HOSPITAL MARKET STRUCTURE AND COST PERFORMANCE: A CASE STUDY

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INTRODUCTION

Neoclassical microeconomics assumes that firm decisions are driven by a profit motive. Market structure plays no role in determining the cost minimizing behavior of firms. Beginning with Berle and Means [1933], however, researchers have seriously questioned whether profit maximization is the sole objective of firms. Instead, contemporary managerial theory posits that unconstrained managers strive to maximize their personal utility at the expense of higher profits. In particular, managers may attempt to enhance their personal utility by expanding the firm beyond the profit-maximizing point [Baunol, 1959] or through discretionary expenditures on items such as executive pay, office furnishings and staff [Williamson, 1965]. Consequently, the managerial theory predicts that market structure does indeed influence cost behavior, since greater competition forces otherwise unconstrained managers to operate more efficiently.

Interestingly, existing studies of the hospital services industry find that increased competition among hospitals generates higher rather than lower costs. This directly conflicts with the neoclassical and managerial theories of firm behavior. A common explanation for this seemingly unusual outcome is that higher costs stem from a more competitive market because hospitals are forced to compete on a non-price basis for scarce physicians, resulting in a "medical arms race" [Robinson and Luft, 1985]. Specifically, a higher level of hospital quality, in the form of advanced medical technology and amenities, is offered as an in-kind payment to physicians in return for admitting their patients. Physicians are free to choose hospitals with the highest in-kind payment since their patients are typically ignorant about the true benefits of medical treatment. Moreover, patients incur relatively low costs from these decisions because of third-party payments. As a result, hospitals are forced to make greater in-kind payments or lose physicians and their patients to competitors.

While empirical studies linking competition to higher hospital costs have illuminated this issue, they are lacking in two important respects. First, these analyses use data for periods prior to the mid-1980s, when third-party payers were less sensitive to hospital prices. After the mid-1980s, prospective payment systems partly replaced cost-based systems and third-party pre-authorization became more common. Thus, hospitals


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may now be reluctant to compete on a non-price basis, since they cannot as easily pass on the increased costs.

Second, existing studies either use a sample of exclusively nonprofit hospitals, or they implicitly assume that all types of hospitals (i.e., for-profit, nonprofit and public) are similarly influenced by market structure. Clearly, it is possible that for-profit, nonprofit and government hospitals respond differently to competitive pressures, especially during more recent periods characterized by an increased sensitivity to hospital prices. For example, a nonprofit hospital may react to more competition by increasing quality rather than lowering prices as a direct consequence of its ownership structure. In particular, the lack of well-defined property rights to residual profits (Stiglitz 1988) might give nonprofit administrators little incentive to engage in price competition. Instead, administrators may be motivated by the prestige associated with super-optimal quality levels. If so, administrators in nonprofit institutions may find quality competition to be more desirable than price competition at the margin.

While publicly-owned hospitals similarly lack a property rights structure, they differ in that they are directly controlled by higher level government, a very influential third-party payer. Clearly, in a price-sensitive environment, higher level government can be expected to more closely monitor the decisions of public hospitals and to carefully scrutinize budgetary requests as a way of containing costs. As a result, market structure may have little or no impact on the cost behavior of public hospitals.

For different reasons, proprietary hospitals may also not compete on the basis of quality during a price-sensitive period. In this environment, the profit incentive would directly control hospital behavior. If for-profit hospitals are unable to cover the costs of enhanced quality, the associated reduction in profits would not be tolerated by residual claimants. Lower profits would also hinder the hospital’s ability to attract financial capital. Hence, the inherent monitoring by investors suggests that performance is not likely to be substantially influenced by market structure, particularly in a period when higher costs cannot be easily passed on to third-party payers.

This analysis re-examines the impact of market structure on hospital cost behavior during periods of increased price-sensitivity, controlling for differences in hospital ownership. The results of this more structured and updated analysis indicate that while the costs of the average nonprofit hospital are directly related to increased competition, the costs of the average for-profit and public hospital are unaffected by market structure.

**EMPIRICAL SPECIFICATION**

To test the effects of market structure on hospital costs, a 1987/88 sample of short-term general hospitals in Texas is used in the empirical analysis. Hospitals within one state are selected to control for the influence of state government tax and regulatory policies. Texas is chosen for three reasons. First, the vast land area in this state provides the requisite large number of local markets for cross-sectional estimation. Second, there is a well-balanced mix of nonprofit (n = 132), for-profit (n = 251) and government (n = 156) short-term general hospitals in Texas. Finally, Texas hospitals are not subject to state rate regulations, potentially giving the representative hospital some control over both price and quality.

It is assumed that each hospital’s objective is to minimize the short-run total variable cost (VC) of producing a particular quantity (QUANT), and quality (QUAL), of hospital services, given fixed inputs of capital (K) and doctors with admitting privileges (DOC). To minimize VC, each hospital chooses the lowest-cost combination of labor and non-labor inputs. Assuming that the hospital is a price-taker in both resource markets, a generalized short-run variable cost function can be written as

\[
VC = V(\text{QUANT, QUAL, WAGE, } P_{\text{lab}}, \text{ K, DOC})
\]

where WAGE and \(P_{\text{lab}}\) represent the price of the labor and nonlabor inputs, respectively.

The quality of hospital services is specified as a function of market structure (MKT), and buyer income (INC). As mentioned above, the theory of nonprice competition predicts that the quality of the hospital services is influenced by market structure. With more intense competition, hospitals have an incentive to offer higher quality, in the form of advanced diagnostic and surgical equipment, and a greater number of operating rooms and patient amenities, as a means of attracting doctors and their patients during a period of price-insensitivity. The inclusion of income follows traditional demand theory, which predicts that the income of actual and potential buyers influences product quality. Substituting these two variables in lieu of hospital quality yields:

\[
VC = V(\text{QUANT, MKT, INC, WAGE, } P_{\text{lab}}, \text{ K, DOC}).
\]

Hospital cost equations have typically been estimated as Cobb-Douglas or translog cost functions. Given that the translog cost function is very expensive in parameters and our for-profit subsample is limited to fifty-five observations (see below), we choose to estimate equation 2 as a Cobb-Douglas function. As one purpose of this paper is to examine whether market structure influences nonprofit, for-profit and government hospital behavior differently, we prefer not to pool these subsamples of hospitals.

For estimation purposes, the total number of inpatient days serves as an output measure of the quantity of hospital services, QUANT. A number of case-mix variables is also specified, given the heterogeneous nature of these services. In addition, an organizational variable (SYST) is specified. Using a logarithmic transformation of the Cobb-Douglas cost function, the exact form of the estimated cost equation is:

\[
\ln VC = b_0 + b_1 \ln \text{PDAYS} + b_2 \ln \text{ACCR} + b_3 \ln \text{SER} + b_4 \ln \text{BRTHS} + b_5 \ln \text{STAY} + b_6 \ln \text{URBAN} + b_7 \ln \text{MKT} + b_8 \ln \text{INC} + b_9 \ln \text{WAGE} + b_{10} \text{SYST} + b_{11} \ln \text{K} + b_{12} \ln \text{DOC}
\]

where:

- \(VC\) = total annual hospital expenses (measure of short-run total variable cost)
- PDAYS = total inpatient days measured by multiplying average daily inpatient census, excluding newborn, by 365 days (output measure of the quantity of hospital services)
- ACCRED = total number of accreditations (proxy for case-mix)
- SER = number of services as indicated by the total number of facilities (proxy for case-mix)
- BRTHS = number of births (proxy for case-mix)
- STAY = average length of stay, computed by dividing total inpatient days by admissions (proxy for case-mix)
same product and geographical markets. The county definition is particularly appropriate
for this study since the county hospital in Texas appears to be the hospital care
provider of last resort.4

The measure of market structure, MKT, is proxied by a Herfindahl-Hirschman
index of market concentration, based on the number of staffed beds of all short-term
general hospitals in the same county.4 According to the nonprice theory of hospital
competition, the sign of the parameter estimate on ln MKT is expected to be negative.
That is, higher values for MKT indicate a more monopolistic market structure and a
smaller number of competing hospital demands for physicians. Thus, each hospital can
attract physicians without offering as much quality as an inducement. Since for-profit
and government hospitals are subject to more direct monitoring in a period of cost
consciousness, market structure is likely to have a greater impact on the costs of
nonprofit hospitals. Thus, for this particular sample, the coefficient on ln MKT should be
larger.

Buyer income, INC, is represented by per capita income in the county. Assuming
that the quality of hospital services is a normal good, a positive parameter estimate is
anticipated for ln INC. Neoclassical theory predicts that a higher wage leads to in-
creased costs, ceteris paribus, so a positive coefficient estimate is expected on ln WAGE.
Since nonlabour inputs are typically purchased in statewide, if not national, markets, and
their prices are relatively constant across hospitals, the price of the nonlabour input (PNN)
is not specified in the regression equation.

The organizational dummy variable, SYSTEM, controls for possible multi-hospital
systems and takes on the value of one if the hospital belongs to a health care system.
A hospital within a multi-hospital chain, signified by SYSTEM = 1, is assumed to face
lower costs because of purchasing discounts on inputs, marketing and managerial
economies and/or system diversification [Arnould and Debrec, 1986]. Thus, the coeffi-
cient estimate on this variable is expected to be negative.

The parameter estimates on ln K and in DOC can be used to determine if hospitals
are employing the optimal long-run capital stock and number of admitting physicians.
Following Cowing and Holtenmann [1983], and using equation 1, a long-run total cost
function can be specified as:

(4) \[ TC = VC^{QUANT}, QUAL, WAGE, P_{\text{PR}}, K, DOC \] + P_{\text{PR}} K + P_{\text{DOC}}

where P_{\text{PR}} and P_{\text{DOC}} are the shadow prices of capital and physician services, respectively.
A necessary condition for long-run cost minimization is that \( \Delta VC^K \) equals \(-P_{\text{PR}}\). This
equality can be derived by taking the first derivative of equation 4 with respect to K and
setting the resulting expression equal to zero. The implication is that the variable cost
savings from substituting one more unit of capital must equal the shadow price of capital
in long-run equilibrium. A nonnegative estimate for \( \Delta VC^K \) is a sufficient condition
for hospitals to be overemploying capital. In the context of the logarithmic function,
overemployment of capital implies that the estimated coefficient on ln K is greater than
or equal to zero. Similarly, the first derivative of the short-run variable cost function
with respect to the number of doctors should equal \(-P_{\text{PR}}\), if the number of admitting
physicians is optimal. Therefore, a nonnegative value for this derivative suggests that
the number of admitting physicians is excessive.

The fixed capital input is measured by the number of hospital beds since data for the
market value of hospital buildings are unavailable. Data for the number of admit-
ting physicians at each hospital are also unavailable, so they are derived by using county figures and the procedure outlined in Appendix A. The basic source of our data is the American Hospital Association’s 1988 Annual Survey of Hospitals. The data in this survey correspond to fiscal year 1987, i.e., year ending Sept. 30, 1988. The data for population, income and number of doctors in the county are for 1985 and are obtained from the 1988 County and City Databook.

**EMPIRICAL RESULTS**

After eliminating the observations with missing data, 278 observations are available. Of these observations, there are 66, 96 and 125 for-profit, nonprofit and government hospitals, respectively. Hospital profiles are depicted in Appendix B. The ordinary least squares technique is used to estimate separate cost equations for each hospital ownership type. The F-statistic for the difference between pooled and unpoled regression equations is F[26,397] = 1.92. Given the five and one percent critical values of approximately 1.57 and 1.88, we reject the null hypothesis of structural homogeneity of parameters across equations. White’s (1960) test finds evidence of heteroscedasticity, so his procedure is used to derive heteroscedastic consistent estimates of the standard errors.

The results of the estimation are given in Table 1. The estimated parameters and corresponding t-statistics (in absolute value) are displayed opposite each explanatory variable in the regression equation. Of particular importance to this research are the results associated with the market structure variable. The underlying hypothesis of this analysis is that for-profit and government hospitals will utilize less cost-enhancing forms of competition than nonprofit hospitals in a time of increased price-sensitivity. Indeed, the empirical results support this hypothesis. The results show that the estimated coefficient on ln MKT is not statistically different from zero in the for-profit and government hospital equations. This provides important evidence that the costs of these two types of hospitals are unaffected by the degree of competition in the market. Conversely, the findings indicate that market structure does affect the behavior of the average nonprofit hospital. Specifically, the statistically significant and negative sign suggests that the presence of more competition, i.e., a lower value for MKT, increases the short-run variable costs, presumably as quality is increased to attract physicians and their patients.

The magnitude of this parameter estimate indicates that a 10 percent increase in market competition leads to a 9 percent increase in the short-run variable costs of production, *ceteris paribus*. While relatively small in percentage terms, the estimated nonprofit cost equation can be fitted with the mean of each independent variable and alternative values for MKT to illustrate that an absolute change in the number of competitors has a sizeable impact on nonprofit hospital costs. For example, suppose the average nonprofit hospital holds monopoly status within a market area (MKT = 1). Calculations indicate that the average variable cost per patient day would be $717 for the representative nonprofit hospital. In contrast, if two hospitals of a similar size operate within a market area (MKT = .5), the average variable cost per patient day is approximately $762 for the typical nonprofit hospital. Finally, if ten hospitals of equivalent size are assumed (MKT = .1), the average variable cost equals $838 per patient day at the typical nonprofit hospital.

<table>
<thead>
<tr>
<th>Independent Variable</th>
<th>For-profit Sample</th>
<th>Nonprofit Sample</th>
<th>Government Sample</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>1.31</td>
<td>3.13**</td>
<td>6.02**</td>
</tr>
<tr>
<td>ln PDAYS</td>
<td>(0.69)</td>
<td>(2.50)</td>
<td>(4.25)</td>
</tr>
<tr>
<td>ln ACCRED</td>
<td>0.47**</td>
<td>0.59**</td>
<td>0.42**</td>
</tr>
<tr>
<td>ln PDAYS</td>
<td>(3.31)</td>
<td>(7.00)</td>
<td>(5.20)</td>
</tr>
<tr>
<td>ln ACCRED</td>
<td>0.72**</td>
<td>-0.15</td>
<td>0.24**</td>
</tr>
<tr>
<td>ln PDAYS</td>
<td>(2.58)</td>
<td>(1.38)</td>
<td>(1.46)</td>
</tr>
<tr>
<td>ln ACCRED</td>
<td>0.11**</td>
<td>0.45**</td>
<td>0.12**</td>
</tr>
<tr>
<td>ln PDAYS</td>
<td>(1.48)</td>
<td>(4.97)</td>
<td>(1.54)</td>
</tr>
<tr>
<td>ln ACCRED</td>
<td>0.012</td>
<td>-0.01</td>
<td>0.02</td>
</tr>
<tr>
<td>ln PDAYS</td>
<td>(0.65)</td>
<td>(1.30)</td>
<td>(0.93)</td>
</tr>
<tr>
<td>ln ACCRED</td>
<td>-0.39**</td>
<td>-0.72**</td>
<td>-0.22**</td>
</tr>
<tr>
<td>ln PDAYS</td>
<td>(2.39)</td>
<td>(6.91)</td>
<td>(1.44)</td>
</tr>
<tr>
<td>ln ACCRED</td>
<td>0.31**</td>
<td>-0.03</td>
<td>0.15**</td>
</tr>
<tr>
<td>ln PDAYS</td>
<td>(3.72)</td>
<td>(0.49)</td>
<td>(2.04)</td>
</tr>
<tr>
<td>ln ACCRED</td>
<td>0.07</td>
<td>0.09**</td>
<td>-0.01</td>
</tr>
<tr>
<td>ln PDAYS</td>
<td>(1.64)</td>
<td>(3.15)</td>
<td>(0.34)</td>
</tr>
<tr>
<td>ln ACCRED</td>
<td>0.37**</td>
<td>0.01</td>
<td>0.15**</td>
</tr>
<tr>
<td>ln PDAYS</td>
<td>(2.64)</td>
<td>(1.16)</td>
<td>(1.35)</td>
</tr>
<tr>
<td>ln ACCRED</td>
<td>0.59**</td>
<td>0.57**</td>
<td>0.33**</td>
</tr>
<tr>
<td>ln PDAYS</td>
<td>(4.42)</td>
<td>(4.55)</td>
<td>(2.08)</td>
</tr>
<tr>
<td>ln ACCRED</td>
<td>-0.05</td>
<td>0.16**</td>
<td>0.19**</td>
</tr>
<tr>
<td>ln PDAYS</td>
<td>(0.61)</td>
<td>(3.81)</td>
<td>(0.19)</td>
</tr>
<tr>
<td>ln ACCRED</td>
<td>0.29**</td>
<td>0.28**</td>
<td>0.59**</td>
</tr>
<tr>
<td>ln PDAYS</td>
<td>(3.16)</td>
<td>(3.38)</td>
<td>(5.58)</td>
</tr>
<tr>
<td>ln ACCRED</td>
<td>0.07</td>
<td>0.07</td>
<td>0.10**</td>
</tr>
<tr>
<td>ln PDAYS</td>
<td>(0.88)</td>
<td>(1.45)</td>
<td>(2.19)</td>
</tr>
<tr>
<td>Adj R²</td>
<td>.95</td>
<td>.98</td>
<td>.96</td>
</tr>
<tr>
<td>F</td>
<td>91.9</td>
<td>451.4</td>
<td>263.1</td>
</tr>
<tr>
<td>N</td>
<td>55</td>
<td>96</td>
<td>125</td>
</tr>
</tbody>
</table>

*** Parameter estimate with the absolute value of the t-statistic in parentheses.
** Significant at the 5 percent level.
* Significant at the 10 percent level.

The findings for the control variables are also interesting and merit discussion. The variable cost elasticity with respect to output is less than one in the three equations and indicates that the average hospital operates with short-run economies of scale. These elasticity estimates imply that a 10 percent increase in total inpatient days raises short-run variable costs by 4.2 to 5.9 percent, *ceteris paribus*. The results also suggest that a greater number of services and accreditations, as well as an urban location, are generally associated with higher costs. Number of births, however, has no statistically
significant impact on costs. Also, a greater average length of stay is associated with lower hospital costs, holding constant the number of inpatient days. The coefficient on \( \text{INC} \) is positive for all samples and statistically significant in two of them. This supports the hypothesis that quality, and therefore hospital cost, is responsive to demand-side conditions. Also, as anticipated, the coefficient estimate on I
\( \text{N} \) \( \text{WAGE} \) is positive and significant for all three samples. However, the same percentage increase in the wage rate generates a greater percentage cost increase for the average for-profit hospital than for either the typical nonprofit or government hospital.

Contrary to expectations, the parameter estimates on \( \text{SYSTEM} \) suggest that a hospital's association with a health care system does not lead to lower costs. In fact, the opposite is suggested for the average nonprofit and public hospital, while no such relationship is found for the typical proprietary hospital. Finally, variable costs are not inversely related to a one unit change in the fixed inputs of capital and doctors, suggesting that the typical hospital is not operating in long-run equilibrium. In fact, the results imply that hospitals are over-employing capital and physicians regardless of the ownership structure. Cowing and Holmström (1983) reach a similar conclusion about the long-run equilibrium behavior of hospitals in New York.

CONCLUSION

Aggregate health care expenditures in the United States increased from 5.3 to 11.6 percent of gross national product during the period 1960 to 1989 (Levit et. al., 1991). A large number of policy-makers have responded to this problem of rising costs by advocating more competition in health care markets. This research, however, sheds some doubt on the appropriateness of this proposal for the hospital services sector.

Specifically, the analysis finds that nonprofit hospitals in Texas react to more intense competition by raising hospital quality and expenditures. Note that this particular finding holds even in a time period characterized by increased sensitivity to hospital costs. The 1987/88 study year follows the implementation of the federal diagnostic related groups (DRG) system, the evolution of health maintenance organizations, and the use of second opinions and third-party pre-authorization. In fact, it may be the piecemeal nature of these cost containment policies that provides nonprofit hospitals with the opportunity to continue to engage in the "medical arms race".

Whether these results can be generalized beyond this sample of Texas hospitals remains unclear at this point, since the market environment and regulatory climate vary so dramatically across states. On the one hand, nonprofit hospitals are much more numerous in other states, representing over seventy percent of all short-term hospitals in the United States. The greater availability and rivalry of these institutions may lead to even higher hospital costs than observed in Texas. On the other hand, hospitals are subject to stringent regulations and rate review in several states (e.g., Connecticut, Massachusetts and Maryland). These regulations may act to control the nonprice competition of nonprofit hospitals. Thus, further studies utilizing other state and/or regional samples may be necessary before any generalizations can be made about the effect of market structure on hospital costs during a period of heightened cost consciousness.

HOSPITAL MARKET STRUCTURE

APPENDIX A
Deriving the Number of Admitting Physicians

The average number of admitting doctors at each hospital, DOC, can be written as:

\[ \text{DOC} = \frac{\text{HADM}/(\text{HADM} + \text{HDOC})} {\text{HADM}} \]

where HADM and HDOC represent the number of hospital admissions and doctors with admitting privileges at each individual hospital, respectively. While data for HADM exist, data for HDOC are unavailable. Therefore, an assumption is made that the ratio of hospital admissions to hospital doctors, HADM/HDOC, is equal to the equivalent county ratio, CADM/CDOC, or:

\[ \text{DOC} = \frac{\text{HADM}/(\text{CADM} + \text{CDOC})} {\text{CADM}} \]

Since data for the number of hospital admissions in the county, CADM, are also unavailable, equation A.2 is rewritten in terms of county population, CPOP, for which data can be found:

\[ \text{DOC} = \frac{\text{HADM} / [(\text{CPOP}/(\text{CADM} + \text{CPOP})] / \text{CDOC}} {\text{CPOP}} \]

It is now assumed that the county admissions to population ratio, CADM/CPOP, is equal to the state admissions to population ratio which is approximately 13. Substituting this figure into equation A.3 provides an estimate of the number of admitting physicians at each individual hospital:

\[ \text{DOC} = (\text{HADM}/(\text{CPOP}/13) + \text{CDOC}) / \text{CPOP}. \]
## APPENDIX B

### Hospital Type Profiles

<table>
<thead>
<tr>
<th>Variable</th>
<th>For-profit Mean (n = 55)</th>
<th>Nonprofit Mean (n = 96)</th>
<th>Government Mean (n = 120)</th>
</tr>
</thead>
<tbody>
<tr>
<td>VC ($)</td>
<td>15,735,527 (14,916,039)</td>
<td>39,417,260 (33,585,906)</td>
<td>11,556,984 (9,289,705)</td>
</tr>
<tr>
<td>PDAYS</td>
<td>12887 (17017)</td>
<td>50731 (69011)</td>
<td>13251 (32078)</td>
</tr>
<tr>
<td>ACCRED</td>
<td>3.0 (6.8)</td>
<td>3.5 (1.3)</td>
<td>2.6 (1.1)</td>
</tr>
<tr>
<td>SER</td>
<td>17 (7.5)</td>
<td>22 (10)</td>
<td>13 (8.2)</td>
</tr>
<tr>
<td>BRTHS</td>
<td>379 (857)</td>
<td>1090 (1351)</td>
<td>522 (1613)</td>
</tr>
<tr>
<td>STAY</td>
<td>5.9 (2.9)</td>
<td>5.6 (1.4)</td>
<td>5.3 (3.0)</td>
</tr>
<tr>
<td>URBAN</td>
<td>0.76 (0.43)</td>
<td>0.72 (0.45)</td>
<td>0.22 (0.41)</td>
</tr>
<tr>
<td>MKT (HII)</td>
<td>0.639 (0.360)</td>
<td>0.420 (0.305)</td>
<td>0.743 (0.317)</td>
</tr>
<tr>
<td>INC ($)</td>
<td>10,513 (2,843)</td>
<td>9,774 (1,864)</td>
<td>8,199 (1,749)</td>
</tr>
<tr>
<td>WAGE ($)</td>
<td>20,813 (3,764)</td>
<td>19,604 (3,500)</td>
<td>17,276 (3,781)</td>
</tr>
<tr>
<td>SYSTM</td>
<td>0.8 (0.6)</td>
<td>0.6 (0.5)</td>
<td>0.2 (0.4)</td>
</tr>
<tr>
<td>K (BEDS)</td>
<td>114 (91)</td>
<td>220 (91)</td>
<td>76 (229)</td>
</tr>
<tr>
<td>DOC</td>
<td>41.9 (53.1)</td>
<td>116.4 (163.6)</td>
<td>34.4 (115.7)</td>
</tr>
</tbody>
</table>

*Mean values with standard deviation in parentheses.*

## REFERENCES


