The Effects of Price Competition and Reduced Subsidies for Uncompensated Care on Hospital Mortality

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Objective. To determine whether hospital mortality rates changed in New Jersey after implementation of a law that changed hospital payment from a regulated system based on hospital cost to price competition with reduced subsidies for uncompensated care and whether changes in mortality rates were affected by hospital market conditions.

Data Sources/Study Setting. State discharge data for New Jersey and New York from 1990 to 1996.

Study Design. We used an interrupted time series design to compare risk-adjusted in-hospital mortality rates between states over time. We compared the effect sizes in markets with different levels of health maintenance organization penetration and hospital market concentration and tested the sensitivity of our results to different approaches to defining hospital markets.

Data Collection/Extraction Methods. The study sample included all patients under age 65 admitted to New Jersey or New York hospitals with stroke, hip fracture, pneumonia, pulmonary embolism, congestive heart failure, hip fracture, or acute myocardial infarction (AMI).

Principal Findings. Mortality among patients in New Jersey improved less than in New York by 0.4 percentage points among the insured ($p = .07$) and 0.5 percentage points among the uninsured ($p = .37$). There was a relative increase in mortality for patients with AMI, congestive heart failure, and stroke, especially for uninsured patients with these conditions, but not for patients with the other four conditions we studied. Less competitive hospital markets were significantly associated with a relative decrease in mortality among insured patients.

Conclusions. Market-based reforms may adversely affect mortality for some conditions but it appears the effects are not universal. Insured patients in less competitive markets fared better in the transition to price competition.

Key Words. Quality of health care, health care markets, hospital competition, mortality

Market-based reforms have been implemented throughout the United States in the past decade to slow recent increases in health care costs, but little is known about how these reforms have affected the quality of care (Gaynor and Haas-Wilson 1999; Gowrisankaran and Town 2003). In the early 1980s, about 30 states employed some form of hospital rate-setting as a cost-containment device, but none of these states except Maryland still uses hospital rate-setting (McDonough 1997). In 1993, New Jersey's Health Care Reform Act (HCRA) replaced an all-payer system that regulated hospital rates with a system that allowed insurers to selectively contract with hospitals and negotiate prices. In addition, HCRA reduced subsidies to hospitals for providing uncompensated care by ending the 19 percent surcharge on all hospital bills that had been earmarked for this purpose (Volpp and Siegel 1993). As a result, total subsidies for uncompensated care in New Jersey fell from \$700 million in 1992, the year before HCRA, to \$500 million in 1993, \$450 million in 1994, \$400 million in 1995, and \$350 million in 1996 (Cantor 1993; Bovbjerg, Cuellar, and Holahan 2000).

The New Jersey policy reform changed hospitals' incentives. Under the rate-setting system, in which rates were set based on incurred costs, hospitals competed for patients by providing more (and presumably better) services. Numerous studies in other contexts have shown that quality competition leads hospitals in more competitive areas to have higher costs and greater availability of technologically sophisticated services and amenities (Held and Pauly 1983; Farley 1985; Robinson and Luft 1985; Robinson and Luft 1987).

Under price competition, by contrast, insurers may negotiate price discounts with hospitals, and hospitals must compete with one another for patients on a combination of price and quality. As price competition intensifies, hospitals may grant discounts to attract and retain patients and may consequently provide lower-quality care. Because the shift to price competition in New Jersey placed greater emphasis on price relative to quality, the likely results were lower prices and lower hospital margins (Zwanziger, Melnick, and Bamezai 2000). Lower margins could have affected the quality of hospital care adversely (Volpp et al. 2003a) because monitoring many of the technical aspects of care is difficult (Arrow 1963; Weisbrod 1989).

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Our previous work demonstrated that uninsured patients with acute myocardial infarction (AMI) admitted to hospitals in New Jersey had a higher mortality rate after HCRA than before HCRA relative to patients in both New York and the Nationwide Inpatient Sample from the Hospital Cost and Utilization Project (HCUP) (Volpp et al. 2003a). Mortality for insured AMI patients, however, did not change. In the present study, we examined the effects of HCRA on mortality for six other common inpatient conditions and how changes in mortality varied in accordance with hospital market conditions.

BACKGROUND

Our underlying conceptual framework is based on the concept that not-for-profit hospitals (all hospitals in New Jersey during this time period were not-for-profits) derive utility from profits, uncompensated care provision, and provision of quality (Newhouse 1970; Sloan and Steinwald 1980; Gruber 1994). Reductions in net revenues, therefore, may lead to reductions in both quality of care for all patients and the quantity and quality of services provided to the uninsured. Because price competition between hospitals depends on both competition between hospitals and the presence of large-volume insurers such as health maintenance organizations (HMOs) that can effectively negotiate price discounts and have strong incentives to do so (Melnick et al. 1992; Dranove and White 1994; Bamezai et al. 1999; Cutler, McClellan, and Newhouse 2000), we hypothesized that any adverse effects on quality of care would be greatest in more competitive markets with higher HMO penetration and least in less competitive markets with lower HMO penetration. In addition, these effects would be expected to be larger in hospitals with greater proportions of uninsured patients, because these hospitals were more affected by reductions in the subsidies for uncompensated care (Cantor 1993).

The ability of insurers to negotiate price discounts in exchange for sending large numbers of patients to particular providers post-rate-setting led to increased competitive pressure on prices that reduced the rate of increase in net revenues in New Jersey relative to New York, putting pressure on hospital margins (Volpp et al. 2003b). As hospital price-cost margins can be used to finance hospital objectives which are not revenue producing such as uncompensated care or quality enhancements above the profit-maximizing level, increased pressure on these margins has implications for the amount of resources available for hospitals to use for these purposes.

The impact of the transition to price competition from hospital rate-setting on hospital margins suggests that hospital outcomes will worsen under

price competition compared with rate regulation and that this effect will be most strongly manifested in more competitive hospital market areas or markets with high levels of HMO penetration. Increased pressure on hospital margins would also imply that the change in outcomes in the transition to price competition will be worse among uninsured patients than among insured patients, as funding for their care, in the absence of explicit subsidies, is provided from hospital margins (Gruber 1994).

METHODS

Study Sample and Control Group

We used hospital discharge data from New Jersey and New York for 1990–1996 to analyze in-hospital mortality among patients with one of seven conditions. We used recommendations from the Agency for Healthcare Quality and Research's (AHRQ's) Quality Indicator report to identify six of these conditions—hip fracture, stroke, AMI, gastrointestinal bleeding, congestive heart failure, and pneumonia—because the agency concluded that mortality rates for these conditions are important identifiers of hospital quality (AHRQ 2003). The seventh condition, pulmonary embolism (ICD-9 CM diagnosis codes 415.1x), met our criteria as a relatively common condition with significant mortality risk that should be affected by hospital quality.

In the first part of the analysis, we examined the changes in overall mortality and the changes in mortality related to insurance status for six of the seven conditions, because we already have reported the changes for AMI (Volpp et al. 2003a). In the second part of the analysis, we used all seven conditions to examine the relationship between changes in mortality and the level of hospital competition and the hospital's proportion of HMO or uninsured patients.

To adjust for improvements in the quality of care over time, we used New York as a control state. New York was chosen because it has a large population; it is an adjacent state; similar patient data were available; there were no major changes in its hospital financing system during the study; and we verified that the mortality time trends prior to the passage of the reform were not statistically different in the two states (Campbell and Stanley 1963).

New Jersey data came from Uniform Billing-82 and Uniform Billing-92 reports, while New York data came from the Statewide Planning and Research Cooperative System (SPARCS) Inpatient Output Data Dictionary. Both states required hospitals to submit data on all patients treated. Because we were

interested in comparing the quality of care received by insured and uninsured patients, we focused on patients under age 65 not enrolled in Medicare. We defined our primary dependent variable as 30-day in-hospital mortality. In addition to mortality, we included data on patient age, sex, race, hospital where treated, length of stay, source and type of admission, discharge date, expected principal payer, principal diagnosis, principal procedure, and any secondary diagnoses and procedures. Finally, the results we present are from estimates of the six or seven conditions pooled together. Because the baseline mortality rates of these conditions differ, we include an indicator variable for each condition to account for any changes in the proportion of patients in the pooled sample with that condition.

There were 710,264 hospitalizations in the two states for non-Medicare patients under the age of 65 with one of the seven study conditions. We excluded patients who were less than 18 years old ($n = 106,266$); transferred out to another hospital ($n = 57,264$) because we could not determine whether the patient subsequently died; stayed in the hospital longer than 30 days ($n = 16,309$); transferred in from a skilled nursing facility ($n = 1,727$) because such patients may be treated less aggressively; were residents of a state other than the one in which they received treatment ($n = 14,752$); had condition-specific exclusions defined by the AHRQ Quality Indicators ($n = 5,159$) or who lacked patient zip code ($n = 9,540$) or market variable data ($n = 2,914$). We also excluded patients who were discharged alive in 1 day or less ($n = 18,172$) because these patients either were miscoded or had extremely mild cases. In addition, we removed AMI admissions with lengths of stay less than 2 days if the patient was transferred in from another hospital or less than 4 days if not transferred in ($n = 8,521$). These exclusions, which had similar rates in both states, left us with 469,629 discharges.

Competition and Payer Mix

We used two different approaches to measuring hospital market areas (Garnick et al. 1987; Baker 2001) to test the sensitivity of our results to market definition. The two types of hospital market areas definitions used were Health Service Areas (HSAs) (Makuc et al. 1991), a standard geographic measure, and hospital-specific markets constructed using patient flow data (Zwanziger and Melnick 1988). HSAs have face validity as representing the geographic area from which hospitals get patients. In addition, the fixed boundaries of HSAs mean that hospital market definition is not affected by changes in patient flow due to patient perceptions of quality. We created a Hirschman-Herfindahl

index (HHI) for each HSA by summing for all HSA residents the squares of each hospital's market share (based on hospital admissions) in 1992, the final year under hospital rate-setting. This allowed us to focus on the hospital market conditions at the time of initiation of price competition.

We also defined hospital-specific markets using patient flow data from 1992. We first calculated zip code-specific HHIs. Then, based on the percentage of patients residing in each zip code who utilized each hospital we constructed a weighted average of the zip code HHIs for each hospital. Defining hospital markets based on patient flow is at least potentially endogenous, as higher quality hospitals may attract patients from greater distances. However, as we are interested in examining how quality changed following implementation of price competition in New Jersey based on the initial hospital market conditions and as we adjust for baseline differences in quality between markets, the endogeneity of hospital markets based on patient flows to quality is less of a concern.

For each hospital, we also calculated the proportion of uninsured and HMO patients in 1992 from the patient discharge data.

Empirical Specification

We used linear probability models to examine how risk-adjusted mortality changed from the pre-reform period to the post-reform period in New Jersey relative to New York. We used an approach similar to that used by McClellan and colleagues in developing the AHRQ Inpatient Quality Indicators by using ordinary least squares regression to examine individual patient mortality risk (Davies et al. 2001). The models allowed baseline mortality to differ between New Jersey and New York and assumed that New Jersey and New York had a common time trend in mortality rates until the implementation of HCRA but not afterwards. To test the validity of New York as a control, we compared the time trends for New Jersey with those of New York before HCRA (from 1990 to 1992) by testing whether the coefficients for the interaction terms for New Jersey combined with 1991 and 1992 were jointly equal to zero.

Our primary study variables were interactions. In the first part of this project, we measured the magnitude and statistical significance of the coefficients on the interactions that measured the relative change in mortality in New Jersey versus New York from 1990–1992 to 1994–1996 by insurance status. Fixed-effects hospital dummy variables were added to examine whether any observed changes were within versus between hospitals.

We controlled for the following comorbidities: types of cancer with a poor prognosis, metastatic cancer, AIDS, chronic pulmonary disease, congestive

heart failure, peripheral vascular disease, severe chronic liver disease, diabetes with end-organ damage, chronic renal failure, nutritional deficiencies, dementia, and functional impairment (Iezzoni et al. 1992; Needleman et al. 2002). Patient gender, age, and age squared also were included in the models.

We tested the appropriateness of excluding elderly and Medicare patients by comparing mortality for under-65, non-Medicare, insured patients with mortality for patients on Medicare or age 65 and older. We determined that they were similar because differences in the relative changes in mortality differences between states were statistically significant only for hip fracture.

In the second part of this analysis we tested whether mortality trends varied with the competitiveness of hospital markets and with the hospital proportion of uninsured patients. We classified hospitals as “more competitive” if they had both an HHI below the New Jersey median (0.052 using HSA markets and 0.302 using markets defined with patient flows) and HMO penetration above the New Jersey median in 1992 (8.0 percent) or “less competitive” if they had both an HHI above the New Jersey median and HMO penetration below the New Jersey median in 1992, with the remaining hospitals serving as the referent group. We regressed 30-day, in-hospital mortality against this measure of hospital market competitiveness while including interactions between the main effects and the state and time dummy variables and controlling for patient risk factors. We also tested whether these results were affected by the percentage of uninsured patients within each hospital in 1992.

To counteract the heteroscedasticity present in all linear probability models, we calculated Huber–White robust standard errors (White 1980), and we also accounted for the clustering of patient discharges within hospitals. Data analysis and model estimation were carried out with the *SAS* 8.2 and *Stata* 7 statistical software packages.

RESULTS

The unadjusted mortality rate improved steadily in both states (Table 1). While there were baseline differences in the number of admissions and proportion of uninsured patients in the two states, the rates of change were similar. Length of stay decreased at similar rates in New Jersey and New York. HSA-based markets had lower HHIs than markets defined using patient flow because there were fewer in each state and thus more hospitals within each market. The New Jersey median HHI was 0.052 in HSA-based markets and 0.302 for the patient flow-based markets.

Table 1: Basic Characteristics of Data Used in the Sample

<i>Basic Characteristics of Patients with Hip Fracture, Stroke, Gastrointestinal Bleeding, Pulmonary Embolism, Congestive Heart Failure, or Pneumonia</i>							
	<i>1990</i>	<i>1991</i>	<i>1992</i>	<i>1993</i>	<i>1994</i>	<i>1995</i>	<i>1996</i>
Admissions (<i>n</i>)							
New Jersey	14,846	14,499	14,766	15,226	16,157	15,437	14,787
New York	35,635	34,995	35,880	39,873	41,982	37,097	36,081
Proportion of uninsured patients (%)							
New Jersey	16.8	17.9	19.1	19.3	18.9	20.0	20.9
New York	10.6	10.1	10.2	10.8	9.9	11.2	13.1
Average length of stay—insured patients							
New Jersey	8.6	8.6	8.5	8.1	7.6	7.1	6.7
New York	8.3	8.3	8.4	8.2	8.0	7.3	7.0
Average length of stay—uninsured patients							
New Jersey	8.5	8.3	8.3	8.1	7.7	7.0	6.6
New York	7.4	7.5	7.4	7.5	7.0	6.5	6.3
Insured mortality rate (%)							
New Jersey	5.0	5.3	4.9	4.9	4.7	4.3	3.9
New York	5.8	6.0	6.0	5.5	5.3	4.9	4.6
Uninsured mortality rate (%)							
New Jersey	6.5	8.0	7.2	5.9	6.8	5.7	4.7
New York	7.7	7.9	7.3	7.5	5.6	5.3	4.8
<i>Basic Characteristics of Markets</i>							
<i>Hirschman–Herfindahl Index</i>	<i>HHI Mean</i>	<i>HHI Min</i>	<i>HHI Max</i>	<i>Number of Markets</i>	<i>Cutpoint</i>	<i>Proportion of Hospitals under Cutpoint</i>	
HSA-based markets							
New Jersey	0.16	0.043	0.343	9	0.052	47/86	
New York	0.19	0.030	0.477	22	0.052	111/224	
Patient flow-based measures							
New Jersey	0.32	0.17	0.69	84	0.302	42/84	
New York	0.29	0.11	0.87	222	0.302	134/222	

HSA, health service area; HHI, Hirschman–Herfindahl index.

The improvements in risk-adjusted mortality that were common to both New Jersey and New York were substantial (regression results from Table 2), as mortality decreased by 1.0–1.7 percentage points per year among the insured and 1.7–2.6 percentage points per year among the uninsured from 1994 to 1996. Mortality decreased less from the pre-HCRA (1990–1992) to post-HCRA (1994–1996) in New Jersey than in New York by 0.4 percentage points ($p = .07$). The magnitude of the difference in intertemporal trends between states was slightly smaller for insured patients (0.4 percentage points, $p = .07$)

Table 2: Overall Effects on Risk-Adjusted Mortality for Patients in New Jersey versus New York for Patients with Hip Fracture, Stroke, Gastrointestinal Bleeding, Pulmonary Embolism, Congestive Heart Failure, or Pneumonia

<i>Hospital Fixed Effects</i>	<i>Insured</i>		<i>Uninsured</i>		<i>Overall</i>	
	<i>No</i>	<i>Yes</i>	<i>No</i>	<i>Yes</i>	<i>No</i>	<i>Yes</i>
New Jersey						
1991	-0.009***		-0.008		-0.007	
1992	0.000	0.000	0.002	0.003	0.000	0.000
1993	-0.003	-0.003	-0.004	-0.004	-0.003	-0.003*
1994	-0.007***	-0.007***	-0.006	-0.006	-0.007***	-0.007***
1995	-0.010***	-0.011***	-0.018***	-0.017***	-0.011***	-0.011***
1996	-0.012***	-0.013***	-0.020***	-0.019***	-0.013***	-0.014***
	-0.016***	-0.017***	-0.026***	-0.026***	-0.017***	-0.018***
Coefficient of the interaction term for being in New Jersey and occurring after the new law	0.004	0.004	0.005	0.005	0.004	0.004
<i>p</i> -value for interaction term						
Number of observations	.073	.078	.371	.408	.069	.083
R^2	318,862	318,862	48,339	48,339	367,261	367,261
	0.106	0.110	0.209	0.206	0.117	0.121
Test of controls (<i>p</i> -value)	.13 (.88)	.13 (.88)	.43 (.65)	.43 (.65)	.17 (.84)	.200 (.82)

Notes: Models adjust for age, gender, cancer with a poor prognosis, metastatic cancer, AIDS, chronic pulmonary disease, coronary artery disease, congestive heart failure, peripheral vascular disease, severe chronic liver disease, diabetes with end-organ damage, chronic renal failure, nutritional deficiencies, dementia, functional impairment, baseline differences in the two states, and intertemporal trends common to both states, and which of the six conditions the patient had.

Asterisks indicate the level of significance:

* $p < .05$,

*** $p < .01$,

*** $p < .0001$.

Test of differences between uninsured and insured is not significant for both fixed-effects and non-fixed-effects models ($p > .10$).

than for uninsured patients (0.5 percentage points, $p = .37$). The difference in the magnitude of the effects between the insured and uninsured was not significant ($p = .9$). Models with fixed-effects hospital dummy variables show no difference in the main coefficients of interest, indicating that observed changes were within the same hospitals as opposed to patients being treated more frequently at lower quality hospitals (Table 2). The F -test of controls indicates that the rate of change in mortality rates was not statistically different in New Jersey and New York for insured patients, uninsured patients, and patients overall during the 3 years before New Jersey's law was implemented in 1993.

There was substantial variation among the six conditions in the degree to which mortality changed in New Jersey relative to New York (Figure 1). There were statistically significant increases in mortality for uninsured patients with congestive heart failure (1.4 percentage points, $p = .04$) and for insured patients with stroke (3.3 percentage points, $p = .05$). The largest increase was for uninsured patients with stroke (5.5 percentage points, $p = .21$), but this was not statistically significant. There were no significant improvements in mortality for patients in New Jersey relative to New York for any of the conditions.

In the second part of the analysis, we included patients with AMI and measured whether the level of market competitiveness for hospital services affected changes in mortality. Among the insured patients, we found that New Jersey patients hospitalized in less competitive markets had 1.2 percentage point smaller increases in mortality relative to New York than patients hospitalized in markets of average competitiveness regardless of whether hospital markets were defined using HSAs or patient flows ($p = .024$ and $.028$) (Figure 2 and Model A in Table 3). In contrast, New Jersey patients hospitalized in more competitive markets had larger increases in mortality relative to New York than patients hospitalized in markets of average competitiveness in HSA-defined markets (0.6 percentage points, $p = .19$) but lower mortality risk in patient flow defined markets (-0.7 percentage points, $p = .14$). Neither of these values was significantly different than zero (Figure 2, Table 3). The difference in the magnitude of the relative change in mortality between more competitive and less competitive markets was significant using the HSA-based model ($p = .004$) but not using the patient flow-based model ($p = .39$). Among uninsured patients, there were no statistically significant relative changes in mortality compared with markets of average competitiveness in either less competitive or more competitive markets, and the magnitude of the coefficients in both types of markets was not significantly different ($p > .6$).

We also examined whether changes in mortality were affected by the percentage of uninsured patients in each hospital during 1992 (Model B in

Figure 1: Change in Mortality in New Jersey versus New York

Models adjust for age, gender, cancer with a poor prognosis, metastatic cancer, AIDS, chronic pulmonary disease, coronary artery disease, congestive heart failure, peripheral vascular disease, severe chronic liver disease, diabetes with end-organ damage, chronic renal failure, nutritional deficiencies, dementia, functional impairment, and intertemporal trends common to both states. Asterisks indicate the level of significance: $*p < .05$.

Table 3). Although the \$350 million reduction in charity care subsidies represented approximately 4.2 percent of hospital net revenues overall, the proportion of uninsured patients was not associated with significant relative increases in mortality either among the insured (0.4 percentage points, $p = .33$), the uninsured (0.3 percentage points, $p = .80$), or overall (0.4 percentage points, $p = .26$). The results were qualitatively similar using patient flow markets. Adding the proportion of uninsured patients changed the significance level of the coefficient that measured the relative change in New Jersey's hospital mortality, but did not change the sign and significance of the

Figure 2: Change in Mortality in Different Markets in New Jersey versus New York—Seven-Condition Sample—Insured Patients Only

Models adjust for age, gender, cancer with a poor prognosis, metastatic cancer, AIDS, chronic pulmonary disease, coronary artery disease, congestive heart failure, peripheral vascular disease, severe chronic liver disease, diabetes with end-organ damage, chronic renal failure, nutritional deficiencies, dementia, functional impairment, baseline differences in the two states, and intertemporal trends common to both states. Level of significance for comparison of low- and average-competition markets: * $p < .05$, ** $p < .01$. Level of significance for comparison of low- and high-competition markets: # $p < .05$, ## $p < .01$.

coefficients that measured the effects of the level of competition on changes in mortality risk. These effects taken together suggest that price competition had more of an impact on hospital mortality post-HCRA than the reductions in subsidies for hospital care for the uninsured. We repeated the analyses reported in Table 3 with hospital fixed effects included. The results were nearly identical to those reported above, implying that the observed differences resulted from within-hospital changes rather than from changes in admission patterns to hospitals with different levels of quality.

Table 3: Change in Mortality in New Jersey versus New York in More Competitive versus Less Competitive Markets—Seven-Condition Sample

	Insured				Uninsured			
	HSA-Based Markets		Patient Flow-Based Markets		HSA-Based Markets		Patient Flow-Based Markets	
	Model A	Model B	Model A	Model B	Model A	Model B	Model A	Model B
Coefficient of the interaction term for being in New Jersey post-reform	0.004 (0.069)	0.003 (0.350)	0.007** (0.006)	0.005 (0.059)	0.009 (0.111)	0.006 (0.545)	0.008 (0.146)	0.006 (0.547)
Coefficient of the interaction term for being in New Jersey post-reform in a less-competitive market	-0.012* (0.028)	-0.012* (0.049)	-0.012* (0.024)	-0.012* (0.036)	0.005 (0.774)	0.007 (0.735)	0.0004 (0.822)	0.005 (0.793)
Coefficient of the interaction term for being in New Jersey post-reform in a more-competitive market	0.006 (0.193)	0.007 (0.132)	-0.007 (0.141)	-0.006 (0.184)	0.017 (0.301)	0.019 (0.234)	0.008 (0.594)	0.008 (0.613)
Coefficient of the interaction term for being in New Jersey post-reform in a hospital with a higher proportion of uninsured patients		0.004 (0.333)		0.003 (0.493)		0.003 (0.802)		0.002 (0.890)

<i>N</i>	410,359	410,359	410,359	410,359	410,359	59,270	59,270	59,270	59,270
Adjusted R^2	0.042	0.042	0.042	0.042	0.042	0.046	0.046	0.046	0.046
<i>F</i> -test (more competitive = less competitive)	8.33	7.99	0.741	0.741	0.741	0.239	0.288	0.031	0.017
<i>p</i> -value	.004	.005	.378	.39	.626	.592	.861	.896	

Notes: Model B differs from Model A by the inclusion of an interaction term of New Jersey post-reform with the proportion of uninsured patients within a given hospital.

Models adjust for age, gender, cancer with a poor prognosis, metastatic cancer, AIDS, chronic pulmonary disease, coronary artery disease, congestive heart failure, peripheral vascular disease, severe chronic liver disease, diabetes with end-organ damage, chronic renal failure, nutritional deficiencies, dementia, functional impairment, baseline differences in the two states, and intertemporal trends common to both states.

P-values added in parentheses below coefficients.

Asterisks indicate the level of significance:

* $p < .05$, ** $p < .01$

HSA, health service area.

Comment

Our findings in this study extend our earlier findings in a similar study of patients with AMI. Our previous work demonstrated that uninsured but not insured patients with AMI admitted to hospitals in New Jersey had a significantly higher mortality following implementation of a law that encouraged price competition among hospitals and reduced subsidies for hospital care of the uninsured (Volpp et al. 2003a). In this study, we found changes in the same direction for patients with six other medical conditions, but the changes were smaller in magnitude and there was no systematic evidence of a larger relative worsening in mortality for the uninsured. The size of the change for insured patients was greater in this study than our previous study (0.4 percentage points versus 0.1 percentage points) but smaller for uninsured patients (0.5 percentage points versus 3.1 percentage points). These changes are smaller than the intertemporal reductions in mortality common to both New Jersey and New York (1.0–2.6 percentage points per year from 1994 to 1996). The combined results from this study and the original AMI study are qualitatively similar to the results from this study, but the addition of patients with AMI makes the relative increase in mortality among the uninsured larger (1.1 percentage points versus 0.5 percentage points) and statistically significant ($p = .05$ versus $p = .37$) (Figure 3).

Based on these findings, we conclude that while the change in the New Jersey law had only a modest effect in increasing overall mortality, it had a real and important effect on mortality for patients with AMI, congestive heart failure, and stroke, especially for uninsured patients with these conditions.

This conclusion is supported by our previous finding that worsening mortality for uninsured patients with AMI in New Jersey was associated with decreased rates of cardiac catheterization and mechanical revascularization (Volpp et al. 2003a), which we believe are markers for changes in hospital care that could explain the increased mortality. In addition, this conclusion is supported by our finding in this study that relative mortality increased less in markets where price competition was weaker. We believe that these findings are plausible parts of a possible explanation for the changes in hospital care that led to increased mortality. More competitive markets put greater pressure on hospitals to lower prices, and lower prices likely led to changes in patient care in attempts to control the cost of care. It is not clear why there was a relative increase in mortality for some conditions but not others. There were no conditions for which there was a relative improvement in mortality for New Jersey compared with New York, so we think the changes measured reflect

Figure 3: Percentage Point Change in Mortality for all Patients

Models adjust for age, gender, cancer with a poor prognosis, metastatic cancer, AIDS, chronic pulmonary disease, coronary artery disease, congestive heart failure, peripheral vascular disease, severe chronic liver disease, diabetes with end-organ damage, chronic renal failure, nutritional deficiencies, dementia, functional impairment, baseline differences in the two states, intertemporal trends common to both states, and, for the pooled estimates, which condition the patient had. Asterisks indicate the level of significance: $*p < .05$.

true changes as opposed to chance observations. The consistency between models with hospital fixed effects and those without rules out the explanation that changes in patient flows to hospitals influenced our results. The proportion of uninsured patients within hospitals did not play a role in determining the increase in post-reform mortality so it appears that the reduction in charity care subsidies was not a significant independent mechanism for relative worsening in mortality.

Insured patients treated at hospitals that faced relatively low competitive pressure had a 1.2 percentage point lower change in mortality rate compared

with the degree of change in average markets in New Jersey relative to New York using both hospital market definitions. Our findings for patients in highly competitive markets, on the other hand, were affected by which definition we used for a market and are thus inconclusive. Among the uninsured, we found no difference in their degree of change in mortality in different markets in New Jersey compared with New York using different hospital market definitions.

Few others have studied how hospital outcomes differ based on market competitiveness. Cross-sectional studies by Shortell and Hughes (1988), Sari (2002), and Mukamel, Zwanziger, and Tomaszewski (2001) found no significant relationships between competition and mortality outcomes. Sari (2002) found that some complications were significantly associated with hospital market share and HMO penetration but that others were not. Kessler and McClellan (2000) demonstrated in a Medicare population of AMI patients that while greater competition was associated with higher costs and had inconsistent effects on quality before 1991, hospital competition was associated with lower costs and improved outcomes after 1990. In their study, the association of competition with lower mortality for Medicare patients post-1990 was seen only in states with above-median HMO enrollment but not in states with HMO penetration below the median. Because their study was based on a national sample, it measured the effect of hospital exit from markets under a mixture of state-level regulatory regimes. In addition, by studying only Medicare patients, they measured the impact of changes in financial incentives different from the ones in New Jersey. Gowrisankaran and Town (2003) found that the associations of outcomes and competition varied by payer type. More competition among HMO patients in Southern California was associated with lower mortality rates for AMI and pneumonia (about 3 percent lower mortality for every 10 percent increase in the HHI), but more competition among Medicare enrollees was associated with a higher mortality rate (1.5 percent higher pneumonia mortality and 2.3 percent higher AMI mortality for every 10 percent increase in the HHI). The difference in mortality they estimate for a 10 percent difference in the HHI are about 0.4 percentage points for AMI and 0.3 percentage points for pneumonia. Our results may differ because a nascent price competitive environment such as that in New Jersey immediately post-HCRA presumably put a much bigger emphasis on price as opposed to quality than might be the case in a market in which price competition has already existed for years, as it had in the Southern California hospitals studied by Gowrisankaran and Town. In addition, we measure within-hospital and within-market changes in outcomes

over time in response to a policy change as opposed to cross-sectional associations.

We examined several other possible explanations for the differential change in mortality rates between New Jersey and New York. Because we observe only in-hospital mortality, higher rates of mortality in New Jersey post-reform could be as a result of larger decreases in length of stay in New York from 1990–1992 to 1994–1996, but we found that average length of stay for both insured and uninsured patients decreased at similar rates in New Jersey and New York (p -value = .08 for the insured, .15 for the uninsured). Market-specific effects could be biased by changes in length of stay if length of stay decreased more in New Jersey than New York in less competitive areas while decreasing more in New York than New Jersey in more competitive areas. However, we found that among each patient group (insured, uninsured), the length of stay changed at similar rates in New Jersey and New York in both less competitive areas ($p > .18$) and more competitive areas ($p > .17$). We are not aware of any policy reforms implemented in New York during this period that could have systematically increased quality of care in New York relative to New Jersey for the conditions we studied. In an earlier work (Volpp et al. 2003a), we showed that the Cardiac Surgery Reporting System introduced in New York in 1989 did not contribute to the observed relative worsening in AMI mortality in New Jersey post-reform, and it is very unlikely that this system would have affected mortality for other conditions. Another potential explanation is that we excluded the hospitalizations of patients transferred to other hospitals. If the proportion of transfers increased in New Jersey relative to New York in less competitive markets during this period, then we would have excluded more patients who were discharged alive, which would have artificially increased New Jersey death rates in these markets. However, there were higher rates of increase in transfers in less competitive markets in New Jersey compared with New York, and in these markets we observed relative decreases in mortality in New Jersey. The fact that we observed higher rates of patients being transferred into hospitals in less competitive markets in New Jersey coupled with the observation that there were no substantial differential changes in the average number of comorbidities of patients transferred in New Jersey and New York across markets suggests that there was no systematic pattern of transferring sicker patients away from New Jersey hospitals in less competitive areas, another possible explanation for the relative decrease in mortality in these markets. Differential changes in admission thresholds could also affect mortality rates, but we found that admission rates (defined as number of admissions from discharge data divided by HSA

population in 1990) changed at about the same rate in more competitive markets in New Jersey and New York (14.4 versus 15.1 percent, respectively). In less competitive markets, admission rates rose 17.6 percent in New York and 11.8 percent in New Jersey, which would bias against finding lower mortality rates in less competitive markets in New Jersey relative to New York.

In assessing these results, other limitations should be kept in mind. We could not directly measure the reductions in subsidies for hospital care for the uninsured and had limited statistical power for market-specific analyses of the effects on the uninsured. Our analysis is limited to seven medical conditions, and thus we do not know whether these results are generalizable to patients with other diagnoses. While mortality is clearly an important outcome measure for these conditions, our analysis is limited by using in-hospital mortality as the sole outcome measure. We adjusted for risk using administrative data and thus did not have detailed clinical information for the adjustments. Risk adjustment was less important in this study, however, because we primarily studied patients with medical conditions that are commonly admitted to the hospital once diagnosed, and we included every eligible admission in both states. The number of admissions for our six conditions increased by 4.9 percent from the pre-reform period (1990–1992) to the post-reform period (1994–1996) in New Jersey and increased 7.6 percent during the same time period in New York. The relative stability supports our assumption that the number of patients admitted from 1 year to the next within the same geographic area was roughly constant, making changes in outcomes as a result of unmeasured differences in admission severity less likely. As New Jersey is a densely populated state, there is not a large range in the degree of hospital market concentration, so the differences we observed within New Jersey could be greater in a state with more variation in the degree of hospital market concentration. New Jersey did not collect information on the transaction prices paid by insurance companies to hospitals, so we could not directly examine cost–quality tradeoffs. Other investigators have shown that this type of reform is likely to cut costs and shrink hospital margins (Zwanziger and Melnick 1988; Dranove, Shanley, and White 1993; Sorenson 2003).

Our results demonstrate that market-based reforms designed to reduce the rate of increase in health care costs may affect mortality among hospital patients, but there was no evidence that the uninsured in New Jersey were disproportionately affected. Further studies should examine the effects on the care provided for different types of conditions and in other policy contexts.

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