Does Removing Certificate-of-Need Regulations Lead to a Surge in Health Care Spending?

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Abstract This study assesses the impact of certificate-of-need (CON) regulation for hospitals on various measures of health spending per capita, hospital supply, diffusion of technology, and hospital industry organization. Using a time series cross-sectional methodology, we estimate the net impact of CON policies on costs, supply, technology diffusion, and industry organization, controlling for area characteristics, the presence of other forms of regulation, such as hospital rate-setting, and competition. Mature CON programs are associated with a modest (5 percent) long-term reduction in acute care spending per capita, but not with a significant reduction in total per capita spending. There is no evidence of a surge in acquisition of facilities or in costs following removal of CON regulations. Mature CON programs also result in a slight (2 percent) reduction in bed supply but higher costs per day and per admission, along with higher hospital profits. CON regulations generally have no detectable effect on diffusion of various hospital-based technologies. It is doubtful that CON regulations have had much effect on quality of care, positive or negative. Such regulations may have improved access, but there is little empirical evidence to document this.

For more than two decades, health care cost containment has been at the forefront of the health policy agenda. However, the approaches used to achieve cost containment have changed. One of the first policies adopted by states (and that for a time was required by federal statute) was certificate-of-need laws (CON). Such laws, which focused on hospitals and nursing homes, were adopted to curb needless duplication of ser-

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vices and consequent excess capacity. At the time, retrospective reimbursement provided guaranteed reimbursement even if facilities operated at well below capacity. Also, given nearly complete insurance coverage for hospitals, competition for patients occurred on a nonprice basis (Robinson and Luft 1987; Dranove, Shanley, and Simon 1992). The hospitals that could offer the most sophisticated range of services and equipment were most attractive to patients and their physicians. The price of such care did not matter, or at least it mattered much less. Competition by service expansion and proliferation of new technology has been termed the "medical arms race." At least in principle, CON regulations could control the medical arms race by requiring that organizations demonstrate need for a facility, service, or equipment before investing in them. Also, in the 1980s, some states expanded CON regulations to control the proliferation of ambulatory care providers that was occurring (Finkler 1985). Other perhaps secondary objectives of CON regulations were to promote access and to promote quality. A less charitable view is that CON regulations sought to establish entry barriers to protect the income of existing providers, especially hospitals (Feldstein 1988; Wendling and Werner 1980).

Several developments have occurred since the late 1960s and early 1970s that have lessened the popularity of CON regulations, especially as they affect hospital care. First, other regulatory mechanisms thought to be more effective in cost containment have been adopted. Primary among these is Medicare's Prospective Payment System (PPS), but some states implemented various forms of regulation of hospital rates and revenue. Although PPS is still in effect, hospital rate-setting remains in only one state.[1] Second, there has been substantial growth in various forms of managed care, stimulated in part by legislation, such as selective contracting laws. Although specific incentives differ, managed care provides incentives for hospitals to be concerned about cost. In this context, there is a perception that CON regulations may not be needed as much as they were previously to control hospital cost growth. As a result of managed care plan growth as well as implementation of PPS, demand for inpatient hospital care has decreased appreciably. Third, as discussed later, a substantial amount of empirical evidence accumulated by the early 1980s indicating that CON regulations were ineffective in cost containment. Research findings per se did not contribute to the demise of CON laws, but such findings probably coincided with

1. At various times, six different states had adopted this approach, with New York being the most recent to abandon it (on 30 June 1996).

experience-based impressions of policy makers and experts in the field. Fourth, the federal law requiring states to have CON regulations expired in 1986. Since then, fifteen states have dropped CON regulations for hospital services; about half of these have retained CON regulations for nursing homes.

Policy makers in many other states have been reluctant to drop CON laws because of a concern that removing them would lead to a surge in health care spending, including both capital expenditures (initially, subsequent to removal of CON laws) followed by increased operating expenses. Some largely anecdotal accounts of surges following removal of CON laws were reported (Simpson 1986; Lewin-ICF 1992b). Although PPS and managed care have changed incentives, these forces may be insufficient to offset the other inflationary factors that preceded these more recent developments. Second, there is concern that without restraint by CON regulations, market forces will exacerbate an existing maldistribution of facilities, thus placing a greater burden on the disadvantaged. Some observers are also worried that for-profit providers would benefit disproportionately from removal of CON regulations. Some view this as troublesome since for-profit facilities may be less willing to provide uncompensated care. Some studies have shown this to be so (see references in Kuttner 1996), but other studies indicate that the contribution to uncompensated or indigent care is about equal, whether measured in terms of the self-pay share of patients, the bad debt–charity care share of charges, or the share of revenue accounted for by Medicaid (see Sloan's 1988 review). Proliferation of low-volume facilities also is a concern on the grounds that high volume is associated with higher quality of care, at least for some procedures (Luft et al. 1990).

Absent from these policy discussions to date has been systematic empirical evidence of the experiences in states that have lifted CON regulations. Did a surge in spending occur? If so, for which types of facilities and services did the surge occur? Did removal of CON regulations open the doors to the for-profits? Conversely, did removal of CON regulations have beneficial effects, such as increasing price competition through promoting growth of managed care, which may have been restrained previously because of CON entry barriers? Compared with other approaches to cost containment, how well do CON regulations perform? This is an old question, but the track record for comparing alternative approaches to cost containment is now far longer than when most studies were conducted during the 1970s and 1980s. Furthermore, it is now possible to follow the experience of states that dropped CON instead of simply com-

paring states with CON to those that had not yet adopted it. Finally, for the first time, a fourteen-year, continuous time series of state per capita health spending data has become available from the U.S. Health Care Financing Administration (HCFA).[2]

This article provides new empirical evidence about these issues with regard to acute care services. In focusing on acute care services, we exclude nursing homes, hospices, and home health care, but we do include ambulatory surgery and visits to physicians' offices as well as to hospitals. Using a state time series of cross-sections, we assess the effects of lifting CON through 1993. The success of CON in cost containment is compared with other approaches. We show that mature CON programs are associated with a modest (5 percent) long-term reduction in acute care spending per capita, but with no significant reduction in total per capita spending. We also found no evidence of a surge in acquisition of facilities or in costs following removal of CON.

Our empirical specification is followed by a discussion of findings on CON, other regulatory programs, competition, control variables on expenditures on acute care services, hospital beds, service intensity, and profitability, diffusion of technology, and industry organization. We then evaluate our results, compare our findings with those from previous studies, and discuss previous research on effects of CON on quality and access. Although we do not present any new direct evidence about quality and access, these issues are clearly germane to states' decisions about whether CON should be retained.

Empirical Specification

Dependent Variables

We specified equations for the following dependent variables. To measure the effects of CON and other factors on per capita health spending, we defined dependent variables for (1) total expenditures on personal health care services; (2) total acute care expenditures (defined as total spending minus nursing and home health expenditures); (3) expenditures on hospital care; and (4) expenditures on physicians' services per person

2. These data have not been published, but can be obtained by sending a blank diskette to Anna Long in the Health Care Financing Administration's Office of National Health Statistics, Office of the Actuary, Room N3-02-02, 7500 Security Boulevard, Baltimore, MD 21244-1805.

for a state's resident population. We also obtained estimates of Medicare spending per elderly enrollee, including total Medicare expenses and Part A and Part B expenditures.[3] Unpublished estimates of personal health care expenditures by state and year in total and by component were obtained from HCFA for 1980–1993.[4] We also analyzed Medicare expenditures for 1980–1993. All monetarily expressed variables were deflated by the all-items Consumer Price Index.

Dependent variables for hospital supply were beds per 1,000 state residents; for service intensity, the dependent variables were expense per adjusted (for outpatient volume) patient day and per adjusted admission; the dependent variable for hospital profits was the ratio of total revenue to total expense. The revenue measure was for funds actually received by hospitals during the fiscal year, not for hospital charges. Data for these dependent variables for 1976–1993 came from the American Hospital Association's *Hospital Statistics* (AHA 1977–1994).

To measure the influence of CON and other factors on the variable diffusion of technology, we defined dependent variables for (1) the number of hospitals with open-heart surgery units (1980–1993), (2) for hospitals with organ transplant units (1980–1993), (3) for hospitals with ambulatory surgery units (1983–1993), and (4) for all ambulatory surgery units, including freestanding facilities, per one million state residents (1983–1993). The different time periods we studied were dictated

3. Our figure for total Medicare per elderly enrollee equals the sum of the per enrollee estimates for Part A and Part B. Given that not all Part A eligibles receive Part B, our figure is slightly different from the HCFA-reported state level estimates of total spending per enrollee who was eligible for either Part A or Part B during the year. This latter figure will fluctuate based on changes in the mix of Part A and Part B eligibles, so we sought a slightly more stable measure that can be interpreted as estimated spending for an elderly enrollee who had enrolled in both Part A and Part B.

4. Most readers may be aware that these HCFA estimates measure spending by place of service, so our measure of spending per state resident is not intended to be an accurate measure of resource consumption by residents in that state, given that many residents may cross state borders to seek care. HCFA is still working on the development of residence-adjusted per capita spending figures. However, even if these were available, we believe they would not have been appropriate for our analysis insofar as the impact of a state's CON should be reflected in all spending within its own borders, not just that of its own citizens. Given that our method in essence measures the influence of various factors on year-to-year changes in per capita spending, the measure we have chosen would be unsuitable only if there were large year-to-year variations in the extent of border-crossing, which seems improbable. On the other hand, we also recognize that if CON regulations had the effect of driving citizens to neighboring states to seek care, our analysis of HCFA data would not be able to detect it. Part of our motivation in also analyzing Medicare spending per eligible person—which is a residence-adjusted measure of spending—was to see whether we got consistent results using both place-of-service and place-of-residence measures of per capita spending.

by data availability.[5] Information on the first three variables came from the *Hospital Statistics* (AHA 1977–1994). Data for the fourth came from the SMG Marketing Group (1984–1995). For the variable industry organization, we defined dependent variables for the for-profit share of hospital beds[6] for 1976–1993 based on *Hospital Statistics* and the HMO enrollments as a fraction of the state population, information taken from the Group Health Association of America's *National Directory of HMOs* (GHAA 1977–1994). We used data for 1976–1993 in our analysis of HMO market share.

Examining Certificate-of-Need Laws

Four binary variables represented certificate-of-need laws: pre-CON—the year before and the first year CON was implemented; young CON—the first two years postimplementation; mature CON—the remaining years CON was in effect; and CON lifted—the first three years after the CON law was dropped. Pre-CON was included to capture anticipatory effects of CON. There is some empirical evidence that hospitals began some capital projects in anticipation of CON (Sloan and Steinwald 1980a). Once enacted, CON laws plausibly had greater effects after they had been in place for a number of years. The variable CON lifted was included to determine whether there was a surge in hospital investment (and consequently in hospital costliness) immediately after CON laws were dropped.

If CON laws constrain hospital investment and cost, the savings may be offset by greater expenditures in other parts of the health care sector, as others have argued (see e.g., Finkler 1987). By including analysis of the ambulatory sector and of total health care expenditures, we were able to examine this possibility.

Program age is only one aspect of CON programs that is heterogeneous. Programs also logically differ in *stringency*, which reflects the scope of coverage and the difficulty applicants have in securing certificates of need. In an alternative specification, we used a CON stringency

5. Because our observational unit was the state, our diffusion measures were based on counts of the number of facilities offering a particular service. At a lower level of aggregation, it would be useful to study whether additional units opened where existing units were, or where the facility was the first of its kind in the area.

6. We recognize that our results might have been somewhat different if we had measured the for-profit share as a percentage of revenues or admissions. Our convention here is typical of previous analyses of CON regulations using state or regional data (see Noether 1988; Lanning, Morrisey, and Ohsfeldt 1991).

measure originally developed by Lewin-ICF (1992a).[7] These measures took account of dollar thresholds used to determine whether a project was subject to CON review, in terms of the scope of specific categories of services subject to review. This produced a continuous numerical score that Lewin-ICF used to categorize states into three mutually exclusive categories: 1 = limited; 2 = moderate; 3 = stringent. These categorical scores were used in our analysis.[8]

Finally, for most of the observational period, states could adopt section 1122 programs at their option. Unlike CON, section 1122 allowed hospitals to make unapproved investments in plant, equipment, and services, but unless approved, there was no Medicare or Medicaid reimbursement for the capital expenditures associated with the projects. The section 1122 variable measured the fraction of hospital revenues from Medicare and Medicaid by state and year, only for the years that section 1122 was in effect in a given state.

Hospital Rate-Setting

An explanatory variable for Medicare Prospective Payment measured the fraction of hospital revenues covered by PPS by state and year. The variable accounts for the years the program was phased in (1984–1987) as well as the fraction of hospital revenue from Medicare by state and year. We also measured the fraction of hospital revenue covered by mandatory rate-setting programs.[9] Following previous work by one of the authors (Sloan 1981), we distinguished between young rate-setting —the first three years of implementation—and mature rate-setting, the remaining years that CON laws were in effect. The variables were defined to reflect the fraction of revenue covered by the program.

7. More recent data for this measure are reported in Lewin-VHI (1995).

8. The Lewin-ICF methodology was not explained in enough detail to replicate the continuous scoring system. Because we had to interpolate figures for 1991 (based on reported figures for 1990 and 1992) and extrapolate to 1993 based on other available information about changes in thresholds, we were able to do so more reliably with the categorical data (whose values tended to be stable over time for any given state) than if we had attempted to replicate the continuous scoring system.

9. Previous work by Sloan (1981) examined a wider range of hospital rate-setting programs, including voluntary and advisory programs. Both theory and most evidence suggest that mandatory prospective rate-setting is the most effective form of hospital rate regulation (Biles, Shramm, and Atkinson 1980; Morrisey, Sloan, and Mitchell 1983; Sloan 1983; Rosko 1989).

Reimbursement

Explanatory variables were included to represent the fractions of hospital revenue that came from Medicare and from Medicaid programs, respectively.

Price Competition

The HMO share—calculated by dividing HMO enrollment by resident population on 1 July of each year—was used to represent the influence of managed care on hospital costs.[10] These data were obtained from GHAA's *National Directory of HMOs*.

Area Characteristics

We controlled for other factors likely to affect the dependent variables: income per capita population (Bureau of Economic Analysis estimates); the ratio of general practitioners to all physicians; the fraction of population over age sixty-five (Bureau of the Census); the population density (Bureau of the Census); and the weekly wage paid to service workers (Bureau of Labor Statistics [BLS] 1976–1994).

Other Explanatory Variables

To capture omitted cross-sectional and intertemporal influences, we included state binary variables and a time trend. To conserve space, coefficients and standard errors on the intercept, state binary variables, and the Voluntary Effort (only included in analysis that spanned the 1970s but not presented because it is no longer of policy interest) are not presented in the tables shown here.[11] To allow us to distinguish between short- and long-run influences on explanatory variables, we included

10. Unfortunately, analogous data on PPO enrollments were not sufficiently reliable to use in our analysis because of changes in definitions over time. HMO share is not a perfect measure of price competition insofar as it does not take into account the nature of plans offered (e.g., group model versus independent practice association) or the aggressiveness of purchasers in the market, which strongly influences the degree to which HMO presence actually affects competition and hospital costs (Robinson 1995; Zwanziger and Melnick 1996). Despite its limitations, HMO share has been shown to be related to price (premium) levels in two different studies (Wholey, Feldman, and Christianson 1995; Feldstein and Wickizer 1995), so in the absence of a better measure, we feel justified in using it.

11. The Voluntary Effort was a voluntary cost-containment effort promoted by the American Hospital Association to diminish support for President Carter's proposed price controls on hospitals. This effort began in December 1977 and lasted until about 1980 (Sloan 1983).

lagged dependent variables. The coefficient on the dependent variable is interpretable as one minus the fraction of the gap between the actual and the equilibrium value of the dependent variable that is closed in a year (λ). Thus, if the coefficient were .8, .2 of the gap would be closed annually. To obtain the long-run influence, the coefficient on an explanatory variable is divided by λ.

Functional Form

With the exception of the HMO share equation, all dependent variables were expressed in natural logarithm form, as were the variables in the other explanatory variables category; all other explanatory variables were entered linearly. Since there were an appreciable number of observations with no HMOs (about one hundred), we estimated the HMO share equation in linear form.

Results

Effects of Certificate-of-Need Laws

Certificate-of-need laws had no effect on total personal health expenditures per capita or on per capita spending on physicians' services (Table 1). For spending on acute care, mature CON had a negative impact that was statistically significant at the five percent level. The long-run effect of mature CON was an almost five-percent reduction in per capita acute care expenditures, which includes ambulatory care as well as hospital expenditures. However, we were unable to detect a statistically significant effect of removing CON on these same expenditures. Surprisingly, in view of this finding, mature CON did not have a statistically significant effect in reducing hospital spending, and in this regression, the coefficient on the variable CON lifted has a negative sign (statistically significant at the 10 percent level).

For Medicare expenditures, the only statistically significant CON coefficients have positive signs. A positive sign on CON lifted suggests a surge in Part A (i.e., hospital expenses), but the positive sign on mature CON in the Part B regression suggests that physicians' services may have substituted for hospital care when the latter was constrained.

On the whole, the section 1122 program seems to have been effective in containing costs. Negative and statistically significant coefficients were obtained in most regressions, but strangely, not in the regression

Table 1 Expenditures on Acute Care Services

	Medical Spending/Pop. (HCFA)				Spending Per Medicare Eligible Age 65+		
	Total Spending	Acute Spending	Hospital Spending	Physician Spending	Total Medicare	Part A	Part B
CERTIFICATE-OF-NEED REGULATION							
Section 1122	−.012[b]	−.018[b]	−.001	−.029[c]	−.049[c]	−.090[b]	.053
	(.005)	(.007)	(.010)	(.015)	(.029)	(.045)	(.063)
Young CON	.006	.001	.0002	−.0001	.002	−.013	.041
	(.006)	(.007)	(.010)	(.015)	(.029)	(.045)	(.064)
Mature CON	−.004	−.009[b]	−.005	.004	.029[c]	−.008	.163[a]
	(.003)	(.004)	(.006)	(.009)	(.017)	(.027)	(.038)
CON Lifted	−.004	−.006[c]	−.010[c]	.003	.032[c]	.017	.143[a]
	(.003)	(.004)	(.006)	(.009)	(.017)	(.026)	(.038)
HOSPITAL RATE-SETTING							
Prospective Payment System (PPS)	.042[a]	.018	.091[a]	.103[b]	−.254[a]	−.401[a]	.169
	(.016)	(.022)	(.031)	(.045)	(.083)	(.128)	(.182)
Young Mandatory Prospective	−.038[b]	−.036[c]	−.063[b]	−.065[c]	.051	−.024	.253
	(.015)	(.021)	(.029)	(.043)	(.082)	(.126)	(.178)
Old Mandatory Prospective	−.011[c]	−.017[c]	−.022[c]	−.027[c]	−.073[b]	−.101[c]	−.052
	(.006)	(.009)	(.012)	(.018)	(.034)	(.053)	(.075)
REIMBURSEMENT							
Medicaid Share	.059[a]	.082[a]	.153[a]	−.039	.125	.330[c]	−.322
	(.022)	(.030)	(.042)	(.063)	(.120)	(.185)	(.261)
Medicare Share	−.179[a]	−.204[a]	−.330[a]	−.092[b]	.008	.124	−.246
	(.017)	(.023)	(.033)	(.047)	(.089)	(.139)	(.193)
COMPETITION							
HMO Market Shares	.033	.011	.041	.031	−.178	−.330[c]	−.420
	(.025)	(.034)	(.049)	(.072)	(.137)	(.208)	(.295)

Table 1 Continued

	Medical Spending/Pop. (HCFA)				Spending Per Medicare Eligible Age 65+		
	Total Spending	Acute Spending	Hospital Spending	Physician Spending	Total Medicare	Part A	Part B
AREA CHARACTERISTICS							
Income Per Capita	.006	−.002	.011	.071[b]	−.249[a]	−.168[c]	−.513[a]
	(.012)	(.016)	(.023)	(.034)	(.065)	(.099)	(.141)
General Practitioner	.061[a]	.089[a]	.088[a]	.019	.442[a]	.521[a]	.599[a]
	(.016)	(.021)	(.030)	(.044)	(.084)	(.129)	(.183)
All Physicians	−.008	-.001	.069[c]	.135[b]	.412[a]	.334[c]	1.081[a]
	(.026)	(.033)	(.046)	(.067)	(.128)	(.197)	(.272)
Elderly	.065[a]	.100[a]	.051	.054	−.085	−.163	.207
	(.021)	(.028)	(.039)	(.059)	(.112)	(.172)	(.243)
Density	−.087[a]	−.127[a]	−.079[a]	.003	−.087	−.112	−.171
	(.016)	(.021)	(.030)	(.045)	(.085)	(.131)	(.186)
Service Wage	.046[a]	.045[b]	−.122[a]	.218[a]	.101[c]	.230[b]	−.053
	(.013)	(.018)	(.025)	(.038)	(.070)	(.108)	(.152)
OTHER							
Lagged Dependent	.847[a]	.815[a]	.732[a]	.508[a]	.458[a]	.358[a]	.105[b]
	(.022)	(.026)	(.030)	(.036)	(.034)	(.044)	(.042)
Time	.008[a]	.012[a]	.016[a]	.034[a]	.035[a]	.041[a]	.068[a]
	(.002)	(.002)	(.003)	(.003)	(.004)	(.006)	(.009)
R^2	.998	.997	.993	.989	.993	.985	.970
R^2(C)	.998	.997	.992	.988	.993	.983	.967
F	4547	2693	1136	770	1259	536	275
N	623	623	623	623	623	623	623

[a] Significant at the 1 percent level (two-tail test).
[b] Significant at the 5 percent level (two-tail test).
[c] Significant at the 10 percent level (two-tail test).

for total hospital spending. The largest negative effect was for Medicare Part A, which was directly affected by section 1122 controls.

Mature CON reduced bed supply by two percent (long-run effect). However, it raised hospital expense per adjusted patient day and per admission, and also increased hospital profitability (Table 2). Lifting CON had no impact on any of these dependent variables. Section 1122 lowered hospital profits, but the magnitude of this effect appears to be implausibly large.

Mature CON or its removal had no effect on diffusion of technology such as open-heart surgery units, organ transplant units, or ambulatory surgery units (Table 3). Availability of organ transplant units rose immediately after the implementation of CON, but this result could reflect the low number of such units in most states. Pre-CON was not included in any of the technology regressions, and young CON was not included in the regressions for ambulatory surgery, because there were no "young" programs during the observational periods for this analysis.

Both mature CON and CON lifted had positive influences on the for-profit share of the hospital market (Table 4). If a policy objective of retaining CON is to keep the for-profit market share in check, the empirical evidence, if anything, suggests that CON has the opposite effect.

Holding other factors constant, none of the CON variables affected HMO market share; however, the signs on the statistically insignificant coefficients are negative, suggesting that CON may have impeded HMO growth. Section 1122 had significantly positive effects on the for-profit share and a positive but insignificant effect on the HMO share.

In an alternative specification of CON, not shown, we examined whether our findings would persist once we had accounted for differences in stringency of CON across different states. The simplest way of measuring stringency is in terms of thresholds for coverage. States with high thresholds have less stringent programs insofar as fewer projects would qualify for review. We analyzed thresholds for capital and major medical equipment separately, and found very few instances in which these had an impact on the many measures examined. States with high capital thresholds (i.e., with less stringent CON) had lower Part B Medicare spending than did states with no CON.

When stringency was defined in terms of the Lewin-ICF categories described earlier, we found that states with limited CON had worse results than states with no CON. Limited CON states had higher hospital spending per capita and higher Medicare Part B spending per person over age sixty-five. For stringent CON, the effect on hospital spending

was not observed. However, in these states too, Part B spending was comparatively high.

Hospital Rate-Setting

Young state hospital rate-setting programs reduced the rate of growth in hospital expenditures overall, and thereby lowered growth rates in both acute care spending and total spending on personal health care services as well (Table 1). The magnitude of effects was lower for the mature programs. There were no statistically significant effects on expenditures for physicians' services. For Medicare, the mature programs had a stronger effect on hospital spending and on total spending. State rate-setting had no statistically significant effects on hospital bed supply, intensity, hospital profitability (Table 2), or on diffusion of technology with the exception of organ transplant units (Table 3).

Although PPS reduced Medicare expenditures through its effect on Part A expenditures, it seems to have had a positive effect on spending overall. These effects are not attributable to a secular trend in expenditures since we included a time trend as a separate explanatory variable. In contrast to state hospital rate-setting, PPS was negatively related to expense per adjusted admission, to expense per patient day, and to for-profit hospital market share, but was positively related to the HMO market share (Table 4).

Price Competition

Holding other factors constant, the HMO market share was associated with lower hospital bed supply, lower expense per adjusted admission, and lower diffusion of open-heart surgery units, but with greater diffusion of organ transplant units. For expenditures, only the effect of HMO share on Part A expenditures is negative and statistically significant at the 10 percent level or better. We split the sample between the periods 1988 and before and 1989 and after (results not presented). The negative effects of HMO share on Part A Medicare, on diffusion of open heart units, and on the number of hospital beds were statistically significant for the earlier but not for the later period. The HMO coefficient on profit was negative and statistically significant at the 10 percent level for the earlier period, but was insignificant for the latter.

Table 2 Hospital Beds, "Intensity," and Profitability

		Intensity		
	Beds per 1,000 Population	Expense per Adjusted Patient Day	Expense per Adjusted Admission	Hospital Profits
CERTIFICATE-OF-NEED REGULATION				
Section 1122	−.0004	−.007	−.002	−.272[b]
	(.008)	(.012)	(.009)	(.130)
Pre-CON	−.002	.007	.003	.263[a]
	(.006)	(.009)	(.007)	(.101)
Young CON	−.007	.006	.007	.256[a]
	(.006)	(.008)	(.006)	(.093)
Mature CON	−.008[c]	.011[c]	.010[b]	.153[b]
	(.004)	(.006)	(.005)	(.069)
CON Lifted	.002	−.001	.004	.018
	(.005)	(.008)	(.006)	(.085)
HOSPITAL RATE-SETTING				
Prospective Payment System (PPS)	−.095[a]	−.125[a]	−.105[a]	−.395
	(.025)	(.035)	(.027)	(.400)
Young Mandatory Prospective	−.005	.027	.038[c]	−.130
	(.018)	(.026)	(.020)	(.382)
Old Mandatory Prospective	.006	−.003	.005	.157
	(.010)	(.014)	(.011)	(.173)
REIMBURSEMENT				
Medicaid Share	.129[a]	.081[c]	.176[a]	−.689
	(.037)	(.053)	(.041)	(.613)
Medicare Share	−.003	.171[a]	.049[c]	2.020[a]
	(.023)	(.034)	(.026)	(.388)
COMPETITION				
HMO Market Shares	−.111[a]	−.003	−.186[a]	−.897[c]
	(.041)	(.054)	(.045)	(.604)
AREA CHARACTERISTICS				
Income Per Capita	−.044[b]	.021	.004	−.019
	(.018)	(.025)	(.019)	(.306)
General Practitioner	.042[b]	.032	.026	−.062
	(.017)	(.024)	(.019)	(.290)
All Physicians	.215[a]	−.002	.097[a]	−1.096[b]
	(.029)	(.044)	(.033)	(.469)
Elderly	.100[a]	−.019	−.070[b]	−.268
	(.026)	(.036)	(.028)	(.414)

Table 2 Continued

		Intensity		
	Beds per 1,000 Population	Expense per Adjusted Patient Day	Expense per Adjusted Admission	Hospital Profits
Density	−.024	−.005	.066[a]	−.125
	(.020)	(.029)	(.022)	(.312)
Service Wage	−.032[c]	.124[a]	.032[c]	1.175[a]
	(.020)	(.028)	(.022)	(.320)
OTHER				
Lagged Dependent	.616[a]	.803[a]	.801[a]	.318[a]
	(.021)	(.023)	(.021)	(.033)
Time	−.007[a]	.009[a]	.006[a]	.075[a]
	(.001)	(.002)	(.001)	(.017)
R^2	.986	.986	.990	.621
R^2(C)	.985	.984	.989	.586
F	818	802	1178	18
N	863	863	863	818

[a] Significant at the 1 percent level (two-tail test).
[b] Significant at the 5 percent level (two-tail test).
[c] Significant at the 10 percent level (two-tail test).

Discussion

The major findings about CON can be summarized as follows: first, we found no surge in expenditures after CON was lifted; second, despite a statistically significant reduction by mature programs on acute spending per capita, there was no corresponding reduction in total per capita spending (apparently due to offsetting expenditures on nonhospital services).

Empirical analysis of CON is an old topic. What is new or relatively new about our analysis is the research on the effects of lifting CON, the broad range of cost-related outcomes of CON studied, and the analysis of CON and other factors on a recently released data base of personal health care expenditures and their components. Particularly given the long history of empirical analysis of CON, it is important to review our evidence in the context of past research. A scorecard of previous studies of the effects of CON is shown in Table 5. Overall, the record for CON as a cost-containment mechanism appears to be mixed at best. If anything, our results provide slight optimism for CON's cost-containing potential relative to some other studies.

To date, only one other study has used the HCFA per capita spending

Table 3 Diffusion of Technology

	Open Heart Units/ Million	Organ Transplant Units/Million	Hospital-based Units/ Million	Total Units/ Million
CERTIFICATE-OF-NEED REGULATION				
Section 1122	−.069[c]	−.084	.001	.005
	(.046)	(.128)	(.022)	(.025)
Young CON	−.005	.235[c]	(—)	(—)
	(.046)	(.141)	(—)	(—)
Mature CON	−.009	−.071	.007	.012
	(.027)	(.078)	(.015)	(.017)
CON Lifted	.022	.019	.007	.021
	(.027)	(.074)	(.012)	(.013)
HOSPITAL RATE-SETTING				
Prospective Payment	.405[a]	−.278	.206[a]	.155[c]
System (PPS)	(.140)	(.407)	(.073)	(.081)
Young Mandatory Prospective	−.082	−1.427[a]	.009	.085
	(.128)	(.345)	(.095)	(.106)
Old Mandatory Prospective	−.031	.050	.022	.034
	(.054)	(.146)	(.028)	(.031)
REIMBURSEMENT				
Medicaid Share	.181	−1.22[b]	−.063	−.003
	(.190)	(.556)	(.102)	(.113)
Medicare Share	−.334[b]	.669	−.022	.023
	(.146)	(.418)	(.095)	(.105)
COMPETITION				
HMO Market Shares	−.495[b]	2.351[a]	−.050	.149
	(.228)	(.645)	(.118)	(.128)
AREA CHARACTERISTICS				
Income Per Capita	.044	.144	−.136[b]	−.113[c]
	(.101)	(.300)	(.056)	(.062)
General Practitioner	.339[b]	.071	.025	−.109
	(.133)	(.469)	(.078)	(.087)
All Physicians	.299[c]	.236	−.043	−.025
	(.197)	(.615)	(.099)	(.109)
Elderly	−.023	.416	.278[a]	−.001
	(.174)	(.560)	(.099)	(.108)
Density	−.117	−.253	−.216[a]	.066
	(.133)	(.416)	(.070)	(.076)
Service Wage	.060	−.755[b]	.041	.080
	(.113)	(.345)	(.059)	(.065)

Table 3 Continued

	Open Heart Units/ Million	Organ Transplant Units/Million	Hospital-based Units/ Million	Total Units/ Million
OTHER				
Lagged Dependent	.543[a]	.409[a]	.477[a]	.639[a]
	(.036)	(.039)	(.043)	(.038)
Time	.006	.036[b]	−.012[a]	.00001
	(.006)	(.017)	(.003)	(.003)
R^2	.931	.750	.988	.981
R^2(C)	.922	.716	.986	.979
F	112	22	532	337
N	617	541	479	479

[a] Significant at the 1 percent level (two-tail test).
[b] Significant at the 5 percent level (two-tail test).
[c] Significant at the 10 percent level (two-tail test).

data to assess the impact of CON. Examining data through 1982, Lanning, Morrisey, and Ohsfeldt (1991) found that after controlling for the fact that per capita spending was significantly different in states which adopted CON early, CON was associated with a 20.6 percent *increase* in hospital spending and a nine percent increase in spending on other health care. The net impact was a 13.6 percent increase in per capita spending on personal health care services. Using data derived from the annual *Hospital Statistics* on per capita hospital spending through 1990 (AHA 1977–1994) and a method that accounted for endogeneity of CON, Antel, Ohsfeldt, and Becker (1995) reported that CON had no impact on this form of spending, although they found that section 1122 reduced hospital spending. Without controlling for the endogeneity of CON, the coefficient on the CON variable was negative but very small, with a t-ratio of −.47. Taking account of endogeneity, the coefficient on CON became positive and statistically significant at the 10 percent level. It is noteworthy that explicitly accounting for CON's endogeneity made it appear to perform *less* well. Salkever and Bice (1976) found no impact of CON on total hospital operating costs per capita. Likewise, an earlier study by the Federal Trade Commission found that CON had no impact on hospital costs, but also found that section 1122 had a negative influence (Sherman 1988). By contrast, in our study, neither mature CON nor section 1122 had an impact on this type of expenditure, although both were associated with lower growth in acute care spending.

Table 4 Industry Organization

	For-Profit Share of Beds	HMO Market Share
CERTIFICATE-OF-NEED REGULATION		
Section 1122	.211[b]	.436
	(.101)	(.364)
Pre-CON	.121	−.279
	(.115)	(.312)
Young CON	.149	−.176
	(.108)	(.285)
Mature CON	.120[c]	−.155
	(.064)	(.213)
CON Lifted	.139[b]	−.335
	(.059)	(.234)
HOSPITAL RATE-SETTING		
Prospective Payment System (PPS)	−.800[b]	1.357
	(.364)	(1.154)
Young Mandatory Prospective	.369	.971
	(.578)	(.875)
Old Mandatory Prospective	−.195	.341
	(.157)	(.444)
REIMBURSEMENT		
Medicaid Share	.329	.938
	(.420)	(1.575)
Medicare Share	.513[c]	3.837[a]
	(.320)	(1.008)
COMPETITION		
HMO Market Shares	.255	(—)
	(.589)	(—)
AREA CHARACTERISTICS		
Income Per Capita	.289	.0001[c]
	(.243)	(.0001)
General Practitioner	.751[a]	−.075[a]
	(.263)	(.024)
All Physicians	.016	−1.247[a]
	(.370)	(.311)
Elderly	−.684[c]	.035[a]
	(.352)	(.053)
Density	.003	−.0002
	(.248)	(.0006)
Service Wage	−.700[b]	.012[a]
	(.294)	(.004)

Table 4 Continued

	For-Profit Share of Beds	HMO Market Share
OTHER		
Lagged Dependent	.585[a]	.879[a]
	(.039)	(.019)
Time	.016	.038
	(.013)	(.028)
R^2	.961	.976
R^2(C)	.955	.974
F	154	463
N	456	815

[a] Significant at the 1 percent level (two-tail test).
[b] Significant at the 5 percent level (two-tail test).
[c] Significant at the 10 percent level (two-tail test).

In our analysis, adoption of CON was certainly exogenous, but eliminating CON may have been endogenous; that is, it was more likely to have occurred in states where legislatures perceived that cost increases were under control without relying on CON. To ascertain whether this was so, we specified CON lifted and the lagged dependent variable as endogenous variables. Instrumental variables excluded from the main equations were the Blue Cross–Blue Shield market share; share of government hospital beds; population; and values of these variables lagged one year. CON lifted, specified as a endogenous variable, had either no effect or a more negative impact on cost than when the variable was assumed to be exogenous. If the above argument held, one would have expected CON lifted to have had a more positive effect on cost when CON lifted was specified to be endogenous.

Further, in analysis not presented, we used a method developed by Hatanaka (1974) to correct for autocorrelated error terms in a pooled time series cross-section. We found some autocorrelation, both negative and positive, but the correction had only minor effects on our results.

Two newer studies by Lewin-ICF (Lewin-ICF and Alpha Center 1991; Lewin-ICF 1992a) took account of differences in CON stringency and found that CON had a negative impact on hospital costs. This evidence conflicts with ours, since, after accounting for stringency, we did not find that CON had a greater cost-constraining influence. On balance, we believe our results merit more confidence since we controlled for many more influences other than CON.

We found that mature CON reduced hospital bed supply per capita

Table 5 Empirical Studies of the Impact of CON on Hospital Costs

Major Impact	Number of Studies Showing: Decrease	No Effect	Increase
Health Spending			
Spending per capita	0	0	1
Hospital expenses per resident	0	3	2
Total hospital costs	2	1	0
Supply/Utilization			
Hospital capital expenditures	2	5	2
Hospital bed supply	2	3	1
Admissions per 1,000	0	2	0
Intensity			
Cost per patient day	2	1	2
Average length of stay	0	2	0
Cost per admission	0	2	6
Resource Mix			
Assets per bed	0	3	1
Labor use per bed	0	1	1
Market Structure			
For-profit share of beds	1	3	1
Public share of beds	1	0	0

population, but could detect no increase in bed supply following removal of CON. The magnitude of the reduction we detected was small—two percent from mature CON. Using an estimate from Ginsburg and Koretz (1983) that a 1 percent reduction in bed supply results in a .4 percent decline in admissions (the predicted reduction in admissions), the 2 percent reduction in supply translates into less than a 1 percent reduction in admissions. For this reason, it may not be surprising that we show only a minor (statistically insignificant) decline in hospital spending.

One of the earliest studies of CON found that CON reduced hospital bed supply, but also led to increased investment per bed (Salkever and Bice 1976, 1979). The result was no net saving on capital expenditures overall—simply a diversion of spending away from beds into other types of capital equipment that, due to less precise standards for judging need, was less well controlled. Sloan and Steinwald (1980b) also found a compensatory response to CON regulation, but it took the form of higher spending on labor rather than greater investment in other forms of capital. Since then, most studies have found that CON had no detectable impact

on hospital bed supply (Eastaugh 1982; Ashby 1984; Lewin-VHI 1995) or on hospital capital spending (PAI-US 1980; Eastaugh 1982; Begley, Schoeman, and Traxler 1982; Ashby 1984; Wedig, Hassan, and Sloan 1989). In fact, only two studies since the landmark study by Salkever and Bice (1976) found evidence that CON reduces bed supply (Joskow 1980; Begley, Schoeman, and Traxler 1982). Whether the true effect of CON is slightly negative or not, there are certainly better ways to control hospital bed supply, in particular by promoting HMO growth. The effect of HMO share on bed supply in our analysis was over ten times that of mature CON.

We found that mature CON increased cost per adjusted patient day and per admission. The mechanism is presumably that cost-increasing investments are unconstrained or, as Sloan and Steinwald found, there is a compensatory response in use of labor, and as a consequence there is an increase in operating costs. Many previous studies have reported results consistent with ours (Salkever and Bice 1979; Sloan and Steinwald 1980a; Sloan 1981; Farley and Kelly 1985; Noether 1988; Anderson et al. 1989; Lewin-ICF and Alpha Center 1991; and Antel, Ohsfeldt, and Becker 1995). Fewer have found no impact (Sloan 1983; Lewin-VHI 1995).

In this study, the now-defunct section 1122 program had no effect on either cost measure, a result consistent with Antel, Ohsfeldt, and Becker 1995; however, Noether (1988) reported that section 1122 reduced cost per admission by seven percent.

We reviewed eight previous studies that examined the impact of CON on diffusion of technology. In nearly seventy separate tests of the relationship between CON and the rate or extent of diffusion contained in these studies, only about one-third found that CON retards diffusion; a few, like our result for organ transplant units, found that CON accelerates diffusion, but the majority found no effect in either direction. None dealt with ambulatory surgery units; we found that CON had no effect on their diffusion.

Taken at face value, these studies suggest that CON appears to have slowed diffusion of the following technologies: hospital-based cardiac catheterization units, CAT-scan units, and MRI units (Lewin-ICF and Alpha Center 1991); open-heart surgery units (Russell 1979; Lewin-ICF and Alpha Center 1991); hip arthroplasty and morbid obesity surgery (Sloan et al. 1986); cobalt therapy (Russell 1979); and nonhospital-based renal dialysis (Ford and Kaserman 1993).

Yet, for the following reasons, even these favorable findings do not provide unambiguous support for the view that CON retards diffusion of expensive technologies. First, there are conflicting results. For example, although Lewin-ICF (1992a and Lewin-ICF and Alpha Center 1991)

found that CON reduced diffusion of MRIs. Teplensky et al. (1995) reported that more stringent CON policies caused an increase in diffusion of such units. Second, some results are counterintuitive. For example, Sloan et al. (1986) reported that CON had no impact on diffusion of coronary bypass graft surgery (CABG) units, a result consistent with the findings reported here. However, the same analysis showed that CON slowed diffusion of hip arthroplasty and morbid obesity surgery. The latter procedures were not subject to CON review, whereas CABG is subject to review in the vast majority of states with CON. Further, explicit guidelines for review had been developed by the agency responsible for federal oversight of state CON programs. No such guidelines existed for the other types of surgery.

There has been comparatively little research on the effect of CON on market structure. Concerns have been expressed that, absent CON, there will be a flood of for-profit entrants. However, the limited empirical evidence suggests no differential effect of CON on for-profit hospitals (Sloan and Steinwald 1980b). Using a time series of state cross-sections, Wedig, Hassan, and Sloan (1989) showed that the for-profit market share was unrelated to CON. In the current study, we found that mature CON stimulated growth of the for-profit hospital market share, and holding other factors constant, that the share was higher during the immediate period after CON was lifted. Rather than confirming the fears of those who favor retaining CON, our result for CON lifted could reflect a spillover from mature CON. This explanation seems especially likely, given the result for mature CON.

Our finding that CON had negative, albeit insignificant effects on HMO market penetration could reflect endogeneity, although this should have been handled by our fixed-effects analysis. That is, states with low HMO market shares may be reluctant to lift CON. We examined HMO market shares in the year that states lifted CON. They ranged from a high of 24.0 percent for California to lows of 1 percent or less for Idaho, New Mexico, South Dakota, and Wyoming. Preferred provider organization (PPO) penetration was also very low in these states (unpublished data from the American Medical Care and Review Association). Clearly, these states had something other than the presence of high HMO or PPO penetration in mind when they dropped CON. In many of the states that lifted CON, the HMO market share was below the national mean. In all of the states, there has been appreciable growth in managed care since they dropped CON.

Unlike research in many areas of health policy, research into CON

effects on acute care costs provides a rather clear answer. CON has not succeeded in cost containment. Other cost-containment programs appear to work better, but even they appear to have lost their effectiveness as they matured. Certainly, from the regression results presented here and from the descriptive evidence we analyzed but have not reported, there is no reason to fear an expenditure surge after CON laws were lifted. But might CON laws be retained for other reasons?

Might CON improve quality of care? It might do this in at least two ways—first, by assuring adequate patient volume and second, by denying entry to facilities that lack the capacity to deliver high-quality care. There is substantial evidence for one aspect of the former, but no "hard" information on the latter.

Luft et al. (1990) compiled an extensive review of the literature on the volume-outcome relationship that we supplemented with our own review of research published in the 1990s. More than one hundred studies have examined the relationship between hospital volume and outcomes, either mortality or complication rates (e.g., infection rates, rates of reoperation), excessive lengths of stay, or other indicators of patient health status. Although the underlying mechanism is not understood, most studies show higher rates of good outcomes in higher volume facilities. By contrast, there are far fewer studies of the relationship between physician volume and outcomes, and for reasons that are also not well understood, the link between volume and outcomes is less clear.

If the relationship between hospital volume and outcomes is accepted as valid, the question remains whether or not CON increases volume. Only one study has assessed the effect of CON on outcomes directly. Analyzing data from nearly 1,000 hospitals, Shortell and Hughes (1988) found that states with more stringent CON policies or more stringent hospital rate-setting experienced higher mortality rates. Although this analysis would suggest that lifting CON may result in favorable effects on mortality, such an inference would be having it both ways. Given that there appears to be no surge in costs following removal of CON, nor much if any effect of mature or stringent CON on hospital costs, nor much if any effect on diffusion of technology, why CON should have an *adverse* impact on mortality defies explanation.

Finally, there is the potential impact of CON on access. The 1974 National Health Planning and Resources Development Act, which mandated that states have CON, contained several provisions designed to promote better access to care. For example, consumer members were

required to outnumber provider members on local planning boards (Sloan 1988). Also, any Health Systems Agency plan that failed to address needs of low-income persons was subject to challenge at a public hearing.

There is a paucity of empirical studies of effects of CON on access to acute care services. One study conducted in Florida reported that a hospital's success in obtaining CON approval was consistently related to the amount of indigent care that it provided (Campbell and Fournier 1993). A study of California hospitals found evidence consistent with the hypothesis that hospital regulators reward large uncompensated care providers with profitable CON licenses, although no CON variables were actually used in estimating the amount of uncompensated care given by providers (Campbell and Ahern 1993).

Even though this information is suggestive, it is difficult to use it as a basis for continuing to support CON. First, it only applies to two states. Second, there must be more efficient ways to promote access than conferring monopoly franchises on facilities. Efforts to promote access are likely to be more productive if they are focused on primary care providers. Lack of adequate and timely primary care has been found to lead to a significant number of avoidable hospitalizations (Billings et al. 1993)

Earlier studies were more favorable than ours to other regulatory programs such as PPS and state hospital rate-setting relative to CON. It is not that CON has become more effective, but rather that the other programs became worse performers in terms of cost containment as the provider community became more familiar with them.

Conclusion

Our empirical analysis of effects of CON on costs revealed that, at best, CON has had a modest cost-containing influence on hospital and other acute care services. We found no evidence for a surge in acquisition of new facilities or in costs following removal of CON. States that lifted CON did not experience a rise in spending on hospital and physicians' services relative to those that retained it. The conclusion of lack of surge even holds for facilities such as ambulatory surgery units that have experienced substantial growth in recent years. It is doubtful that CON has had much of a positive or negative influence on quality of care. CON may have improved access, but the empirical evidence for this is quite meager.

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