Do Physician Incentives Affect Hospital Choice? A Progress Report*

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Abstract

The US health reforms of March 2010 introduce new provisions for physicians providing Medicare and Medicaid services to be given financial incentives to control costs. Physician payment mechanisms generating similar incentives are currently used by some health maintenance organizations in California. We describe an ongoing research project in which we investigate physician responses to these payment schemes. The question is whether patients whose physicians have incentives to control hospital costs are admitted to lower-priced hospitals than other patients, all else equal. We provide an initial analysis of California hospital discharge data from 2003, documenting evidence consistent with this hypothesis.

1 Introduction

The Affordable Care Act of March 2010 introduces comprehensive health reforms that expand health insurance coverage, subsidize premiums and increase consumer choice. The costs of these provisions are partially offset by increased taxes and fees on various entities (including new Medicare taxes on high-income brackets and fees on medical devices and pharmaceuticals). In the long term, however, many policy-makers believe that cost controls rely on health insurance programs such as Medicare and Medicaid moving away from traditional fee-for-service payment systems, which reward providers that generate high service volume, towards systems that encourage them to use resources efficiently while still providing high-quality services. The Act begins this shift by introducing provisions to make providers who are organized as Accountable Care Organizations (ACOs) eligible, from 2012 onwards, to share in any cost savings they achieve for the Medicare and Medicaid programs. The regulations pertaining to ACOs have not yet been determined: for example the threshold above which savings will be counted, and the share of those savings that will go to providers, are still uncertain. The overall concept, however, is clear. A set of providers,

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encompassing both primary and specialty care, will share in financial incentives to deliver cost-effective, high-quality services. A second relevant provision that will be piloted under the Act is payment bundling. Under these pilot arrangements Medicaid providers will receive bundled payments that pull together fees for the components of a particular episode of care: for example, combining the obstetrician's and the hospital's payments for a labor and birth episode into a single fee that is shared by the providers.[1] The goal here is to control costs by coordinating the number and cost of services provided per patient episode.

The health policy literature has noted that similar cost-control mechanisms are currently utilized by some health maintenance organizations (HMOs) for privately insured enrollees in California.[2] However, relatively little is known about physician responses to these incentive schemes. Previous papers document low costs in HMOs compared to other types of insurers but very few provide information on the mechanisms used to reduce costs. This paper describes an ongoing research project in which we investigate one potential cost-control mechanism. We use hospital discharge data from California to ask whether patients whose insurers give physicians incentives to control hospital costs are admitted to lower-priced hospitals than other patients. This particular mechanism is important for two reasons. The first is that hospital costs make up over 30% of national health spending (Kaiser Family Foundation, 2007 data). Second, if hospital admissions are affected by their prices this has implications for their incentives in, and hence the likely outcome from, their negotiations with HMOs. This in turn affects their incentives both to invest in costly new technologies and to engage in market-changing activities such as mergers since both are likely to impact these negotiations.

The analysis in this paper provides an initial comparison of the prices of hospitals chosen when insurers utilize physician payment schemes that generate cost-control incentives, to those of hospitals used when such incentives do not exist. Our results indicate that patients enrolled in insurers that make a high proportion of capitation payments to physicians are admitted to lower-priced hospitals than same-severity patients who are enrolled in low-capitation insurers. This evidence is consistent with the idea that physician groups receiving capitation payments from HMOs respond to incentives to reduce hospital costs by referring patients to low-priced hospitals.

We begin in Section 2 by describing the relevant features of the California health insurance market, comparing it to the provisions introduced more broadly by the 2010 health reforms. Section 3 briefly describes the relevant previous literature and Section 4 introduces our dataset. In Section 5 we present our preliminary analysis comparing the prices of hospitals utilized by patients whose insurers have different physician payment schemes. Section 6 discusses the potential implications of this descriptive analysis and outlines our agenda for future research.

[1] Medicare already bundles payments to hospitals through the D.R.G. (Diagnosis Related Group) payment system. Bundled payments to all payers for larger episodes of care have been piloted in the Medicare program, for example in the Medicare Participating Heart Bypass Center demonstration (see Cromwell et al 1998) and in the Acute Care Episode demonstration, currently being run, which expands the model to other types of discharges.

[2] See, for example, Hammelman et al (2009).

2 The Market: Why Should Price Affect Hospital Choices?

We consider the California medical care market, and in particular the market for privately insured women admitted to hospital for a labor / birth episode in 2003. We focus on women enrolled in health maintenance organizations (HMOs).[3] The process by which a patient chooses a hospital involves multiple players. The HMO contracts with a network of providers, including primary care physicians, specialists such as obstetricians, and hospitals. Enrollees in the HMO receive all services from within the HMO's network; unlike enrollees in other types of insurer, such as preferred provider organizations, there is no option to pay an out-of-pocket fee and receive treatment outside the network. Each patient chooses an obstetrician subject to this constraint. The obstetrician is affiliated with a small number of hospitals in the network and will refer the patient to one of those hospitals to give birth. The patient's choice of obstetrician is informed by the list of affiliated hospitals, which is public information. The HMO influences the obstetrician's choice by choosing the hospitals to include in the network, through direct financial incentives described below, and by making physicians' promotion on the pay scale contingent (formally or informally) on their management of costs. The choice of network hospitals is not very restrictive; for example Ho (2006) documents that on average over 80% of hospitals were included in each insurer's network in a sample of 43 large markets (including seven California markets) in 2003. The physician, rather than the insurer, is therefore arguably the primary decision-maker in the hospital referral. However, the financial and other incentives imposed by the HMO may have a substantial effect on physician referral choices. This is the issue analyzed in our ongoing research. We note that if the quality of hospital services varies directly with the hospital's price, the physician / insurer making the hospital referral faces a trade-off between price and the impact that quality might have on the doctor's own as well as the HMO's reputation and hence desirability. The reputational problem may be less severe for less serious conditions, such as normal pregnancy, where the quality of the hospital is less likely to affect outcomes.

The direct financial incentives utilized in California are based on capitation contracts under which physicians bear financial risk for services they provide and also for hospital services provided to their patients. Our analysis focuses on six of the seven largest HMOs in California, covering 57% of the HMO market in 2002. These HMOs contract on a non-exclusive basis with medical groups or independent practice associations (IPAs).[4] Both types of physician groups tend to be very large, covering 50,000 lives and containing between 200 and 300 physicians on average. Physicians in medical groups are either employees or partners of the group. IPAs are organized differently: they are administrative organizations that contract with independent physicians or clinics and sign network contracts with health plans on behalf of their physicians. They exist primarily to negotiate

[3] HMOs covered 21.4 million Californians, or about 63% of the population, as of December 2002. See Baumgarten (2004) for a detailed description of the California Health Care Market in 2002/3.

[4] We exclude Kaiser Permanente, the largest HMO with 30.5% of the market, because the prices paid by this vertically integrated insurer to its hospitals are not observed in our data. Kaiser uses a different model of physician organization from the other large HMOs: the HMO contracts exclusively with two particular medical groups, paying physicians a salary.

and manage capitation contracts for their member physicians. In our data 73% of payments made to these physician groups are capitation payments; the remainder are fee-for-service payments.

The extent of financial risk passed to the physician group varies across capitation contracts. Rosenthal et al (2001) and Robinson and Casalino (2001) provide detailed information about the types of contracts that existed in California in the period we consider. Just under 20% of capitation contracts were global capitation arrangements under which the payment made to the physician group covered all services required by their patients including hospital stays. The remainder were professional service capitation arrangements under which the payment covered the costs of services provided by the physicians within the group, and sometimes also the costs of ancillary services such as outpatient medical tests, but not the cost of inpatient hospital stays. Physician groups with global capitation contracts have a clear incentive to refer their patients to lower-cost hospitals: in essence they take on some of the role of the health insurer, bearing direct financial risk for services provided outside the group.[5] The link to inpatient hospital costs is less clear for groups paid through professional services capitation; however, 90% of these non-global capitation payment systems also include "shared risk arrangements" under which inpatient cost savings relative to a pre-agreed spending or utilization target are shared between the physician group and the HMO. All physician groups receiving capitation payments are therefore likely to have incentives to reduce the costs of inpatient stays. Fee-for-service payment schemes do not include shared risk arrangements or any similar incentives to control hospital costs.

If capitation arrangements are to influence hospital referral choices, however, cost-control incentives must be passed from the physician group to the individual physician. Table 1 summarizes results from Rosenthal et al (2002) which surveys a large number of physician groups in California regarding individual physician compensation schemes. Most physician groups (both IPAs and medical groups) utilize either capitation-based compensation, cost of care bonuses or profit sharing arrangements (or some combination of the three) for their member physicians. Grumbach et al (1998) report similar findings and note that IPAs that are paid on a fee-for-service basis make fee-for-service payments to their member physicians: that is, there is no disconnect between the payment arrangement between the HMO and the physician group and that passed on to individual physicians. Overall the implication for our project is that obstetricians in physician groups that receive capitation payments from HMOs have incentives to be affiliated with and refer patients to low-priced hospitals, whereas obstetricians in groups receiving fee-for-service payments usually have no such incentives. We now proceed to investigate their responses to these incentives.

We note that there are several dimensions on which the incentives generated by the California medical care system are similar to those introduced by the 2010 health care reforms. Capitation payments are similar in some respects to the payment bundling to be piloted in the Medicaid program. Both are intended to reduce the incentives, generated by fee-for-service payment systems, to provide more services than necessary. Both reward physicians for referring patients to lower-

[5] Following a wave of medical group bankruptcies in the 1990s, the state imposed strict financial reporting requirements and criteria for the financial health of these groups. Physician groups are required to hold a license, called a limited Knox Keene license, in order to accept global capitation payments.

priced hospitals. The difference is that bundled payments address these incentives within an episode of care while capitation payments address them both within and across episodes (presumably generating longer-term incentives). The Accountable Care Organizations set up by the reforms are also likely to generate incentives to control hospital costs by referring patients to low-priced hospitals. We therefore expect our analysis to be informative regarding the impact of the reforms on hospital inpatient costs. However we note that the physicians currently choosing to practice in groups receiving capitation payments represent a selected sample that is potentially pre-disposed towards responding to financial incentives. If true, and if there is no equilibrium change in the response function of agents as a result of the health care reforms, we would expect our results to represent an upper bound on the response of the universe of physicians.

3 Previous Literature

A number of health policy papers, some of which have already been cited, describe the financial arrangements health plans make with physicians, often based on survey data and often focused on California (see, for example, Rosenthal et al (2001 and 2002) and Grumbach et al (1998a. and b.). Glied (2000) summarizes previous papers assessing whether managed care plans reduce utilization and/or costs compared to other insurers. Her summary suggests that HMOs reduce inpatient admissions and costs, although interpreting the results of the studies is often difficult because, for example, physician and patient preferences over intensity of treatment may differ across types of insurer.[6] There are a few more recent studies that consider similar questions. For example, Cutler et al (2000) compare the treatment of heart disease in HMOs and traditional insurance plans in Massachusetts and find that HMOs have 30% to 40% lower expenditures. Virtually all the difference comes from lower unit prices rather than differences in actual treatments. Escarce et al (2001) studies an HMO in Michigan offering both an HMO and POS product and finds that the HMO, which requires referrals for specialty care, has lower physician and drug expenditures than the POS plan which does not. Gaynor, Rebitzer and Taylor (2004) look in more detail at how HMOs achieve cost savings. They analyze physician responses to group-based financial incentive contracts within a single HMO. They find that spending on medical utilization increases with the size of physician groups receiving group-based incentives. That is, spending is negatively correlated with the intensity of incentives to limit these expenditures. The correlation is greater for outpatient expenditures than for inpatient expenditures. However, there is little if any analysis of the mechanism used to reduce costs (for example, whether physicians move patients from high-cost to lower-cost providers).

[6] Gosden et al (1999) and Armour et al (2001) review the literature on the effects of financial incentives on physician behavior and come to similar conclusions.

4 The Data

We utilize hospital discharge data covering all discharges for privately insured HMO enrollees from hospitals in California in the year 2003. The source is the state's Office of Statewide Planning and Development (OSHPD). We link this to hospital characteristics data from the American Hospital Association 2003 and to the State of California Department of Managed Health Care Annual Financial Reporting Forms for 2003; these include financial and enrollment information for all HMOs in California. We do not observe details of the physician or physician group making the hospital referral, nor of the types of payments (global capitation, other capitation or fee-for-service) received by each physician group. We do, however, observe each patient's zip code, demographic characteristics, diagnosis, health insurer and the hospital chosen.[7] We also know the percent of each insurer's payments to primary physicians (including obstetricians) that are made on a capitated basis. The proportions vary substantially across the carriers in our sample, from 38% for Blue Cross to 97% for Pacificare. We therefore investigate physician responses to incentives by comparing patient admission patterns across insurers utilizing different payment arrangements. The initial analysis in this paper and the model estimated in our broader project investigate whether patients enrolled in high-capitation insurers are admitted to low-price hospitals, all else equal. The data limitations require us to make an assumption: that patients in high-capitation insurers have obstetricians whose physician groups receive a high proportion of capitated payments. If we also assume that any cross-insurer differences in price sensitivity, conditional on patient severity and location, are caused by differences in physician responses to financial incentives rather than by other unobserved differences across insurers, then we can draw conclusions about the physician referral behavior generated by different compensation schemes.

We do not observe the price charged to the insurer by the hospital. Instead our data includes the list price for every discharge at the level of the patient. As noted in Melnick (2004), this is similar to the "rack rate" published by hotels. Every year each hospital publishes a schedule of list prices for its services but few if any patients actually pay these list prices. Each insurance company has a contract with each provider in its network that defines a discount from the list price for its enrollees. We observe the average negotiated discount at the hospital level. The analysis in this paper uses two price measures. The first is the observed list price multiplied by 1 minus the average discount. The second takes account of expectational error caused by the fact that hospital admissions are made on the basis of ex ante expectations regarding prices, but we observe in the data their ex post realizations. We define groups of similar-severity patients based on characteristics known at the time of admission and use the average price for each group in place of the observed list price. Details of the definition of similar-severity patients are provided below.

Tables 2-4 provide summary statistics on the dataset. Table 2 sets out the sample size and proportion of capitated payments for the six insurers included in the analysis, ordered by decreasing percent capitated payments. The first column lists the number of labor discharges included in our

[7] We have a Private Use version of the data in which patient zip code, age, race and gender are not masked.

analysis for each carrier, ranging from 6,291 for Aetna to 25,038 for Blue Cross. Column 2 lists the percent of each HMO's primary services that are capitated.[8] There is considerable dispersion across insurers. Table 3 provides summary statistics on the discharges in the dataset. Our sample contains 88,157 patients and 195 hospitals. 27% of discharges are from teaching hospitals. The average price paid (approximated as observed list price*(1-average discount) is $4,317 for labor admissions. The average length of stay is 2.5 days. The table also records means for three potential measures of outcomes: death while in hospital, transfer to an acute care setting (at this hospital or a different hospital) and transfer to a special nursing facility (again at either this or a different hospital). These are useful inputs to our initial investigation of the patterns in the data. The average probability of each event is low: 0.01% for death, 0.3% for acute care transfer and 1.5% for transfer to a special nursing facility.

Table 4 investigates the variation in price and in outcomes across patient ages and severities. We use the Charlson score (Charlson et al, 1987) as our severity measure here; it will be an input to our overall measure of patient severity later in the analysis. This score assigns integer-valued weights (from 0 to 6) to comorbidities other than principal diagnosis where higher weights indicate higher severity. The weights are summed to generate a single integer-valued index. For example, patients with comorbidities indicating that they have diabetes or mild liver disease would receive a Charlson score of 1; those with renal disease or any malignancy would have a Charlson score of 2; those with a metastatic solid tumor or AIDS would have a Charlson score of 6. A patient with both diabetes and renal disease would have a score of 3. The index was developed by physicians and is widely used to measure severity based on diagnoses listed in patient records. Table 4 indicates an intuitive relationship between both age and expected severity (as measured by the Charlson score) and outcomes. Women giving birth who are aged over 40 have a significantly higher price, significantly higher probability of acute care transfer and also a slightly higher probability of transfer to a special nursing facility although the latter is not significant at $p=0.05$. Women with higher Charlson scores in our data have significantly higher prices and significantly higher probabilities of adverse outcomes than women with lower Charlson scores.

5 A Comparison of Hospitals Utilized Across Insurers

We now provide an initial analysis of the importance of price in the hospital choice process, and the impact of insurer capitation payments on this relationship. We begin by categorizing patients by severity. Our definitions are based on advice from obstetrical experts at Columbia Presbyterian Hospital. We define two women as having the same severity if they are in the same age group and have the same principal diagnosis, the same Charlson score and the same diagnosis generating

[8] Capitation payments for primary professional services are defined in the HMO Annual Financial Statements as "capitation costs incurred by the reporting entity to primary care physicians, dentists and other professionals for the delivery of medical services". They include capitation payments to obstetricians. The statements also record capitation payments to other medical professional services, including support personnel such as nurses, ambulance drivers and technicians.

the Charlson score.[9] There are 254 populated severity groups in the data (across all insurers). Both principal diagnosis and Charlson score are based only on diagnoses known on admission. We define the expected list price as the average list price for the particular hospital over patients with the same severity by this definition. Our second (preferred) measure of price is this expected list price multiplied by 1 - the average discount at the hospital level (our first is the observed list price normalized by the same discount measure).

If physicians respond to the financial incentives generated by capitation payments, we expect patients enrolled in higher-capitation insurers to be admitted to lower-priced hospitals, all else equal. We investigate this possibility by comparing the ratio of the observed price to the average for same-severity patients for insurers with different capitation levels within a market. First we generate the price ratio measure: for each price variable this is defined as

$$p_i^{ratio} = \frac{p_i}{\bar{p}_{s_i}}$$

where p_i is the relevant price variable for patient i and $\bar{p}_{s_i}$ is the average of that variable for same-severity patients across all hospitals in the sample.[10] A price ratio greater than 1 implies that the patient is admitted to a hospital with a relatively high price compared to other same-severity patients in the sample; a ratio less than 1 implies that the patient's hospital has a relatively low price given her severity. We then estimate the following regression equation:

$$p_i^{ratio} = \alpha.percent\_cap_{\pi_i} + \gamma x_{m_i} + \varepsilon_i$$

where π_i is patient i' HMO and m_i is her market. That is, we regress the price ratio measure on the proportion of the insurer's payments to primary physicians that are capitated and market fixed effects.[11] If insurer capitation payments influence physician referrals, we expect α to be negative. The results for our two price measures are reported in Table 5. The coefficient on the capitation variable is negative and statistically significant in both regressions. It is more significant for the measure based on the expected list price, and the R^2 is substantially higher, as we would expect since admissions are based on an ex ante expected notion of price. To interpret the magnitudes of the coefficients, consider the implications of the regressions for two insurers whose capitation rates differ by 40 percentage points (approximately the difference between Blue Shield and Pacificare in our data). The regression based on expected list price predicts that, if insurer 1 has the higher capitation rate, its hospital prices relative to the average for same-severity patients will be 2% lower

[9] The age categories we use are 11-19, 20-39, 40-49 and 50-64. There are 21 principal diagnosis categories including, for example, "normal delivery", "previous Cesarean Section", and "early labor". Similar results are generated using other age categories such as 11-19, 20-29, 30-39 and 40-64.

[10] Using an average for same-severity patients across hospitals within the same market generated very similar results.

[11] Markets are defined as Super-Health Service Areas (SuperHSAs). There are ten of these in California. They are based on the fourteen Health Service Area (HSA) regions used by the state for health planning; we follow Baumgarten (2004) by combining three regions in the Bay Area into a single market and the Los Angeles and Orange HSAs into another single market.

than those for insurer 2 on average.[12]

6 Discussion and Conclusion

These initial analyses provide suggestive evidence on the differences in referral decisions across obstetricians facing different financial incentives. They are consistent with the idea that obstetricians who benefit from controlling inpatient costs may refer patients to lower-priced hospitals than their peers. They also raise interesting questions regarding the trade-offs made between hospital characteristics and price. Table 6 reports the results of a regression where the dependent variable is the price ratio measure defined using our preferred price variable (expected list price normalized by the discount) and the independent variables are hospital characteristics and market fixed effects. For profit hospitals and hospitals offering transplant services have significantly higher prices, relative to the average for same-severity patients, than other hospitals. Hospitals with more nurses per bed have higher prices, as do larger hospitals. Labor services have a negative coefficient in the price regression. The data are consistent with the hypothesis that prices may be influenced by hospital reputation, perhaps related broadly to quality but not specifically to labor services. It is possible that obstetricians contracting with high-capitation insurers direct labor / birth patients away from hospitals that have high prices on average but towards those whose characteristics are specifically related to quality of care for their diagnoses. We are pursuing this idea in more detail in our ongoing research.

However, simple regressions like those conducted here cannot provide more than suggestive evidence. We have not modelled the trade-offs made between price and other hospital characteristics in the hospital choice equation or allowed these trade-offs to differ across different types of patient. We have also made no attempt to account for factors, such as the distance between the patient's home and the hospital, that are known to be important determinants of hospital choice (see for example Ho (2006)). Finally we have not modeled the variation in discounts across insurers for each hospital; this too may confound our descriptive analyses. A full hospital choice model is needed to understand these issues.

We are developing a series of models to estimate the preferences of the composite agent, including patient, insurer and referring physician, that makes the hospital choice. The utility equation we estimate can be thought of as a weighted sum of the patient's and providers' utilities, where the price term enters only through physician and insurer preferences; the patient does not observe this price and therefore is assumed not to react to it. The other variables included are the distance from the patient's home (which is likely to affect only the patient's preferences) and hospital quality measures (which probably affect the utility of all three agents). It is important to include extensive controls for hospital quality to avoid price endogeneity problems. For example, quality may vary across diagnoses within a hospital, so interactions are needed between hospital characteristics and patient diagnoses or other measures of severity. Alternatively, the insurer / physician's preference

[12] This magnitude is obtained by multiplying the difference in average capitation rates in percentage points by the estimated coefficient.

for quality of the hospital may depend on the severity of the patient's illness. We begin by estimating multinomial logit demand models, following the previous literature on hospital choice but acknowledging that this methodology can only imperfectly address the price endogeneity issues faced in our application. We then develop a new methodology based on moment inequalities to address these issues. The methodology requires only weak assumptions on the form of the interactions between hospital quality and patient severity and only non-parametric assumptions on the unobservables. Our preliminary results suggest that, when the appropriate steps to model the hospital choice process have been taken, physician capitation payments are found to have a significant effect on referral decisions.

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Table 1: Compensation Schemes and Bonuses/Withholds from Primary Care Physicians in California

Method	Medical Groups	Independent Practice Assocns
Capitation-based compensation	21%	87%
Salary	41%	0%
Fee-for-service	39%	13%
	All Physician Groups	
Cost of care bonuses	17%	
Profit sharing	48%	

Notes: All data in the table is reported in Rosenthal et al (2002). The authors surveyed physician organizations covering approximately 87% of all Californians enrolled in managed care plans (excluding Kaiser). Data were collected through structured interviews between May 1999 and June 2000. The paper does not provide a breakdown of the data on cost of care bonuses and profit sharing by type of physician organization.

Table 2: Summary Statistics by Insurer

	Labor discharges	% Primary Capitation
Pacificare	15,479	0.97
Aetna	6,291	0.91
Health Net	16,950	0.80
Cigna	8,097	0.75
Blue Shield	16,302	0.57
Blue Cross	25,038	0.38

Notes: Data on the six insurers included in our analysis. "Labor discharges" is the number of discharges in the data sample used in our analyses. "% Primary Capitation" is the percent of payments to primary providers made on a capitated basis in 2003 (source: State of California Department of Managed Health Care Annual Financial Reporting Forms, 2003).

Table 3: Summary Statistics by Discharge

	Mean	Std. Devn.
Number of patients	88,157	
Number of hospitals	195	
Number of insurers	6	
Teaching hospital	0.27	
List price	$13,312	$13,213
List price*(1-discount)	$4,317	$4,596
Length of stay	2.54	2.39
Died	0.01%	0.004%
Acute transfer	0.3%	0.02%
Special Nursing Transfer	1.5%	0.04%

Notes: Summary statistics for dataset comprising private enrollees of the six largest HMOs excluding Kaiser who are admitted for labor-related diagnoses. "Died" is the probability of death while in hospital, "Acute Transfer" the probability of transfer to an acute care setting (in this or a different hospital) and "Special Nursing Transfer" the probability of transfer to a special nursing facility (again at this or a different hospital). "Std Devn" for "Died", "Acute transfer" and "Special Nursing Transfer" are calculated under the assumption that the 0/1 variable is binomially distributed.

Table 4: Prices and Outcomes by Patient Type

		N	Price*(1-disc)	Acute Transfer	Special Nursing
Age					
	<40	84130	4269 (4488)	0.3% (0.0%)	1.49% (0.0%)
	>40	4027	5310 (6373)	0.5% (0.1%)	1.54% (0.2%)
Signif diff			0.000	0.009	0.797
Charlson					
	0	86326	4276 (4501)	0.3% (0.0%)	1.5% (0.0%)
	1	1753	6079 (7060)	0.6% (0.2%)	2.3% (0.4%)
	>1	78	10022 (15186)	5.1% (2.5%)	12.8% (3.8%)
p value (0 to 1)			0.000	0.005	0.003
p value (1 to >1)			0.000	0.000	0.000

Notes: Means of outcome variables for different patient types. See notes to Table 2 for variable definitions. Standard deviations in parentheses; for Acute Transfer and Special Nursing these are standard errors calculated assuming that the 0/1 variables are binomially distributed. Charlson scores assign weights to comorbidities (known on admission to hospital) other than principal diagnosis where higher weight indicates higher severity. Value 0-6 are observed in the data. "p value" states the probability of obtaining a test statistic at least as extreme as the one that was observed, under the null hypothesis that the means in the two samples are the same; these are the results of a t-test for price*(1-discount) and a z-test assuming two binomial distributions for Acute Transfer and Special Nursing.

Table 5: Relation between Insurer Capitation and Hospital Prices

	List price ($000)*(1-$\delta_h$)		E(list price)*(1-δ_h)	
	Coefficient	S.E.	Coefficient	S.E.
Percent capitation	-0.051**	0.014	-0.052**	0.007
Market fixed effects				
Napa-Sonoma	0.123**	0.024	0.133**	0.012
Bay Area	0.234**	0.013	0.232**	0.007
Sierra Nevada	-0.073**	0.020	-0.077**	0.010
Central Coast	-0.429**	0.031	-0.431**	0.016
Central Valley	-0.351**	0.019	-0.353**	0.010
Santa Barbara	-0.579**	0.019	-0.583**	0.010
Los Angeles-Orange	-0.413**	0.012	-0.416**	0.006
Inland Empire	-0.418**	0.015	-0.421**	0.008
San Diego	-0.384**	0.014	-0.382**	0.007
Constant	1.282**	0.014	1.284**	0.007
N	88,157		88,157	
R^2	0.075		0.245	

Notes: We report the results of two regression analyses. We calculate a price ratio measure as $p_i^{ratio} = \frac{p_i}{\bar{p}_{s_i}}$, the ratio of the observed price for a particular patient to the average price for same-severity patients across the sample. Severity is defined by principal diagnosis, age, Charlson score and diagnosis inputs into the Charlson score. In column 1 "price" is the observed list price multiplied by 1 minus the observed average discount at the hospital level. In column 2 we use a measure of expected list price, the average list price across same-severity patients, in place of the observed value. We regress the price ratio measure on the insurer's percent of payments to primary physicians that are capitated and market fixed effects. Markets are Super-Health Service Areas as defined in Baumgarten (2004); the omitted market is North/Sacramento. Standard errors are reported in columns 2 and 4; ** indicates significance at p=0.05.

Table 6: Regression of Price Ratio on Hospital Characteristics

	Coefficient	S.E.
Number of beds	0.001**	0.000
RNs per bed	0.517**	0.036
Teaching hospital	-0.026**	0.004
For profit hospital	0.076**	0.004
Offers transplants	0.155**	0.005
Neonatal IC Unit	0.038**	0.007
Labor services	-0.122**	0.008
Constant	0.952**	0.008
Market fixed effects	Yes	
N	88,157	
R^2	0.346	

Notes: Regression of the price ratio measure $p_i^{ratio} = \frac{p_i}{\bar{p}_{s_i}}$, defined as for Table 5 (where p_i is the expected list price*(1-average hospital discount)), on hospital characteristics. Standard errors are reported in column 2; ** indicates significance at p=0.05 and * significance at p=0.10.