NBER WORKING PAPER SERIES

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Working Paper 21513
http://www.nber.org/papers/w21513

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
August 2015

We would like to thank participants at the Stanford GSB Economics Brown Bag seminar series for helpful comments. Jack Boeglin and Tarun Narasimhan provided excellent research assistance. All errors are our own. The views expressed herein are those of the authors and do not necessarily reflect the views of the National Bureau of Economic Research.

At least one co-author has disclosed a financial relationship of potential relevance for this research. Further information is available online at http://www.nber.org/papers/w21513.ack

Does Health Plan Generosity Enhance Hospital Market Power?
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NBER Working Paper No. 21513
August 2015
JEL No. I11

ABSTRACT

We test whether the generosity of employer-sponsored health insurance facilitates the exercise of market power by hospitals. We construct indices of health plan generosity and the price and volume of hospital services using data from Truven MarketScan for 601 counties from 2001-2007. We use variation in the industry and union status of covered workers within a county over time to identify the causal effects of generosity. Although OLS estimates fail to reject the hypothesis that generosity facilitates the exercise of hospital market power, IV estimates show a statistically significant and economically important positive effect of plan generosity on hospital prices in uncompetitive markets, but not in competitive markets. Our results suggest that most of the aggregate effect of hospital market structure on prices found in previous work may be coming from areas with generous plans.

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Introduction

Since Feldstein's (1970, 1973) early work, economists have been interested in whether the generosity of health insurance coverage facilitates the exercise of market power by sellers of health services. Even a perfectly competitive insurance market makes demand for insured services more inelastic by protecting consumers against financial risk. More inelastic demand for services, in turn, makes it profitable for providers to charge higher prices. The problem is that each individual consumer creates a negative externality for the others when deciding how generous of a health plan to buy because he does not account for the market-wide consequences of his actions on prices.

Assessing the importance of this phenomenon is essential to current health policy debates. Market power is endemic in the health sector. Health services are highly complex, differentiated products that are not priced transparently. Because many providers are non-profit organizations, courts have been reluctant to attack their anticompetitive practices on the grounds that they would use their pricing power to subsidize charitable activities (Havighurst and Richman 2011). The Affordable Care Act has given hospitals and physicians powerful incentives to integrate into Accountable Care Organizations, which has the potential to further enhance their bargaining power with private purchasers (Rosch 2011; Bacher et al. 2013). Taken together, these factors have given some providers, especially large hospital systems, growing power to obtain steep payment increases (Ginsburg 2010; Berenson et al. 2012). This has wide-ranging effects beyond allocative inefficiency, including redistribution from patients to providers, reduced insurance coverage rates, and public-sector deficits.

Yet, despite this, few studies have investigated whether the generosity of insurance leads to higher prices in imperfectly competitive markets for insured services. Although some work has estimated the effect of the design of public-program and prescription drug insurance on drug prices, the results from these papers do not necessarily apply beyond the particular settings that they examine. In addition, none of these papers investigate the extent to which coverage generosity interacts with market power in markets for hospital services, even though lack of hospital competition poses a serious public policy problem (Berenson, Ginsburg, and Kemper 2010).

One reason for the lack of work on this topic may be the difficulty in identifying a causal effect. Although more health insurance may enable providers to raise prices, higher prices may also affect the scope of insurance. The direction of the reverse causality is theoretically indeterminate. On one hand, high prices might increase the demand for insurance; high prices mean that there is more financial risk associated with a given episode of illness. On the other hand, high prices might reduce the demand for insurance; high prices increase the cost of moral hazard. Because low deductibles and coinsurance rates increase demand for health services that have relatively low value (Newhouse and the Insurance Experiment Group 1993), the deadweight loss from more generous insurance increases as prices rise.

We use an insight from Freeman (1981), updated by Buchmueller, DiNardo, and Valetta (2002), to solve this endogeneity problem. Numerous studies have shown that industry and union membership affects the generosity of fringe benefits, including health insurance. We use unique information on the union status and industry composition of insured workers in 601 large US counties to identify the effect of health plan generosity on hospital prices and utilization. We measure generosity by the average actuarial value of coverage for a sample of large self-insured employers. We estimate the effect of county-average actuarial value on the price and quantity of hospital services by instrumental variables, holding constant county- and time-fixed effects and time-varying county characteristics, using county-level unionization and industry structure as instruments.

I. Previous Literature

In an early empirical paper, Feldstein (1970) estimated the relationship between insurance coverage rates and physician prices based on national aggregate data from 1948-1966. He finds that physician prices rise, holding constant other factors, when the fraction of households with private insurance coverage rises. He later went on to demonstrate that reductions in insurance coverage might actually be welfare-improving because the utility loss from decreased risk protection might be more than counterbalanced by gains from lower prices and the reduced purchase of excess care (Feldstein 1973). Chiu (1997) and Vaithianathan (2006) extend this work to show more generally the conditions under Feldstein's result obtains.

A more recent wave of empirical work examines how reimbursement for prescription drug spending affects drug prices. Pavcnik (2002) evaluates a policy experiment from Germany, in which the country's statutory health insurance scheme implemented reference pricing, and finds that making consumers responsible for the full cost of drugs above the reference price rather than just a flat fee led to price reductions of 10 to 26 percent. Duggan and Scott-Morton (2006, 2010, 2011) estimate the effect of

Medicaid and Medicare Part D market shares on drug prices, and show that prices are significantly positively related to Medicaid's market share, but negatively related to Medicare' Part D's market share. As they point out, the effect of Medicaid market share is consistent with Feldstein's (1970) classical result.

Although these studies have provided many important insights, they focus on how particular changes to benefit design affects drug prices. As the differences between the findings of the Duggan and Scott-Morton papers show, each paper's result depends at least in part on the specific features of the insurance programs that they analyze. As such, they do not directly identify the more general effect of plan generosity on health spending: as private health plans become more generous, does the market respond as it did to an expansion of Medicaid, or to the rollout of Medicare Part D? In addition, none of the studies examine how generosity interacts with the competitiveness of markets for hospital services, which are characterized by a very different price determination process and represent a far larger share of overall health spending than drugs.

Our paper is most closely related to new work by Robinson and Brown (2013) and Gowrisankaran, Nevo, and Town (2013). Robinson and Brown (2013) evaluate the impact of reference pricing on the use of and prices paid for knee and hip replacement surgery. They compare enrollees in the California Public Employees Retirement System (CalPERS) to enrollees in another plan that did not use reference pricing, after versus before the change in CalPERS benefit design. They found that CalPERS members shifted from high- to low-priced facilities, and that the prices paid by CalPERS members at high-price facilities declined significantly, relative to enrollees in the comparison group. Gowrisankaran, Nevo, and Town (2013) specify and estimate a model of the bargaining process between hospitals and managed care organizations (MCOs). Based on data from four large MCOs in Northern Virginia from 2003-06, they find that MCOs' bargaining leverage partially (but not completely) offsets the price insensitivity induced by insurance. These papers suggest that consumer cost-sharing is likely to affect prices in a broader setting, the question to which we now turn.

II. Data and Variable Construction

We use data from Truven MarketScan from approximately 2.1 million hospital claims from privately-insured individuals enrolled in a fee-for-service self-insured employer-sponsored health plan between 2001 and 2007. These claims include only the facility portion of the payment; professional claims (i.e., from physicians) are excluded from our analysis. Though (as we discuss below) these data are not representative of the entire U.S. population, the areas they span are sufficient to characterize patterns of hospital price variation.

For each claim, we analyze what is commonly referred to as the allowed amount—the amount that the plan allows the hospital to be paid for the service, after the application of contractual discount provisions and other plan rules, but before adjustment for patient copayments or deductibles. The hospital may receive part of this amount from the insurance plan and part of it from the patient in the form of copayments or deductibles. The allowed amounts are not charges or a function of charges. Instead, they are the actual transaction payments under contracts with health plans, including payments made by both the patient and the insurer.

We use the allowed amounts to calculate indices for county i during year t of the price per hospital admission (P_{it}), the number of hospital admissions per enrollee (Q_{it}), the coefficient of variation of hospital prices (CV_{it}), and the actuarial value of inpatient coverage (AV_{it}), where AV_{it} is the share of inpatient spending paid by plans (rather than enrollees or other sources) on behalf of enrollees who live in county i during year t. AV_{it} therefore captures the average generosity of health plans among enrollees in a county/year, weighted by enrollees' spending. We normalize P_{it}, Q_{it}, and CV_{it} by the national average value of each variable in each year so that each has mean 1 in every year by construction; this enables us to interpret percentage-point changes in the value of each variable as percent changes. Thus, if N_{it} is the number of hospital admissions in MarketScan for county i during year t, P_{it} is defined as follows:

$$P_{it} = \frac{1}{\overline{P}_t} \sum_{\substack{hospital \\ admissions\, l}} allowed\ amount_{lit} / N_{it}, \qquad \overline{P}_t = \sum_{\substack{hospital \\ admissions\, l, \\ counties\, i}} allowed\ amount_{lit} / N_t,$$

with Q_{it} and CV_{it} defined similarly.

We use MarketScan information on enrollees to calculate county/year level characteristics of the insured population, including its age and gender distribution (with indicator variables for the proportion of the population aged 18-34, 35-44, 45-54, and 55-64, omitted group age < 18), industry of employment (with indicator variables for mining/manufacturing/transportation/services and retail trade(omitted group all other industries), and proportion union.[1]

[1] Ideally, we would obtain data on proportion union from an external source that captured the unionization rate in the county as a whole. However, the only annual source of union coverage data of which we are aware is the Current Population Survey, and the public-use version of it does not contain county identifiers. We used the MarketScan unionization data because we could not merge county-level restricted-use Current Population Survey to our (county-level restricted-use) MarketScan data.

We follow the approach in Kessler and McClellan (2000) to measure the county/year density of several hospital characteristics, including ownership status (for-profit or nonprofit, omitted group public), size (<100 beds, >300 beds, omitted group 100-300 beds), system membership, teaching status, and bed capacity. We define the density of each hospital characteristic H in county i at year t, Z_{it}^H, as

$$Z_{it}^{H} = \sum_{\substack{j\ serving \\ county\ i}} c_{ijt} \times \sum_{\substack{k\ admitting \\ to\ j}} b_{kjt} \times \sum_{\substack{j\ serving \\ zipcode\ k}} a_{jkt} AHA_{jt}^{H} ,$$

where j and k index hospitals and zip codes, respectively; a_{jkt} is the share of elderly Medicare patients who live in zip k admitted to hospital j; b_{kjt} is the share of patients admitted to hospital j who live in zip k; c_{ijt} is the share of patients who live in county i admitted to hospital j ; and AHA_{jt}^H is an indicator variable that is equal to 1 if the hospital has characteristic H according to the American Hospital Association (AHA) survey.[2] The b_{kjt}-weighting in Z_{it}^H assumes that the characteristics of hospital j's market depends on the weighted average of all of the zip-code patient residence areas that it serves; the c_{ijt}-weighting defines a county's characteristics as the weighted average of all of the hospitals that serve patients who live in county i. To measure the extent of hospital market power, we construct a Hirschman-Herfindahl index (HHI) of hospital services analogous to Z_{it}^H, where

$$HHI_{it} = \sum_{\substack{j\ serving \\ county\ i}} c_{ijt} \times \sum_{\substack{k\ admitting \\ to\ j}} b_{kjt} \times \sum_{\substack{j\ serving \\ zipcode\ k}} a_{jkt}^{2} .$$ [3]

Finally, to obtain information on other time-varying characteristics of counties, we use the Area Resource File (ARF, for population, the number of Medicare

[2] a_{jkt} , b_{kjt} , and c_{ijt} are derived from 100% MEDPAR inpatient claims files, matched with fee-for-service Medicare enrollment files.
[3] In what follows, we treat market structure (Z_{it}^H and HHI_{it}) as exogenous.

beneficiaries, the number of physicians, and median household income) and the Medicare hospital wage index (to measure hospitals' labor costs).

III. Models

Our basic model specifies P_{it}, Q_{it}, and CV_{it} as a function of county- and year-fixed-effects; whether the average health plan's actuarial value is above the median or 75th percentile level, $HIAV_{it}$; whether the hospital market is uncompetitive (above the median HHI, $HIHHI_{it}$); hospital market characteristics Z_{it} ; and other time-varying county characteristics X_{it}:

$$\begin{matrix} P_{it} \\ Q_{it} \\ CV_{it} \end{matrix} \quad = \quad \alpha_i + \theta_t + \beta HIAV_{it} + \lambda HIHHI_{it} + Z_{it}\gamma + X_{it}\delta + \varepsilon_{it}. \qquad (1)$$

We also estimate a model with an interaction between $HIAV_{it}$ and $HIHHI_{it}$ to test whether the effect of plan generosity varies in different market environments:

$$\begin{matrix} P_{it} \\ Q_{it} \\ CV_{it} \end{matrix} \quad = \quad \alpha_i + \theta_t + \beta HIAV_{it} + \lambda HIHHI_{it} + \pi(HIAV_{it} * HIHHI_{it}) + Z_{it}\gamma + X_{it}\delta + \varepsilon_{it}. \qquad (2)$$

We begin by estimating (1) and (2) by ordinary least squares. However, as discussed above, although $HIAV_{it}$ might affect P_{it}, P_{it} might also affect $HIAV_{it}$. For this reason, we estimate (1) and (2) by instrumental variables (IV). We identify the interaction effect in (2) using the method in Wooldridge (2010, Section 21.4): we regress $HIAV_{it}$ on the (excluded) unionization rate and industry dummies and all the other exogenous variables; construct the fitted value of $HIAV_{it}$, $HIAVHAT_{it}$; and then estimate (2) using $HIAVHAT_{it}$ and $HIAVHAT_{it}$*$HIHHI_{it}$ as instruments.

As we discuss below, the consistency of our IV estimates depends on the assumption that the unionization rate and industry structure only affect prices through plan generosity. Because the most likely channel through which these variables affect hospital prices directly is labor costs, we control explicitly for this factor in X_{it} with the Medicare hospital wage index. The Medicare wage index is based on hospitals' Medicare cost reports, Occupational Mix surveys, payroll records, contracts, and other wage-related documentation. In computing the wage index, CMS derives an occupation-weighted average hourly wage for each county in each year (total wage costs divided by total hours for all hospitals in the area) and a national average hourly wage in each year (total wage costs divided by total hours for all hospitals in the nation). A county's Medicare wage index value is the ratio of the area's average hourly wage to the national average hourly wage.

IV. Results

Table I presents descriptive statistics on the variables used in our analysis.[4] The left panel of the table presents means and standard deviations for the control variables that we derived from Medicare, the AHA survey, and the ARF. The first column presents means and standard deviations for the control variables for all US counties that had the controls present in every year 2001-07. As the table shows, we had control variables present for 2,454 of the approximately 3,100 US counties with residential population, covering around 94 percent of the US population (= 279.7 million / 296.4 million total population in 2005 [not in any table]). The second column presents means and standard

[4] We also present the average price per admission, number of admissions per enrollee, and coefficient of variation of prices for background although we do not use any of these variables directly.

deviations for the subset of counties that had at least 100 MarketScan enrollees and a price index between 0.5 and 2 in every year 2001-07. Although we only have valid MarketScan data for around a quarter of the counties with control variables (601 / 2,454), these counties cover around 60 percent of the US population. Comparing the first and the second columns shows that analysis counties are representative of the country as a whole in most (although not all) dimensions. Analysis counties are larger on average (300,000 population as compared to 114,000), and more likely to be in midwestern and southern states, but with mostly similar health care market characteristics. Although analysis counties have a higher density of for-profit hospitals and have slightly fewer beds (reflecting their regional distribution), they are similar in terms of the number of Medicare beneficiaries, physicians per capita, wage index, and market competitiveness (HHI for the analysis counties is 0.455 [standard deviation 0.122], compared to 0.481 [standard deviation 0.151] for the US as a whole [not in any table]).

The right panel of the table presents means and standard deviations for variables derived from MarketScan. The indices of price, quantity, and price variation are all 1 by construction. The remainder of the right panel shows how the MarketScan sample resembles and differs from the US workforce as a whole. Our computed actuarial value is virtually identical to that computed by the Actuarial Research Corporation for self-insured employers based on the 2005 National Compensation Survey (0.867 versus 0.886; Yi and Mays 2009). The unionization rate is slightly higher than the 2005 coverage rate in the CPS for private nonagricultural employees (0.104 versus 0.085; Hirsch and Macpherson 2005), reflecting the relatively greater unionization of (large) self-insured employers. The industry distribution does not reflect the US workforce; it is much more highly concentrated in manufacturing, reflecting the composition of firms that report to MarketScan.

Tables II and III show the source of identification of our IV models: differences in trends in unionization and industry of employment across counties. Table II presents trends in health plan generosity, hospital market, and demographic characteristics in counties that have declining versus rising or stable unionization and manufacturing employment from 2001-07. The generosity of health plans in our sample was rising over this period, reflecting national trends that have been documented elsewhere (e.g., Gabel et al. 2006). The first column shows that the proportion of MarketScan enrollees in the 235 counties with declining unionization (accounting for around two-thirds of MarketScan enrollees in 2007) who had generous health plans rose by 17.2 percentage points. By comparison, the proportion of enrollees in the 366 counties with rising or stable unionization who had generous plans rose by 40.8 percentage points. Similarly, the proportion of enrollees in the 354 counties with declining manufacturing employment (accounting for around 84 percent of MarketScan enrollees) who had generous plans rose by 23.9 percentage points, but the proportion in the 247 counties with rising or stable manufacturing employment rose by 48.5 percentage points.

Trends in hospital market and demographic characteristics are much more similar across type of counties than trends in health plan generosity. For example, in counties with declining unionization, hospital market concentration (as measured by the HHI) declined by 0.062; in counties with rising or stable unionization, market concentration declined by 0.045. Although counties with declining unionization were significantly larger, trends in population growth in these places were within one percent of the trend in places with rising or stable unionization.

Table III presents the coefficients on the excluded instruments from the first-stage regressions of plan generosity. Counties with rising rates of unionization show differential growth in the proportion of enrollees with generous plans, as do counties with rising rates of employment in manufacturing and services, and falling rates of employment in retail trade. Although the F-statistic on the excluded instruments in the 75^{th} percentile model is above the rule-of-thumb threshold of 10 suggested by Staiger and Stock (1997), we investigate the possibility of bias due to the weakness of our instruments in greater detail below.

Table IV presents estimates from models (1) and (2) of the effect of health plan actuarial value and hospital market competitiveness on hospital prices. The left panel of the table presents OLS estimates; the right panel presents IV estimates. The first two columns of each panel present estimates that define a county/year as high-actuarial-value if the enrollment-weighted value is above the median (= 0.894); the second columns define a county/year as high-actuarial-value if the value is above the 75th percentile (= 0.918).

According to the OLS estimates, more generous health plans lead to *lower* hospital prices, holding constant county- and year-fixed effects and other time-varying characteristics of health care markets; there is no significant interaction between actuarial value and hospital market competitiveness. The IV estimates, in contrast, show no significant average effect of actuarial value on prices but significantly *higher* prices in uncompetitive markets. Counties with high plan generosity and uncompetitive hospital markets have approximately 8.6 percent higher prices than counties with low generosity and competitive markets (model (2c)). The total effect of hospital market competitiveness on prices in counties with generous plans -- the sum of the direct effect of hospital HHI and the interaction between actuarial value and HHI -- is 6.2 percent (standard error 2.6 percent). By comparison, the effect of hospital market competitiveness on prices in counties with low plan generosity is small and statistically insignificant. The magnitudes of the interaction and total effects grow as the cutoff for high generosity is raised from the median to the 75th percentile. Counties with above-75th-percentile plan generosity and uncompetitive hospital markets have approximately 10.6 percent higher prices than counties with below-75th-percentile generosity and competitive markets, with a total effect of hospital market competitiveness conditional on high plan generosity of 9.7 percent (standard error 4.2 percent, model (2d)).

Table V presents estimates from models (1) and (2) of the effect of health plan actuarial value and hospital market competitiveness on hospital admissions per enrollee. The interaction between actuarial value and market competitiveness is small and generally statistically insignificant in both OLS and IV models. In contrast, the average effect of plan generosity is large and statistically significant, especially in the IV models. Using the median as the cutoff for high plan generosity, counties with high generosity have approximately 18.9 percent more hospital admissions than counties with low generosity (standard error 6.5 percent, model (1c)); when the cutoff for high generosity is raised to the 75th percentile, the effect remains approximately the same.

Table VI presents estimates from models (1) and (2) of the effect of health plan actuarial value and hospital market competitiveness on the coefficient of variation of hospital prices. The OLS estimates suggest that, on average, more generous health plans lead to more variation in hospital prices; the confidence intervals around the IV estimates of the average effect are so large that they include both the OLS estimates and the null hypothesis. However, counties with above-75th-percentile plan generosity and uncompetitive hospital markets have approximately 18.8 percent more price variation than counties with below-75th-percentile generosity and competitive markets (standard error 8.2 percent, model (2d)).

We estimate additional models to investigate the validity of our results. First, we reestimate model (2) with limited-information maximum likelihood (LIML) to explore the extent of bias due to the weakness of our instruments, because LIML is less subject to bias than two-stage least squares (2SLS) in the presence of overidentifying restrictions (Angrist and Pischke 2009). We find that LIML and 2SLS estimates are virtually identical, and that formal tests of instrument strength reject the null hypothesis of weak instruments.[5] We also calculate Hansen's J statistic to test the validity of the LIML model's overidentifying restrictions, which fails to reject the null hypothesis that the instruments are invalid.[6] Second, we reestimate model (2) using the 75th percentile HHI as the threshold above which a hospital market is anticompetitive. We find that the estimates in this model are very similar to those reported in Table IV.[7]

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[5] For example, the LIML estimate of π corresponding to Table IV, model (2c) is 0.087 (standard error 0.034). In this model, the Kleibergen-Paap (2006) rk Wald statistic is 4.74, which just exceeds the Stock-Yogo (2005) 5% critical value of 4.72 for bias equal to 10% of the estimator's size.

[6] $\chi_{(2)}^2 = 3.89$, $p = 0.143$.

[7] For example, the estimate of π corresponding to Table IV, model (2c) is 0.090 (standard error 0.047), with an estimate of $\pi + \lambda$ of 0.061 (standard error 0.030).

V. Conclusion

To what extent does the generosity of health insurance coverage facilitate the exercise of market power by producers of health services? Although it has been more than 40 years since Feldstein first provided suggestive macroeconomic evidence that it does, and identifying the importance of this effect has become crucial to health policy, only a handful of papers have sought to investigate this hypothesis. To date, the literature has mostly focused on the effects of public insurance program and prescription drug benefit design. This work has contributed much to our understanding of the problem, but the more general question remains unanswered.

This paper seeks to fill this gap. It analyzes data from Truven MarketScan from approximately 2.1 million hospital claims from privately-insured individuals enrolled in a fee-for-service self-insured employer-sponsored health plan between 2001 and 2007. It uses data on actual transaction payments to hospitals, including payments made both by the patient and by the insurer. It constructs three sets of key variables at the county-year level: the average price of a hospital admission; the generosity of private health insurance (equal to plans' average actuarial value); and measures of the characteristics of hospital markets, including market competitiveness. It uses an insight from labor economics to identify the causal effect of generosity, and the interaction of generosity with hospital market competitiveness, on hospital prices: that the union status and industry of workers is correlated with the generosity of their benefits. Under the assumption that unionization and industry structure in a county is otherwise uncorrelated with hospital prices, conditional on observable characteristics, this insight identifies the effect of generosity, even though the process determining plan generosity might itself depend on the level of prices for health services.

We find a statistically significant and economically important effect of plan generosity on hospital prices in uncompetitive markets. Defining a county as "high generosity" when the weighted average actuarial value of plans in it is above the median, counties with generous plans and uncompetitive hospital markets have approximately 8.6 percent higher prices than counties with low generosity and competitive markets, holding constant county- and time-fixed effects and other time-varying characteristics of counties, including their hospital market characteristics. Defining a county as high generosity when the weighted average value of plans in it is above the 75th percentile leads the estimated interaction between generosity and competitiveness to rise to 10.6 percent. These findings suggest that most of the aggregate effect of hospital market structure on prices found in previous work may be coming from areas with generous plans.

We also find substantial effects of plan generosity on the hospital admissions rate. Using the median as the cutoff for high plan generosity, counties with high generosity have approximately 18.9 percent more hospital admissions than counties with low generosity. We find no significant interaction between generosity and hospital market competitiveness on the admissions rate.

These findings are consistent with other work, like the RAND Health Insurance Experiment, that seeks to estimate the price elasticity of demand for medical care. To see this, convert the variation in actuarial values in our sample into percent changes in out-of-pocket spending at the mean: the difference in actuarial value between plans in counties above versus below the median is 11 percentage points (in counties with above-median generosity, plans have actuarial value of 92.2 percent; in counties with below-median generosity, plans have actuarial value of 81.2 percent [not in any table]). Because this amounts to a 82.7 percent decrease in the patients' share of spending (= 0.11 / 0.133 = 0.11 / (1 - 0.867) [table I]), it implies an elasticity of demand for hospital admissions of 0.23 (= 0.189 [table V] / 0.827). This is of the same order of magnitude as, although larger than, the arc elasticity of demand for hospital care of 0.17 reported in Manning et al. (Table 5.1, 1988); the difference is not surprising, given that the use of actuarial value to approximate the price of medical care generally leads to upward bias in the implied elasticity (Newhouse et al. 1980).

The key issue in our paper is whether the instruments are conditionally exogenous. We show (Table II) that the generosity of employer-sponsored insurance is increasing more slowly in counties with declining unionization and manufacturing employment, but is there some other channel through which unionization and industry affect hospital prices? To make the assumption underlying our models as plausible as possible, we control for the best available measure of input prices to the hospital production process: the Medicare wage index. The wage index is designed to measure how geographic differences in the cost of labor -- including those due to geographic differences in unionization and industry structure -- affect the cost of producing hospital services, independent of hospitals' market power. We show that the wage index has a large and significant impact on hospital prices, and the estimated elasticity of hospital prices with respect to the index of around 0.4 is approximately equal to labor's share of hospitals' accounting costs over our study period (American Hospital Association 2010).

We take three other approaches to investigate the validity of our instruments. If the estimated effect of generosity on prices is truly causal, it should be more pronounced in less competitive markets -- which it is. In contrast, the estimated effect of generosity on quantity should be independent of or declining in the extent of hospital market power. Generosity leads to increased quantity through moral hazard, whether or not hospital markets are competitive. If anything, when markets are uncompetitive (and prices are high), generosity might lead to a smaller effect on quantity. We therefore estimate models of both price and quantity that include both a direct effect of plan generosity and an interaction between generosity and hospital market power. We find that generosity has a large and significant average effect on the hospital admissions rate, but no interaction effect. Taken together, these findings rule out the hypothesis that unionization and industry structure are picking up some other unobserved time-varying characteristic of counties -- such as a taste for intensive medical care by patients or physicians -- that would affect prices and admissions rates in the same way.

It is still possible that unionization and industry structure interact with hospital market competitiveness to affect prices through some other channel -- such as rent-sharing with hospital staff. To investigate this hypothesis, we reestimated model (2c) including in X_{it} an interaction between the wage index and $HIHHI_{it}$; this approach would control for rent-sharing as long as it was proportional to the wage level. Adding the interaction effect as an exogenous variable has virtually no effect on the estimated effect of $HIAV_{it}*HIHHI_{it}$, which increases to 0.085 (standard error 0.035).

It is also possible that our results are due to the interaction of some other (unobserved) dimension of plan generosity, such as the breadth of provider networks, with hospital market competitiveness; this would be the case if unionization-induced increases in actuarial value were associated with increases in network breadth. Even if this were the case, however, it would only affect the interpretation of our estimates as a causal effect of cost sharing; it would not lead to false rejection of the null hypothesis that generosity, defined broadly, interacts with market competitiveness to increase prices. And as Gowrisankaran, Nevo, and Town (2013) observe, the bias could go the other way, if cost sharing and nonprice efforts to exercise bargaining leverage are substitutes in equilibrium.

Finally, we investigate one potential mechanism through which plan generosity and hospital market power might interact. If plan generosity increases prices by dampening consumers' incentive to search, then it should affect the variation in prices as well as the level. We therefore test whether plan generosity and hospital market competitiveness interact to affect a county-level measure of price dispersion. We find that it does. Although there is no statistically significant interaction between generosity and hospital market competitiveness when the cutoff for a generous plan is the median, there is a significant effect when the cutoff is for a generous plan is the 75th percentile. Counties in the top quartile of generosity that have uncompetitive hospital markets have approximately 21.7 percent more variation in prices, as defined by the ratio of the standard deviation of prices to the mean, holding constant county- and time-fixed effects and other time-varying characteristics of counties. This is evidence that, at least in uncompetitive markets with generous plans, reduced search is one reason that hospitals are more able to exercise market power.

These results have important implications for policy. At current levels, increasing the generosity of employer-sponsored insurance has three effects: a greater volume of hospital care, and higher levels of and variation in hospital prices in uncompetitive markets. Policies that seek to maintain or expand current levels of plan generosity therefore need to balance its potential benefits against the consequences that we document. Our results also suggest that there may be advantages to targeting antitrust enforcement at hospital markets with particularly generous employer-sponsored insurance, since it appears to be these places where hospitals are most able to exercise market power.

Nonetheless, our paper has some limitations. Because we do not observe enrollees health outcomes or satisfaction, we can not reject the hypothesis that the higher prices paid and extra hospital admissions attributable to plan generosity were "worth it." We also can not rule out the possibility that rent sharing with unionized workers (or some other unobserved process) is responsible for our results, to the extent that it would not be picked up by the interaction between market competitiveness and the overall wage level. Finally, we do not explore how generosity affects spending on physician or other outpatient services. Further investigation of these questions is an important topic for future research.

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Table I: Descriptive Statistics for Variables Used in Analysis

Medicare, AHA, and ARF variables	All US mean (sd)	Sample mean (sd)
Population (1,000,000)	0.114	0.300
	(0.336)	(0.626)
Medicare beneficiaries /pop	0.139	0.131
	(0.037)	(0.035)
Physicians/1,000 pop	2.365	2.431
	(1.491)	(1.313)
Median income (100,000 $)	0.472	0.511
	(0.124)	(0.137)
Medicare wage index	1.011	0.990
	(0.159)	(0.118)
Northeast region	0.190	0.080
Midwest region	0.225	0.281
South region	0.356	0.503
West region	0.229	0.136
Hospital capacity index	0.975	0.912
	(0.491)	(0.404)
For-profit	0.130	0.158
	(0.182)	(0.203)
Non-profit	0.754	0.718
	(0.256)	(0.273)
<100 bed hospital	0.094	0.075
	(0.153)	(0.111)
>300 bed hospital	0.497	0.534
	(0.235)	(0.208)
Teaching	0.303	0.313
	(0.241)	(0.238)
System	0.641	0.679
	(0.253)	(0.242)
HHI	0.481	0.455
	(0.151)	(0.122)
# counties	2,454	601
2007 population	279,690,260	189,058,012
2007 MarketScan enrollees		9,266,549

MarketScan variables	Sample mean (sd)
Price index	0.999
	(0.228)
Price per admission ($)	11579
	(2703)
Quantity Index	1.000
	(0.188)
Admissions per enrollee	0.056
	(0.011)
Price variation index	1.000
	(0.278)
Coefficient of variation of prices	2.364
	(0.657)
Actuarial value	0.867
	(0.093)
Unionization rate	0.104
Mining/manufacturing/transportation	0.290
Retail trade	0.050
Finance/insurance/real estate	0.055
Services	0.064
Industry not specified	0.541
Age 18-34	0.197
Age 35-44	0.177
Age 45-54	0.210
Age 55-64	0.161
Female	0.518

Notes: All US includes counties with county and market characteristics in every year 2001-2007; statistics are population weighted. MarketScan counties include those with > 100 enrollees and price index between 0.5 and 2.0 in every year 2001-2007; statistics are enrollment weighted.

Table II: Trends in Health Plan Generosity
Counties with Rising vs. Falling Unionization and Manufacturing Employment, 2001-07

	Unionization Rate		Manufacturing Employment	
	Declining 2001-07	Rising or Stable 2001-07	Declining 2001-07	Rising or Stable 2001-07
Proportion of enrollees with high-generosity plans				
2001	0.522	0.256	0.432	0.269
2007	0.694	0.664	0.671	0.754
2001-07	+0.172	+0.408	+0.239	+0.485
HHI				
2001	0.478	0.524	0.500	0.518
2007	0.416	0.479	0.429	0.474
2001-07	-0.062	-0.045	-0.071	-0.044
Household income (100,000 $)				
2001	0.463	0.408	0.455	0.404
2007	0.578	0.522	0.570	0.505
2001-07	+24.9%	+28.0%	+25.3%	+25.0%
Population (1,000,000)				
2001	0.472	0.125	0.371	0.101
2007	0.570	0.150	0.450	0.121
2001-07	+20.8%	+20.0%	+21.3%	+19.8%
# counties	235	366	354	247
# enrollees, 2007	6,192,714	3,073,835	7,755,577	1,510,972

Table III: First Stage Estimates of the Effect of Instruments on Proportion of Enrollees with Generous Health Plans

Unionization rate	0.262 *	0.425 ***
	(0.154)	(0.114)
Mining/manufacturing transportation/services	0.209 **	0.082
	(0.106)	(0.093)
Retail trade	-0.154	-0.270 ***
	(0.137)	(0.090)
Cutoff for high actuarial value	Median	75th percentile
F(3, 3579) for instruments	5.23	10.80

Notes: N = 4,207 = 601 counties x 7 years. Models also include county- and year-fixed effects plus county and market controls in Table 1. Heterscedasticity-consistent standard errors in parentheses. *, **, *** denote statistical significance at the 10, 5, and 1 percent levels, respectively.

Table IV: Effect of Health Plan Generosity and Hospital Market Competitiveness on Hospital Prices

	(1a)	(2a)	(1b)	(2b)	(1c)	(2c)	(1d)	(2d)
High actuarial value* high HHI (π)		0.019		0.035		0.086 ***		0.106 **
		(0.018)		(0.021)		(0.033)		(0.047)
High actuarial value (β)	-0.053 ***	-0.063 ***	-0.032 ***	-0.050 ***	-0.003	-0.088	0.054	-0.044
	(0.010)	(0.015)	(0.012)	(0.017)	(0.096)	(0.091)	(0.087)	(0.071)
High HHI (λ)	0.011	0.003	0.011	0.004	0.013	-0.024	0.016	-0.009
	(0.020)	(0.022)	(0.021)	(0.021)	(0.020)	(0.025)	(0.020)	(0.023)
Hospital wage index	0.446 ***	0.438 ***	0.415 ***	0.402 ***	0.443 ***	0.411 ***	0.488 ***	0.410 ***
	(0.144)	(0.144)	(0.145)	(0.145)	(0.133)	(0.135)	(0.152)	(0.143)
Effect of high HHI conditional on high actuarial value ($\pi+\lambda$)		0.022		0.039		0.062 **		0.097 **
		(0.022)		(0.026)		(0.026)		(0.042)
Cutoff for high actuarial value	Median	Median	75th percentile	75th percentile	Median	Median	75th percentile	75th percentile
Estimator	OLS	OLS	OLS	OLS	IV	IV	IV	IV

Notes: See table III.

Table V: Effect of Health Plan Generosity and Hospital Market Competitiveness on Hospital Admissions

	(1a)	(2a)	(1b)	(2b)	(1c)	(2c)	(1d)	(2d)
High actuarial value* high HHI (π)		0.005		-0.024 *		0.029		0.054
		(0.010)		(0.014)		(0.024)		(0.037)
High actuarial value (β)	0.008	0.005	0.032 ***	0.044 ***	0.189 ***	0.160 ***	0.234 ***	0.184 ***
	(0.005)	(0.008)	(0.008)	(0.012)	(0.065)	(0.058)	(0.063)	(0.048)
High HHI (λ)	-0.002	-0.004	-0.001	0.005	0.005	-0.008	0.011	-0.001
	(0.011)	(0.011)	(0.011)	(0.011)	(0.013)	(0.016)	(0.012)	(0.012)
Hospital wage index	0.054	0.050	0.082	0.091	0.042	0.031	0.254 **	0.214 **
	(0.072)	(0.072)	(0.071)	(0.071)	(0.088)	(0.085)	(0.099)	(0.087)
Effect of high HHI conditional on high actuarial value ($\pi+\lambda$)		0.001		-0.020		0.021		0.053
		(0.013)		(0.015)		(0.021)		(0.034)
Cutoff for high actuarial value	Median	Median	75th percentile	75th percentile	Median	Median	75th percentile	75th percentile
Estimator	OLS	OLS	OLS	OLS	IV	IV	IV	IV

Notes: See table III.

Table VI: Effect of Health Plan Generosity and Hospital Market Competitiveness on Hospital Price Variation

	(1a)	(2a)	(1b)	(2b)	(1c)	(2c)	(1d)	(2d)
High actuarial value* high HHI (π)		0.050 *		0.039		0.042		0.188 **
		(0.030)		(0.036)		(0.059)		(0.082)
High actuarial value (β)	0.039 **	0.014	0.043 **	0.023	-0.201	-0.242	-0.001	-0.173
	(0.017)	(0.027)	(0.020)	(0.033)	(0.158)	(0.161)	(0.125)	(0.118)
High HHI (λ)	0.009	-0.012	0.010	0.002	0.000	-0.018	0.008	-0.036
	(0.031)	(0.034)	(0.032)	(0.033)	(0.033)	(0.042)	(0.030)	(0.034)
Hospital wage index	0.808 ***	0.788 ***	0.847 ***	0.832 ***	0.824 ***	0.809 ***	0.810 ***	0.673 ***
	(0.213)	(0.214)	(0.212)	(0.212)	(0.211)	(0.214)	(0.222)	(0.209)
Effect of high HHI conditional on high actuarial value (π+λ)		0.039		0.041		0.024		0.151 **
		(0.036)		(0.043)		(0.048)		(0.073)
Cutoff for high actuarial value	Median	Median	75th percentile	75th percentile	Median	Median	75th percentile	75th percentile
Estimator	OLS	OLS	OLS	OLS	IV	IV	IV	IV

Notes: See table III.