entry draft in hockey.

The plan of the paper is as follows. Summary data on recent changes in the level and dispersion of players' salaries are presented in the next section. This is followed by a section on the estimation of marginal revenue products. The remaining sections of the paper include comparisons of marginal revenue products and salaries, the effects of free agency, definitions and measurement of competitive balance, and the effect of the entry draft on competitive balance. The final section includes a summary and concluding comments.

RECENT TRENDS IN PLAYERS' SALARIES

Figure 1 is a histogram of the salaries of 520 players in the NHL in 1995/96. As is evident from the histogram, the distribution of players' salaries is highly skewed to the right. The ratio of the mean of the distribution ($964,421) to its median ($680,892) is 1.42. By comparison, the ratio of the mean earnings to the median earnings for male full-time, year-round workers in the United States in 1995 was 1.28 [U.S. Bureau of the Census 1995, 38]. The distribution of hockey players' salaries is therefore more skewed than the distribution of earnings in the general population of workers.

Average salaries of players in the NHL from 1984/85 to 1995/96 are given in Table 1 along with the growth of these salaries. Although the average salaries in hockey have lagged behind those of other professional sports, especially basketball and baseball, the recent growth in hockey salaries has been exceptional. As a comparison, the increase in the average NHL salary in 1991/92 of 40.3 percent is roughly the same as the percentage increase in average baseball salaries in the year immediately following free agency in 1976 (47.7 percent) or immediately following the collusion rulings in 1991 (42.5 percent).[2] Unlike baseball, a full explanation for the explosion of hockey salaries cannot be found in changes in the rules for free agency or salary arbitration because no major changes were made to these institutional arrangements over the time period in question. Part of the increase in the average salary is due to the increase in average team revenues, as shown in the last column of Table 1. Other explanations might be the change in union leadership, public disclosure of salaries, a brief strike in April 1992, and an increase in the union-negotiated minimum salary in 1992/93 from $25,000 to $125,000.

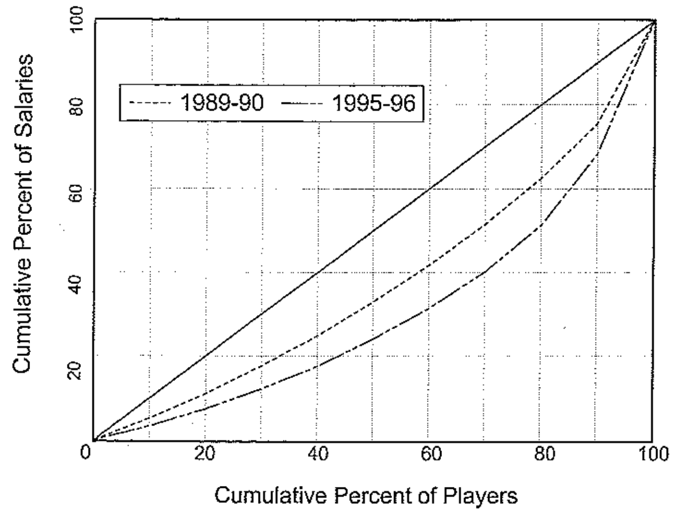
TABLE 1
Growth in Players' Salaries and Team Revenues

Season	Average Player Salary[a] ($1,000)	Average Player Salary[b] ($1,000)	Average Team Revenue[c] ($ millions)
1984-85	149		
1985-86	159 (6.7%)		
1986-87	173 (8.8%)		
1987-88	184 (6.4%)		
1988-89	201 (9.2%)		
1989-90	232 (15.4%)	225	21.0
1990-91	263 (13.4%)	—	24.7 (17.7%)
1991-92	369 (40.3%)	349 (24.6%)	26.1 (7.7%)
1992-93	463 (25.5%)	—	28.3 (7.1%)
1993-94	558 (20.5%)	616 (32.8%)	31.4 (10.8%)
1994-95*		—	28.0 (-10.9%)
1995-96		964 (25.2%)	42.3 (51.1%)

Numbers in parenthesis are annual percentage increases from previous measurement.
* Season interrupted by lockout.
a. National Hockey League, as reported in *The Wall Street Journal*, 15 November 1994;
b. Author's calculations, based on 20 highest-paid players on each team;
c. *Financial World*, various issues.

Another salient feature of the distribution of salaries in the NHL is its marked increase in dispersion over time. The increase in dispersion is measured using individual data for four seasons: 1989/90, 1991/92, 1993/94, and 1995/96.[3] To make the comparisons as valid as possible, only the 20 highest-paid players on each team were included in the calculations. Each of the five univariate measures of dispersion presented in Table 2 shows a steady increase in inequality from 1989/90 to 1995/96 and, given the standard errors of these measures, these increases are statistically significant. This increase in inequality is also apparent in the Lorenz curves plotted in Figure 2.

Is the increase in inequality due to an increase in inequality between teams or within teams? This question can be answered by decomposing the mean logarithmic deviation into the sum of a between-team measure of inequality and a within-team



measure.[4] Since the proportion of players on each team is constant, changes in the inequality measure are not contaminated by changes in the relative sizes of the groups, as is usually the case with decomposition analysis. The decompositions of the mean logarithmic deviation, which are also presented in Table 2, show that within-team inequality accounts for between 79 percent and 87 percent of total inequality in players' salaries. Moreover, most of the increase in inequality from 1989/90 to 1995/96 is due to increases in the within-group component. This suggests that the cause of the increased dispersion in players' salaries will be found in the behavior within teams rather than in market forces which might have disparate impacts on the teams.

It is generally agreed that the increased dispersion observed in baseball players' salaries is due to the system of veteran free agency and salary arbitration that has evolved in that sport.[5] Given the substantial differences in the two labor markets, it seems unlikely that similar explanations can be invoked to account for the increase in dispersion in Table 2. Explanations that might be appropriate are: a reduction in fighting and violence in the games [Scully 1995, 66], the public disclosure of players' salaries accompanied by sharp increases in rookies' salaries, and a relative increase in the demand for the most highly skilled players.

ESTIMATING MARGINAL REVENUE PRODUCTS

The technique for estimating the marginal revenue products of individual players is similar to the procedure used in baseball by Scully [1974; 1989] and Zimbalist [1992]. The marginal revenue product of a player is estimated as the product of marginal revenue and marginal product. Marginal revenue is derived from a team revenue equation, which depends on the team's performance during the regular season and in the playoffs. Marginal product is derived from a team production function

TABLE 2
Measures of Dispersion in Players' Salaries

	20 Highest-Paid Players Per Team			
	1989-90	1991-92	1993-94	1995-96
Gini Coefficient	.262	.322	.357	.397
	(.024)	(.017)	(.019)	(.013)
Variance of Logarithm	.178	.271	.337	.427
	(.024)	(.024)	(.026)	(.027)
Theil's Entropy	.162	.201	.249	.281
Measure	(.051)	(.028)	(.043)	(.021)
Atkinson's	.112	.153	.185	.221
Index (ϵ = 1.0)	(.035)	(.032)	(.033)	(.030)
Mean Logarithmic	.118	.166	.205	.250
Deviation	(.026)	(.018)	(.023)	(.017)
Between-Teams	.023	.021	.043	.038
Within-Teams	.095	.145	.162	.212
Number of Players	420	440	520	520

Standard errors in parentheses.

that incorporates player performance at each of the three positions: forward, defense, and goaltender. Players are divided into two groups, roster players and reserve players, and the marginal revenue products and marginal salaries of the roster players are computed relative to the performance and salaries of the reserve players.

The team revenue equation is specified as a fixed-effects regression model in which the dependent variable is team revenue in a given year and the independent variables measure team performance, a linear time trend, a dummy variable for the lockout in 1994/95, and team fixed-effects dummy variables. Team revenue is defined as total revenues from gate receipts, broadcast media, stadium revenues, and revenues from licensing and merchandise. Broadcast revenues include national, local, cable, and pay-per-view television and radio. Stadium revenues include suites, luxury seating, concessions, parking, and venue advertising. All revenue data are provided by annual surveys of the teams conducted by *Financial World*, covering the seven seasons from 1989/90 to 1995/96.[6] In the regression analysis, the dependent variable team revenue (*TR*) is expressed in thousands of U.S. dollars.

Team performance is measured by the winning percentage in regular season games (*PCT*) and by the number of playoff game appearances (*POG*). *PCT* is defined as 1000 × (2 × games won + games tied)/(2 × games played). Both *PCT* in the current season and *PCT* in the previous season are included as regressors in the revenue equation. A linear time trend (*TREND*) is defined as 1 in 1989/90, 2 in 1990/91, and so on up to 7 in 1995/96. The coefficient of *TREND* thus measures the increase in team revenues over time that is independent of team performance and is experienced by all teams. *LOCKOUT* is a dummy variable equal to one in the lockout year, 1994/95. The team dummy variables measure all other influences of *TR* that vary across teams but do not vary over time. Examples of variables that do not vary much over time and are thus captured in the fixed-effects coefficients are market size, per capita income of the team's location, size of the team's arena, the number of professional sports franchises in the area, and whether or not the team is located in Canada.

This specification of the team revenue function is similar in many respects to revenue functions used in the baseball literature.[7] Both Scully [1989] and Zimbalist [1992] use lagged values of *PCT* in the revenue equation. None of the baseball studies use playoff performance measures, but that is probably not a serious misspecification because few teams make the playoffs in baseball and their playoff performance is therefore captured adequately by performance during the regular season. The situation is quite different in hockey, where 16 teams are involved the playoffs and the number of playoff games for a team can be as high as 35 percent of regular season games. Since revenue from playoff games in hockey can be a significant portion of total team revenues, omission of *POG* could result in a serious specification error.[8] None of the baseball studies use fixed-effects models to capture the differences among teams that do not vary over time.

The estimated team revenue function, apart from the team fixed effects, is:

$$(1)\quad \hat{TR}_{i,t} = a_{oi} + \underset{(4.867)}{8.954}\, PCT_{i,t} + \underset{(4.339)}{12.074}\, PCT_{i,t-1} + \underset{(51)}{293}\, POG_{i,t} + \underset{(158)}{3476}\, TREND_t - \underset{(783)}{9409}\, LOCKOUT_t .$$

Estimates of the standard errors of the coefficients are given in parentheses. All of the coefficients have the anticipated signs and all one-sided tests are significant at the five percent level. There were 161 observations in the combined time-series cross-section data set and the coefficient of multiple determination (R^2) was .953.

The coefficient of *POG* implies that an additional playoff game is worth about $293,000 to each team with a standard error of $51,000. Since playing sites alternate in the playoffs and virtually all of the revenues accrue to the home team, it follows that the value of a home playoff game is about $586,000 with a standard error of $102,000. This estimate is roughly consistent with a simple calculation based on an average ticket price ($40) times the capacity of the arena (15,000). Finally, the coefficient of *TREND* implies a trend increase in team revenues of about $3.476 million per year with a standard deviation of $158,000. This result indicates the rapid growth in popularity the sport has realized over the seven years of the sample. The coefficient of *LOCKOUT* implies an average loss in team revenues during the lockout of $9.409 million with a standard deviation of $783,000.

It is assumed that the number of playoff games (*POG*) is determined by *PCT* so that players affect team performance only by their impact on *PCT*. The relationship between *PCT* and *POG* was estimated by fitting the following Tobit (censored) regression model:[9]

$$POG_{i,t} = 0 \quad \text{if} \quad POG^*_{i,t} \leq 0 \quad \text{and}$$

$$POG_{i,t} = POG^*_{i,t} \quad \text{if} \quad POG^*_{i,t} > 0, \quad \text{where}$$

$$POG^*_{i,t} = \beta_0 + \beta_1 PCT_{i,t} + \beta_2 PCT^2_{i,t} + \varepsilon_{i,t}$$

and $\varepsilon_{i,t}$ are independent normal error terms with mean zero and standard deviation σ. $POG^*_{i,t}$ can be interpreted as a latent variable indicating whether or not the team made the playoffs. The Tobit model was estimated using the team data for the seven years and the following results were obtained:

$$(2)\quad \hat{POG}^*_{i,t} = \underset{(21.60)}{-106.43} + \underset{(.079)}{.358}\, PCT_{i,t} - \underset{(.000071)}{.000266}\, PCT^2_{i,t} \qquad \hat{\sigma} = 6.934.$$

From equations (1) and (2) the marginal revenue associated with a one-unit change in *PCT* is given by:

$$(3)\quad MR_{i,t} = 8.954 + 12.074/(1+r) + 293\,[.358 - .000532\, PCT_{i,t}]\, \Phi(z)$$

where r is the discount rate (.03) and $\Phi(z)$ is the cumulative distribution function of the standard normal distribution with

$$z = (-106.43 + .358\, PCT_{i,t} - .000266\, PCT^2_{i,t}) / 6.934.$$

The output of the team (*PCT*) is assumed to depend on the playing performance at each of the three positions: forwards, defensemen, and goaltenders. The playing performance of a team's forwards is measured by their goals per game (*GPGF*) and by their assists per game (*APGF*). The contribution of defensemen is measured by their points (goals plus assists) per game (*PPGD*) and by their penalty minutes per game (*PMPGD*). The performance of goaltenders is measured by their save percent (*SVP*), which is 1000 times the ratio of saves to shots attempted by the opposing team. The following equation was estimated using data on all NHL teams from 1989/90 to 1995/96:

$$(4)\quad \hat{PCT}_{i,t} = -4240 + \underset{(25.65)}{73.38}\, GPGF_{i,t} + \underset{(16.09)}{34.92}\, APGF_{i,t} + \underset{(10.49)}{83.28}\, PPGD_{i,t} - \underset{(2.258)}{5.633}\, PMPGD_{i,t} + \underset{(0.351)}{4.792}\, SVP_{i,t}$$

NOBS = 166 SER = 53.17 R^2 = .772

All of the coefficients in equation (4) are significant at the 5 percent level and all have the signs one would expect *a priori*.[10]

Equation (4) can be viewed as a production function in which *PCT* depends on the following player performance variables: *GPGF*, *APGF*, *PPGD*, *PMPGD*, and *SVP*. The marginal revenue product of player j on team i in year t is given by

$$(5)\quad MRP_{j,i,t} = MR_{i,t} \times \Delta_j PCT_{i,t}$$

where $\Delta_j PCT_{i,t}$ is the difference between $PCT_{i,t}$ with player j and $PCT_{i,t}$ without player j. $PCT_{i,t}$ without player j is defined by considering the performance of roster players relative to reserve players. The *MRP*s of roster payers are computed by assuming that the roster player would be replaced by a reserve player and the performance of that reserve player would be equal to the average of all reserve players in that position.[11] Since all of the performance measures for forwards and defensemen are on a per game basis, it is not necessary to adjust the data for the number of games played.

Marginal revenue products are calculated for all roster players in the 1993/94 season. Roster players are defined as the 20 highest-paid players on each team by position (i.e., the 12 highest-paid forwards, the 6 highest-paid defensemen, and the 2 highest-paid goaltenders). The reserve players are all other players for which salary and performance data are available. For 1993/94, there were 520 roster players and 279 reserve players. Summary statistics for these players are given in Table 3.

To illustrate the procedure for computing *MRP*s, consider a roster forward who played 65 games during the regular season and had 20 goals and 30 assists. His impact on team *PCT* would be through the performance variables *GPGF* and *APGF*. Using the coefficient estimates of equation (4) and *GPGF* and *APGF* for reserve players in Table 3, we calculate $\Delta_j PCT_{i,t}$ as

$$73.38[20 - (.1261)(65)] / 84 + 34.92 [30 - (.1717)(65)] / 84 = 18.14.$$

If it is further assumed that this player's team had a winning percentage of .500 this year, his $MR_{i,t}$ from equation (3) is

$$8.954 + 12.074/1.03 + 293 [.358 - .00532(500)] (.809) = 44.35.$$

It follows from equation (5) that the *MRP* of this player is (18.14) (4.35) = 804.51 or $804,510. Similar procedures are followed to compute the *MRP*s of defensemen and goaltenders.

In each case, it is implicitly assumed that the productivity of the other players is not affected when a roster player is replaced by a reserve player. This lack of interdependence among players is a critical assumption and is likely to lead to an underestimate of *MRP*. On the other hand, the procedure also fails to account for the fact that the performance per game of reserve skaters is likely to be lower because they do not receive as much playing time per game as roster players. This is likely to cause the estimates of *MRP* to be overestimated. The two biases thus work in opposite directions and have a net effect of offsetting each other.

RELATIONSHIP BETWEEN MARGINAL REVENUE PRODUCTS AND SALARIES

Using the procedure described in the previous section, an estimate of the marginal revenue product was computed for each of the 520 roster players in 1993/94. Marginal salary (*MSAL*) is defined as the difference between a roster player's salary and the average salary of the reserve players at that position. *SURPLUS*, defined as

TABLE 3
Summary Statistics for NHL Players, 1993-94

A. Forwards				
	Roster Players		**Reserve Players**	
Variable	**Number**	**Mean**	**Number**	**Mean**
Salary ($1,000)	312	623.91	151	199.53
Goals per game (*GPGF*)	312	.2382	155	.1261
Assists per game (*APGF*)	312	.3268	155	.1717
B. Defensemen				
	Roster Players		**Reserve Players**	
Variable	**Number**	**Mean**	**Number**	**Mean**
Salary ($1,000)	156	576.60	95	223.42
Points per game (*PPGD*)	156	.3590	96	.2143
Penalty Minutes per game (*PMPGD*)	156	1.226	96	1.262
C. Goaltenders				
	Roster Players		**Reserve Players**	
Variable	**Number**	**Mean**	**Number**	**Mean**
Salary ($1,000)	52	661.86	22	214.98
Save percentage (*SVP*)	52	.896[a]	28	.889[a]

a. Computed as a weighted average with number of shots against as weights.

MRP − *MSAL*, measures the extent to which the player is "underpaid" or "overpaid" in the 1993/94 season.

Summing *SURPLUS* over all 520 roster players yields a total of $5.0 million, which was only 1.6 percent of total salaries in 1993/94. Using the terminology of Zimbalist [1992], there is very little *monopsonistic exploitation* in the aggregate. There is, of course, considerable variation in *SURPLUS* across players, as would be expected given the crudeness of our estimation procedure and the fact that salaries are determined *ex ante* and *MRP*s *ex post*.

Summary statistics for *MRP*, *MSAL* and *SURPLUS*, by position, are given in Table 4. Aggregating by position, *SURPLUS* for forwards, defensemen, and goaltenders are −$3.1 million, $4.1 million, and $4.0 million, respectively. While the aggregates by position are interesting, care should be exercised in drawing conclusions from this data. It is likely that the differences by position are primarily due to the procedures used to measure player performance. Comparisons among players at a given position are probably more interesting and they reveal a considerable amount of dispersion and skewedness.

The simple correlation coefficients between *MRP* and *MSAL* of .52 for forwards and .47 for defensemen are higher than one might expect. Recall that the *MRP*s are based on actual performance during the 1993/94 season and salaries are set before the season begins. Even if salaries were based on expectations of *MRP*, there are bound to be significant deviations because owners' expectations may not be accurate.

TABLE 4
MRP, MSAL, and SURPLUS, by Position
(in thousands of dollars)

	Forwards	Defensemen	Goaltenders
MRP			
Mean	414	379	525
Median	279	161	251
Std Dev	548	644	1469
MSAL			
Mean	424	353	447
Median	201	229	273
Std Dev	661	372	450
SURPLUS			
Mean	-10	26	78
Median	-50	-92	37
Std Dev	599	574	1373
Correlation Between MRP and MSAL	.52	.47	.36

Younger players may perform better than expected and veteran players may miss games due to injuries. Variation would also result from long-term contracts and differing degrees of bargaining power under the reserve system.

Table 5 gives the average *SURPLUS* by salary quintile for each position. For forwards and defensemen, average *SURPLUS* is positive for the lowest salary quintile, negative for the second quintile, positive for the third and fourth quintiles, and then negative for the highest quintile. For goaltenders, those in the lowest and highest salary quintiles have negative average *SURPLUS* whereas those in the middle three salary quintiles have positive average *SURPLUS*. These patterns suggest that forwards and defensemen in the second and highest salary quintiles and goaltenders in the lowest and highest quintiles are "overpaid" relative to the players in the other salary quintiles.

What relationships would one expect to observe between *MRP* and salary for all players under a player reservation system? Three possibilities are considered here.[12] The first is a human-capital/life-cycle relationship in which salary is held below *MRP* early in the player's career and is above *MRP* later in the player's career. The lower salary early in the career reflects training expenses incurred by the firm and the sequencing of pay is designed to increase the productivity of younger players. The extent to which *MRP* deviates from salary throughout the life cycle depends on how much of the training is team-specific, as compared to general training. This hypothesis was tested by regressing *SURPLUS* on years of experience (*YRS*) and years of experience squared for each position. For forwards, the results are

TABLE 5
MRP, MSAL, and SURPLUS, by Salary Quintile
(in thousands of dollars)

	Salary Quintile	Average MRP	Average MSAL	Average SURPLUS
A. Forwards				
	1	165	50	115
	2	61	145	-83
	3	250	230	20
	4	599	452	147
	5	992	1279	-287
B. Defensemen				
	1	213	64	150
	2	44	156	-112
	3	250	241	9
	4	503	408	95
	5	885	944	-59
C. Goaltenders				
	1	-168	53	-220
	2	542	202	340
	3	411	296	115
	4	939	662	278
	5	1113	1177	-63

$$\widehat{SURPLUS} = -51.33 + \underset{(25.53)}{55.51\ YRS} - \underset{(2.06)}{5.69\ YRS^2} \quad (6)$$

NOBS = 312 SER = 591.93 $R^2 = .030$.

Although the fit as measured by the coefficient of determination is not very good, the F-test of overall significance yields a prob-value of .009. The estimated coefficients are both significantly different from zero at the five percent level and they produce the hypothesized pattern of positive *SURPLUS* early in the career, rising at a decreasing rate and becoming negative at about nine years of service. The results for defensemen and goaltenders are not, however, supportive of the human-capital hypothesis; the prob-values for the F-tests of overall significance were .738 and .900 for defensemen and goaltenders, respectively. The data thus support the human-capital/life-cycle hypothesis for forwards but not for defensemen or goaltenders.

A second possibility is a rank-order hypothesis in which competition for a starting position on the team is viewed as a rank-order tournament. In this view, the best players at each position are overpaid in order to induce higher productivity in the other players. This implies that *SURPLUS* is negative for the highest-paid players on each team and that *SURPLUS* increases at a decreasing rate as the rank of the player increases from the top to the bottom on the team. *RANKSAL* is defined as the salary ranking of the player on his team where 1 has the highest salary, 2 the next highest, etc. The hypothesis is tested for forwards and defensemen by regressing *SURPLUS* on *RANKSAL*, *RANKSAL* squared, and team dummy variables. The results for forwards, apart from the team dummy variables, are as follows:

$$(7) \quad \widehat{SURPLUS} = a_i + \underset{(112.83)}{124.68\, RANKSAL} - \underset{(3.14)}{8.46\, RANKSAL^2}$$

$$NOBS = 312 \qquad R^2 = .122$$

For defensemen, the estimated equation is

$$(8) \quad \widehat{SURPLUS} = a_i + \underset{(122.67)}{47.35\, RANKSAL} - \underset{(17.30)}{11.66\, RANKSAL^2}$$

$$NOBS = 156 \qquad R^2 = .148$$

The rank-order hypothesis implies that the derivative of surplus with respect to *RANKSAL* (the coefficient of *RANKSAL* plus two times the coefficient of $RANKSAL^2$) is positive and that the coefficient of $RANKSAL^2$ is negative. Both of these conditions are satisfied by the point estimates of equations (7) and (8). The coefficient estimates are significant in equation (7) but not in equation (8). We therefore accept the rank-order hypothesis for forwards but not defensemen.

The test of the rank-order hypothesis for goaltenders is more straightforward because there are only two roster goaltenders per team. The rank-order hypothesis implies that the *SURPLUS* of the starting goaltender is less than the *SURPLUS* of the backup goaltender. This hypothesis is tested using the Mann-Whitney-Wilcoxon test statistic, which is distributed as a standard normal variate under the null hypothesis that the surplus for the starter and the backup were from the same distribution. Since the computed value of the test statistic was 0.37, which is not significant, the data fail to support the rank-order tournament hypothesis for goaltenders.

A third hypothesis concerning the relationship between *SURPLUS* and the salaries of players is provided by Telser [1995]. He views the bargaining between players and owners in a bilateral monopoly situation as playing the *ultimatum game*. The ultimatum game is a two-person game in which a fixed amount of money is distributed among the players. One player offers a portion of the fixed sum to the second player who can either accept or refuse the offer. If the offer is rejected both players receive nothing. Although rational behavior implies that the final offers will be a small fraction of the initial amount, experimental evidence shows offers averaging from 40 to 60 percent of the initial amount. Telser notes that these experiments are performed with small amounts (usually $10) and suggests that if larger amounts were considered, the proportion of the initial amount going to the second player should fall as the size of the initial amount increases. Telser tested this proposition using

TABLE 6
Tests of the Ultimatum Game Hypothesis
Ratio of Salaries to Gross Marginal Revenue Product

Quintile	Mean	Standard Error
A. Forwards		
1	13.49	46.00
2	1.119	.684
3	.810	.389
4	.657	.395
5	.736	.628
Sample Size = 276	F = 4.28	Prob Value=.0022
B. Defensemen		
1	6.121	3.739
2	1.609	1.028
3	.894	.665
4	.494	.226
5	.477	.263
Sample Size = 127	F = 46.38	Prob Value=.0001
C. Goaltenders		
1	1.253	.497
2	.776	.505
3	.308	.124
4	.234	.082
5	.194	.182
Sample Size = 34	F = 12.45	Prob Value=.0001

the estimates of *MRP*s for baseball players obtained by Scully [1974] and found that the ratio of salary to net *MRP* does indeed fall as the *MRP* increases.

A test of the ultimatum game hypothesis for the NHL is conducted by computing the ratio of player salary to gross marginal revenue product for each roster player.[13] The players were then ranked by the value of the gross marginal product and a one-way analysis of variance was performed on the ratios with classes being the five quintiles for that position. The results presented in Table 6 support Telser's prediction that the means of the ratios should fall as the *MRP* quintiles increase. The decrease in the ratio is monotonic across all quintiles except the 5th quintile for forwards and the F-test clearly rejects the hypothesis of equal ratios for all three positions.

In summary, of the three hypotheses considered, only the ultimatum game hypothesis is supported by the data for all three positions. The human-capital/life-cycle hypothesis and the rank-order hypothesis are consistent with the data for forwards but not for the other two positions. While the results tend to view the ultimatum game hypothesis more favorably than the others, the testing procedures are not powerful enough to determine which of the three hypotheses is to be preferred for forwards. The three hypotheses are not mutually exclusive and the best conclusion to draw from the analysis is that there is some support for all three hypotheses, at least for forwards.

THE EFFECT OF FREE AGENCY ON PLAYER SURPLUS

In 1993/94 the National Hockey League had six different categories of free agency. These categories were defined by age, years in the league, salary, and type of contract signed. Although players in each of these categories are referred to as free agents, the extent to which players were free to negotiate with other teams varied considerably among categories. The largest group of free agents (Group I) could not change teams unless the team holding their contract received compensation in the form of cash and/or future picks in the entry draft. The compensation schedule was clearly designed to discourage the movement of highly-paid players. For example, the compensation for a player earning $400,000 or more[14] was $100,000 and two first-round draft picks, each of which must be one of the top seven draft picks. If the new club failed to deliver both draft picks, it would have to pay an additional $100,000 and forfeit the next five first-round picks!

Of the 520 roster players in 1993/94, only 48 were free agents prior to the season. Since there were not enough free agents to consider each group of free agents separately, all free agents were grouped together in the analysis. To estimate this effect of free agency, a regression equation was computed with *SURPLUS* as the dependent variable and the following independent variables: a free agency dummy variable (equal to one if the player was a free agent in 1993/94), marginal revenue product, marginal revenue product squared, a dummy variable for election to the all-star team in a prior season, dummy variables for positions (defense and goaltender), number of games played in 1993/94, number of years in the league, and number of years in the league squared, as well as interaction terms for these variables. Team fixed-effect variables were also included and they were selected using the backward elimination method with a ten percent level of significance.

The estimates of this regression equation are not reported here because many of the coefficients were not significant at the ten percent level. Table 7 reports the estimates of two regression equations in which only the significant variables are included as independent variables. The first equation includes team fixed effects while the second does not.[15] Both equations lend some support to the hypothesis that free agency depresses a player's *SURPLUS*, other things being equal. The estimated equations imply that free agency reduces *SURPLUS* by $61,000 for the equation with team fixed-effects and by $87,000 for the equation without team fixed effects. The standard errors of these coefficients, however, are such that the one-sided prob-values under the null hypothesis of a zero coefficient on the free agency variable are .20 and .12, respectively. These cross-section results thus offer only limited support to the proposition that free agency effectively transfers wealth from the owners to the players.

TABLE 7
Effects of Free Agency on SURPLUS

Dependent Variable = SURPLUS (in $1,000)

Independent Variable	With Team Fixed Effects	Without Team Fixed Effects
FREE AGENT	−60.77	−86.57
	(72.58)	(73.76)
MRP	0.7520	0.7545
	(0.0308)	(0.0313)
ALL STAR	−771.90	−826.44
	(104.25)	(105.80)
YEARS	−13.33	−14.98
	(5.41)	(5.51)
DET	−200.86	—
	(108.83)	
FLA	205.06	—
	(108.69)	
LA	−378.64	—
	(109.40)	
NYR	−294.91	—
	(109.03)	
STL	−196.45	—
	(109.29)	
R^2	.557	.533
Sample size	520	520

Standard errors are in parentheses.

Why is the measured effect of free agency on *SURPLUS* so weak? It may be due to the fact that so few of the players were free agents in 1993/94 and that all six types of free agents were grouped together in the analysis. It could also be the case that free agency is simply not an effective means of transferring wealth from owners to players because the player reservation system is so strong.

COMPETITIVE BALANCE

In assessing the effects of institutional changes in a sports league it is important to measure how the degree of competitive balance in the league changes over time. This section offers measurements of changes in competitive balance in the NHL using two different statistics: dispersion in regular season winning percentages and the number of playoff games. Both trends and correlations over time are considered.

The dispersion in the winning percentages in a season is calculated as the ratio of the standard deviation of the winning percentages among the teams to the standard deviation that would prevail if all teams were of equal playing strength. The denominator of this ratio is determined by taking into account the three possible outcomes of a regular season game: win, loss, or tie. Under equal playing strength, the probability of a win is equal to the probability of a loss. Using the trinomial distribution, the standard deviation of the winning percentage under equal playing strength is $[(1 - p)/4n]^{1/2}$ where p is the probability of a tie and n is the number of games in the season. To measure competitive balance, we therefore need an estimate of the probability of a tie. The historical proportion of ties in regular season games provides one estimate. From 1979/80 to 1982/83 the proportion of ties was .173 and from 1983/84, when overtime games were introduced, to 1995/96 the proportion of ties was .117. The problem with using historical proportions is that they clearly underestimate the true probability of a tie because all of the games during the regular season are not between teams of equal strength.

An alternative method of estimating the probability of a tie under equal playing strength is to assume that the number of goals scored by a team in a game follows a Poisson process and is independent of the number of goals scored by the opposing team.[16] This procedure requires an estimate of the parameter (mean) of the Poisson distribution. In an effort to include only evenly matched contests, the data set consists of the scores of the games in all Stanley Cup series during the past ten years for teams that played six or seven games. There were 862 such games and the mean number of goals scored per game was 3.158.[17] The resulting Poisson distribution fit the data very well[18] and implied a 0.162 probability of a tie.[19]

The ratios of the standard deviation of the winning percentage to the standard deviation expected under equal playing strength from 1979/80 to 1998/99 are given in Table 8 and are plotted in Figure 3. This measure registers a slight increase in competitive balance from 1982/83 to 1987/88, a reduction in competitive balance from 1987/88 to 1991/92, and then another increase from 1992/93 to 1998/99. A linear regression of the three-year moving average of the ratio on a time trend and the number of teams in the league yields

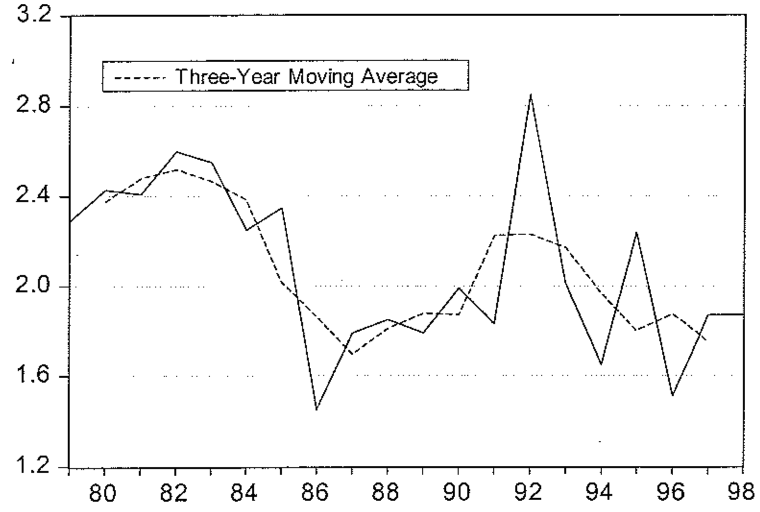
$$\text{(9)} \quad \widehat{RATIO} = \underset{(1.58)}{.896} - \underset{(.039)}{.075}\, TREND + \underset{(.084)}{.088}\, TEAMS$$

where $R^2 = .742$, the sample size is 17, the estimated first-order serial correlation is .625, and standard errors are in parentheses. The coefficient of *TREND* is significantly negative at the five percent level. The ratio, therefore, is on a discernible downward trend over the 19-year period. According to this measure, the NHL has become more competitive over time.

The cyclical pattern in the ratio in Figure 3 has no simple explanation. It is not correlated with changes in the collective bargaining agreement (1982, 1984, 1986, 1992), the change in leadership of the players' union (1991), the addition of expansion teams (1991, 1992, 1993), or major changes in the playing rules.

The second measure of competitive balance is the total number of playoff games played in a season. At the end of the regular season of play, the top 16 teams play a single-elimination tournament to determine the winner of the Stanley Cup. The tournament has four possible rounds and the team that wins four out of a possible

FIGURE 3
Change in Competitive Balance
SD(PCT) to SD(PCT) Under Equal Playing Strength



seven games advances to the next round. The total number of playoff games thus varies from a minimum of 60, where one team wins the first four games of each round, to a maximum of 105, where all rounds go to seven games. A small number of playoff games is associated with unequal playing strength and a lack of competitive balance. Since playoff games have no ties, it is relatively easy to compute the distribution of the total number of playoff games when all teams are of equal playing strength and the probability of any team winning a given game is .5. Under equal playing strength, the total number of playoff games has a mean of 87.1875 and a standard deviation of 3.926.

The data series for total playoff games only extends back to 1985/86 when a seven-game series was introduced for all rounds of the playoffs. The number of playoff games per season is given in Table 8 and is plotted in Figure 4. There is a slight downward trend and all of the observations are within the two standard deviation band that would be expected under equal playing strength. Using playoff games as a measure of competitive balance, we find a slight improvement in competitiveness and cannot reject the hypothesis of equal playing strength among the teams making the playoffs.

Competitive balance can also be measured by the correlation between a team's winning percentage or the number of playoff games in one season and its winning percentage or playoff games in previous seasons. A lack of competitive balance in the league would be indicated by a high correlation coefficient, and one would expect the correlation coefficient to fall as the lag length increased.

The correlation coefficients for winning percentages, which are given in Table 9, show a strong correlation for one- and two-year lags. Of the 19 one-year lag coefficients, only 1, (the 1991/92 year), does not have a significantly positive correlation at the ten percent level. Of the 18 two-year lag coefficients, 7 are not significantly posi-

TABLE 8
Measures of Competitive Balance

Season	Std Dev	Ratio	Playoff Games
1979-80	.1117	2.29	
1980-81	.1185	2.43	
1981-82	.1172	2.41	
1982-83	.1265	2.60	
1983-84	.1305	2.55	
1984-85	.1154	2.25	
1985-86	.1205	2.35	
1986-87	.0742	1.45	87
1987-88	.0916	1.79	83
1988-89	.0948	1.85	82
1989-90	.0919	1.79	85
1990-91	.1019	1.99	92
1991-92	.0937	1.83	86
1992-93	.1421	2.85	85
1993-94	.1003	2.01	90
1994-95	.1089	1.65	81
1995-96	.1133	2.24	86
1996-97	.0763	1.51	82
1997-98	.0943	1.87	82
1998-99	.0946	1.87	86

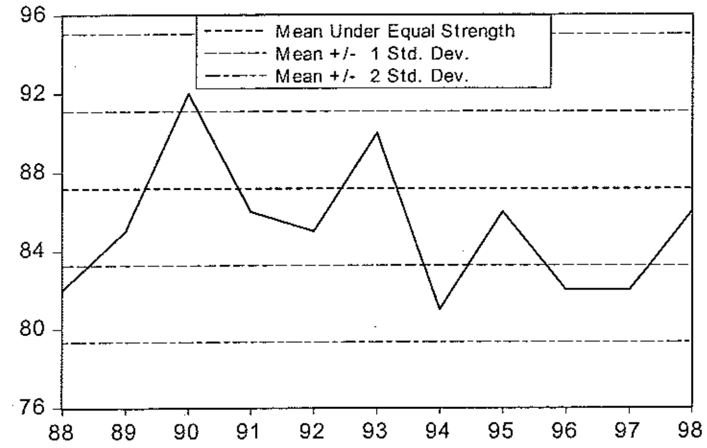
Ratio is the standard deviation of the winning percentage divided by the standard deviation under equal playing strength.

tive. Ten of the 17 three-year lag coefficients are not significantly positive at the ten percent level.

The lag correlations for playoff games, which are given in Table 10, are much weaker than those for winning percentages. Only 4 of the 12 lag coefficients are significantly positive at the five percent level; only 1 of the 11 two-year lag coefficients is significant; and none of the 10 two-year lag coefficients is significant. Since the same teams (usually expansion teams) tend to miss the playoffs year after year, it is not surprising that there is more balance using playoff games than winning percentages.

The results on competitive balance in the NHL are mixed. Using winning percentage as a measure, the long-term trend is toward *more* competitiveness accompanied by strong correlations in winning percentages over time, at least for one- or two-year lags. Playoff games provide no evidence of imbalance: the data show no statistically significant deviation from equal playing strength and few significant correlation coefficients over time. The general conclusion that emerges from these calculations is that no significant change in playing strengths has developed among the teams in the NHL. On balance, the increased revenues in the league, the escalation of salaries, or the addition of new teams in the United States did not result in any major changes in relative playing strengths among the teams.

FIGURE 4
Number of Playoff Games



Do these measures of competitive balance throw any light on the validity of the invariance proposition with respect to the player reservation system? If one accepts that the reserve system has been gradually liberalized over time, these results do lend support to the invariance proposition, (i.e., changing the institutional rules has not had an effect on competitive balance in the league). On the other hand, one could argue that the competitive balance has not changed because in the rules governing the reserve system have not changed. Perhaps the safest conclusion to draw is that the institutional arrangements have not changed sufficiently to permit a test of the invariance proposition with respect the player reservation system.

THE IMPACT OF THE ENTRY DRAFT ON COMPETITIVE BALANCE

Of the institutional arrangements that are intended to maintain competitive balance in a sports league, the most easily identified and measured is the entry draft. The NHL entry draft is conducted in June following the conclusion of the season and the draft pick order is in reverse order of finish in the regular season.[20] Since players eligible for the entry draft are primarily 18-year-olds who play in the minor leagues or college for three or four years, the lag time between the entry draft and the expected effect on team performance is long. To test the relationship between the entry draft and competitive balance,[21] the winning percentage of a team (*PCT*) at time t was regressed on a distributed lag of its draft pick order (*DP*)[22] in years $t-1$ to $t-6$ and a distributed lag on *PCT* from $t-1$ to $t-3$. The regression also included dummy variables for each team, selected using a backward elimination method at a ten percent level. The only distributed lag coefficients that were significant were the coefficient of the draft pick in $t-4$ and the coefficients of PCT_{t-1} and PCT_{t-2}. The equation was reestimated to include only the significant distributed lag coefficients. The estimated equation was:

TABLE 9
Correlations of Winning Percentages with Previous Seasons

	Pearson Correlation Coefficients of PCT_t and		
	PCT_{t-1}	PCT_{t-2}	PCT_{t-3}
1980-81	.732[a]	—	—
1981-82	.495[b]	.559[b]	—
1982-83	.713[a]	.428	.608[a]
1983-84	.793[a]	.699[a]	.327
1984-85	.654[a]	.629[a]	.461[c]
1985-86	.677[a]	.679[a]	.644[a]
1986-87	.618[a]	.683[a]	.457[c]
1987-88	.630[a]	.354	.522[b]
1988-89	.649[b]	.434[c]	.227
1989-90	.480[b]	.450[b]	.291
1990-91	.618[a]	.589[a]	.375
1991-92	.425	.291	.306
1992-93	.518[b]	.124	−.204
1993-94	.596[a]	.347	.216
1994-95	.622[a]	.677[a]	.106
1995-96	.787[a]	.487[b]	.659[a]
1996-97	.433[c]	.367	.182
1997-98	.495[b]	.204	.466[b]
1998-99	.704[a]	.534[a]	.052

a. Indicates that the coefficient is significantly positive at a 1 percent level.
b. Indicates that the coefficient is significantly positive at a 5 percent level.
c. Indicates that the coefficient is significantly positive at a 10 percent level.

$$(10)\quad \hat{PCT}_{t,i} = a_i \underset{(.698)}{- .778\,DP_{t-4,i}} \underset{(.051)}{+ .440\,PCT_{t-1,i}} \underset{(.052)}{+ .214\,PCT_{t-2,i}}$$

where a_i is the fixed team effect for team i, R^2 = .388, the sample size is 350, and standard errors are in parentheses. Since the low draft picks produce the better players, the expected sign of the draft pick coefficient in equation (10) is negative. The estimated coefficient of DP in equation (10) is negative but not significant at a 5 percent level.

In an alternative specification, equation (10) was estimated with $DP_{t-4,i}$ replaced by the averages of three past draft pick orders, (e.g., $(DP_{t-3,i} + DP_{t-4,i} + DP_{t-5,i}) / 3$). The results of these regressions are reported in Table 11. The estimated coefficients of the average draft pick orders have the expected negative signs in all three cases. The estimated coefficient is not significant for the average of lags 2 through 4, but the coefficients for the average of lags 3 through 5 and for lags 4 through 6 are significantly negative at the five percent level. The results of these tests thus provide some support for the hypothesis that the entry draft is effective at maintaining competitive balance in the league.

TABLE 10
Correlations of Number of Playoff Games with Previous Seasons

	Spearman Correlation of POG_t and		
	POG_{t-1}	POG_{t-2}	POG_{t-3}
1987-88	.331	—	—
1988-89	.174	-.001	—
1989-90	.231	.366	−.015
1990-91	.451[b]	.482[b]	.274
1991-92	.511[b]	.361	.251
1992-93	−.127	-.035	−.282
1993-94	.149	.174	.159
1994-95	.602[a]	−.003	.117
1995-96	.385	-.025	.188
1996-97	.245	.217	−.194
1997-98	.412	−.033	.083
1998-99	.486[b]	.370	.026

a. Indicates that the coefficient is significantly positive at a 1 percent level.
b. Indicates that the coefficient is significantly positive at a 5 percent level.
c. Indicates that the coefficient is significantly positive at a 10 percent level.

SUMMARY AND CONCLUSIONS

Most of the published studies using hockey data have been concerned with salary determination and discrimination against Francophone players.[23] The focus of this paper is different in that it attempts to use the hockey data to test some propositions from the literature on the economics of professional team sports. The objective has been to do for hockey what has already been done for baseball and, to a lesser degree, football and basketball.

One of the main features of the paper is its methodology for estimating marginal revenue products of the players. The estimates of *MRP* lead to several conclusions. First, the aggregate *SURPLUS* of only 1.6 percent of salary suggests that, in the aggregate, the player reservation system does not result in significant monopsonistic exploitation. Looking at the data by position and salary rank, forwards and defensemen in the first, third and fourth salary quintiles generate positive *SURPLUS* while those in the second and highest salary quintile have negative *SURPLUS*. Goaltenders in the middle three quintiles have positive *SURPLUS* while those in the lowest and highest salary quintiles have negative *SURPLUS*. Second, regarding the relationship between *MRP* and *MSAL*, the data support the ultimatum game hypothesis for all three positions and the human-capital/life-cycle hypothesis and the rank-order hypothesis for forwards. There is no support for the human-capital/life-cycle hypotheses or the rank-order hypothesis for either defensemen or goaltenders. Third, the evidence suggests that free agency reduces the surplus that a player generates, but the effect is not highly significant.

TABLE 11
Effect of Draft Pick on Winning Percentage

Dependent Variable = PCT_t
Time Period: 1979-80 to 1998-99

	ADP_t = Average of Draft Picks in Years		
	$t-2, t-3, t-4$	$t-3, t-4, t-5$	$t-4, t-5, t-6$
ADP_t	0.463	−1.440	−1.740
	(.927)	(0.877)	(0.875)
PCT_{t-1}	.450	.411	.437
	(.052)	(.053)	(.055)
PCT_{t-2}	.210	.213	.207
	(.055)	(.054)	(.055)
R^2	.387	.384	.344
Sample size	350	324	298

Numbers in parentheses are standard errors.

Although the results on *SURPLUS* are interesting and, in many respects, supportive of the theoretical arguments in the literature, some caveats about the procedure for estimating *MRP*s are in order. It is difficult to assess the validity of the *MRP* estimates because there is no standard on which to judge them. If the player market were perfectly competitive, the judgment could be based on how close *MRP*s were to *MSAL*. But given the restrictions on player mobility and the fact that salaries are set prior to the season, one would not expect a close relationship. Anyone familiar with the game of ice hockey would point out omitted variables and mis-specifications associated with the estimation procedure (e.g., forwards do not contribute to team defense, coaching does not matter, and play during power plays and short-handed situations are not considered explicitly). Perhaps most importantly, the technology implied by the procedure assumes that player performance is separable and additive and that there are no interactions among players in the production of team performance. Nevertheless, there is some comfort in the fact that equations (1) through (3) fit the data well and have coefficients that are significant and of the expected signs.

The other major innovation of the paper is the procedure for measuring competitive balance in the league. If the focus is on the dispersion of winning percentages during the regular season, the conclusion is that there is a slight increase in competitive balance over time along with high correlations in winning percentages from year to year. Using the number of playoff games as a measure, there is also a slight improvement in competitive balance over time, but one cannot reject the hypothesis of equal playing strength among the playoff teams and the correlations in playoff games from year to year are very low.

The data on competitive balance were intended as a means of assessing the validity of the invariance proposition. If it is assumed that the player reservation system has been generally liberalized over time, then the relative constancy of the measures of competitive balance lend support to the invariance proposition. But this is not a very strong result because the rules governing player movement have changed very little. The other labor market institution considered is the reverse-order-of-finish entry draft. Here the data show that the order in the entry draft influences subsequent team performance, a result that is not supportive of the invariance proposition. On balance then, the empirical evidence presented in this paper does not lend support to the invariance proposition.

In an early test of the invariance proposition in baseball, Daly and Moore [1981] argued that teams had strong financial incentives to develop mechanisms for maintaining competitive balance in the league. The empirical results presented here can be viewed as supporting the Daly-Moore view in the case of hockey. It appears that the entry draft has been an effective means of maintaining competitive balance, and one could argue that NHL teams have managed to find other ways to equalize playing strength in the face of rising player salaries, a stronger union, and a gradual liberalization of the player reservation system.

Finally, what does the analysis say about the sharp increases in the average salary and dispersion of salaries from 1989/90 to 1995/96 that were presented in Tables 1 and 2? The estimates of the revenue equation (1) imply that winning regular season games and making playoff appearances significantly increases team revenues. Since team performance is fairly similar, it is not that difficult for most teams to become competitive. The high degree of competitive balance combined with a substantial financial payoff to winning teams contributes to the bidding up of players' salaries, in spite of the player reservation system. Further, since the higher paid players have higher *MRP*s, it makes sense for the owners to pay relatively more to the better players, thus increasing the dispersion in the salary distribution over time. An owner's assessment of being "only one player away from winning the Stanley Cup" may be an accurate reflection of market forces.

NOTES

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1. See Fort and Quirk [1995] for details and references to the literature.
2. The increases in baseball salaries were computed from data provided in Zimbalist [1994, 85].
3. Sources for the salary data are as follows: 1989/90, Rodney Fort; 1991/92, Toronto Star, November 1, 1991; 1993/94, *The Hockey News*, June 10, 1994; 1995/96, Richard Crowe. Salary includes base salary plus bonuses and deferred payments. Salaries reported in Canadian dollars were converted to U.S. dollars using the average exchange rate for September prior to the start of the season.
4. The mean logarithmic deviation is $\log(\Sigma x_i / n) - (\Sigma \log x_i)/n$. According to Bourguignon [1979], the mean logarithmic deviation is the only population-weighted decomposable measure of inequality which is continuous, differentiable, symmetric, homogenous of degree zero with respect to salaries, and satisfies the Pigou-Dalton condition.
5. Quirk and Fort [1992, 235-239]; Scully [1995, 76]; and Porter and Scully [1996, 160].
6. The sample does not extend beyond 1995/96 because Financial World did not conduct surveys after 1995/96.
7. Sommers and Quinton [1982] include an interaction term of *PCT* times population to capture the importance of market size on the marginal effect of team performance on revenues. This interaction

term in the hockey revenue equation was statistically significant at the 10 percent level, but the interaction term was not included in the final specification because the coefficient was not large enough to make a difference in the estimates of marginal revenue product. The impact of market size is therefore picked up in the team fixed-effects coefficients.

8. Fort and Quirk [1995] distinguish between their winning percent model and an alternative model in which team performance is based solely on championships won. Including both *PCT* and *POG* as regressors in the revenue equation can be viewed as capturing features of both models.
9. See McDonald and Moffitt [1980] for the interpretation of the coefficients in this model.
10. The coefficient of penalty minutes (*PMPGD*) could be either positive or negative. In the case of forwards, one might expect a positive coefficient in that more penalties would be a sign of aggressive and intense play. For defensemen, however, penalties are more likely to be an indication of poor defensive play.
11. Scully [1974; 1989] essentially assumes that the marginal product of the replacement player is zero. The procedure adopted here follows Zimbalist [1992] in assuming that the performance of a replacement player is equal to the average performance of replacement players at that position.
12. Scully [1995] suggests that the first two hypotheses might be appropriate for the market for professional athletes.
13. Gross marginal revenue product is defined as an adjusted *MRP* plus the average salary of the reserve players in that position. The *MRPs* are adjusted by multiplying the *MRP* by 84 and dividing by the number of games played. This has the effect of making the *ex post MRP* closer to the *ex ante MRP*. Players for which gross marginal product was negative were excluded from the analysis.
14. For comparison, note that the median salary for roster players in 1993/94 was $450,000.
15. An F-test of the hypothesis that the coefficients of the five team fixed-effects in the first equation are zero yielded a test statistic of 5.62 which is significant at a one percent level.
16. I am indebted to Robin Lock for suggesting this procedure. His analysis of data from college hockey games lends support to the Poisson approximation and the stochastic independence of goals scored by the two teams in a game.
17. Goals that were scored after the first 10 minutes of the first overtime period were not counted.
18. The chi-square goodness of fit test statistic was 4.65. The critical value for the test at a 10 percent level of significance was 10.645.
19. For seasons before the introduction of overtime games in 1983/84, the probability of a tie under equal playing strength was set equal to $.162 \times (.173/.117) = .240$.
20. Starting in 1996, the first four places in the draft order are determined by lottery. This is intended to reduce the likelihood that teams will reduce playing intensity at the end of the season to secure a higher draft pick.
21. The general procedure adopted here follows Grier and Tollison [1994] who used data from the National Football League.
22. The draft pick order was determined according to the team's order of finish; it is not the order in which the team actually picks; this might be different because of trades.
23. See Lavoie and Grenier [1992], Lavoie, Grenier, and Coulombe [1992], and McLean and Veall [1992] for references. An exception is Jones and Walsh [1987] who attempted to measure the extent of monopsonistic exploitation using data from the 1977-78 season when the World Hockey Association was operating.

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