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AN EVALUATION OF MEDICAID SELECTIVE CONTRACTING IN CALIFORNIA*

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This study used 1982–1986 data on 262 private community hospitals to evaluate the effects of selective contracting for inpatient services by California's Medicaid program. Selective contracting by Medicaid significantly reduced the rate of inflation in average costs per admission and per patient day, while slightly increasing average lengths of patient stays. Private sector contracting also reduced cost inflation rates significantly and caused small, non-significant, reductions in lengths of stays. Hospital savings in 1986 due to Medicaid selective contracting were $836 million, 7.6% of what hospital expenditures would have been in the absence of contracting.

1. Introduction

Rapid increases in expenditure levels for Medicaid programs have placed increasing amounts of pressure on state budgets over the past two decades. Individual states have experimented with a wide variery of Medicaid cost-control strategies, including both prospective payment and utilization review. Very little empirical evidence has accumulated concerning the effectiveness of these policy experiments. This paper reports on a study of the effects of one particularly important policy experiment, the adoption by California of a selective contracting program for Medicaid inpatient care.

The growth in the size of the state's Medicaid program ('Medi-cal') during the 1970s made it a highly visible and vulnerable target for cutbacks when the California state budget entered a period of severe strain in the early 1980s. In a series of quasi-secret meetings that excluded the medical establishment, the legislature hammered out a package that allowed an appointed governmental body (initially an individual 'czar' and subsequently a commission) to negotiate contracts with individual hospitals for fixed daily

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rates, regardless of diagnosis [Bergthold (1984)]. This was a form of prospective payment, but one that differed from the regulatory versions of prospective payment being developed elsewhere. The principle was that the state would not need to mandate particular payment levels; rather, individual hospitals would offer low rates in exchange for higher volumes of patients. Medi-Cal beneficiaries would be restricted to hospitals who were awarded contracts, except in emergency situations.

As a major step in the direction of price competition in the hospital sector, the California Medicaid experiment and the private sector selective contracting policy that accompanied it have generated considerable interest among policy analysts [Melia et al. (1983); Petersdorf (1983); Iglehart (1984)]. A study sponsored by the National Governors' Association evaluated the experiment in term of its effects on the state budget, and reported that the 1982 law saved the state $419 million in the next two years [John et al. (1985)]. It has never been clear, however, whether the budgetary savings accruing to the State of California have resulted from genuine reductions in the rate of medical care cost inflation due to a changed set of incentives, or rather whether they resulted from a shift of charges from the Medi-Cal program to other public and private payers. Nor has it been ascertained whether the Medicaid selective contracting program reduced expenditures for other third party payers as an indirect effect of the Medicaid program's influence on hospital efficiency.

This paper evaluates the influence of the Medi-Cal selective contracting program on hospital performance, as measured by rates of change in costs per admission, costs per patient day, and average length of stay. A varying-parameter cost function is estimated that allows the analysis to measure changes in the structural parameters linking the distribution of payer shares to changes in hospital costs and utilization. Given the clearly non-random manner in which individual hospitals choose to pursue or not pursue Medi-Cal contracts and in which the state chooses to grant or withhold contracts, we control for regression-to-the-mean effects in cost inflation rates for individual hospitals.

We also investigate the determinants of whether individual hospitals are awarded Medi-Cal contracts and the influence of such an award on the subsequent flow of Medi-Cal patients. We utilize data on 262 private hospitals from 1982, the year immediately prior to the implementation of selective contracting, through 1986.

2. Methods

2.1. A varying parameter model of cost inflation

Conventional models assume that the level of costs at any one point in time depends on the levels of output quantities and input prices, which vary across time, and on behavioral parameters, which do not vary across time. These models generate specifications for cost inflation equations in which the rate of change in costs is dependent on the rate of change in output quantities and input prices, but not upon changes in the behavioral parameters. When applied to the hospital industry, these models have usually been expanded to include independent variables other than output quantities and input prices. The behavioral parameters linking the independent variables to the dependent variable continue to be treated as invariant with respect to time, however.

Fixed-parameter cost function models are inappropriate for the study of recent developments in the hospital care industry, since changes in the behavioral parameters themselves are of key importance. Central to the philosophy underlying selective contracting is the principle that the impact on hospital behavior and costs of treating Medi-Cal patients would change when the reimbursement mechanism changed from cost reimbursement to negotiated per diem rates.

This study specified a varying-parameter cost function

$$C_{it} = \alpha_0 + \alpha_{1t} P_{it} + \alpha_{2t} M_{it} + \beta_0 Y_t + \beta_1 W_{it} + \beta_2 H_{it} + \beta_3 Z_i + u_{it}. \tag{1}$$

Here the costs in the ith hospital in time t, (C_{it}), depend on a vector of third party payer patient shares (P_{it}), a measure of local market structure (M_{it}), a vector of output quantities and input prices (W_{it}), a vector of hospital and patient severity variables that change over time (H_{it}), a vector of hospital and patient mix variables that vary among hospitals but not over time (Z_i), a period effect (Y_t) that affects all hospitals similarly, and stochastic error term (u_{it}). The parameters on the third party payer share variables (α_{1t}) and market variables (α_{2t}) are time dependent but the other parameters are not.

The determinants of the rate of hospital cost inflation in this model can be studied by taking the first differences of (1) at times $t+1$ and t.

$$\begin{aligned} \Delta C_i = {} & \beta_0 + \alpha_{1t+1} \Delta P_i + \Delta\alpha_1 P_{it} + \alpha_{2t+1} \Delta M_i \\ & + \Delta\alpha_2 M_{it} + \beta_1 \Delta W_i + \beta_2 \Delta H_i + \Delta u_i. \end{aligned} \tag{2}$$

Here $\Delta C_i = C_{it+1} - C_{it}$ is the rate of change in costs; ΔP_i, ΔM_i, ΔW_i, and ΔH_i are similarly defined first differences. The influence of time invariant hospital and patient mix characteristics (Z_i) has been eliminated. This is particularly important since many of these characteristics cannot be observed. Of particular concern for cross-sectional studies of hospital costs are unobserved patient severity characteristics. These vary enormously among hospitals at any one point in time but typically do not vary as much for any one institution across relatively short time periods. The intercept α_0 in eq. (1) was not time dependent and is also differenced away. The coefficient β_0 on the period effect Y_t in eq. (1) serves as the intercept in eq. (2) since $(Y_{t+1}-Y_t)$ is identical for all hospitals and functions as a constant. We exclude ΔM_i, since no local hospital markets changed substantially over the 1982–1986 period. We include ΔP_i, the change in Medicaid and Medicare patient shares, as well as P_{it}, the 1982 levels of these variables. Inclusion of the 1982 levels of P_{it} and M_{it} permits the analysis to measure changes in the behavioral relationship between payer mix, market structure, and hospital performance ($\Delta\alpha_1$ and $\Delta\alpha_2$).

The variables describing the distribution of patients by third party payer (percentage of discharges reimbursed by Medicaid and Medicare, respectively, with privately insured and uninsured constituting the comparison category) are interpreted differently in a first difference equation such as (2) than in a levels equation such as (1). In principle, the distribution of patients by payer could influence cost levels either of two ways. The Medicaid patients themselves could differ from other patients in terms of case mix severity, in which case the Medicaid share variable would measure the marginal cost of treating a particular type of patient. Alternatively, the different reimbursement and utilization review policies pursued by Medicaid could influence hospital behavior, even in the absence of case mix differences. This latter possibility is based on models of the hospital that portray the institution as increasing expenditures per patient and number of patients treated up to the point where revenues are exhausted [Feldstein (1971); Newhouse (1970); Held and Pauly (1983)]. In this view, a hospital with a greater proportion of patients covered by a less generous insurance plan would behave differently and exhibit a different expenditure function than an otherwise comparable hospital with a smaller proportion of patients covered by that plan.

In a cross-sectional equation such as (1), it would be impossible to separate out the case mix-related effect of having a high Medicaid share from the effect of Medicaid reimbursement policies on hospital expenditures. In a first-differences equation such as (2), however, the association between Medicaid patient share and case mix has been eliminated. Time-invariant relative differences in case mix severity (Z_i) have been differenced away. As a control for changes among hospitals in relative case mix severity over time, we include in the empirical formulation of (2) several variables measuring important case mix changes that would be correlated with Medicaid enrollment status (percentages of inpatient days accounted for by obstetric and pediatric care, respectively) plus other case mix variables. As these variables constitute time-dependent hospital characteristics, they are conceptualized as falling within the ΔH_i vector in (2). To the extent that Medicaid case mix changed between 1982 and 1986 within (as distinct from across) these categories in different ways across hospitals, the Medicaid payer share variables in (2) will continue to reflect case mix as well as program reimbursement effects.

2.2. Regression-to-the-mean effects

A potential bias enters into the estimation of eq. (2) due to the nonrandom manner in which hospitals choose to seek a Medi-Cal contract and in which the state chooses to award such a contract. Given that marginal costs are typically below average costs in the hospital industry, average costs will be higher in institutions with greater excess capacity, other things equal. These hospitals may be precisely those most anxious to obtain a Medi-Cal contract so as to reduce excess capacity. While experiencing high average costs, they might be willing to offer low per diem prices. On the other hand, if all hospitals set their initial bids at approximately the level of their average costs, and if the state awarded contracts on the basis of the lowest bids in each area, then hospitals with low initial average costs would be especially likely to obtain contracts.

This matter is of concern for the analysis of the effects of selective contracting on subsequent cost inflation rates if initial cost levels (at time t) were due to random disturbances in addition to structural characteristics of the hospital. Random disturbances are transitory in nature and decline over time in a process of 'regression to the mean'. If Medi-Cal contracts were disproportionately obtained by initially high cost hospitals, then eq. (2) would overestimate the independent impact of selective contracting in reducing the subsequent rate of cost inflation. On the other hand, if Medi-Cal contracts were disproportionately awarded to initially low cost hospitals, then eq. (2) would underestimate the true effect of selective contracting. More generally, regression-to-the-mean effects in hospital costs will bias parameter estimates for any panel study if stochastic shocks in one period are correlated with hospital characteristics in the succeeding period.

The most straightforward way to model the influence of transitory disturbances on hospital cost inflation is to assume that the error term in eq. (1) is characterized by a first order autoregressive process.

$$u_{it+1} = \rho u_{it} + \varepsilon_{it+1}, \quad 0 < \rho < 1. \tag{3}$$

The effect of a transitory disturbance on costs in the subsequent period is less than that on costs in the present period, but not zero. A white noise term ε_{it}, uncorrelated across hospitals and time periods, also influences costs. The error term in eq. (2) is

$$\Delta u_i = (\rho - 1) u_{it} + \varepsilon_{it+1}, \tag{4}$$

and the estimated parameter on the change in Medi-Cal patient share (ΔP_i) is biased.

$$\mathrm{E}(\hat{\alpha}_{1t+1} - \alpha_{1t+1}) = \theta(\rho - 1), \tag{5}$$

where θ is the correlation between P_{it+1} and u_{it}. (P_{it} and u_{it} are uncorrelated.) If Medi-Cal contracts were subsequently obtained by hospitals with initially high costs, then $\theta > 0$, while if contracts were subsequently obtained by low cost hospitals then $\theta < 0$.

Following Dranove and Cone (1985), we first estimate a cross-sectional cost equation using data at time t and include the estimated residuals $\hat{u}_{it}$ from this regression in the first difference specification of eq. (2).

$$\begin{aligned}\Delta C_i = {} & \beta_0 + \alpha_{1t+1}\Delta P_i + \Delta\alpha_1 P_{it} + \Delta\alpha_2 M_{it} \\ & + \beta_1 \Delta W_i + \beta_2 \Delta H_i + (\rho - 1)\hat{u}_{it} + \varepsilon_{it+1}.\end{aligned} \tag{6}$$

Since $\hat{u}_{it}$ is an unbiased estimate of u_{it}, eq. (6) produces unbiased estimates of all parameters.

2.3. Determinants of Medi-Cal utilization

We investigated the association between hospital costs in one period and Medicaid utilization in the subsequent period in two steps. One equation modeled the probability a hospital is awarded a Medicaid contract as a function of actual minus expected costs in time t plus other factors. The second equation modeled the Medicaid share in time $t+1$ (P_{it+1}) as a function of whether or not the hospital received a contract, the actual minus expected costs variable, plus other factors.

$$\text{contract} = 1/(1 + \exp - (P_{it}\gamma_{11} + \hat{u}_{it}\gamma_{12} + M_{it}\gamma_{13} + S_{it}\gamma_{14})). \tag{7}$$

$$P_{it+1} = \gamma_{20} + P_{it}\gamma_{21} + \hat{u}_{it}\gamma_{22} + M_{it}\gamma_{23} + S_{it+1}\gamma_{24} + \text{contract}\ \gamma_{25} + v_{it+1}. \tag{8}$$

The market characteristics variable M_{it} is included given the concern on the part of the Medi-Cal program that a sufficient number of hospitals receive contracts in each area (California Department of Health Services 1988). Two hospital characteristics, teaching status and ownership status, are included (S_{it}) to investigate whether these structural characteristics were systematically associated with hospital contracting strategy. Medicaid share in time t (P_{it}) is included in (7) and (8) given Medi-Cal's explicit attempt to minimize disruption of practice patterns by favoring hospitals with high existing Medicaid patient loads.

3. Data

This study uses 1982–1986 data on 262 private short-term hospitals in California. Cost data are obtained from the Annual Survey of Hospitals, conducted by the American Hospital Association. The structure of local hospital markets was measured using patient discharge abstract data, which include the patient's ZIP code of residence, from the California Health Facilities Commission. Physician and demographic data describing the environment in which each hospital operates were obtained from the 1982–1986 Area Resource Files, compiled by the Bureau of Health Resources, U.S. Department of Health and Human Services. These variables from the Area Resources File are measured at the county level. Information on whether the hospital was located in a geographical area subject to the selective contracting program and, if so, whether the hospital was awarded a contract, was obtained from the California Office of Statewide Health Planning and Development.

The functional form chosen for the analysis was a Cobb–Douglas version of the modified translog hospital cost function proposed by Breyer (1987), with first differencing and regression-to-the-mean characteristics as described in eq. (6). Costs are measured in terms of the difference in logarithms of average expenditure per patient discharge and per patient day from 1982 to 1986. Hospitals that closed over the four year period are excluded.

Output quantities are measured in terms of the changes in the logarithms of beds, inpatient surgical cases per year, outpatient surgical cases per year, and outpatient visits per year. In order to control for scale effects, these four output measures are divided by annual discharges. An explicit justification for this approach to functional form is provided by Breyer (1987). A Cobb–Douglas rather than translog version is used since the availability of only one input price eliminates the principal advantages of the translog functional form. Input prices are measured in terms of the change in the logarithm of annual earnings for non-physician hospital staff. Case mix changes within individual hospitals are measured in broad terms via the 1982–1986 change in the percentage of annual inpatient days accounted for by each of six classifications: adult medical and surgical care, pediatrics, obstetrics, other acute, intensive care (including neonatal intensive, burn, and cardiac intensive), and subacute care. Binary variables were included for membership in the Council of Teaching Hospitals and for-profit status, respectively. As additional, indirect, controls for differences among hospitals in input prices and the intensity of hospital care practice styles, population per square mile and the ratio of active physicians to area population were included. These two variables are measured at the county level and were obtained from the 1982 and 1986 Area Resource Files.

The impact of state Medicaid policies is measured using the percentages of discharges reimbursed by the program. Selective contracting for Medi-Cal patients was limited to those California areas where the state government believed the local hospital market was structurally competitive. Several counties experimenting with mandatory capitation programs for Med-Cal beneficiaries were also exempted from the selective contracting program [California Department of Health Services (1988)]. Hospitals in Health Facility Planning Areas not subjected to selective contracting were therefore excluded from this study. We also excluded public hospitals. As part of the 1982 Medi-Cal reforms, large numbers of Medically Indigent Adults were transferred from the Medi-Cal program to the countries (which run the public hospitals). For these hospitals, statistics on changes between 1982 and 1986 in Medicaid-insured patients seriously misrepresent actual changes in patient mix. Since hospitals owned by the Kaiser Permanente Health Maintenance Organization do not serve Medi-Cal patients except in emergency situations, they were also excluded.

The structure of the local market in which each hospital operates was measured using patient flow data. Under state law, all California hospitals must submit to the state abstracts of the medical record for every patient discharged. These abstracts contain the ZIP code of the patient's home. We aggregated the 3.5 million abstracts to hospital ZIP code pairs. These 85,361 pairs were sorted by hospital and then by number of patients from each ZIP code going to each hospital. Each hospital's market area was defined as including those ZIP codes sufficient to account for 75% of the hospital's discharges in 1983, with ZIP codes being added to the market area in declining order of size. To limit each hospital's market area to those ZIP codes where the hospital had a significant presence, we excluded ZIP codes in which the hospital did not account for at least one percent of the patients admitted to any hospital in 1983. With this exclusion, the mean percent of a hospital's admissions contained within its market area was reduced to 71.5%.

We used the extent of overlap in market area as the basis for our measure of the amount of potential competition. The extent of market area overlap between two hospitals, *A* and *B*, was measured as the percent of hospital *A*'s patients that came from the part of *A*'s market area that overlapped with *B*'s market area. This was calculated by summing the percent of hospital *A*'s admissions that came from each of the ZIP codes in its market area that overlapped with *B*'s market area. Hospital *A* considers hospital *B* to have competitive potential if *A* received at least five percent of its admissions from their common market area and hospital *B* had at least a 5% market share in this area of overlap. We then counted the number of competitors, as defined, faced by each hospital in the state. This ranged from 0 to 43, with a median

Table 1
Descriptive statistics on dependent and independent variables used in 1982–1986 panel regressions.

Variable	Mean	Std. Dev.	Source
Change in log cost per admission (1982–1986)	0.360	0.233	AHA
Change in log cost per day (1982–1986)	0.413	0.232	AHA
Change in length of stay (1982–1986)	−0.322	1.226	AHA
Medi-Cal share (1982)	14.64	10.52	AHA
Change in Medi-Cal share (1982–1986)	−2.48	9.27	AHA
10–20 competitors	0.237	0.426	CHFC
21 or more competitors	0.286	0.453	CHFC
Medicare share (1982)	39.94	10.28	AHA
Change in Medicare share (1982–1986)	1.81	7.88	AHA
Change in population density	1.45	1.80	ARF
Change in physician density	0.0091	0.0074	ARF
Change in log hospital wages	0.299	0.236	AHA
Change in log beds per admission	0.119	0.289	AHA
Change in log inpatient surgery per admission	−0.168	0.323	AHA
Change in log outpatient surgery per admission	0.998	1.160	AHA
Change in log outpatient visits per admission	−0.070	0.798	AHA
Change in percentage medical–surgical adult	−5.77	17.86	AHA
Change in percentage pediatrics	0.015	2.378	AHA
Change in percentage obstetrics	0.581	2.549	AHA
Change in percentage other acute	2.274	8.711	AHA
Change in percentage intensive care	0.988	8.053	AHA
Contract status (yes = 1)	0.744	0.436	CHFC

Note: AHA = American Hospital Association, annual survey of hospitals, 1982 and 1986; CHFC = California Health Facilities Commission, patient discharge abstracts 1983; ARF = Area Resource Files 1982 and 1986.

of 10. We used this measure in the form of two binary variables that take the value one if the hospital has 11–20 or 21–43 competitors, respectively, and zero otherwise. Hospitals with 0–10 competitors form the comparison category. These three categories were adopted after initial experimentation with a finer mesh of market structure categories revealed no significant differences among markets within each category.

We experimented with various versions of the market structure measure. To obtain a narrower measure limited to the most important competitors each hospital faces, we restricted the definition of competitors to those pairs of hospitals 25% of whose patients came from overlapping ZIP codes. We also recalculated our broad and narrow measures of market structure using the ZIP codes accounting for 90% rather than 75% of each hospital's patients in 1983. Table 1 presents descriptive statistics for the dependent and independent variables used in the analysis.

4. Cost function results

The conversion from cost-based reimbursement to selective contracting on a prospective per diem basis exerted an important influence on hospital costs and average length of patient stay. Table 2 presents ordinary least squares coefficient estimates and standard errors (in parentheses) for the 1982–1986 rates of change in average costs per admission, average costs per patient day, and average length of stay. Hospitals with Medicaid patient shares one standard deviation (10.5 percentage points) above the mean experienced inflation rates in costs per admission 8.2% below the inflation rate in otherwise comparable hospitals with a Medicaid share one standard deviation below the mean ($p<0.001$). An equivalent two standard deviation difference in Medicaid shares was associated with a 10.1% lower inflation rate in costs per day ($p<0.001$). The difference between these two effects is due to the tendency for hospitals contracting on a per diem basis with Medi-Cal to maintain longer average lengths of stay than comparable hospitals with fewer Medi-Cal patients. A two standard deviation difference in Medicaid patient share was associated with a 1.9% increase in length of stay over the 1982–86 period.[1] This effect is not statistically significant.

As discussed in the Methods section, the coefficients on the 1982 Medicaid patient share variable in the 1982–1986 panel regressions presented in table 2 represent the 1982–1986 changes in the behavioral association between Medicaid patient share and the three dependent variables ($\Delta\alpha_1$). The coefficients on the 1982–1986 change in Medicaid patient share are estimates of the state of the behavioral association at the end of the period ($\alpha_{1,86}$). Together, these coefficients portray the structural association between Medicaid patient share and the three measures of hospital performance without the confounding influence of unobserved time-invariant hospital and patient characteristics (which have been differenced away). The association between Medi-Cal share and hospital performance measures in 1982 ($\alpha_{1,82}$) can be derived by subtracting $\Delta\alpha_1$ from $\alpha_{1,86}$.

Using this methodology to derive $\alpha_{1,82}$, the coefficient on Medicaid patient share changes from 0.00412 in 1982 to 0.00037 in 1986 for cost per admission. Prior to the implementation of selective contracting in 1983, participation in the Medi-Cal program exerted a cost-increasing influence on hospital behavior. The magnitude of this effect declined dramatically after the implementation of selective contracting. For costs per day, the coefficient changes sign as well as magnitude, from 0.00391 in 1982 to -0.00069 in 1986. As mentioned above, this is due to the impact on average length of stay. The length of stay coefficient changed sign as well as size, from -0.01258 to 0.00604 between 1982 and 1986.

In order to obtain insights into the extent of the bias introduced into cross-sectional analysis by unobserved hospital and patient characteristics

[1]These point estimates of percentage effects for the cost data are derived from the logarithmic results in table 2 using the transformation $(\exp \alpha_1)-1$. For the length of stay regression, the coefficient in table 2 is divided by the mean length of stay in 1982 (6.59 days) and then multiplied by 100.

Table 2
Determinants of changes in costs per admission, costs per day, and length of stay, 1982–1986 (Standard errors in parentheses).

	Change in log costs per admission	Change in log costs per day	Change in length of stay
Medi-Cal share (1982)	−0.00375 (0.00112)	−0.0046 (0.00138)	0.00604 (0.00654)
Change in Medi-Cal share (1982–1986)	0.00037 (0.00123)	−0.00069 (0.00152)	0.01862 (0.00716)
1982 regression residuals	−0.4225 (0.0491)	−0.3819 (0.0629)	−0.3048 (0.0466)
10–20 Competitors	−0.0579 (0.0236)	−0.0489 (0.0291)	−0.1592 (0.1388)
21 or more competitors	−0.0668 (0.0227)	−0.0685 (0.0280)	−0.0019 (0.1335)
Medicare share (1982)	−0.0016 (0.0011)	0.0002 (0.0013)	−0.0148 (0.0063)
Change in Medicare share (1982–1986)	0.0016 (0.0014)	0.0013 (0.0017)	0.0022 (0.0082)
Change in population density	0.0086 (0.0052)	0.0160 (0.0064)	−0.0498 (0.0305)
Change in physician density	1.2555 (1.3395)	1.9635 (1.6510)	−4.2838 (7.8661)
Change in log hospital wages	0.2535 (0.0399)	0.2818 (0.0488)	−0.4096 (0.2331)
Change in log beds per admission	0.4222 (0.0345)	0.1803 (0.0419)	1.9390 (0.2004)
Change in log inpatient surgery per admission	−0.0190 (0.3010)	−0.0450 (0.0372)	0.2186 (0.1769)
Change in log outpatient surgery per admission	−0.0124 (0.0086)	−0.0171 (0.0105)	0.0963 (0.0503)
Change in log outpatient visits per admission	0.0470 (0.0117)	0.0653 (0.0145)	−0.0820 (0.0691)
Change in percentage medical-surgical adult	0.0017 (0.0007)	0.0049 (0.0009)	−0.0251 (0.0042)
Change in percentage pediatrics	0.0100 (0.0041)	0.0195 (0.0051)	−0.0770 (0.0243)
Change in percentage obstetrics	−0.0014 (0.0038)	0.0187 (0.0047)	−0.1250 (0.0222)
Change in percentage other acute	0.0044 (0.0013)	0.0051 (0.0017)	−0.0159 (0.0080)
Change in percentage intensive care	0.0043 (0.0013)	0.0079 (0.0017)	−0.0289 (0.0079)
Intercept	0.3563 (0.0517)	0.3654 (0.0638)	0.0891 (0.3033)
Adjusted $R2$	0.62	0.38	0.50
N	262	262	262

Table 3

Cross-sectional estimates of cost and length of stay functions, 1982 and 1986 (Standard errors in parentheses).

	Cost per admission		Cost per day		Length of stay	
	1982	1986	1982	1986	1982	1986
Medi-Cal share	0.0029	−0.0015	0.0011	−0.0039	0.0167	0.0229
	(0.0014)	(0.0013)	(0.0013)	(0.0015)	(0.0088)	(0.0079)
10–20 competitors	0.0998	0.0746	0.0550	0.0280	0.2711	0.2750
	(0.0339)	(0.0302)	(0.0323)	(0.0354)	(0.2111)	(0.1900)
21 or more competitors	0.0517	0.0205	0.0339	−0.0102	0.1819	0.3002
	(0.0329)	(0.0288)	(0.0313)	(0.0337)	(0.2050)	(0.1813)
Adjusted R^2	0.57	0.62	0.42	0.48	0.64	0.69
N	262	262	262	262	262	262

Note: These regressions also include the output quantity, input price, patient care mix, area demographic, Medicare, and hospital ownership and teaching status variables presented in table 2.

(Z_i), we also measured the effects of the Medi-Cal program changes by estimating 1982 and 1986 cross-sectional versions of eq. (1) and then subtracting the parameters $(\alpha_{1,t+1}-\alpha_{1,t})$. The cross-sectional analyses slightly over-estimated the size of the Medi-Cal program changes on average costs. According to the cross-sectional regressions (presented in table 3), a two standard deviation difference in Medicaid patient share was associated with a 9.7% decrease in the rate of inflation in costs per admission, and a 11.1% decrease in the rate of inflation in costs per day. The cross-sectional approach correctly estimates the effect of Medi-Cal program changes on average length of stay, with a two standard deviation difference in Medi-Cal patient shares causing a 2.0% increase. The extent of the differences between the panel and cross-section results is not large, suggesting that unobserved hospital and patient characteristics were not strongly correlated with Medicaid patient share in these data.[2]

Hospitals whose costs in 1982 were above expected levels experienced significantly lower rates of cost inflation over the next four years than hospitals whose costs in 1982 were at expected levels, consistent with the hypothesis that regression-to-the-mean effects play an important role in hospital cost dynamics. The 1982 cost per admission regression residuals variable had a sample standard deviation of 0.20. A two standard deviation difference in 1982 residuals measures the difference between hospitals with

[2]Bias in the estimation of Medicaid effects also enters due to the correlation between unobserved characteristics and any included variables in eq. (1), not just between the unobserved characteristics and the Medicaid patient share per se. Apparently these second order effects were not substantial.

costs in 1982 at the high end of the distribution, after controlling for the measured determinants of cost, and hospitals at the low end of the cost distribution. Multiplied with the estimated parameter on the residuals variable in the 1982–1986 regression reported in table 2, this implies that hospitals with exceptionally high (higher than predicted by the model) costs in 1982 had rates of cost inflation per admission over the next four years 18.4% lower than hospitals with exceptionally low costs in 1982.[3] Strong regression-to-the-mean effects are also observed in the cost per day and length of stay equations. The autocorrelation parameter ρ is estimated to be 0.58, 0.62 and 0.70 for costs per admission, costs per day, and average length of stay, respectively.

Hospitals in more competitive local markets, measured in terms of the number of other hospitals with overlapping market areas, experienced rates of cost increase lower than did hospitals in less competitive markets. No significant association was observed between market structure and changes in average length of patient stay. Results comparable to the ones presented in table 2 were also obtained using the alternative measures of local market structure (described in the Data section). The two area physician and demographic variables, change in population density and change in active physicians per 1,000 county residents, are both positively and significantly associated with rates of change in average costs per admission and per day but not with changes in average lengths of stay.

Hospitals with large percentages of patients covered by the federal Medicare program experienced rates of inflation in costs per admission modestly lower than did hospitals with proportionately fewer Medicare patients, other things equal. A two standard deviation (20.6 percentage point) difference in Medicare patient share was associated with a 3.3% decrease in inflation, a statistically insignificant amount. The influence of Medicare's Prospective Payment System on costs per admission was due solely to its effect on average length of stay. This contrasts with the Medi-Cal program effect, which worked largely via reductions in inflation rates of costs per day (with relatively little change in length of stay). Hospitals with Medicare patient shares one standard deviation above the mean experienced rates of decrease in average length of stay 4.6% greater than otherwise comparable hospitals with Medicare patient shares one standard deviation below the mean.

Hospitals facing high rates of growth in wages experienced high rates of growth in average costs. The coefficient on the outpatient visits variable is significantly positive. The dependent variables in the cost functions are total costs per patient admission and per patient day, where total costs include

[3]This point estimate of the percentage effect is derived from the logarithmic data as described in the first footnote.

Table 4
The influence of selective contracting on hospital utilization by Medi-Cal patients (Standard errors in parentheses).

	Probability of contract award	1986 Medi-Cal patient share
Contract awarded	–	5.753
		(1.187)
1982 Medi-Cal share	0.147	0.494
	(0.030)	(0.050)
1982 regression residuals	−1.567	−4.087
	(0.786)	(2.472)
10–20 competitors	0.385	−0.180
	(0.417)	(1.241)
21 or more competitors	0.014	0.021
	(0.368)	(1.164)
Investor-owned	−0.161	−2.030
	(0.332)	(1.036)
Council of teaching hospitals	0.249	1.135
	(0.842)	(2.163)
Intercept	−0.645	1.494
	(0.426)	(1.287)
−2 log likelihood function	249	–
Adjusted R^2	–	0.43
N	262	262

expenditures in the outpatient departments as well as in the inpatient departments. Changes in volumes of inpatient and outpatient surgery are not associated with cost inflation rates, after controlling for changes in bed capacity, outpatient visits, and total inpatient admissions (implicitly controlled for in the denominator of all the utilization variables).

Hospitals whose case mix changed in favor of adult medical and surgical, pediatric, other acute care, and intensive care treatment reported higher rates of inflation in costs per admission than did hospitals whose case mix changed in favor of subacute care (the comparison category). The five case mix change categories all produced significant positive coefficients in the average cost per day and significant negative coefficients in the length of stay regressions due to the exceptionally long lengths of stay and low costs per day on the part of patients needing subacute care, the comparison category.

5. Utilization results

Table 4 presents parameter estimates for the contracting and 1986 Medicaid patient share eqs. (7) and (8). Medicaid patient share in 1982 and the excess of actual over predicted average costs per admission in 1982 strongly influenced the likelihood that a hospital was awarded a contract. A two standard deviation difference in 1982 Medicaid share was associated with a 0.447 greater probability of obtaining a contract. Hospitals with actual minus predicted costs one standard deviation above the mean reduced their probability of obtaining a contract by 0.089 compared to hospitals with adjusted costs one standard deviation below the mean.[4] Neither the structure of the local market nor the two hospital characteristics (teaching status and ownership status) was associated with probability of obtaining a contract.

The second column of table 4 presents ordinary least squares estimates of the determinants of Medi-Cal patient share in 1986, several years into the selective contracting program. Hospitals holding a Medi-Cal contract had significantly larger Medi-Cal patient shares than did hospitals without such contracts, even controlling for Medi-Cal share in 1982. Conversely, both 1982 Medi-Cal share and 1982 actual minus expected costs significantly influence 1986 Medi-Cal share even after controlling for presence of a contract. The continuing influence of cost levels on Medi-Cal patient share is interesting because Medi-Cal beneficiaries face no direct incentives to choose low-cost institutions from among the set of contracting hospitals. This association may be due to referral patterns developed by physicians treating large numbers of Medi-Cal patients. Investor-owned hospitals had a smaller dependence on Medi-Cal patients than did private nonprofit hospitals, even though they were no less likely to obtain a contract.

6. Estimation of program-related savings

While the selective contracting program had important effects on the California state budget, an evaluation from a social perspective must include its effects on hospitals. The results presented in tables 2–4 indicate that the selective contracting program has influenced both the distribution of Medi-Cal patients across hospitals and the nature of the behavioral response by hospitals to serving Medi-Cal patients. Both of these effects need to be considered in estimating the program's overall effect on the hospitals studied. This section combines the results presented in the previous sections into a single dollar estimate of the effect of Medicaid selective contracting on expenditures in California's private hospitals in 1986, three years into the program.

[4]These present estimates were derived using the partial derivative of the logistic function

$$\frac{\partial P}{\partial x}=\beta P(1-P);\; P=1/[1+\exp(-X\beta)]$$

where β is the logistic parameter on x and P is the probability of obtaining a contract, evaluated at the sample mean of the vector of independent variables $\bar{X}$.

Table 5
Estimated 1986 savings due to the Medi-Cal selective contracting program (dollars).

	Mean	Median	Minimum	Maximum	Sum
Savings per admission	447	348	62	2,952	-
Total savings	3,190,740	1,733,230	31,313	29,161,625	835,974,000
Percentage saved	7.17	6.05	1.37	23.24	7.56

Actual average costs in 1986 for the ith hospital were

$$C_{i,86} = P_{i,86}\alpha_{86} + X_{i,86}\gamma + u_{i,86}, \tag{9}$$

where $P_{i,86}$ is the percentage of discharges reimbursed by Medi-Cal and $X_{i,86}$ includes all the other independent variables in eq. (1). In the absence of selective contracting, costs would have been

$$\mathrm{E}_{i,86} = P_{i,82}\alpha_{82} + X_{i,86}\gamma + u_{i,86}, \tag{10}$$

which, after rearranging terms, is

$$\mathrm{E}_{i,86} = C_{i,86} - (\Delta P_i \alpha_{86} + P_{i,82}\Delta\alpha). \tag{11}$$

The expected cost per admission was calculated for each hospital using the coefficients in the first column of table 2, and then was retransformed from logarithmic to dollar units using Duan's (1983) nonparametric 'smearing' method.[5] Total predicted costs were obtained by multiplying predicted average cost per admission by number of 1986 admissions. Total actual costs for 1986 were then subtracted from total predicted costs to measure hospital-specific savings associated with the Medi-Cal selective contracting program. Percentage savings were calculated by dividing savings by predicted 1986 costs (i.e., actual costs plus savings).

Descriptive statistics on the distribution of Medicaid-related hospital savings per admission and in total are presented in table 5 both in dollar terms and in terms of percentages. Savings ranged from a low of 1.4% to a high of 23.2%, with a median of 6.1%. Twelve hospitals experienced savings greater than 15%. All of these institutions had Medi-Cal patient shares in 1982 of at least 30%; all were awarded contracts. In dollar terms, savings ranged from \$0.03 million to \$29.2 million, with a median of \$1.7 million. As

[5]The conventional formula for the retransformation of predicted values from logarithmic to natural units is $\exp(X\beta + \sigma^2/2)$. For non-lognormally distributed variables, Duan (1983) shows that the nonparametric retransformation $\exp(X\beta)^*\phi(X)$, where $\phi(X)$ is the sample average of the exponentiated least squares residuals, has smaller average squared prediction error than the conventional formula.

a group, California hospitals saved an estimated $836 million dollars in 1986 compared to what they would have spent in the absence of the Medicaid selective contracting program.

7. Discussion

The quantitative assessment of policy changes such as the shift to selective contracting by California's Medicaid program is beset by numerous econometric problems. Many key variables, especially those pertaining to patient case mix severity, are unmeasurable and also correlated with the observed variables of interest. By design, selective contracting only occurs among pairs of economic agents when each perceives a personal benefit: self-selection issues are therefore important. Finally, these policy initiatives change the behavioral association between exogenous and endogenous variables, implying that parameter values are themselves changing over time. This study has developed a varying-parameter cost function model with regression-to-the-mean controls in order to deal with these methodological pitfalls.

The most important findings of the study, for policy purposes, is that the market-oriented strategy adopted by the Medi-Cal program and private health insurers have partially achieved their goal of reducing the rate of increase in hospital costs. Hospitals with larger Medi-Cal patient shares and hospitals in structurally competitive local markets reported lower rates of cost inflation over the four years following the 1982 reforms than did otherwise similar hospitals with smaller Medi-Cal shares and fewer local competitors. The total savings due to Medi-Cal contracting were estimated to equal 7.6% of what private California hospitals would have spent in the absence of the selective contracting program. This figure assumes that the statistical model has adequately controlled for differences among hospitals in the 1982–1986 rate of change in case mix severity for those patients covered by Medi-Cal. Time-invariant case mix differences between Medi–Cal beneficiaries and non-beneficiaries were eliminated by differencing and the focus on rates of change rather than levels of costs. Broad changes over time in case mix between Medi-Cal beneficiaries and non-beneficiaries were captured by the variables measuring changes in the distribution of patients across adult medical and surgical care, obstetrics, pediatrics, other acute care, intensive care, and subacute care. Nevertheless, the possibility always remains that some of the effects we have ascribed to cost-control efforts were due to residual unobserved differences in case mix severity.

The findings presented in this paper are broadly consistent with those obtained by Zwanziger and Melnick (1988) using different data and a different econometric approach. Zwanziger and Melnick use 1980–1985 quarterly financial data from the California Health Facilities Commission to estimate a total cost function. The effect of Medi-Cal selective contracting is captured via a series of interaction terms between Medi-Cal patient share and year dummy variables. Local market structure is measured in a manner different from ours, but uses the same basic ZIP code data from the patient discharge abstracts. In terms of our model, Zwanziger and Melnick estimate a pooled cross-section time series version of eq. (1). They find that the parameters on the Medi-Cal patient share and local market structure variables change over time in the anticipated direction. While hospitals with high proportions of Medi-Cal patients and many local competitors had relatively high costs in 1980, they had relatively low costs in 1985, with most of the change occurring after 1982.

Hospital selective contracting has often been viewed by the state officials as a halfway house on the road to full capitation [Halfon and Newacheck (1986)]. While the enabling legislation covered outpatient as well as inpatient care, selective contracting was never extended beyond the hospital sector. This is a limitation of increasing importance, given the rapid growth in both prices and utilization of outpatient services. The expected follow-up legislation to the 1982 reforms, i.e., the transition to capitation, also has not occurred. This caution is due both to the relative success of hospital selective contracting in reducing expenditure inflation and to doubts that linger from California's disastrous leap into Medicaid capitation in the early 1970s.

Short term savings in the health care system are to be made most easily through reductions in the utilization of high cost services, especially through the substitution of outpatient for inpatient modes of care and also through reductions in average lengths of stay. As argued by Schwartz (1987), however, these savings are necessarily limited, since a floor of medically necessary inpatient days per capita exists below which further economies will result in significant adverse outcomes. Over the long term health care cost control will require limitations on the rate of increase in the intensity of services used per case. The reductions in hospital cost inflation rates achieved by the Medi-Cal program came via changes in the level of expenditures per patient day rather than through changes in the number of patient days per admission. A more detailed set of data would be necessary in order to examine the micro changes in hospital practices that have occurred over the course of the 1980s. More years of experience will be necessary in order to discern whether the effects reported here mark a fundamental turning point in hospital cost dynamics or merely a transitory dip in the steep upward gradient.

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