

THE PRICING OF SPORTS EVENTS: DO TEAMS MAXIMIZE PROFIT?*

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A model of price setting behaviour by National Hockey League teams based on the assumption of profit maximization is developed, estimated, and tested. The model implies parameter restrictions across equations of a two-equation simultaneous nonlinear econometric model, tested by a likelihood ratio test, and implies restrictions on the first and second derivatives of the revenue function, tested with Wald tests. The results in large measure support the hypothesis that hockey teams are profit maximizers, in contrast to some suggestions in the literature. The analysis provides an attractive example of the potential of sports data for testing behavioural hypotheses in economics.

I bought the team out of love of the game and pride in the city
and not for profit....

You're kidding!

You guessed, eh.

Harv Antoine, *Apocryphal Northern Tales*

INTRODUCTION

THE LONGSTANDING debate over whether firms are profit maximizers has been given new life by recent evidence that both buyers and sellers are influenced by the perceived fairness of prices. Okun [1981], in particular, has argued that the threat of withdrawal of patronage can serve to punish firms who set prices in excess of those perceived by customers to be fair (warranted by costs). This enforces an implicit contract at prices below short run profit maximizing levels. Kahneman, Thaler and Knetsch [1986a, b], generalizing from the results of an extensive series of surveys, have gone further and argued that perceptions of fairness affect pricing on a much wider scale and do so even if the means of enforcement are not available. While this leads them to question the relevance and scope of profit maximization as a behavioural assumption, their case is far from conclusive.

Although buyers may express a dislike for profit maximizing prices and while suppliers may deny that they are motivated by profit maximization, it is not clear what this means for their actual behaviour. By their very nature, no number of surveys can resolve the issue. Consider the case of professional sports which is cited by both Okun and Kahneman *et al.* Despite their

* We would like to thank our colleague Serge Nadeau for his comments on an earlier draft of this paper.

protestations will sports fans really deny themselves attendance at a game of a favoured (winning) team for this reason alone? Will teams really forego opportunities for more revenue? There is sufficient evidence—largely impressionistic but highly suggestive nevertheless—to indicate that the answer to both questions is no. For example, ticket prices in most North American professional sports leagues vary widely and seem to have more to do with the nature of the local market than they do with costs. Nor is it evident that attendance responds to price increases in a way that differs fundamentally from other goods. It is also apparent that, with respect to the pricing of different categories of seats or the pricing of season tickets and other ticket packages, teams engage in rather sophisticated practices that seem to be geared to extract surplus from their fans. In short it is not at all clear that teams have been inhibited from pricing so as to maximize profits.

Our intention in this paper is to shed some light on the issue of pricing motivation by examining the ticket pricing behaviour of teams in one particular sports league, the National Hockey League (NHL). This is of interest from two points of view. First, it provides a vehicle for introducing and implementing a number of procedures for testing price setting behaviour using firm level data. Sports leagues—and the NHL is a prime example¹—provide one of the few instances of something close to an unregulated local monopoly for which data on price, sales, product attributes and market characteristics are available for individual agents (the teams). This allows us to extend the methods that have been developed for testing the optimizing behaviour of perfectly competitive agents and to overcome the aggregation problems that plague most of those studies (and which may partially explain their frequent rejection of the restrictions implied by optimizing behaviour).²

Second, professional team sports are industries in which there has been an active debate over the appropriateness of the profit maximization assumption³—not only by those concerned with general motivational issues, as discussed above, but also by those supporting and opposing changes in public policy toward the industry. Briefly, the argument is that if teams are not profit maximizers, and the suggested alternate motivations have ranged from 'satisficing' through 'love of the game' to 'civic pride', the application of standard economic policy models to sport is also inappropriate.⁴ Hence, current public policy which might be justified by standard analysis—regulating monopoly, promoting entry, the application of antitrust laws in

¹ For a discussion of the particular features of the NHL see Jones [1969] and Jones and Ferguson [1988].

² See Nicol [1989], for example, for a discussion of some of the aggregation and conditioning problems that arise in applied demand analysis.

³ See the survey in Cairns *et al.* [1986, 7–10].

⁴ See, for example, the discussion of alternate motivations in Markham and Teplitz [1981, 25–31] for baseball (their study was commissioned by Major League Baseball) and the references in Jones [1969, note 3] for hockey.

general—may be ill-founded. Alternatively, if profit maximization holds, then standard analysis and the conventional nostrums presumably apply.

The central issue in our analysis is the narrow one of whether NHL teams act as profit maximizers in setting seat prices. In section I we outline a simple short run model of ticket pricing and introduce our method for testing whether prices are profit maximizing. The test procedure is based on the observation that if firms do act as profit maximizers then (i) the resulting price behaviour is restricted by the nature of the firms' demand and cost conditions and (ii) the appropriate first and second derivative conditions must be satisfied. The former is examined by testing cross-equation restrictions on parameters and the latter by numerically evaluating and testing the derivatives. The empirical implementation and results are presented in sections II and III. We find that, even in the context of this simple model, we fail to reject the hypothesis of profit maximization.

I. THE MODEL AND THE TEST METHOD

Our model of teams' ticket pricing is a simple one. We suppose that at the opening of a season each team considers their fans' willingness to pay for attendance at a representative game. It then maximizes its profits for the season by setting a single (average) seat price to maximize its gate receipts from the representative game. In making this choice it is constrained by the capacity of its arena and hence it may choose to price at a point at which the arena sells out.

Implicit in this approach are the assumptions that:

- (i) costs which vary with attendance are small so that profit maximization coincides with revenue maximization,
- (ii) games are sufficiently homogeneous that the depiction in terms of a representative game does not do undue violence to a team's actual calculations,
- (iii) a single price is sufficient to describe a team's choice alternatives.

The first assumption does seem to be in accord with the facts—most costs of mounting a game are independent of how many people show up. However, the other two warrant more extended comment.

To some extent games do differ in their appeal to fans depending, for example, on the day of the week or the quality of the visiting team.⁵ Although such variations can be treated as deviations about an average game, the principal difficulty in doing so is the existence of the capacity constraint imposed by arena size. If average attendance below capacity masks a large

⁵ For a discussion of the significance of these factors for attendance in hockey see Jones [1984], and Jones and Ferguson [1988].

number of sellouts then our model would misrepresent the nature of the choice before such a team. Fortunately for us our sample is one in which teams can be clearly divided into those that sell out (or come quite close) for almost all games within a season and those that rarely (if ever) reach the capacity of their arena.⁶

With respect to the single price assumption, we can only express regret that despite repeated attempts we were unable to obtain data on pricing and sales by type of ticket.⁷ As indicated in the introduction, it is our conjecture that teams use ticket bundling and the structure of seat prices to extract surplus from their fans and that they do so quite effectively. This limited our ability to deal with all of the relevant aspects of their behaviour and forced us to treat tickets as a composite good.

Given these assumptions we can now proceed with a more formal description of the model. The demand for attendance at the representative game of a team is described by an (inverse) demand function

$$(1) \quad p = f(A, \mathbf{z}; \theta)$$

where p is the team's average ticket price, A is average attendance per home game within the season, $\mathbf{z}$ is a vector of attributes of the team and its home city and θ is a vector of parameters. The team's ticket revenue is

$$(2) \quad R(A, \mathbf{z}; \theta) = Af(A, \mathbf{z}; \theta)$$

and, as described above, we suppose that they choose a level of attendance A and corresponding price so as to

$$(3) \quad \max_A R(A, \mathbf{z}; \theta) \quad \text{s.t. } A \leq C$$

where C is their arena capacity. The associated Kuhn-Tucker condition

$$(4) \quad \frac{\partial R(A, \mathbf{z}; \theta)}{\partial A} \geq 0, \quad \frac{\partial R(A, \mathbf{z}; \theta)}{\partial A} (C - A) = 0, \quad C - A \geq 0$$

then characterizes the team's behaviour. In particular, if the revenue function is concave in A , then (4) is sufficient for A (and $p = f(A, \mathbf{z}; \theta)$) to be profit maximizing. More importantly, local concavity is necessary if the capacity constraint is not binding.

For the purpose of testing the hypothesis of profit maximization one critical point is that the parameters of the inverse demand function recur in

⁶ See Table II for identification of the teams which sold out in each of the seasons. In general, this appears to reflect the dominance of fans' desire for a winning home team and in the Canadian case (note in Table II that only two of the Canadian teams failed to sell out) seems to reflect the significance of hockey in Canadian culture. For supporting evidence for an earlier period, see Jones [1984] and Jones and Ferguson [1988].

⁷ The only price we could obtain was average price data for the three seasons 1981-1983 (see footnote 9).

(4). More specifically, the choice of a functional form for (1) also determines the form and parameterization of (4). The assumed behaviour imposes restrictions and a test of these restrictions can then be used to test the behavioural hypothesis.

Such tests are performed on restrictions across equations and consequently the first inequality in (4) cannot be represented directly. Our solution was to use

$$(5) \quad \frac{\partial R(A, z; \theta)}{\partial A} (C - A) = 0$$

as the equation representing the teams' choices. The estimation of a system consisting of (1) and (5) can then serve as the basis for a likelihood ratio test of the restrictions, and in addition numerical evaluation and testing of $\partial R / \partial A$ and of $\partial^2 R / \partial A^2$ can be used to determine if the conditions (4) are satisfied and whether the revenue function is concave in A for those teams that do not sell out.

Although we are concerned with a very specific instance here, this method is clearly linked to the broader literature on tests of behavioural hypotheses. The procedure is similar in form to the familiar parametric tests of the (cost minimizing or profit maximizing) behaviour of perfectly competitive firms, in so far as it uses tests of cross-equation restrictions and numerical checks on derivatives to determine if the implications of the assumed behaviour are satisfied.

II. EMPIRICAL IMPLEMENTATION

Estimation requires, of course, that the vector of home city and team attributes z be specified, as well as a functional form for the inverse demand function $f(A, z; \theta)$. The following list of attributes was used.⁸

- z_1 Population of the team's home city
- z_2 Per capita income of the team's home city
- z_3 Dummy = 1 if the home city is in Canada
- z_4 Number of 'superstars' on the team

⁸ These attributes and variations on them have been used in virtually all demand studies of professional team sport using time series and/or cross section data. For specific examples see Noll [1974], Jones [1984], Jones and Ferguson [1988], and for overviews see Cairns *et al.* [1986, 12–27] and Schofield [1983]. The attribute z_1 is from the *Statistical Abstract of the United States*, 1985 for the US teams; and interpolated between the *Census of Canada* 1981 and 1986 for the Canadian teams. The attribute z_2 is from the US Department of Commerce, *Survey of Current Business*, April 1985 for US teams; and extrapolated from Statistics Canada, *Income Estimates for Sub-Provincial Areas*, 1983 for Canadian teams. The observations for z_4 were determined subjectively, by the authors. The attribute z_5 was computed from the daily standings of the teams during each of the three seasons as published in the *Victoria Daily Times*; and z_6 was obtained from the *NHL Guide*, 1981–1983.

- z_5 The team's average rank in the League over the current season
 z_6 The team's rank in the League at the end of the previous season

Data were collected for the 1981, 1982, and 1983 hockey seasons, the only ones for which data on ticket prices are available.⁹ Since the NHL consisted of 21 teams at this time, our data set therefore contains 63 observations describing team behaviour.

As our theory does not suggest a functional form for the inverse demand function it seems desirable to choose as general a form as possible so as not to unduly restrict the demand relationship. For the most part we experimented with empirical models based on inverse demand functions of the form

$$(6) \quad f(A, \mathbf{z}; \boldsymbol{\alpha}, \boldsymbol{\beta}, \gamma) = a(\mathbf{z}; \boldsymbol{\alpha}) + b(\mathbf{z}; \boldsymbol{\beta}) A^\gamma$$

with a number of particular forms chosen for $a(\mathbf{z}; \boldsymbol{\alpha})$ and $b(\mathbf{z}; \boldsymbol{\beta})$. In addition, we experimented with alternate ways of representing the first derivative

$$(7) \quad \frac{\partial R}{\partial A} = a(\mathbf{z}; \boldsymbol{\alpha}) + (\gamma + 1)b(\mathbf{z}; \boldsymbol{\beta}) A^\gamma$$

Two issues governed our ultimate choice of specification—the computational problems that frequently arise in obtaining maximum likelihood estimators when using complicated nonlinear forms and the need to normalize (7) when substituted in (5). In estimation we considered linear, log-linear, and CES forms for $a(\mathbf{z}; \boldsymbol{\alpha})$ and $b(\mathbf{z}; \boldsymbol{\beta})$. It was found that nonlinear forms invariably led to intractable convergence problems in the application of nonlinear estimation procedures, and so in the end we settled on linear specifications. Normalization of the parameters in (7) was treated by substituting from (6) and representing (7) as

$$(8) \quad \frac{\partial R}{\partial A} = p + \gamma b(\mathbf{z}; \boldsymbol{\beta}) A^\gamma$$

The empirical model then consisted of¹⁰

$$(9) \quad p_i - \boldsymbol{\alpha} \cdot \mathbf{z}_i - (\boldsymbol{\beta} \cdot \mathbf{z}_i) A_i^\gamma = \varepsilon_{1i}$$

$$(10) \quad (p_i + \gamma(\boldsymbol{\beta} \cdot \mathbf{z}_i) A_i^\gamma)(C_i - A_i) = \varepsilon_{2i}$$

where i refers to an observation for a particular team in a particular season.

This model represents a system of two stochastic simultaneous nonlinear equations. Although the equations have been derived in a deterministic setting, for all the usual reasons we do not expect that the data should fit them

⁹ The ticket price data was obtained from the *Toronto Globe and Mail* (February 25, 1985, p. 20) and from information supplied by Mr Al Strachan of the *Toronto Globe and Mail*. The data was placed in real exchange rate adjusted terms.

¹⁰ In the following, the first entry in each $\mathbf{z}_i$ is the number one so that the first terms in $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are intercepts.

exactly. This consideration has been treated in the usual way, by the introduction of additive random disturbances which we assume to be normally distributed. As the equations are derived from a common underlying behavioural model the disturbances are in part generated from common sources and so should be treated as being contemporaneously correlated. Accordingly the appropriate estimation methodology is nonlinear full information maximum likelihood, specifying price p and attendance A as endogenous and capacity C and the z variables as exogenous.¹¹ This is a standard estimation procedure available in a number of econometrics packages: our own implementation made use of the FIML routine of mainframe TSP.

III. THE RESULTS

The model was estimated both in the restricted form above, in which the cross-equation restrictions on γ and β are imposed, and in the unrestricted form consisting of (9) and

$$(11) \quad (p_i + \lambda(\delta \cdot z_i)A_i^{\lambda})(C_i - A_i) = \varepsilon'_{2i}$$

Coefficient estimates for the restricted and unrestricted models are presented in Table I. Using the estimates from the restricted model the numerical values of $\partial R/\partial A$ and $\partial^2 R/\partial A^2$ were computed for each of the 63 observations and these are given in Table II. A discussion of the results of each table follows.

In Table I, although the coefficient standard errors are reported for completeness we regard them as being of comparatively little interest because the testable predictions of our theory are in terms of the cross-equation restrictions and the first and second order conditions rather than in terms of the coefficients individually. That is, we do not regard the individual coefficients of the model as having strong intuition associated with them with respect to magnitude or sign in the way that is common in many other econometric applications. Hence inferences with respect to the coefficients individually will not play a role in our statistical analysis of the model.

It is, however, notable that in both the restricted and unrestricted models many coefficients are statistically insignificant. In the context of the restricted model this is likely due in large measure to our specification (6) which allows both the intercept and slope of the inverse demand function to shift with our list of z variables. Given this specification it is hardly surprising that

¹¹ Although (10) is an implicit equation and cannot be normalized on any single variable, this does not represent any difficulty in estimation. The data on attendance was obtained from the individual game attendance figures published in the *Victoria Daily Times* over the three seasons. The stated capacity, as published in the *NHL Guide* for the respective seasons, was used for C in most instances. In some cases the average attendance exceeded the stated capacity and in those cases the observed average attendance was used for C . This was also done for those teams that attained or exceeded capacity for almost all games, but whose average attendance was slightly less than the stated capacity.

TABLE I
PARAMETER ESTIMATES

<i>Restricted Model (Log of Likelihood Function = -952.95)</i>					
α_0	-829.8283 (834.337)	β_0	1.1953 (5.4349)	γ	0.6893 (0.4311)
α_1	-0.1689 (0.08008)	β_1	0.2639e-3 (0.1152e-2)		
α_2	0.2231 (0.10023)	β_2	-0.1818e-3 (0.8674e-3)		
α_3	-213.0697 (261.168)	β_3	0.6129 (2.7321)		
α_4	31.8027 (119.699)	β_4	-0.4663e-1 (0.2146)		
α_5	36.8637 (22.2257)	β_5	-0.7746e-1 (0.3356)		
α_6	4.2888 (14.6054)	β_6	-0.1269e-1 (0.5641e-1)		

<i>Unrestricted Model (Log of Likelihood Function = -946.29)</i>							
α_0	-782.6429 (0.5038e-1)	β_0	0.8229 (41.875)	δ_0	0.6206 (35.031)	γ	0.7076 (385.11)
α_1	-0.2017 (100.06)	β_1	0.2333e-3 (92.125.0)	δ_1	0.1438e-3 (50.836.1)	λ	0.7328 (553.19)
α_2	0.2033 (699.99)	β_2	-0.1297e-3 (577.630.0)	δ_2	-0.9855e-4 (462.316.7)		
α_3	876.1851 (0.2953e-1)	β_3	-0.8490 (25.572)	δ_3	0.3942 (10.078)		
α_4	-69.3134 (0.9074e-1)	β_4	0.1228 (76.291)	δ_4	-0.4663e-1 (21.911)		
α_5	-7.4069 (0.6848)	β_5	0.2667e-2 (526.32)	δ_5	-0.4904e-1 (587.91)		
α_6	-6.3058 (0.7668)	β_6	0.5420e-2 (614.49)	δ_6	-0.7452e-2 (609.43)		

The standard errors are reported in parentheses.

the role of the variables is not well determined, as is reflected in some relatively large standard errors. As our interest is not with respect to the coefficients individually, however, this does not represent a difficulty; of much more importance and value is that the inverse demand function has been relatively generously parameterized so that the problem of implicit restrictions being placed on the demand relationship by an unduly restrictive functional form is mitigated.

An additional possible explanation for some large standard errors is that the inclusion of some variables may be unnecessary. As emphasized earlier, however, our variable set is well established in the literature. It is an elementary result in econometric theory that the inclusion of irrelevant explanatory variables is a far less serious specification error than the

TABLE II
THE DERIVATIVES OF THE REVENUE FUNCTIONS

	1981-1982		1982-1983		1983-1984	
	$\partial R/\partial A$	$\partial^2 R/\partial A^2$	$\partial R/\partial A$	$\partial^2 R/\partial A^2$	$\partial R/\partial A$	$\partial^2 R/\partial A^2$
Calgary	1267.8	-0.150	1331.9	-0.164	138.3	-0.128
Flames*	(0.000)	(0.000)	(0.000)	(0.000)	(0.424)	(0.000)
Edmonton	722.0	-0.056	566.6	-0.073	845.9	-0.067
Oilers*	(0.047)	(0.027)	(0.063)	(0.005)	(0.033)	(0.018)
Montreal	1302.0	0.008	1257.7	-0.003	716.7	-0.044
Canadiens*	(0.000)	(0.848)	(0.000)	(0.929)	(0.009)	(0.005)
New York	1504.3	0.030	1555.5	0.027	1339.2	-0.001
Islanders*	(0.000)	(0.633)	(0.000)	(0.682)	(0.000)	(0.999)
New York	849.9	-0.012	932.2	-0.015	1434.0	-0.008
Rangers*	(0.058)	(0.762)	(0.055)	(0.727)	(0.044)	(0.902)
Quebec	624.0	-0.041	-88.4	-0.095	640.5	-0.055
Nordiques*	(0.029)	(0.131)	(0.553)	(0.007)	(0.005)	(0.000)
Toronto Maple	356.1	-0.068	167.8	-0.090	456.3	-0.082
Leafs*	(0.068)	(0.000)	(0.350)	(0.000)	(0.040)	(0.000)
Chicago Black	415.8	-0.084	597.8	-0.046	101.1	-0.085
Hawks**	(0.000)	(0.000)	(0.073)	(0.017)	(0.437)	(0.000)
Philadelphia	245.1	-0.058	418.6	-0.055	340.2	-0.068
Flyers**	(0.168)	(0.000)	(0.087)	(0.000)	(0.108)	(0.000)
Boston Bruins	536.9	-0.083	540.1	-0.086	540.3	-0.096
	(0.000)	(0.000)	(0.010)	(0.000)	(0.020)	(0.000)
Buffalo Sabres	238.2	-0.066	164.8	-0.097	292.5	-0.079
	(0.167)	(0.000)	(0.990)	(0.000)	(0.052)	(0.000)
Detroit Red	-155.1	-0.131	-286.2	-0.141	-615.5	-0.125
Wings	(0.038)	(0.000)	(0.000)	(0.000)	(0.000)	(0.010)
Hartford	-181.6	-0.169	-46.2	-0.197	-63.7	-0.185
Whalers	(0.000)	(0.000)	(0.596)	(0.000)	(0.420)	(0.000)
Los Angeles	443.2	-0.102	432.2	-0.097	357.0	-0.131
Kings	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Minnesota	145.7	-0.088	385.6	-0.085	76.3	-0.120
North Stars	(0.251)	(0.000)	(0.043)	(0.000)	(0.571)	(0.000)
Colorado	400.8	-0.202	-191.9	-0.168	-117.7	-0.202
Rockies	(0.000)	(0.000)	(0.008)	(0.000)	(0.166)	(0.000)
Pittsburgh	182.7	-0.115	379.3	-0.168	550.6	-0.198
Penguins	(0.061)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
St. Louis Blues	-41.4	-0.100	-76.5	-0.133	-32.0	-0.132
	(0.771)	(0.000)	(0.209)	(0.000)	(0.678)	(0.000)
Vancouver	602.5	-0.086	426.7	-0.098	476.5	-0.109
Canucks	(0.000)	(0.000)	(0.015)	(0.000)	(0.010)	(0.000)
Washington	-78.4	-0.184	198.8	-0.150	362.5	-0.163
Capitals	(0.251)	(0.000)	(0.142)	(0.000)	(0.007)	(0.000)
Winnipeg Jets	380.3	-0.082	539.2	-0.087	425.9	-0.118
	(0.029)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)

* Sold out in all three seasons.

** Chicago sold out in 1982-1983 and Philadelphia sold out in all but 1983-1984.

Values in parentheses are the marginal significance levels of a Wald χ^2 test of the hypothesis that the derivative is equal to zero.

exclusion of relevant variables. Hence in our model we prefer to err on the side of the less serious specification error by using the full set of variables that are established in the literature, rather than to omit variables which other

studies have found to be important and risk a more serious specification error.

For the unrestricted model standard errors are large in part simply because, once the restrictions are relaxed, we are estimating 23 parameters on the basis of 63 observations. It is therefore not at all surprising that the estimated coefficients are not well determined, and this is reflected in the large standard errors/low *t*-ratios. Again, this poses no difficulty for our analysis because nothing of importance rests on the coefficient estimates of the unrestricted model; only the loglikelihood function value for the unrestricted model is of interest, in connection with the likelihood ratio test of the cross-equation restrictions. Before turning to this test, however, let us consider our results on the first and second order conditions.

Table II offers valuable insights into the validity of our model because the model implies the following restrictions on the first and second order conditions for the individual teams.

- (i) First order conditions:
 - (a) $\partial R/\partial A = 0$ for non-sellout teams.
 - (b) $\partial R/\partial A \geq 0$ for sellout teams.
- (ii) Second order necessary conditions:
 - (a) $\partial^2 R/\partial A^2 \leq 0$ for non-sellout teams.
 - (b) $\partial^2 R/\partial A^2$ unrestricted for sellout teams.

Table II reports the numerical values of the first and second derivatives of the revenue function of each team in each season and, in parentheses, the marginal significance level of a Wald test for the hypothesis that each derivative is equal to zero.

The implications of the model for the *second order conditions* are very strongly supported by our estimation results. Numerical evaluation reveals that for *all* non-sellout teams the second derivative is negative. Furthermore the Wald test indicates that this negativity is statistically significant (at conventional significance levels) *in every case* for the non-sellout games. This regularity is so strong that it might lead one to wonder whether it doesn't represent some bias in the test statistic rather than a confirmation of the model; but this suspicion is dispelled by the fact that for sellout teams the second derivative is sometimes positive, and even when negative it is sometimes not significantly different from zero—often with a very large marginal significance level. Hence the second order condition results are entirely in accordance with the predictions of the model.

Results associated with the *first order condition* are somewhat more mixed. For non-sellout teams, where marginal revenue should not differ significantly from zero, there is in fact a fair incidence of rejection of this hypothesis, although in 16 of the 39 non-sellout cases marginal revenue is not significantly different from zero (at a 5% significance level). This leads us to some comments below on the general issue of the reliability of Wald tests. It is

however notable that for sellout teams marginal revenue should be non-negative and this is violated in only a single instance (Quebec in 1982–83 and even here the violation is not statistically significant), whereas negative values for marginal revenue are more common for the non-sellout teams. This is supportive of the theory. Another encouraging aspect of the results is that rejection of the hypothesis that marginal revenue equals zero tends to be more common for sellout teams than for non-sellouts (although this result depends somewhat on the significance level chosen), as the model would lead one to expect.

Our inferences with respect to the first and second order conditions have been based on Wald test statistics. Wald tests are the natural ones to apply in this context because they are computed on the basis of the estimated model where the restrictions to be tested have not been imposed. Given that the first order condition is often rejected for non-sellout teams when the theory implies that it should not be, it may be useful to note that Wald test results must be qualified by a number of considerations generic to Wald testing.

- (a) The three standard asymptotic test criteria are the Lagrange Multiplier (LM), Likelihood Ratio (LR), and Wald (W) test statistics. For linear restrictions these are related by the well known Berndt–Savin [1977] inequality.

$$LM < LR < W.$$

Although the first and second order conditions for our model represent nonlinear restrictions, the Berndt–Savin inequality nevertheless suggests strongly that Wald tests are the most likely to yield rejections of hypotheses. This feature of the Wald statistic may be reflected in the results of Table II; in particular, it is possible that the frequent rejections of the first order condition for the non-sellout teams is a statistical artifact rather than a deficiency of the model. This problem is most likely to be apparent, of course, in analyses such as ours which are based on comparatively limited samples, so that the small-sample behaviour of the statistic is not well approximated by its asymptotic distribution.

- (b) Monte Carlo studies such as those of Gallant [1976, 1977] suggest that the finite sample distribution of the Wald statistic is less closely approximated by its asymptotic distribution than is the case for the Likelihood Ratio or Lagrange Multiplier statistics. Hence Wald test results are relatively unreliable.
- (c) The Wald test statistic is not invariant to reparameterization of the model. "With the same data and an equivalent model and hypotheses, two investigators could obtain different values of the test statistic." (Burguete, Gallant, and Souza [1982, p. 185]) This element of arbitrariness is not present in the other two test statistics.

Given these deficiencies of Wald testing, it is desirable that not all our

inferences with respect to our model make use of this procedure. Fortunately the other restrictions to be tested, the cross-equation restrictions, are most naturally tested with a likelihood ratio test, and it is to this that we now turn.

Returning to Table I, a likelihood ratio test of the cross-equation restrictions is based on a comparison of the loglikelihood function values associated with the restricted and unrestricted models. The computed value of the likelihood ratio test statistic is $2(-946.92 - (-952.95)) = 13.32$, which for the eight cross-equation restrictions being tested does not represent a rejection at conventional significance levels. (For example $\chi^2_{0.05} = 15.507$.)

In summary, to a very considerable extent our empirical results are consistent with the economic behaviour hypothesized by the model. This is especially noteworthy because of the simplicity of the model and the frequency with which similar tests of optimizing behaviour have failed. Our results are in striking contrast to the rejections of the restrictions implied by micro-theoretic models of agent behaviour which so often arise in the application of asymptotic test procedures in other areas of applied econometrics. For example, as documented by Deaton [1986], the homogeneity and symmetry restrictions of consumer theory are routinely rejected in applied demand analysis. It is also common to find that second derivative conditions, such as the negative semidefiniteness of the Slutsky matrix, are not satisfied. The comparative success of our results may be due, in part, to our use of data for individual agents rather than the aggregated data commonly employed in other studies.

IV. CONCLUSION

The above analysis is of interest both methodologically and from the point of view of substantive policy implications. Our econometric results, although based on a fairly simple model which abstracts from some of the complexities of actual team behaviour, offer considerable support for the profit maximizing behaviour hypothesized. These results have made use of both Wald and likelihood ratio tests and have involved the testing of cross-equation restrictions and hypotheses concerning the first and second order conditions for profit maximization, as well as the numerical evaluation of these first and second order conditions. Our results are particularly encouraging given the negative results which so often arise from analogous tests of agent behaviour in other areas of the application of econometrics to microeconomics, such as production and demand analysis. We have conjectured that one explanation for the comparative success of our results may be the availability of data on individual agents (the teams)—a level of disaggregation which is often not available in the industry or commodity data sets normally used to investigate hypotheses concerning microeconomic agent behaviour. Indeed it is not uncommon to find aggregation issues cited

as a likely cause of the failure of such empirical work, and hence the increasing use of micro data by researchers.

It is evident from our results that ticket pricing by NHL teams may well be consistent with profit maximization so that, in contrast to the suggestions of Okun, Kahneman *et al.*, and others, the possibility that the behaviour of sports teams may be motivated by much the same considerations as are true for economic agents generally should not be dismissed. It follows that there may be no basis for an exception to the standard policy prescriptions.

Of at least as much interest as its implications for this latter issue, however, is that our analysis provides an attractive example of the potential of sports data for the implementation of tests of agent behaviour. In short, it is our view that sports data represents a potentially rich but previously unrecognized and little-tapped source of micro-level industry/firm/product data suitable for testing a wide range of hypotheses—a source which may compare favourably in cost and quality with other types of disaggregated data.

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