Nurse Staffing, Quality, and Financial Performance

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In examining the relationship among nurse staffing, quality of care, and financial performance, prior empirical studies used competing measures and applied different levels of analysis. Using longitudinal data from 1990 through 1995, our study applied a dynamic econometric model to evaluate whether hospitals that changed their nurse staffing and quality of care affected their financial performance. Sampling 422 hospitals over this study period, we found a statistically significant increase in operating costs when registered nurse levels increase, but no statistically significant decrease in profit. Higher levels of non-nurse staffing caused higher operating expenses, as well as lower profits. Key words: *nurse staffing, quality, financial performance, mortality.*

THROUGHOUT the 1990s, hospitals experienced significant variability in financial performance. At the start of the 1990s, increasing competition, shrinking government payments, and rising levels of managed care penetration reduced revenues of hospitals and eroded their profit margins such that 32 percent of all hospitals operated at a loss in 1990.[1] Between 1992 and 1995, however, average operating margin showed a steady increase. Factors cited for the improved performance included decreased patient length of stay (LOS), reduced operating expenses, lower staffing levels, and higher government payments. For example, staffing levels declined from 5.21 full-time equivalents (FTEs) per adjusted discharge in 1992, to 4.70 FTEs per adjusted discharge in 1996.[2] Overall, hospital FTE employees fell in both 1993 and 1994, declining 2.8 percent.[3] Approximately 20,000 hospital FTEs were cut in 1995.[4] With nursing personnel comprising approximately 30 to 40 percent of overall hospital FTE personnel and approximately 30 percent of the hospital budget,[5] strategies to improve hospital financial performance

Acknowledgement: Supported by grant number HS10153 from the Agency for Healthcare Research and Quality.

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J Health Care Finance 2003;29(4):54–76

frequently involved hospital downsizing and reductions in nursing staff.[6] Underlying these administrative actions was a belief that reducing nurse staffing would reduce labor costs, which would in turn be reflected in improved hospital financial performance.

Policymakers and health care executives are concerned about the relationship between nurse staffing and financial performance for several reasons. First, as previously mentioned, labor costs account for more than half of a hospital's operating expenses[7] and the nurse workforce continues to constitute more than one-third of its labor force.[8] Second, with a significant nursing shortage looming on the horizon, labor costs, particularly higher wages needed to attract and retain nurses as well as the costs for temporary nurses, may place a substantial burden on acute care hospitals.[9] Third, the implementation of minimum staffing ratios, now mandated in California and under consideration in at least 18 states, may place further pressure on the operating cost and profit margins of hospitals. Yet, no empirical analyses have been conducted to determine whether changes in nurse staffing levels affect the financial performance of hospitals. Therefore, the primary aim of this study was to evaluate the effects of change in nurse staffing on changes in hospital financial performance during the years of 1990 through 1995.

Since hospitals were reducing costs to raise their profit margins, quality of patient care may have been compromised to achieve this financial outcome. Conversely, hospitals may have attempted to raise their quality of care, thereby improving their reputation in order to attract patients and earn higher profits. Therefore, we also examined whether changes in quality of care were related to changes in hospital financial performance. Further, based on differences in competitive factors in markets characterized by high managed care penetration versus those in low managed care markets, we examined whether the relationships between change in nurse staffing, change in quality, and change in financial performance differed based on the extent of managed care penetration. Finally, we examined whether these relationships differed based on hospital ownership.

We improve on prior analyses in three areas. First, we evaluate both the effects of change in nurse staffing and change in quality of care on the change in financial performance by controlling for output quantities, input prices, market conditions, population characteristics, and hospital factors during the period of 1990 through 1995. Second, we improve on the empirical rigor of previous studies by using panel data as well as a dynamic econometric model rather than a static cross-sectional model, and by controlling for simultaneity between financial performance, staffing, and number of inpatient days.

Literature Review

Despite its importance, research examining the relationship of nurse staffing and hospital financial performance is descriptive in nature and relatively sparse. Flood and Diers[10] compared two nursing units and found that the short-staffed unit had higher costs and patient LOS. Glandon *et al.*,[11] found that nursing units with higher levels of registered nurse (RN) staff also had higher nursing expenses. Bloom, Alexander, and Nuchols[12] found that a richer nursing skill mix (*i.e.*, a higher ratio of RNs to total nursing staff) was not associated either with personnel costs per adjusted admission or with total operating costs per adjusted admission. In terms of total hospital FTEs, Hadley *et al.*,[13] found that declining staffing levels were associated with decreasing costs.

More recently, Spetz, Seago, and Coffman[14] examined the potential costs of implementing minimum nurse staffing ratios in California and found that, under various scenarios (depending on the actual ratios mandated, the size of the hospital, and the hospital's current staffing level), the statewide average cost of new staffing per hospital would be between $198,880 and $1,311,946. Revised estimates by Spetz[15] based on newly published final staffing ratios, predict that, on average, implementing these ratios will cost hospitals $217,000 per year. These estimates are undoubtedly biased downward because they assume that RN wages will not change from their 1999 to 2000 levels. With some evidence that nurses are receiving pay increases double the inflation rate,[16] these figures demonstrate the stark reality of the tight coupling between nurse staffing and hospital costs. However, these prior studies do not directly address the change in level of staffing and the change in financial performance. Thus, one of the research aims of this study is to implement an estimation technique that examines changes over time in nurse staffing and changes in costs and operating profit margin, while controlling for simultaneity between financial performance and staffing levels.

In terms of quality of care and costs, the prior empirical findings have been mixed. In a study of 12 diagnosis-related groups (DRGs), Garber *et al.*,[17] found no relationship between quality of care measures and costs for 9 out of 12 DRGs. For three DRGs, they found higher costs of patient care resulted in higher quality of care. Fleming[18] followed the Granneman *et al.*,[19] cost function and found that quality of care measured by readmission index had a convex linear relationship with cost. That is, improved outcomes were seen in hospitals with average levels of quality as costs fell, while outcomes worsened in hospitals with either very low or high levels of quality as costs fell. Schultz *et al.*,[20] conducted a cross-sectional analysis of acute-care hospitals in California to examine the relationship between financial factors and the rate of mortality caused by acute-myocardial infarction (AMI). They found operating expenses per day had a positive relationship with AMI mortality.

In terms of quality of care and profits, Epstein *et al.*,[21] measured quality of care by readmission rate and LOS and found lower profits associated with higher readmission rates and that shorter LOS was related to higher profits. However, Harkey and Vraciu[22] as well as Cleverley and Harvey[23] found a positive relationship between quality and profits.

A second research aim of this study is to evaluate the effect of changes in quality of care on financial performance. Unlike previous studies, however, our empirical specification assumes that quality of care in year $t-1$ affects financial performance in year t. There are two reasons for this approach. First, to assume that financial performance depends on current quality requires the assumption that we could estimate how the financial performance changes when the nurse staffing changes, *holding quality of care constant*. Previous work on quality of care and staffing levels indicates that quality of care is affected by staffing levels.[24] Therefore, such an assumption is invalid. Second, perceptions of quality may affect choice of hospital and hence financial performance, but the effect is not instantaneous. Hence, we assume that the lagged value of quality of care affects the current value of financial performance.[25]

Since managed care penetration and hospital ownership differences may influence quality of care, nurse staffing, and financial performance, the final research aim of this study is to examine whether the structure of the relationship differs by ownership or the level of managed care penetration. Several studies have examined the effect of managed care penetration and competition on hospital financial performance. Reduction in hospitals costs and greater cost savings have been reported in markets with large health maintenance organization (HMO) market share and growth in the number of HMOs, although high HMO penetration did not lower cost growth in highly concentrated markets.[26] In markets with fewer hospital competitors and high HMO penetration, Connor, Feldman, and Dowd[27] found that hospitals had higher prices and costs, while in markets with a greater number of hospital competitors and high HMO penetration, growth in hospital costs was constrained.

The influence of managed care penetration on nurse staffing is unclear. Evaluating Western New York hospital data, Brewer and Frazier[28] found that the presence of managed care did not affect the level of RN staffing. Using data from California, Spetz[29] found that HMO penetration also had no affect on hours worked by RNs, but did have a significant negative effect on hours worked by LPNs and aides. Buerhaus and Staiger[30] concluded that employment growth and earnings for RNs in the hospital sector had slowed significantly in states with high managed care enrollment.

Finally, there are major policy concerns about the effect of ownership status on quality of care and costs. McClellan and Staiger[31] found that, on average, for-profits had higher mortality among elderly patients with heart disease, but much of that difference was associated with their location. More recently, Sloan, Picone, and Taylor *et al.*,[32] found that Medicare payments were higher for patients admitted to for-profit hospitals than those admitted to nonprofit facilities while no quality differences (measured in terms of survival, changes in functional and cognitive status, and living arrangements) were found. Using the ratio of actual LOS divided by expected LOS as a measure of quality of care, Mark, Harless, McCue, and Xu[33] found that LOS ratio at nonprofits changed more slowly over time than LOS ratio at for-profits. They also found that increases in the proportion of Medicare and Medicaid discharges decreased the LOS ratio for nonprofits while adding HMOs in a market decreased LOS ratio for for-profits. In terms of hospital ownership and costs/profits, studies have found that freestanding for-profits incur higher operating costs,[34] higher administrative costs,[35] and greater variation in profits.[36]

Overall, these prior empirical studies reflect competing measures and different levels of analysis in examining the relationship between nurse staffing, quality of care, and financial performance. By using panel data over time, we fill the gap in the literature by evaluating the effects of change in nursing staffing and quality of care on change in financial performance.

Methodology

Sample. We sampled 422 hospitals that provided data for the years 1990 through 1995 from the Healthcare Cost and Utilization Project (HCUP) National Inpatient Sample (NIS). These 422 hospitals represent 49 percent of the HCUP base year sample and are located in 11 states (Arizona, Colorado, Florida, Illinois, Iowa, Massachusetts, New Jersey, Oregon, Pennsylvania, Washington, and Wisconsin). Because of the inability to match hospitals across all data sets, we eliminated six hospitals. Two more hospitals were eliminated because some data were provided only for a system rather than an individual hospital, and two others were dropped because revenue information was missing from all Centers for Medicare and Medicaid Services (CMS) files. To assure stability in our quality measure, we excluded 230 observations from 51 hospitals in years in which they had 15 or fewer expected mortalities. An additional 16 observations from 7 hospitals were excluded because of staffing outliers.[37] Finally, there were operating margin values for eight hospitals that were so extreme (*e.g.*, operating margin less than –80 percent); they were also excluded as outliers.

Data. We accessed data from the following sources: the area resource files, American Hospital Association (AHA) annual survey, CMS (formerly HCFA) minimum cost and capital file, CMS provider of services file, CMS case mix index file, CMS Online Survey Certification and Reporting (OSCAR) system files, and HCUP files.

In the CMS minimum cost and capital files, most hospitals had reporting periods different than calendar years (and some hospitals had reporting periods covering a period less than 365 days). To match appropriately data from CMS reports and calendar year data on quality of care and other variables, we converted CMS data to calendar year equivalent data using weighted averages with the weights depending on the number of days falling in a particular reporting period and the number of days covered by the report.[38]

Empirical Methods

In this study, we follow the basic empirical specification of prior cost and profit studies[39] and make several extensions to allow an analysis of the effect of staffing and quality of care on operating margin and on operating expenses. These prior studies used a production-theoretic framework to specify their expense and profit functions. Under this specification, expense and profits are functions of hospital outputs, the price of inputs, as well as hospital, market, and demographic factors. To this standard model we make four changes. First, we include the additional variables of RN, licensed practical nurse (LPN), non-nurse staffing, and the lagged value of the mortality ratio (actual mortalities/expected mortalities) to measure quality of care. Second, recognizing that staffing levels and patient LOS may be simultaneously determined with financial performance, we apply instrumental variable estimation. Third, to avoid omitted variable bias caused by (unmeasured) hospital-specific effects, we use panel data. Fourth, the model proposed is dynamic rather than static so that a hospital's past history may influence current financial performance through the lagged value of the dependent variable. Variable definitions and sources of data are displayed in Figure 1. The measures of the dependent and independent variables are presented here.

Profit and Expense Variables

Two primary financial performance measures, operating margin and the natural log

Figure 1. Variable Definitions, Property, and Sources of Data

Variable	Definition and Property*	Source of Data
Financial Performance		
Operating expense	Total hospital operating expense	HCFA cost and capital files
Operating margin	(Net patient revenue–operating expense)/net patient revenue	HCFA cost and capital files
Hospital Output		
Inpatient days	Total hospital inpatient days	HCFA cost and capital files
Discharges	Total hospital discharges	HCFA cost and capital files
Outpatient visits	Total hospital outpatient visits	AHA Survey
Input Price		
Input price	Average payroll costs per FTE	AHA
Hospital Characteristics		
Case mix index	Complexity of Medicare cases treated	HCFA case mix index files
High-tech services (Saidin index)	High-tech services provided	AHA survey
Payer mix	Medicare + Medicaid discharges/total discharges	HCFA cost and capital files
Beds	Number of open and operating beds	AHA survey
Nonprofit ownership	1 = not-for-profit 0 otherwise	HCFA cost and capital files
For-profit ownership	1 = for profit 0 otherwise	HCFA cost and capital files
Location	Within an MSA or not	HCFA cost and capital files
System affiliation	System affiliated or not	AHA survey
Market Characteristics		
HSA hospital use	Inpatient days/1,000 HSA population	Area resource files
Herfindahl index	Sum of squared market shares in an HSA	AHA survey
Number of HMOs	Number of HMOs in an HSA	Wholey/InterStudy
HMO penetration	% HMO enrollment as % of total HSA population	Wholey/InterStudy
Population Characteristics		
Variable and Source*	***Definition***	***Source***
Unemployment rate (1)	HSA unemployment rate	ARF
Per capita income (1)	HSA per capita income	ARF
Population	HSA population	ARF
Percent over age 65 (1)	HSA population over age 65	ARF
Staffing		
RN FTES	RN FTEs staffing	AHA survey, OSCAR files
LPN FTES	LPN FTEs staffing	AHA survey, OSCAR files
Non-nurse FTES	Non-nurse FTEs staffing	AHA survey, OSCAR files
Quality Measures		
Mortality	Observed/expected mortality	HCUP, MEDSTAT

of operating expenses, are used in this study. To measure the profits generated from operations, we employ the operating margin ratio, which measures the income earned from patient operations, and is defined as operating income as a proportion of net patient revenues.[40] Inadequate operating profits may hamper hospitals' ability to generate sufficient funds to pay for health care services and accumulate the earnings to provide new services and finance plant and equipment. Since hospitals can affect their operating profits by changing their operating expenses, we provide complementary results for a model that uses the natural log of operating expenses as a dependent variable. To normalize the operating expense measure across hospitals of varying size, we use the natural log of the operating expenses (measured in constant 1990 dollars) so that coefficient estimates measure the effect of a one-unit change in a regressor on the proportional change in operating expenses.

Hospital Output

Consistent with the standard model of cost and profit, we include several measures of hospital output to account for the (nonlinear) effect of inpatient and outpatient volume on financial performance: total hospital discharges, inpatient days, outpatient visits, and their squares, and cross-products.[41] Previous cross-sectional studies assumed that these output measures are exogenous.[42] Although we also assume that discharges and outpatient visits are exogenous, we weaken the assumption concerning inpatient days by allowing for inpatient days and financial performance to be simultaneously determined. Given discharges, average LOS is determined by inpatient days. Hence, treating inpatient days as endogenous is equivalent to treating average LOS as endogenous.

Input Price

The CMS wage index is routinely employed to measure variation in hospital wages in a given year, but is not a valid measure of changes in wages for a given hospital over time. To construct a measure of average wages that is valid over time, we use AHA data on total FTEs and total payroll expenses at acute care hospitals to calculate the average cost per FTE in the geographic area where a hospital is located.[43] We exclude all hospitals with a long-term care facility and hospitals that had extraordinarily large percentage changes (greater than 100 percent or less than –50 percent) in total FTEs, RN FTEs, and total payroll. We then estimate the payroll cost per FTE had hospitals maintained the same percentage of RN FTEs to total FTEs and LPN FTEs to total FTEs as they did in 1990 (the first year of the sample). The payroll cost per FTE is assigned to a particular hospital based on whether it was in or out of a metropolitan statistical area (MSA). For hospitals in an MSA, the payroll cost per FTE is assigned based on the average cost for hospitals in the MSA.[44] For hospitals outside of an MSA, the payroll cost per FTE is assigned based on the state average for hospitals outside of MSAs. Finally, we use the consumer price index to express cost per FTE in constant 1990 dollars.

Hospital Factors

Several hospital operational and structural measures are employed. In contrast to a simple approach of counting the number of high technology services offered, which assumes that each extra service contributes the same "amount" of additional complexity, this study measured high technology services using a "Saidin index."[45] The index measures the weighted sum of the number of technologies and services available in a hospital, with the weights being the proportion of hospitals in the country that do *not* possess the technology or service. As a result, the index increases more when a hospital adds technologies that are relatively rare than with the addition of technologies that are commonplace.[46] Payer mix measures the proportion of discharges that are Medicare and Medicaid. Lillie Blanton *et al.*,[47] found that a larger proportion of Medicare and Medicaid patients served by a hospital indicated greater likelihood of hospital closure. Following Alexander and Morrisey,[48] we use bed size to measure capacity. Hospital characteristics also include beds, ownership, location, and system affiliation measures. Finally, we control intensity of resource use by including the Medicare case mix index. This index measures the complexity of a hospital's Medicare cases and often is used by studies as a proxy for overall hospital case mix.[49]

Market Factors

We use the health service areas (HSAs) approach developed by Makuc *et al.*,[50] in which counties are aggregated into geographic regions based on flows of inpatient hospital admissions to define the market. By definition, each HSA includes at least one hospital. Wholey[51] suggests that aggregation to the HSA level corrects for over- or under-allocation errors in county-level aggregation. Hospital concentration is measured by the Herfindahl index and calculated from hospital admission data from AHA. HMO penetration is measured as the percentage of an HSA's population enrolled in HMOs, while the number of HMOs represents the number of HMOs competing in the market.[52]

Population Factors

Other socioeconomic factors can influence the demand and the ability to pay for hospital services, which, in turn, could affect hospital profits. These measures include proportion of population over 65 in the HSA, unemployment rate in the HSA, natural log of population in the HSA, and natural log of per capita income in the HSA.[53] The natural log transformation is used to allow for the existence of a nonlinear relationship, but do so parsimoniously.

Hospital Staffing

Data on RN FTEs, LPN FTEs, and nonnurse FTEs (total FTEs minus RN FTEs and LPN FTEs) were derived from the AHA database. In developing the staffing measures, several complications arose because of changes in the AHA database. From 1990 through 1993, the AHA annual survey required hospitals to report staffing separately by hospital unit and nursing home/long-term care unit. After 1993, the reporting was for the total facility only. To calculate hospital RN and LPN FTEs, we use the OSCAR system for nursing homes, which allows us to subtract nursing home staffing from total facility staffing to arrive at hospital staffing. In our empirical specification, staffing levels enter using the natural log transformation. This allows us to properly analyze hospitals with quite different staffing levels since the natural log transformation allows us to make comparisons based on a proportional change in FTEs (*e.g.*, the change in operating margin per 1 percent increase in RN FTEs). Further, recognizing that the effect of increasing RN FTEs on operating margin may depend on the level of LPN FTEs (and vice versa), we also include an interaction variable equal to the product of the natural log of RN FTEs and the natural log of LPN FTEs.

Quality of Care

To measure quality of care, we include a risk-adjusted mortality measure—the mortality ratio. Risk adjustment is conducted by using Medstat's disease staging methodology[54] and is selected over other available severity adjustment systems because of its demonstrated performance[55] and the ability to apply it directly to HCUP data. These data were used to compute a predicted probability of death and the probabilities were summed over a hospital's discharges to produce an estimate of the number of deaths "expected" for that hospital. The mortality ratio then combines information about observed and expected quality of care in a ratio. The standardized mortality ratio equals observed in-hospital deaths divided by the expected number of in-hospital deaths. Thus, a mortality ratio greater than 1.0 indicates that the actual number of deaths exceeds the expected number, while a mortality ratio less than 1.0 indicates that the actual number of deaths was less than expected.

Empirical Specification and Analytic Approach

We apply two analytical approaches that are critical to assessing the effects of changes in nurse staffing and change in quality of care on the change in financial performance. First, we control for two types of omitted variable bias so that our coefficient estimates should come much closer to the actual policy-impact of changes in the regressors on financial performance. Previous research[56] that evaluated nurse staffing skill mix and financial performance used cross-sectional data. This type of analysis assumes that hospitals are *homogeneous* and differ only by the explanatory variables included in the model. There are, however, likely to be *unmeasured* attributes (*e.g.*, goodwill in the community, favorable location) that affect financial performance, but are omitted from the models.

When these hospital-specific traits are correlated with the explanatory variables in the model, their exclusion leads to omitted-variable bias. With panel data, we can control for such time-invariant unmeasured attributes by including a separate intercept for each hospital.[57] Furthermore, any study using cross-sectional data necessarily assumes a static model. The current value of the dependent variable is explained by the contemporaneous values of the explanatory variables in the model. The implicit assumption is that the dependent variable adjusts instantaneously to changes in the values of the explanatory variables in the current period. A more complete empirical model should recognize that contemporary circumstances and the history of circumstances determine current financial performance. A parsimonious way to generalize from a static model to a dynamic model is to account for the influence of the past through the lagged value of the dependent variable. The coefficient for the lagged dependent variable indicates the extent to which a shock to financial performance in year t is transmitted to financial performance in subsequent years.

Second, we recognize that the level of staffing and the number of inpatient days may be simultaneously determined with financial performance. As explained previously, since average LOS equals inpatient days divided by discharges, if discharges are exogenous, then control over inpatient days is equivalent to control over average LOS. Previous studies[58] of nurse staffing and financial performance assumed that staffing was exogenous to financial performance. These models produce biased estimates of the effect of staffing, however, if staffing level depends on financial performance. For example, suppose there is a negative shock to operating margin (or, equivalently in this example, a positive shock to operating expenses) in a given year. If the decrease in profit results in staffing cuts in the same year, then financial performance and staffing are simultaneously determined. An estimation technique that does not account for this simultaneity (*e.g.*, use of ordinary least squares[59]) will result in estimates of the effects of staffing on financial performance that are biased toward zero. Similarly, financial performance and number of inpatient days may be simultaneously determined. Instrumental variable estimation using the information in the first differences of the variables allows us to obtain consistent estimates of the parameters in our empirical model.

Recall that to control for hospital heterogeneity, we include a separate intercept for each hospital. A standard method for estimating such a model is the "within group" or "fixed effects" estimator wherein the time-series mean for each cross-sectional unit is subtracted from each variable (this is equivalent to ordinary least squares with dummy variables for each hospital). This estimation method, however, requires that all regressors be strictly exogenous. That is, uncorrelated with the error term in *all* time periods. However, the regressors staffing and inpatient days are *endogenous* rather than exogenous. The regressor—the lagged value of the dependent variable—is also not strictly exogenous since its value is clearly correlated with the lagged value of the error term (*i.e.*, the regressor $y_{i,t-1}$, the lagged dependent variable for hospital i in year $t-1$, is correlated with the error term $u_{i,t-1}$). One can, however, eliminate the hospital specific intercepts by taking first differences. That is, instead of using the level values (*i.e.*, $y_{i,t}$ and $x_{i,t}$), we consider the change in the variables from one year to the next (*i.e.*, $(y_{i,t} - y_{i,t-1})$ and $(x_{i,t} - x_{i,t-1})$). Importantly, if we analyze the model in first differences, then we can apply instrumental variable estimation using as instruments the twice-lagged values of the regressors suspected of being simultaneous with financial performance, as well as the levels of the other regressors that we assume to be strictly exogenous.[60]

To illustrate the choice of instruments, take the natural log of RN FTEs as an example. Our first difference model specifies that (operating $\text{margin}_{i,t}$ – operating $\text{margin}_{i,t-1}$) depends on, among other first differenced regressors, (ln (RN $\text{FTEs}_{i,t}$) – ln (RN $\text{FTEs}_{i,t-1}$)), and that the error term is $(u_{i,t} - u_{i,t-1})$. Since ln (RN $\text{FTEs}_{i,t}$) is correlated with $u_{i,t}$, we cannot apply OLS to the first differenced model, but we can apply instrumental variable estimation using ln (RN $\text{FTEs}_{i,t-2}$) as an instrument for (ln (RN $\text{FTEs}_{i,t}$) – ln (RN $\text{FTEs}_{i,t-1}$)).[61] Further, the first differenced equation includes first differences of other variables that are assumed to be strictly exogenous, such as (HMO $\text{penetration}_{i,t}$ – HMO $\text{penetration}_{i,t-1}$). Since the level value HMO $\text{penetration}_{i,t}$ and similar strictly exogenous variables are *not* included as regressors, the levels of such variables also may serve as instruments. Finally, we suspect that the variance of the error term is inverse to the size of the hospital, and so apply weighted least squares, as in Bazzoli *et al.*,[62] with weights equal to the square root of average bed size.

Results

Figure 2 presents descriptive statistics on all variables included in the model for the six-year period. Note that in our sample, on average, RN FTEs make up 24.24 percent of total hospital FTEs, but the range of values goes from 8.2 percent to 42.99 percent.[63]

Figure 3 presents results for both the operating margin and natural log of operating expense models for the base specification (column 1 and column 3) and the final specification (column 2 and column 4). Figure 3 indicates the coefficients (and standard errors in parentheses under the coefficients) for the particular variables of interest in this study (the lagged dependent variable, staffing variables, and mortality ratio) and, with the exception of the variables measuring output, all other variables that were statistically significant in at least one of the models. The other regressors included in the models are listed in the footnote to the table and the complete set of coefficient estimates are provided in the Appendix. F-statistics testing the overall significance of each of the models also are included at the bottom of Figure 3.

The base and final specifications in Figure 3 reflect the specification tests undertaken to detect whether structural differences existed in the operating margin model

Figure 2. Summary Statistics for Sample Hospitals, 1990–1995

Variable	Mean	Standard Deviation	Minimum	Maximum
Operating margin	–0.09	10.59	–70.81	45.31
Operating expenses[a]	53,635	67,876	1,535	612,924
RN FTEs	193.5	250.3	5.1	2,425.5
LPN FTEs	29.3	41.0	0.0	800.0
Non-nurse FTEs	559.3	692.1	28.5	6,206.1
Mortality ratio	1.07	0.23	0.33	2.01
Payer mix percentage	55.80	13.78	9.28	100
System	0.46	0.49	0	1
Inpatient days (in thousands)	39.91	45.83	1.26	395.43
Discharges (in thousands)	7.19	7.63	0.23	55.41
Outpatient visits (in thousands)	75.85	81.02	0	892.88
Case mix index	1.29	0.21	0.87	2.08
Saidin index of high-tech services	3.29	2.33	0.26	10.81
Number of beds	194.30	177.23	19	1311
Public	0.17	0.38	0	1
Profit	0.14	0.34	0	1
Number of HMOs in HSA	7.41	6.10	0	23.64
HMO penetration in HSA	16.80	12.52	0	63.83
Herfindahl index in HSA	18.29	16.13	1.16	100
Average payroll cost per FTE in HSA[b]	28,685.9	4,800.8	18,191.1	40,226.8
Percentage of population 65 or older in HSA	15.07	4.25	8.71	32.42
Unemployment rate in HSA	6.58	2.20	2.52	16.63
Population in HSA (in thousands)	1,647	2,516	11	11,703
Per capita income in HSA[b]	19,621	3,989	11,216	36,074
Inpatient days per 1,000 population in HSA	875.5	311.6	152.7	2,331.3

[a]In thousands of constant 1990 dollars.
[b]In constant 1990 dollars.

Figure 3. Results for Operating Margin and Natural Log of Operating Expenses[a]

Regressors	Base (1) opm[b]	Final (2) opm[c]	Base (3) ln (operexp)[b]	Final (4) ln (operexp)[c]
Dependent variable$_{t-1}$	0.280***	0.280**	0.289*	0.222*
	(0.071)	(0.090)	(0.124)	(0.102)
(low HMO penetration) × (dependent variable$_{t-1}$)		-0.337*		0.104
		(0.137)		(0.127)
ln (RN FTEs$_t$)	-1.099	-12.519	0.324**	0.253*
	(12.312)	(12.237)	(0.114)	(0.106)
ln (LPN FTEs$_t$)	-18.549	-28.017	0.040	-0.038
	(18.570)	(17.043)	(0.171)	(0.149)
ln (RN FTEs$_t$) × ln (LPN FTEs$_t$)	3.252	4.580	-0.012	-0.002
	(2.966)	(2.734)	(0.027)	(0.024)
ln (non-nurse FTEs$_t$)	-13.067	-20.934*	0.174	0.181*
	(10.275)	(9.743)	(0.090)	(0.086)
Mortality ratio$_{t-1}$	-0.521	-0.464	-0.026	-0.019
	(2.841)	(2.774)	(0.026)	(0.025)
Payer mix$_t$	-0.093	0.233*	0.001	0.001
	(0.068)	(0.101)	(0.001)	(0.001)
(high bed quartile) × payer mix$_t$		-0.593***		0.001
		(0.137)		(0.001)
System$_t$	-3.596*	-3.240*	-0.001	0.005
	(1.588)	(1.600)	(0.014)	(0.013)
Number of HMOs in HSA$_t$	-0.489	-0.302	-0.001	-2.5E-4
	(0.347)	(0.347)	(0.003)	(0.003)
HMO penetration in HSA$_t$	-0.381*	-0.308*	-0.001	-0.001
	(0.155)	(0.153)	(0.001)	(0.001)
(number of HMOs$_t$) × (HMO penetration$_t$)	0.029*	0.023	2.2E-5	6.8E-7
	(0.012)	(0.012)	(1.1E-4)	(1.0E-4)
ln (payroll cost per FTE in HSA$_t$)	-18.469	-21.111	0.230*	0.246*
	(11.766)	(11.631)	(0.114)	(0.107)
Number of observations	1235	1235	1235	1235
F-test of significance of overall model	4.87	4.63	7.83	8.20
degrees of freedom	(36;1,199)	(43;1,192)	(36;1,199)	(43;1,192)
p-value	1.3E-18	6.8E-20	4.1E-35	3.6E-41

[a]Standard errors in parentheses under coefficients.
[b]Other variables included in the model (but not reported here): squares and cross-products of inpatient days, discharges, and outpatient visits; case mix index, Saidin index of high-tech services, number of hospital beds, Hirfindahl index in HSA, percentage of population age 65 and older in population in HSA, unemployment rate in HSA, natural log of per capita income in HSA, natural log of inpatient days per 1,000 population in HSA, and dummy variables for ownership (public hospitals, for-profit hospitals), and year (1992 to 1995).
[c]In addition to the control variables listed previously, this model included five additional control variables representing the interaction of a dummy variable for the set of hospitals in the highest bed quartile interacted with each of inpatient days, inpatient days,^{2} discharges, discharges,^{2} and (inpatient days) × (discharges).
*Indicates coefficient significant at 0.05 level.
**Significant at 0.01 level.
***Significant at 0.001 level.

by hospital ownership, location in an MSA, HMO penetration, and hospital size.[64] For example, in the test for structural differences by ownership, we estimated a model that allowed different coefficients for for-profit hospitals, and then tested the null hypothesis that these different coefficients for for-profit hospitals were all equal to zero (*i.e.*, that there were no structural differences for the model of operating margin between for-profit and not-for-profit hospitals).

In the test for structural differences by ownership, we could easily accept the null hypothesis that for-profit hospitals had the same coefficients as not-for-profit hospitals ($F_{34,1166} = 0.58$, $p = 0.97$). There also were no structural differences for hospital located in an MSA ($F_{34,1165} = 1.00$, $p = 0.47$).[65] We did, however, observe structural differences for hospitals in low HMO penetration areas and by hospital size. To test for differences by HMO penetration, we separated hospitals into quartiles based on the average HMO penetration in the HSA where the hospital was located. No structural differences ($F_{34,864} = 1.13$, $p = 0.28$) were evident between hospitals in the highest HMO penetration quartile (average HMO penetration ≥ 27.85 percent) and hospitals in the middle two quartiles of HMO penetration (average HMO penetration between 7.236 percent and 27.85 percent). However, a structural difference did exist between hospitals in the lowest HMO penetration quartile and hospitals in the highest three quartiles of HMO penetration ($F_{34,1165} = 1.84$, $p = 0.002$). The source of this difference was the coefficient for the lagged dependent variable, which was significantly lower for hospitals in the lowest quartile of HMO penetration. This variable aside, one can easily accept the hypothesis that the coefficients for all other variables are the same between the lowest HMO penetration quartile and the top three HMO penetration quartiles ($F_{33,1165} = 1.20$, $p = 0.20$). Hence, our final specification allows for a different coefficient for the lagged value of operating margin for hospitals in the lowest HMO penetration quartile.

Structural differences ($F_{34,1164} = 1.69$, $p = 0.008$) were also detected for hospitals in the largest quartile of bed size (average number of beds ≥ 256). The structural differences by size involved a coefficient for the payer mix percentage and the coefficients for inpatient days and discharges and their squares and cross product. We can easily accept the null hypothesis that all the other coefficients are the same between hospitals in the largest size quartile and other hospitals ($F_{28,1164} = 0.95$, $p = 0.54$).

Consider now the findings concerning the effect of changing staffing FTEs on operating margin and natural log of operating expense. In the final specification, we find that only non-nurse FTEs have a statistically significant effect on both operating margin, as shown in Figure 3 (column 2) and operating expenses (column 4). For operating margin, the coefficient for ln (non-nurse FTEs) is −20.934 ($p = 0.03$) indicating that a 1 percent increase in non-nurse FTEs decreases operating margin by $(20.93/100) \approx .21$ percentage points. In the final specification for natural log of operating expenses, the coefficient is 0.181 indicating that a 1 percent increase in non-nurse FTEs increases operating expenses 0.18 percent. For RN FTEs, however, we find a statistically significant effect on operating expense, but no statistically significant effect on operating margin. In the final specification, the coefficient for ln (RN FTEs) in the operating expense model indicates that a 1 percent increase in RN FTEs increases operating expenses by approximately 0.25 percent ($p = 0.014$).[66]

In addition to adding staffing variables and the lagged dependent variable to the standard model of operating expense and profit, we also considered the impact of the lagged value of quality of patient care as measured by the mortality ratio. In the operating margin model, the sign of the coefficient suggests that increases in the mortality ratio in year $t-1$ decreases operating margin in year t, but the effect is not measured with any precision.

The coefficient for the system affiliation variable is statistically significant and suggests that hospitals that affiliate with systems suffer a decrease in operating margin of 3.24 percentage points. This result serves well as an illustration of the information that is used to construct the coefficient estimates. In our sample, the mean operating margin for hospitals affiliated with a system (2.06) is higher than the mean operating margin for unaffiliated hospitals (−1.1), and this same pattern holds in each year of the sample. Given that the level of operating margin is higher for affiliated hospitals, how could we obtain a coefficient for a system that is negative and statistically significant? Recall that we are regressing *change* in financial performance on the *change* in the regressors. Hence, for the system dummy variable, the coefficient estimate uses only the information on hospitals that switched to being affiliated with a system (or vice versa).

We find that HMO penetration decreases operating margin, but the size of the effect depends on the number of HMOs in the HSA. The effect is larger when the number of HMOs is smaller. When the number of HMOs is at the 25th percentile value for our sample (2.31 HMOs), the marginal effect is –0.25 percentage points operating margin per one percentage point increase in HMO penetration. When the number of HMOs is at the median value in the sample (5.79), the marginal effect is –0.17 percentage points operating margin. When the number of HMOs is at the 75th percentile value (10.98), the marginal effect is only −0.05 and the marginal effect is zero—just below the 90th percentile value of the number of HMOs.

Finally, note that the coefficient for the natural log of payroll cost per FTE in the operating expense model is 0.246 ($p = 0.012$) suggesting that a 1 percent increase in real payroll cost per FTE increases operating expenses 0.25 percent. The coefficient for this variable in the operating margin model is –21.11 ($p = 0.07$), suggesting that a 1 percent increase in real payroll cost per FTE decreases operating margin 0.21 percentage points.

The impact of the structural differences is examined by comparing several coefficients in the base and final specifications. For example, in the base specification for operating margin, the coefficient for payer mix percentage is negative, but statistically insignificant. In the final specification (see column 2), however, we find two effects, opposite in sign, for hospitals of different size. In the smallest three quartiles of hospital size, our coefficient estimate suggests that operating margin increases by 0.23 percentage points per one percentage point increase in payer mix (p-value = 0.02).[67] In the largest quartile of hospital size, we find a *difference* in the coefficient for payer mix of –0.593 (p = 1.5E-5). Hence, for hospitals in the largest size quartile, our model suggests that operating margin decreases by 0.36 ($p = 0.233 - 0.593$) percentage points per one percentage point increase in payer mix. In the natural log of operating expense model, the coefficients for payer mix are not statistically significant suggesting that the effect of payer mix on operating margin comes entirely or almost entirely from patient revenues.[68]

Structural differences for hospitals in the lowest quartile of HMO penetration also are immediately apparent in the difference in coefficients for the lagged value of the dependent variable (which indicates the dynamic effect—the influence of the past on present value of the dependent variable). For hospitals in the highest three quartiles of HMO penetration, the coefficient on the lagged value of operating margin (see column 2) is 0.280 indicating that a change in operating margin is transmitted from one period to the next. Suppose there is a shock to operating margin in year t and operating margin falls by one percentage point. Our model suggests that operating margin in year t + 1 will be 0.28 percentage points lower. However, this dynamic effect is absent for hospitals in the lowest quartile of HMO penetration where the coefficient on the lagged dependent variable is .337 ($p = 0.014$) *lower* than for other hospitals, so the coefficient on the lagged value of operating margin is nearly zero ($0.28 - 0.337 = -0.058$). Note that this structural difference for hospitals in the lowest HMO penetration quartile appears to be limited to operating margin. In the operating expense model, the coefficients indicating the dynamic effect are of quite similar size to the coefficient in the operating margin model, but there is no significant difference in the coefficient for hospitals in the lowest HMO penetration quartile.

Discussion

Using longitudinal data from 1990 through 1995, our study applied more careful econometric techniques to evaluate whether changes in nurse staffing and in quality of care affected hospitals' financial performance. We found that hospitals experienced increased operating costs when they increased RN staffing. However, we found no statistically significant effect of RN staffing on profit margins. At a minimum, these results call into question the idea that a route to greater profitability is through cuts in RN staffing.

The effect of change in nurse staffing on operating costs and margin can be calculated. Consider a hospital with the median level of 8,900 adjusted discharges (case mix adjusted). Suppose that this hospital had RN and LPN staffing at the 25th percentile level for our sample: 10.45 RN FTEs per 1,000 adjusted discharges (93 RN FTEs) and 1.14 LPN FTEs per 1,000 adjusted discharges (10 LPN FTEs). Our coefficient estimates imply that if such a hospital were to increase RN FTEs to 112 to reach the median level of RN FTEs per 1,000 adjusted discharges (12.60), operating margin would fall by 0.37 percentage points while operating expenses would increase 4.6 percent. Alternately, suppose that RN and LPN staffing were at the median values (19 LPN FTEs, 2.09 LPN FTEs per 1,000 adjusted discharges) and the hospital increased RN FTEs to 135 so as to reach the 75th percentile level of RN FTEs per 1,000 adjusted discharges (15.17). Because the RN-LPN interaction term was quite small in the operating expenses model, it would still indicate a 4.6 percent increase in cost. For the operating margin model, when the RN-LPN interaction coefficient is taken into account, operating margin would increase by 0.18 percentage points. Although the effect of RN staffing on operating margin was not statistically significant, it is interesting to note that, for hospitals with less than the median number of LPN FTEs, adding RNs is predicted to decrease operating margin, while for hospitals with more than the median number of LPN FTEs, adding RNs is predicted to improve margins. In the operating expenses model, the effect of adding RNs on cost was slightly lower the greater the number of LPN FTEs. The pattern of coefficients in our models indicates that LPNs and RNs are complements, not substitutes. It also may be that in hospitals with fewer RN FTEs, turnover is high and overtime use is extensive—costs that are reduced when there are more RN FTEs. Nursing turnover is costly for hospitals. Current estimates are that it costs $46,000 on average to replace one medical/surgical nurse and about $64,000 to replace a critical care nurse. High turnover rates also may affect hospital profitability through higher average costs per discharge, and decreased return on assets.[69]

However, hospitals adding higher levels of administrative and operational support staff incurred higher operating expenses, as well as lower profits. This outcome suggests that hospitals were unable to generate higher revenues to cover higher costs from adding support staff and as a result experienced a decline in profits. We also found that the change in quality of care did not have a statistically significant effect on either costs or profits. Consequently, hospitals that expended resources to improve quality of care were unable to improve their profitability or decrease the operating costs.

In terms of the managed care market, we found no significant relationship between costs and HMO penetration; however, our findings did report decreasing operating profits for hospitals located in markets with increasing HMO penetration. This outcome confirms Thorpe, Seiber, and Florence's[70] findings that increasing penetration reduces hospital profits. Other studies reported similar significant effects of HMO penetration on other financial measures, but not profit. Connor, Feldman, and Dowd[71] found decreasing hospital revenues and costs for hospitals in markets with increasing HMO penetration. McCue, Clement, and Luke[72] assessed local hospital alliances and found lower revenues for markets with high HMO penetration. We also found that increasing the number of HMOs in markets with higher HMO penetration reduced this decline in hospitals profits. Thus, increasing competition among health plans may have allowed hospitals to negotiate favorable contracts.

During our study period, payments increased for both Medicare and Medicaid patients, and contributed to a rise in total hospital profit margins to 5.8 percent in 1995 from 3.6 percent in 1990.[73] Our results also found an increase in profit margins for hospitals serving a great proportion of government payers. However, profit margins declined for larger hospitals with an increasing proportion of Medicare and Medicaid patients. Evidently, lower profits among these larger facilities suggested that they were serving complex Medicare and Medicaid patients whose costs exceed these rising payments.

Finally, we reported declining profits for hospitals that were acquired by a multi-hospital system. Since we do not know the length of time that the acquired facility was part of the system, we presume that these facilities were newly acquired. This outcome supports the hospital acquisition literature[74] that found lower profits, or in some cases financial losses for newly acquired hospitals relative to non-acquired ones.

We incurred several limitations in terms of measurements and econometric concerns. First, in measuring nurse staffing, we were confronted with the switch by AHA in 1993 from hospital unit RNs to total facility RNs. Second, in developing our cost input price measure, we did not have specific salary data for RN, LPN, and other FTEs to develop a more valid measure. Instead, availability of data limits us to an aggregate cost per hospital FTE over all types of workers. Third, we controlled for possible endogeneity of staffing and inpatient days, but we assumed that other hospital variables are strictly exogenous (discharges, outpatient visits, Saidin technology index, etc.). Fourth, our operating margin ratio was highly variable, which may stem from different interpretations of accounting standards by different hospitals. Finally, we lacked a theoretical model indicating how to incorporate quality of care into a model of cost or profit. Our specification seems reasonable, but future theoretical work could help guide empirical work of this nature.

In terms of future research, the results of this study also provide a baseline view of how changes in nurse staffing and quality of care affect changes in financial performance prior to the substantial fiscal Medicare payment reform of the Balanced Budget Act of 1997 (BBA). Even with the restoration of $17 billion through the Medicare Balanced Budget Refinement Act of 1999 (BBRA), payment reductions in BBA had a significant effect on the financial performance of hospitals by reducing their total profit margins to 2.8 percent in 1999 from 6.1 percent in 1996. In the late 1990s, other financial pressures, such as losses on hospital-owned physician practices, increased pharmaceutical prices, inability to staff beds due to nursing and other staff shortages, and higher payroll costs threaten further the financial viability of acute care hospitals. Future research should assess how hospitals alter their nurse staffing and quality of care in reaction to these recent financial pressures.

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41. *See* G. Bazzoli *et al.*, and T. Granneman *et al.*, *supra* n.39.
42. *Id.*
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64. We performed these tests only for the operating margin model because operating margin was the measure of financial performance of greatest interest, but we provide estimates for the same specification using natural log of operating expense as the dependent variable so that we can ascertain whether changes in operating margin were driven by changes in operating expenses.
65. There is a difference in the denominator degrees of freedom in the F-statistic because when testing for structural differences for for-profit hospitals, we cannot also estimate a dummy variable for for-profit hospitals (as is included in the base specification). Hence, there is one more denominator degree of freedom in the test for structural differences by ownership than in other tests.
66. Note that our specification includes an interaction for RN and LPN FTEs, so the effect of increasing RN FTEs by 1 percent depends on the level of ln (LPN FTEs), but in the operating expense model the coefficient for the interaction term is so small that it only impacts the marginal effect at the third decimal point. Although the effect of RN FTEs is statistically insignificant in the operating margin model, the estimates in Figure 3 should be consistent estimates of the effects nevertheless. Moreover, the pattern of the marginal effects of RN staffing is of interest and so we describe briefly those effects. Our specification includes an interaction term so that the effect of an increase in RN FTEs depends on the level of ln (LPN FTEs). The consequence is that the marginal effect of a 1 percent increase in RN FTEs on operating margin changes from negative to positive as the level of ln (LPN FTEs) increases; at the 25th percentile value of LPN FTEs in our sample (8.2 LPN FTEs), we find that a 1 percent increase in RN FTEs decreases operating margin by 0.03 percentage points; at just below the median value of LPN FTEs (15.5) the marginal effect is zero, and at the 75th percentile value of LPN FTEs (34.8) the margin effect is positive–a 1 percent increase in RN FTEs increases operating margin 0.04 percentage points. Recall that we instrument for all the staffing variables and the consequence of instrumental variable estimation to get consistent estimates of the effects is higher standard errors of the coefficients. We note, however, that instrumental variable estimation is necessary since staffing is expected to be simultaneously determined with financial performance. In fact, if we were to assume that staffing was strictly exogenous and fail to instrument, then the coefficient estimates would indeed be biased toward zero. In the operating expense model, the coefficient for ln (non-nurse FTEs) would be less than half the size of the coefficient in Figure 3, and the coefficient for ln (RN FTEs) would be only one-fourth the size of the coefficient in Figure 3. In the operating margin model, the coefficient for ln (non-nurse FTEs) would be positive instead of negative, and the coefficient for ln (RN FTEs) would still be negative but only one-fifth the magnitude of the coefficient in Figure 3.
67. Recall that we measure the payer mix variable as a percentage, so, for example, the mean of payer mix is 55.80.
68. The differences in coefficients for hospitals in the highest bed quartile for the discharges, inpatient days, their squares, and cross product. The differences in these coefficients may be seen in the Appendix, which gives the complete set of coefficient estimates.
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Appendix

Results for Operating Margin and Natural Log of Operating Expenses: Complete Set of Coefficient Estimates[a]

(Version for appendix, published)

Regressors	(1) opm	(2) opm	(3) ln (operexp)	(4) ln (operexp)
Dependent $variable_{t-1}$	0.280*** (0.071)	0.280** (0.090)	0.289* (0.124)	0.222* (0.102)
(low HMO penetration) × (dependent $variable_{t-1}$)		−0.337* (0.137)		0.104 (0.127)
ln (RN $FTEs_t$)	−1.099 (12.312)	−12.519 (12.237)	0.324** (0.114)	0.253* (0.106)
ln (LPN $FTEs_t$)	−18.549 (18.570)	−28.017 (17.043)	0.040 (0.171)	−0.038 (0.149)
ln (RN $FTEs_t$) × ln (LPN $FTEs_t$)	3.252 (2.966)	4.580 (2.734)	−0.012 (0.027)	−0.002 (0.024)
ln (non-nurse $FTEs_t$)	−13.067 (10.275)	−20.934* (9.743)	0.174 (0.090)	0.181* (0.086)
Mortality $ratio_{t-1}$	−0.521 (2.841)	−0.464 (2.774)	−0.026 (0.026)	−0.019 (0.025)
$Payermx_t$	−0.093 (0.068)	0.233* (0.101)	0.001 (0.001)	0.001 (0.001)
(high bed quartile) × $payermx_t$		−0.593*** (0.137)		0.001 (0.001)
$System_t$	−3.596* (1.588)	−3.240* (1.600)	−0.001 (0.014)	0.005 (0.013)
Number of HMOs in HSA_t	−0.489 (0.347)	−0.302 (0.347)	−0.001 (0.003)	−2.5E-4 (0.003)
HMO penetration in HSA_t	−0.381* (0.155)	−0.308* (0.153)	−0.001 (0.001)	−0.001 (0.001)
(number of $HMOs_t$) × (HMO $penetration_t$)	0.029* (0.012)	0.023 (0.012)	2.2E-5 (1.1E-4)	6.8E-7 (1.0E-4)
ln (payroll cost per FTE in HSA_t)	−18.469 (11.766)	−21.111 (11.631)	0.230* (0.114)	0.246* (0.107)
Inpatient $days_t$	−0.328 (0.286)	−0.589 (1.180)	0.001 (0.003)	0.004 (0.010)
Inpatient $days_t^b$	0.001* (0.001)	0.010 (0.015)	−6.7E-6 (4.7E-6)	5.2E-5 (1.3E-4)
$Discharges_t$	0.072 (0.528)	8.088* (4.082)	0.012* (0.005)	0.093** (0.036)
$Discharges_t^b$	0.014 (0.011)	−0.234* (0.093)	−2.9E-5 (1.0E-4)	−0.002* (0.001)
Outpatient $visits_t$	−0.002 (0.036)	0.016 (0.037)	−1.5E-4 (3.3E-4)	−1.0E-4 (3.2E-4)
Outpatient $visits_t^b$	−6.6E-6 (5.7E-5)	−2.4E-5 (5.6E-5)	2.8E-7 (5.1E-7)	2.5E-7 (4.9E-7)
(inpatient $days_t$) × ($discharges_t$)	−0.007 (0.004)	−0.042 (0.084)	2.4E-5 (4.4E-5)	−0.001 (0.001)
(inpatient $days_t$) × (outpatient $visits_t$)	2.9E-5 (0.001)	2.4E-5 (0.001)	6.3E-6 (8.0E-6)	5.1E-6 (7.5E-6)
($discharges_t$) × (outpatient $visits_t$)	−0.001 (0.005)	−0.002 (0.005)	−3.8E-5 (4.3E-5)	−3.4E-5 (4.2E-5)
(high bed quartile) × (inpatient $days_t$)		0.224 (1.032)		−0.003 (0.009)
(high bed quartile) × (inpatient $days_t^b$)		−0.008 (0.015)		−5.9E-5 (1.3E-4)
(high bed quartile) × ($discharges_t$)		−8.859* (4.056)		−0.087* (0.035)
(high bed quartile) × ($discharges_t^b$)		0.262** (0.093)		0.002* (0.001)

continues

Appendix (Continued)

Regressors	(1) opm	(2) opm	(3) ln (operexp)	(4) ln (operexp)
(high bed quartile) × (inpatient days$_t$) × (discharges$_t$)		0.034 (0.084)		0.001 (0.001)
Case mix index$_t$	−4.896 (7.356)	−4.649 (7.273)	0.098 (0.067)	0.087 (0.065)
Saidin index of high-tech services$_t$	0.896 (0.492)	0.928 (0.482)	−0.001 (0.004)	−0.001 (0.004)
Number of beds$_t$	0.010 (0.014)	0.011 (0.014)	1.1E-4 (1.2E-4)	8.5E-5 (1.2E-4)
Public$_t$	18.835 (10.200)	15.073 (9.952)	−0.081 (0.100)	−0.075 (0.095)
Profit$_t$	−7.148 (5.304)	−10.157 (5.420)	−0.063 (0.048)	−0.079 (0.045)
Herfindahl index in HSA$_t$	−0.473 (0.365)	−0.440 (0.364)	0.001 (0.003)	0.001 (0.003)
Percentage older in HSA$_t$	0.103 (2.263)	1.094 (2.243)	0.004 (0.021)	0.001 (0.020)
Unemployment rate in HSA$_t$	−0.582 (0.621)	−0.870 (0.614)	−0.006 (0.006)	−0.007 (0.005)
ln (population in HSA$_t$)	−7.087 (37.836)	−6.930 (36.762)	−0.318 (0.342)	−0.450 (0.330)
ln (per capita income in HSA$_t$)	19.701 (16.255)	13.571 (16.025)	0.103 (0.151)	0.100 (0.143)
ln (inpatient days per 1,000 population in HSA$_t$)	11.671 (7.069)	12.820 (7.182)	−0.042 (0.064)	−0.058 (0.063)
1992	2.571* (1.285)	2.788* (1.287)	0.022 (0.015)	0.030* (0.013)
1993	1.749 (1.916)	2.544 (1.989)	0.025 (0.021)	0.035 (0.020)
1994	1.314 (2.707)	1.891 (2.902)	0.009 (0.030)	0.020 (0.029)
1995	2.717 (3.663)	2.837 (3.916)	0.025 (0.037)	0.033 (0.037)
Number of observations	1235	1235	1235	1235
F-test of significance of overall model	4.87	4.63	7.83	8.20
degrees of freedom	(36;1,199)	(43;1,192)	(36;1,199)	(43;1,192)
p-value	1.3E-18	6.8E-20	4.1E-35	3.6E-41

[a]Standard errors in parentheses under coefficients.
*Indicates coefficient significant at 0.05 level.
**Significant at 0.01 level.
***Significant at 0.001 level.
[b]Squared value.