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HOSPITAL PERFORMANCE

PRICE COMPETITION AND HOSPITAL COST GROWTH IN THE UNITED STATES (1989–1994)

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SUMMARY

In recent years, most health care markets in the United States (US) have experienced rapid penetration by health maintenance organizations (HMOs) and preferred provider organizations (PPOs). During this same period, the US has also experienced slowing health care costs. Using a national database, we demonstrate that HMOs and PPOs have significantly restrained cost growth among hospitals located in competitive hospital markets, but not so in the case of hospitals located in relatively concentrated markets. In relative terms, we estimate that HMOs have contained cost growth more effectively than PPOs. Copyright © 1999 John Wiley & Sons, Ltd.

KEY WORDS — managed care; hospital competition; hospital costs

INTRODUCTION

The share of US GDP devoted to health care grew from 7.1% in 1970 to 13.6% in 1995 [1], far exceeding the experience of other OECD countries even after accounting for differences in per-capita income [2]. Many stakeholders are vitally interested in reducing the rate of health expenditure growth in the US. These include large employers facing rapidly growing health benefit costs; Federal, State, and County governments trying to provide social services with shrinking resources; and less well-off Americans who have seen their incomes stagnate, even decline, since the early 1970s [3].

In 1982 the State of California adopted a landmark legislation designed to control health expenditures. By allowing health insurance plans to selectively contract with hospitals, the new law intended to stimulate price competition in the health sector. This law was widely emulated throughout the US. As a result, market-based managed care health plans have become an important new force in the US health care system. It is estimated that more than half of the total US population (58% in 1995) is enrolled in some form of managed care plan—either health maintenance organizations (HMOs) or preferred provider organizations (PPOs) [4].

Despite the rapid growth of managed care in the US, to date there has been very little empirical research regarding its effects at the national level. Available studies are generally regional in focus, often limited to data from a single state such as California. Although these studies indicate that managed care reduces hospital cost growth, ana-

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lysts question the wide applicability of these regional studies [5]. The analyses presented here bridge this research gap by addressing three interrelated issues. First, using a national database, we test whether managed care has attenuated hospital cost growth broadly in the US. Second, we explore whether HMOs and PPOs vary in their ability to attenuate hospital cost growth. Finally, we explore the importance of the interactive relationship between managed care's market share and hospital competition on hospital cost growth.

CONCEPTUAL FRAMEWORK

The purchase of health care involves a complex set of transactions among many parties. For example, the market for hospital services involves at least four players (patient, physician, hospital, and insurance plan), with employers now playing an increasingly important role as well. In the traditional fee-for-service era, competition among hospitals was based predominantly on non-price factors. Physicians were reimbursed on a fee-for-service basis, a regime that did not foster cost-effective utilization of health care resources. Hospitals competed by duplicating costly amenities and services (non-price competition) to attract physicians and patients [6–8]. With such incentives in place, it was no surprise to find health care costs to be positively correlated to the level of non-price competition prevailing in a market [9].

The growth of managed care reverses many of the cost-increasing incentives present in the traditional fee-for-service model [10]. First, selective contracting sets in motion price competition among providers who offer price and other concessions in order to obtain access to a plan's subscribers. Depending upon the level of competition in insurance markets, these price concessions are passed on to consumers as plans jockey for market share. Employers that join a purchasing coalition or improve price sensitivity among their employees can leverage interplan competition more effectively to obtain lower premiums. Employers can foster price sensitivity by providing employees with clear information about price, benefits, and quality on each plan, and by limiting the employer's premium contribution to the cost of covering essential benefits. Second, managed care plans reimburse providers so as to put the provider, at least partially, at risk for excess utilization reversing incentives embedded in traditional indemnity insurance plans. Price concessions and risk bearing in turn puts pressure on providers to reduce costs and to improve their efficiency. The degree of pressure brought to bear by HMOs and PPOs differs because of the way they are structured. HMO's are supposedly more effective at cost control because they rely on capitated payments and often distribute a share of the inpatient care savings back to the physicians. Thus hospitals and physicians contracting with a HMO have strong incentives to reduce utilization. PPOs control hospital costs through less aggressive techniques, such as relying on selective contracting to obtain price discounts, and on utilization review to prune cost-ineffective care.

The extent to which managed care reduces costs in practice is largely an empirical question. Many countervailing forces can limit the level of bargaining power plans enjoy *vis-à-vis* providers. The level of price competition that selective contracting can introduce in a given market depends upon several factors. These include: (1) the number of competing hospitals and plans; (2) the elasticity of health care demand with respect to price, quality and distance; and (3) patient loyalty to a particular hospital or plan. If plans assemble a network only on the basis of price, they risk losing subscribers because of either unacceptable quality or access. Over time, providers of medical care (both doctors and hospitals) can also consolidate themselves into larger groups to better manage risks associated with capitated reimbursement and to improve their bargaining position. The net impact of these changes is difficult to predict. Larger units may be more efficient and, therefore, able to offer lower prices; but their negotiating position is also stronger, tending to raise health expenditures.

PREVIOUS STUDIES

Most previous studies have examined California's hospital industry, the state with the longest managed care history in the US, and fortuitously a good state-level system for collecting detailed hospital data. These studies demonstrate that stimulation of price competition in California has significantly slowed the rate of hospital cost growth [11–15]. Melnick *et al.* [16] find that managed care plans can effectively negotiate lower prices through selective contracting, with price being inversely related to the level of hospital competition.

Robinson [12] specifically analyses the impact of HMO enrollment on hospital cost growth in California by comparing general acute-care hospitals located in areas with the most and least HMO penetration (upper and lower quartile of the hospital sample). In 1993, HMO enrollees accounted for 10.7% of all admissions in the least penetrated markets and 31.4% in the most penetrated markets. Between 1983 and 1993, Robinson found that hospital costs increased by 61% and 109%, respectively, in the most and least penetrated markets, equivalent to (nominal) annual growth rates of 4.9% and 7.6%, respectively. In other words, even when HMOs account for only about one-third of hospital admissions the impact on cost trends is both substantial and sustained.

Inter-state comparisons also lend support to the effectiveness of managed care in restraining health expenditure growth. Melnick and Zwanziger [17] compare growth in hospital costs and per-capita health expenditures between California and rate-regulated states such as New Jersey, New York, Massachusetts and Maryland. Between 1980 and 1991, they estimate that per-capita (total) health expenditures grew by 39% in California, 86% in New Jersey, 85% in New York, 70% in Massachusetts, 59% in Maryland, and 63% overall for the entire US; hospital cost trends follow a similar pattern. These long-run trends challenge Robinson and Luft's [18] earlier finding that rate-regulated states slowed hospital cost growth somewhat more effectively than California—their analyses were limited to a very short time period (1982–1986) when selective contracting and managed care were still nascent forces. These previous studies taken together suggest that price competition can slow the rate of health care cost growth. In this paper we test whether this finding holds for the entire US.

METHODS AND DATA

We use a national database to estimate the interdependent effects of managed care penetration and hospital competition on US hospital cost growth between 1989 and 1994. To date, hospital cost studies have principally relied on measures of hospital market structure while other studies have included measures of only managed care penetration, but none have explored the important interactive effects between these two phenomena [19]. Our model includes both types of measures as well as their interactions. The impact of managed care is allowed to depend upon its organizational form (HMO versus PPO) and upon a plan's ability to generate price competition among hospitals, in part a function of hospital market structure. The model also separately accounts for areas subject to hospital rate-regulation during the study period.

MODEL SPECIFICATION AND ESTIMATION

Hospital costs are conceptualized as a time-series, cross-section model as shown below (Equation (1)). The model controls for changes in hospital output, input prices, demand, hospital ownership and urban/rural location, and Medicare and Medicaid fiscal pressure. External influences on a hospital such as fiscal pressure from government payers (Medicare and Medicaid), managed care and hospital competition are allowed to have time-varying coefficients. Hospital ownership and urban/rural indicators are also allowed to have time-varying coefficients; the former capture the effect of changes in operating objectives across for-profit, not-for-profit and public hospitals while the latter capture unmeasured input price changes over time. In other words, the relationship between hospital costs, output, input prices and demand is expected to stay steady over time in the absence of changes in the external environment.

$$\ln E = \alpha_t + \beta(\ln(O, P, I)) + \gamma_t(X, FR, FD, H, M) + \mu + \varepsilon \ldots \quad (1)$$

where

E	annual operating hospital costs
O	output (discharges, casemix, visits, teaching intensity)
P	input prices (relative wage index)
I	demand (per-capita income)
X	hospital characteristics (ownership, urban/rural location)
FR	Medicare fiscal pressure
FD	Medicaid fiscal pressure

H Hirschman–Herfindahl index (HHI) of hospital market structure
M indicator variables for HMO penetration, PPO penetration and hospitals under rate-regulation
μ hospital-level fixed effects
ε random error $\sim N(0, \sigma^2)$.

Managed care's effect on hospital costs is captured by categorizing hospitals according to whether they operated in markets with high, medium or low levels of HMO and PPO penetration or whether they were subject to state-level hospital rate-regulation programmes. The model includes interactions between the HHI and these categorical variables as well. Hospital markets are nested within managed care markets (geographically the latter tend to be far bigger). Therefore, in a hospital-level model it makes more sense to categorize the managed care penetration measures and to interact these categories with continuous hospital-level HHI measures.

Taking differences between the end-year (1994) and base-year (1989) eliminates the fixed effects, leading to Equation (2). Since the dependent variable is logarithmically transformed before taking differences, it can also be interpreted as the logarithm of the ratio of end-year to base-year total hospital costs.

$$\begin{aligned}\Delta\, Ln\, E = {} & \Delta\alpha + \beta(\Delta\, Ln(O, P, I)) \\ & + \gamma_{94}(\Delta(X, FR, FD, H, M)) \\ & - \Delta\gamma((X, FR, FD, H, M)_{89}) + \Delta\varepsilon \qquad (2)\end{aligned}$$

where $\gamma_{94} = \gamma_{89} - \Delta\gamma$.

Independent variables that do not change significantly over time are dropped prior to estimation (that is, Equation (2)'s middle term associated with the 1994 vector of coefficients). For example, urban/rural location is time invariant. Very few hospitals in the sample switch ownership during the analysis period (approximately 6%), so the results are insensitive to whether base- or end-year ownership is used in the model. Due to data limitations, our measures of Medicare and Medicaid fiscal pressure and hospital competition (HHI) also do not change over time. Our previous research on California's hospital markets suggests that the HHI changes very slowly over time—for California hospitals, the 1989 HHI was correlated approximately 95% with the 1994 HHI—so lack of data about HHI changes at the national level is not a serious shortcoming.

The final model also excludes changes in managed care penetration over time. Managed care penetration data necessary for the analyses are available only for the period 1989–1992. We estimated a 1989–1992 difference model including changes in managed care penetration between these two years, but these turned out to be statistically insignificant. In other respects, however, the 1989–1992 difference model provides qualitatively similar results compared to the 1989–1994 difference model which omits managed care penetration changes over time. This finding is not altogether surprising given that managed care penetration in 1989 and 1992 are correlated approximately 90%. Thus, use of only a single year of managed care penetration data is unlikely to introduce significant errors.

Length of stay and number of operating beds is excluded from the model because changes over time are potentially endogenous—the impact of managed care on visits and discharges has been far more equivocal [20]. If managed care generates savings only by reducing admission rates on a per-capita basis, then hospital cost growth controlling for discharges should exhibit no relationship with managed care. It follows that any observed relationship between hospital cost growth and managed care controlling for discharges probably understates the full impact of managed care on *per-capita* hospital cost growth; in other respects inference at the hospital level should remain valid. The biases one has to worry about in a hospital-level analysis are due to managed care's impact on length of stay or substitution of outpatient for inpatient care. For this reason, length of stay is omitted from the model. We also re-ran the final model without visits and the paper's key results remained unchanged.

The difference model is estimated using weighted least squares because error variance declines with hospital size. A variance function relating the logarithm of absolute residuals to the logarithm of base-year hospital discharges is estimated in order to derive the weights. Carroll and Ruppert [21] discuss advantages of variance functions estimated using absolute residuals, namely, much less susceptibility to outliers. The results are not sensitive to whether weights are applied or ignored. Ramsey's [22] RESET test does not reject (at a 5% level of significance) the null hypothesis of no omitted variables bias.

DATA AND SAMPLE CHARACTERISTICS

Our national hospital-level database combines information from several sources. Data on hospital total annual operating costs, discharges, teaching intensity and other hospital characteristics were obtained from the annual Medicare cost reports. Reporting cycles vary across hospitals, so data from successive cost reports were linked, then annualized. Differences in casemix and input prices across hospitals are captured through the Medicare casemix and relative wage index, obtained from the Health Care Financing Administration (HCFA). The relative wage index is a cross-sectional deflator. Teaching intensity is measured as the ratio of interns and residents to total hospital beds, also obtained from HCFA. The Area Resource File provides data about per-capita income by county. Data sources underlying the market structure measures are described along with the method used for their construction.

The original Medicare cost report file contains a universe of 5418 and 4974 hospitals for 1989 and 1994, respectively. As a result of either out-of-range or incomplete data, our edited Medicare cost report analysis file contains 4382 and 3904 hospitals for 1989 and 1994, respectively. Because of mergers and closures, the universe of short-term general hospitals has contracted over time. Most hospital-level data are drawn from the Medicare cost reports. Outpatient visits, however, are not reported to HCFA but are available from the American Hospital Association's (AHA) annual surveys. Merging information between the AHA and Medicare databases causes additional sample attrition due to an incomplete crosswalk between the AHA and Medicare hospital identification numbers. The final regression model is based on 3456 hospitals for which we have complete data in the base- and end-years.

We compared the analysis sample to the base- and end-year hospital universe by region, ownership and bed size. The sample appears similarly weighted for both the base- and end-years, but somewhat under-weighted in the case of for-profit hospitals, partially explaining the slight under-weighting of the pacific census region where for-profit hospitals are disproportionately located. If this under-weighting introduces any bias at all, it is likely to be toward a null finding, because the under-weighted areas have been subject to higher levels of managed care penetration.

KEY VARIABLE CONSTRUCTION

Dependent variable

The change in annual hospital operating costs between the base- and end-year is used as the dependent variable.

HMO and PPO penetration

Separate HMO and PPO penetration measures are constructed from plan-level enrollment and market-definition data obtained from SMG Marketing, Inc. for 1989–1992. SMG Marketing, Inc. surveys both HMOs and PPOs and collects detailed information about enrollment levels and geographic areas served by each plan. Because private insurance is mostly employment-based, total plan enrollment is allocated to each county in the plan's market in proportion to the total population employed in the government, utilities, manufacturing and transportation sectors. For PPOs we have information only about the number of policies issued, not total covered lives. To estimate the total PPO-enrolled population, we assume that each PPO policy on average covers 2.3 individuals based upon aggregate published data [23]. Any error in this assumption affects only the estimated magnitude of PPO penetration, not the relative ranking of different geographic areas with respect to this measure. HMO and PPO enrollment is then aggregated across counties included in an MSA, and divided by the MSA's population to obtain MSA-level penetration measures. Each county in the MSA receives the same penetration score, minimizing the impact of errors in the allocation of plan enrollment to specific counties. Similarly, HMO and PPO enrollment is aggregated across non-MSA counties in each state and divided by the total population in these counties to generate state-specific, non-MSA penetration measures.

Our MSA-level estimates of HMO penetration compare very well with other similar estimates previously published in the literature [24]. The population-weighted correlation between the two series of MSA-level HMO penetration estimates works out to approximately 83%. In any event, measurement error in our managed care penetration estimates is likely to generate a downward bias in the estimated impact of price competition on hospital cost growth.

States with rate-regulation programmes are accounted for separately. New York and Maryland adhered to rate-regulation for all payers throughout the study period, while New Jersey adhered to rate-regulation for the greater part of the analysis period (until 1992) [25]. Thus, New Jersey is also included in the rate-regulated group. In the past, many have advocated rate-regulation as a way of containing health care costs, making the experience of rate-regulated states an important benchmark by which to judge the experience of other parts of the US.

Hospital market structure

We construct a hospital-level HHI from the 1989–1990 Medicare discharge data (MEDPAR) using actual zip code level patient flow data to define hospital markets [11,26]. Five steps are required to calculate each hospital's HHI. First, all DRGs (diagnostic related groups) are categorized into 48 separate service categories based on the type of physician that typically treats a patient in a given DRG. Next, a hospital's market area by service is identified using patient origin data—zip code areas (ZCAs) are included in the market if they contribute at least 3% of the hospital's discharges for that service. Third, competing hospitals are identified on a zip code level basis as those facilities that draw significantly—at least 3% of a ZCA's total discharges for a given service—from at least one of the ZCAs in the marker hospital's market area. Fourth, an HHI is calculated for each ZCA service combination. Finally, the degree of competition facing each hospital is captured by estimating a weighted average of the ZCA service HHIs in its market, with the proportion of patients it draws from each ZCA service combination serving as the weight.

Because Medicare discharges account for approximately 40% of total hospital admissions, we validated these measures in five states (California, New Jersey, Florida, Michigan, Wisconsin) for which all-payer discharge data are also available. The all-payer and Medicare HHIs are highly correlated within and across these five states with Michigan exhibiting the highest correlation (approximately 94%). Across all five states taken together the correlation works out to approximately 88%. In other words, the high correlation is not caused by the inclusion of California and Florida (with large Medicare eligible populations) among the validation states. Measurement error in measures of hospital market structure is once again likely to introduce a downward bias in the impact of this variable on hospital cost growth.

Medicare fiscal pressure

It is not easy to assess the level of Medicare fiscal pressure a hospital was facing during the analysis period because of the complexity and evolution of the Prospective Payment System (PPS) over time. At the inception of this programme (1984) it was well-known that hospitals ultimately would be reimbursed according to a national rate adjusted for patient casemix and regional wage rate differences. But this goal was approached gradually in the interim, with hospitals being reimbursed with a blend of the national rate and their pre-PPS historical costs. We capture Medicare fiscal pressure through an index that ranks hospitals according to how profitable their Medicare business would have been at the inception of the PPS had a national rate gone into effect immediately. This index is created as follows. First, for each hospital we estimate the average Medicare cost per discharge, normalized for casemix and wage rate differences in 1985. We then calculate a discharge weighted average of these averages separately for large urban hospitals and the remaining group in order to estimate the national reimbursement rate (PPS uses a different national rate for these two hospital groups). Any difference between our estimated and actual national average rates does not affect the relative ranking of hospitals. For each hospital we simulate what its total payments would have been in 1985 had it been reimbursed according to the estimated national average rate, taking into account casemix and wage rate differences. The difference between actual and simulated payments divided by total net patient revenues in 1985 then yields the Medicare fiscal pressure variable.

Medicaid fiscal pressure

Medicaid fiscal pressure is also difficult to capture because of considerable variation in the generosity of Medicaid reimbursement policies across states and over time. Loprest and Gates [27] report a state-level Medicaid hospital payment index that measures the fraction of hospital Medicaid costs

Table 1. Means and S.D.s of the model variables

	1989		1994	
	Mean	S.D.	Mean	S.D.
Hospital Costs	33.9 (M)	46.5 (M)	50.3 (M)	70.7 (M)
Discharges	6067.6	6647.9	6080.6	6869.4
Medicare casemix index	1.198	0.174	1.243	0.222
Outpatient visits	51390.2	67624.5	71482.4	90192.3
(Interns + residents)/beds	0.023	0.076	0.027	0.089
Available beds	189.9	173.6	187.1	171.4
Relative wage index	1.086	0.183	1.098	0.216
County per-capita income	15819.5	4126.4	19698.5	5005.8
Medicare pressure index	−0.041	0.092	—	—
Percent Medicaid days	0.120	0.118	—	—
HHI	0.378	0.130	—	—
Medicaid payment index <0.75) indicator	0.299	—	—	—
Medicaid payment index (≥ 0.75 & <1.00) indicator	0.656	—	—	—
Not-for-profit indicator	0.597	—	—	—
For-profit indicator	0.108	—	—	—
Sole community provider indicator	0.136	—	—	—
Rural location indicator	0.001	—	—	—
High HMO penetration ($>7\%$) indicator	0.404	—	—	—
Medium HMO penetration (2–7%) indicator	0.264	—	—	—
High PPO penetration ($>7\%$) indicator	0.314	—	—	—
Medium PPO penetration (2–7%) indicator	0.478	—	—	—
Regulated state indicator	0.072	—	—	—

that were reimbursed by state Medicaid programmes on average in 1990. In some states such as California Medicaid programmes reimbursed only approximately 67¢ on the dollar, while in New Jersey, Maryland and Arizona they reimbursed roughly dollar for dollar. In our model, we capture the impact of Medicaid fiscal pressure in two ways. First, we group states into three categories based upon the generosity of their reimbursement policies, and include indicator variables for these categories. Second, we include interactions between the proportion of total hospital days accounted for by Medicaid beneficiaries and these state-grouping indicator variables in order to capture differences in fiscal pressure across hospitals. The model uses percent Medicaid days in the base-year (1989). Because short-term changes in percent Medicaid days are small, substituting with end-year (1994) data generates qualitatively similar results.

RESULTS

Table 1 presents descriptive statistics for the model variables while Table 2 presents the regression output; the dependent variable is the 1989–1994 difference in the logarithm of annual hospital operating costs. Many of the estimated coefficients are significant and all of the significant coefficients are in the expected direction. For example, changes in output and per-capita income are all strongly and positively related to changes in total hospital costs. The estimated coefficient for inpatient discharges is 0.402, while the estimated coefficients for casemix and outpatient visits are 0.191 and 0.065, respectively. Teaching output, measured as interns and residents per bed, also has a positive and significant estimated coefficient of 0.283.

The estimated coefficient for the relative wage index, although negative, is small and insignifi-

Table 2. Hospital Cost Growth Model (1989–1994)

Covariate	Coefficient	S.E.
Δ ln(discharges)	0.402**	0.011
Δ ln(Medicare casemix index)	0.191**	0.031
Δ ln(outpatient visits)	0.065**	0.005
Δ ln(1 + ((interns + residents)/beds))	0.283*	0.122
Δ ln(relative wage index)	− 0.035	0.039
Δ ln(county per-capita income)	0.156**	0.033
Not-for-profit indicator	− 0.041**	0.006
For-profit indicator	− 0.133**	0.009
Sole community provider indicator	0.031**	0.008
Rural location indicator	− 0.045	0.066
Medicare pressure index	− 0.282**	0.029
Medicaid payment index (< 0.75) indicator	− 0.109**	0.019
Medicaid payment index (≥ 0.75 & < 1.00) indicator	− 0.085**	0.019
Percent Medicaid days × Medicaid payment index (< 0.75)	0.178	0.129
Percent Medicaid days × Medicaid payment index (≥ 0.75 & < 1.00)	0.200	0.128
Percent Medicaid days	− 0.145	0.124
HHI	− 0.041	0.070
HHI × high HMO penetration (> 7%) indicator	0.253**	0.064
HHI × medium HMO penetration (2–7%) indicator	− 0.064	0.068
HHI × high PPO penetration (> 7%) indicator	0.247**	0.097
HHI × medium PPO penetration (2–7%) indicator	− 0.014	0.066
HHI × high HMO penetration indicator × high PPO penetration indicator	− 0.097	0.102
HHI × regulated state indicator	0.005	0.088
High HMO penetration (> 7%) indicator	− 0.166**	0.027
Medium HMO penetration (2–7%) indicator	− 0.002	0.030
High PPO penetration (> 7%) indicator	− 0.080†	0.043
Medium PPO penetration (2–7%) indicator	0.004	0.030
High HMO penetration indicator × high PPO penetration indicator	0.042	0.042
Regulated state indicator	− 0.101**	0.038
Intercept	0.528**	0.038
Adjusted R^2	0.434	
n	3456	

** Significant at 1% level.
* Significant at 5% level.
† Significant at 10% level.

cant. This finding is not surprising because HCFA's relative wage index is a cross-sectional adjustor, not a time-series deflator. The underlying time trend is captured by the intercept.

The estimated coefficients for ownership reveal differential cost growth patterns. Relative to the omitted category (government hospitals), for-profit hospitals and not-for-profit hospitals showed much slower growth in costs. The estimated coefficient for the for-profit hospital category is − 0.133, which translates into 12.2% ($e^{-0.133} - 1$) lower cost growth compared to government hospitals, while the estimated coefficient for non-profit hospitals is − 0.041. Hospital cost growth in sole community hospitals was slightly higher than government hospitals, as indicated by an estimated coefficient of 0.031.

Medicare fiscal pressure is strongly and inversely related to hospital cost growth. For every 1 percentage point of total net patient revenues potentially threatened by the PPS, the ratio of end-year to base-year total hospital costs is lower by 0.282%. Medicaid fiscal pressure also appears to have forced hospitals to control their costs; estimated coefficients are negative and significant for states reimbursing less than 100% of Medicaid costs.

Several variables are included in the model to capture the main and interactive effects of managed care and competition on hospital cost growth. The direct effect of hospital competition, captured by the HHI, is statistically insignificant. To fully understand the effects of hospital competition, however, it is necessary to take into account HHI's interactions with the other variables. The model includes six such interaction terms of which two are highly significant. The estimated coefficient for the interaction of HHI with the high HMO and high PPO penetration indicator is 0.253 and 0.247, respectively, but the three-way interaction between these variables is not significant. HHI's interaction with the rate-regulated state indicator also has a small and statistically insignificant coefficient.

Five categorical measures of managed care penetration are also included in the model. The estimated coefficients for the high HMO and high PPO penetration indicators are both negative and significant: -0.166 and -0.080, respectively. The interaction between these two indicator variables is not statistically significant. Finally, the estimated coefficient for the rate-regulated state indicator is negative and significant with an estimated value of -0.101.

When interactions between the managed care penetration indicators and the HHI are significant (high penetration areas), then the HHI level in part drives how effective HMOs and PPOs are at controlling cost growth. At the *limiting* case of HHI equal to zero (highly competitive hospital markets) the HHI interactions also reduce to zero—under these conditions high HMO penetration reduces hospital cost growth by the greatest amount (15.3%) followed by high PPO penetration (7.7%). And since the HHI's interactions with the high HMO and high PPO indicators have roughly equal coefficients, it follows that in the high penetration areas, HMOs exercise greater cost control than PPOs regardless of the HHI level. Hospital competition does not appear to play a significant role in the rate-regulated areas. Finally, the model tests whether high levels of HMO and PPO penetration interact to generate a level of cost control that exceeds the sum of their independent effects. This appears not to be the case—interactions between the high HMO and high PPO penetration indicators as well as their further interaction with the HHI are all statistically insignificant.

DISCUSSION AND CONCLUSIONS

Our results suggest that managed care's effects, first observed in California, are now spreading throughout the US. In other words, managed care's expansion in the US is changing the underlying market dynamic by placing greater importance on price competition instead of the traditional quality and service-based (non-price) competition.

Hospitals in high managed care penetration areas (that is, areas with high HMO or high PPO penetration) displayed a significantly lower rate of cost growth between 1989 and 1994. Our findings, however, also suggest that the ability of managed care plans to slow hospital cost growth is critically dependent upon the level of hospital competition. The estimated managed care effect—that is, the difference in hospital cost growth between high and low managed care penetration areas—is statistically significant only in the case of hospitals located in highly competitive hospital markets.

This study also provides preliminary empirical evidence on the relative effectiveness of HMOs compared to PPOs in controlling hospital cost growth. It has been hypothesized that HMOs exercise greater cost control relative to PPOs for several reasons. First, HMOs typically construct more restrictive provider networks endowing the HMO with greater leverage in negotiating contracts with health care providers. Second, some HMOs pay physicians and hospitals on a capitated basis while PPOs rely almost exclusively on price discounts. Capitation endows physicians with stronger incentives to reduce hospital costs than discounted fees. Our study, although unable to separate the independent effects of alternative payment mechanisms, does provide an aggregate estimate of the relative effect of HMO versus PPO penetration on hospital cost growth. For example, at a theoretical *limiting* case of HHI equal to zero (highly competitive markets), hospital cost growth was 15.3% less in high HMO penetration areas compared to medium or low HMO penetration areas, and 7.7% less in high PPO penetration areas compared to medium or low PPO penetration areas.

Our findings also highlight the important interactive effect between hospital competition and managed care penetration. The estimated coefficient on the HHI is not significant, but its interactions with the high HMO and high PPO penetration indicators are highly significant. In other words, greater hospital competition is effective only in areas with high levels of managed care penetration and *vice versa*. As the intensity of hospital competition declines, the effects of managed care penetration also decline and, in fact, are no longer significant in the more concentrated markets. The estimated interactive effects, however, may be subject to biases that deserve additional research. For example, our cost model does not explicitly control for the technological sophistication or quality of care in a given hospital, except by way of the fixed effects. Our measure of hospital competition is also only based upon Medicare patients. To the extent Medicare patients prefer technologically sophisticated hospitals, and technological sophistication is related to the level of managed care penetration, it is possible that our empirical estimates of the interactive effects are somewhat biased.

The substantial growth in managed care enrollment in the US combined with our findings of significant cost savings have important implications for the future growth of managed care. Managed care's expansion in the US has been remarkable. More than 140 million Americans are now covered by either HMOs or PPOs, representing more than half of the total or two-thirds of the insured population. Moreover, there are now very few areas in the country where managed care has yet to penetrate. Our findings on the substantial cost savings from managed care growth, and particularly HMOs, suggest there will be strong economic incentives to maintain and expand managed care. Taken together, these trends indicate that managed care is a national phenomenon that will become the dominant form of health insurance in almost every part of the US. Several factors, however, may affect cost savings under managed care in the future. First, it has been suggested that managed care plans target high cost areas for early entry and expansion; by implication, later expansion into less costly areas may yield lower savings. Second, there appears to be a trend toward the introduction of additional state-level legislation to regulate managed care plans. Many of these laws could reduce innovation and reduce price competition leading to higher cost growth in the future. Third, mergers among health care providers could accelerate in the future. Should these trends continue, managed care plans' ability to generate price competition will be severely hampered leading to higher rates of cost growth in the future.

Many areas in the managed care and competition literature could benefit from additional research. For example, it is important to note that our findings pertain only to hospital costs. Hospital costs comprise the single largest component of total health care expenditures in the US (approximately 40%). To the extent managed care reduces hospital cost growth by shifting care from hospital to non-hospital settings, the net savings from managed care may be lower. Future studies in this area should, therefore, attempt to incorporate broader measures of health expenditures while assessing overall health system performance.

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