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ABSTRACT

During the 1990s US healthcare markets underwent a significant transformation. Managed care rose to become the dominant form of insurance in the private sector. Also, a wave of hospital consolidation occurred. In 1990, the mean population-weighted hospital Herfindahl-Hirschman Index (HHI) in a Health Services Area (HSA) was .19. By 2000, the HHI had risen to .26. This paper explores whether the rise in managed care caused the increase in hospital concentration. We use an instrumental variables approach with 10-year differences to identify the relationship between managed care penetration and hospital consolidation. Our results strongly imply that the rise of managed care did not cause the hospital consolidation wave. This finding is robust to a number of different specifications.

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Introduction

During the 1990s the landscape of US healthcare markets underwent a dramatic transformation. Managed care in one of its many forms displaced indemnity insurance to become the dominant form of insurance in the private sector (Glied, 2001). Over the same period, a wave of hospital mergers, acquisitions and hospital system expansions occurred. In 1990, the mean, population weighted, hospital Herfindahl-Hirschman Index (HHI) in a Health Services Area (HSA) was .1903. By 2000, the HHI had risen to .2566.[1] The HHI measures the level of 'competition' in a market with 0 being perfect competition and 1 being pure monopoly.

Understanding the determinants of hospital concentration allows the development of policies to manage hospital competition. Hospital market power impacts the functioning of health care markets in several important ways. First, as hospitals garner more market power, their bargaining position with payers may improve and that can increase the cost of hospital care for the privately-insured population. Research shows that increases in hospital concentration lead to higher hospital prices (Gaynor and Vogt, 2001, and Dranove and Saitherwaitte, 2001, provide excellent reviews of this literature). Increased hospital market power is also associated with decreased hospital quality (Kessler and McClellan, 2002; and Gowrisankaran and Town, 2003). Finally, there is some evidence that decreases in hospital competition decrease access to health services for underserved populations (Aizer, Currie and Morretti, 2004).

[1] The unweighted HHI is .31 in 1990 and .37 in 2000.

It is conventional wisdom that the rise of managed care precipitated the hospital consolidations and concentration in the 1990s.[2] Graphical analysis is consistent with the conventional wisdom. Figure 1 graphs the annual mean population-weighted levels of hospital HHI and HMO penetration across HSAs.[3] Two hospital HHI measures are graphed, overall concentration and concentration due to changes in hospital ownership structure (i.e. mergers, acquisitions and hospital system formation and expansion).[4] Several features of this graph are noteworthy. First, hospital concentration and HMO penetration share a common trend throughout most of the 1990s. Only in the late 1990s was there a break in this trend, as HMO penetration experienced a slight decline while hospital concentration continued to increase. While this correlation is suggestive, it does not prove that there is a causal link between HMO penetration and hospital concentration. The second noteworthy feature of Figure 1 is that most of the increase in hospital concentration occurred in the mid-1990s. Finally, of the .063 overall increase in the population weighted HHI, .053 is due to mergers, acquisitions or hospital system expansion. The remaining .01 change in HHI is a consequence of hospital exits and changes in the distribution of beds across hospitals.

This paper tests the proposition that the rise of managed care caused hospitals to consolidate in the 1990s. Our models explicitly account for the possibility that markets are systematically heterogeneous and that this heterogeneity may bias cross-sectional estimates of the parameters of interest. We also use an instrumental variables (IV)

[2] For example, Evans Cuellar and Gertler (2005) say, "Hospital consolidation is likely a response to managed care,." (p.214). Also, see Dranove, Simon and White (2002), Employee Benefit Research Institute (1999), Czajkowski (1999), Grembowski et al. (2002), McCue et al. (1999) and Hollis (1997).

[3] Our measure of hospital concentration includes direct ownership of hospital assets as well as membership in a hospital system.

[4] The exact calculation of the latter measure is discussed in the Methods section.

approach to account for the possibility that managed care penetration may be endogenous.

Our estimates indicate that the rise in managed care did *not* cause the increase in hospital concentration. In every specification we estimated, the coefficient of managed care penetration is negative and in many of the specifications it is significant. That is, if anything, our estimates indicate that the rise of HMOs reduced hospital concentration.

The next section discusses some of the reasons why managed care might lead to hospital consolidation and reviews the relevant literature. The following sections present the methods, the data, the results and the discussion of our findings.

Hospital Consolidation and Managed Care

Why Managed Care May Cause Hospital Consolidation

There is a widespread belief that the wave of hospital consolidation that occurred over the 1990s was driven by the rise of managed care. There are at least three reasons why this might be so. First, managed care may reduce the demand for hospital beds thus creating excess capacity in the market. Second, managed care may change the bargaining power of hospitals relative to health insurers. Third, the value of contracting with an integrated hospital system may be greater for managed care organizations than indemnity insurers.

One of the theories underlying managed care is that by monitoring and controlling health care use, insurers can reduce health care expenditures and perhaps increase enrollee health. In the RAND health insurance experiment, enrollment in the prepaid Group Health Cooperative of Puget Sound reduced the likelihood of a hospital admission

by 35% compared with a fee-for-service population (Manning et al., 1987).[5] However, as an explanation of hospital consolidation this finding needs embellishment. Most economic models of mergers predict that the incentive to merge increases with the demand for the product (e.g. Deneckere and Davidson, 1985). The intuition behind this result is straightforward—the larger the market the larger the profit gains from market power.

However, it is possible to conceive of circumstances in which a decline in demand for inpatient services leads to hospital consolidation. If demand falls far enough so the market can no longer support the old number of hospitals under the old ownership structure, then there may be an incentive to merge. According to neo-classical economic theory, if the reduction in demand leads to lower prices and if price falls below average variable cost, the market will need to remove the capacity in some way. Hospital closure is one way to reduce capacity. However, hospital assets have "high specificity" and few alternative uses. These features of the hospital market may drive hospitals to seek to combine operations in order to gain market power and thereby raise price above average cost. These combinations may also allow hospitals to achieve efficiencies and reduce average costs, thereby keeping price above average costs. Hospitals may prefer to merge and perhaps lose valued autonomy rather than close. This explanation requires that hospitals value autonomy as well as profits. This is a plausible assumption given the hospital industry's historical basis as a decentralized, community-based system. If autonomy was not valued, the hospitals would have been better off by merging earlier. Another motivation for consolidation is that merging hospitals may wish to internalize

[5] While managed care may reduce demand, the advancement of medicine has surely had a larger effect on the demand for inpatient days. According to our estimates, from 1990 to 2000 average inpatient days per-capita declined by 54%.

externalities. In the face of excess capacity hospitals may want to reduce capacity by closing down wings or converting parts of their hospital to an outpatient or skilled nursing facility. But, there may be an externality problem. If a hospital reduces the number of beds, then all hospitals in the market may benefit. A merger or hospital system expansion may allow hospitals partially to internalize that externality.

The second reason that increases in managed care might trigger hospital consolidation is based on the ability of managed care organizations (MCOs) to bargain effectively with hospitals over prices. The rise of managed care may change the price elasticity of demand for hospital services. By selective contracting, MCOs increase their bargaining leverage with hospitals vis-à-vis indemnity plans. However, in economic theory the effect of increasing the price elasticity of demand on the incentive to consolidate is ambiguous. Mergers in markets with intermediate levels of price elasticity are more likely to increase market power than those in markets with either low or high elasticity. Thus, the common belief that changes in bargaining power between payers and hospitals increase the incentive to merge is not necessarily true.

The third possible reason why managed care may increase the incentive for hospitals to consolidate is that by forming an integrated delivery system the hospitals are better able to manage patient care (Shortell, et al., 1996). This ability to better manage patient care may make the organization better able to engage in capitation arrangements with health plans. The formation of these integrated delivery systems requires hospitals to form both vertical (with physician groups) and horizontal (with other hospitals) arrangements so that they can manage care for a geographically dispersed population. That is, in order for the managed care organizations to achieve efficiencies they need to

write 'high-powered' contracts and hospitals need to integrate to be able to respond effectively to the incentives contained in these contracts. The development of integrated delivery systems is a weak explanation for the link between managed care and hospital consolidation. Capitated hospital arrangements never became widespread and integrated delivery systems ultimately "didn't work." (Burns and Pauly, 2002).

Empirical Evidence

There are two studies of the relationship between managed care penetration and hospital market competition. Dranove, Simon and White (2002) (herafter, DSW) use data on physicians' reports of managed care revenue and find that the change in hospital concentration in 68 large MSAs between 1981 and 1994 is positively correlated with the level of managed care concentration in 1993/1994. This research has some limitations. First, the time period of the analysis is only of modest interest because most of the hospital consolidations and most of the increase in managed care in the 1990s occurred after 1994. Much of the variation in the variables of interest is missing. Second, DSW focus on large metropolitan areas in which hospital mergers are much less likely to lead to significant increases in market power. The mean end-of-period Herfindahl-Hirschman Index they report is .126 – well below concentration levels that would give the Federal Trade Commission or the Department of Justice cause for concern.[6] Third, the measure of managed care penetration is the mean percent of physician practice revenues from managed care. This measure of managed care penetration, based on the physician services factor markets, is not widely used in research. It is also a function of physician prices in a location and thus may be more subject to endogeneity problems than simple

[6] It is also likely that antitrust markets are much smaller than the large MSAs and thus they likely have significant measurement error.

managed care penetration rates. However, a potential advantage of DSW's measure is that it includes all forms of managed care, not just HMOs. The more commonly used measure is penetration based on the managed care product market.[7] Fourth, DSW regress the change in concentration on the end-of-period *level* of managed care penetration. They argue that this is a reasonable approach because managed care penetration in most cites was close to zero at the beginning of their sample. However, in this type of statistical analysis, the difference between being 'close' to zero and being exactly zero matters. In this circumstance, failure to difference a variable, even if it is close to but different from zero, can still lead to significant biases.[8] According to InterStudy data, 62% of MSAs with over 800,000 population (roughly DSW's sample) had managed care penetration greater than 5% in 1985 and 37% had HMO penetration greater than 5% in 1981.[9] We replicate and discuss the DSW results in greater detail in the Results section.

Chernew (1995) analyzes the relationship between the change in the number of hospitals and managed care penetration in 175 large MSAs between 1982 and 1987. He finds a negative and significant relationship between managed care penetration and the number of hospitals in an MSA. Like DSW, Chernew's analysis occurs before the rise in managed care. Also, Chernew's analysis does not explicitly focus on changes in market

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[7] For example, Baker and Brown (1999) use HMO penetration data from Group Health Association of America as a measure of managed care penetration.

[8] We performed the following Monte Carlo simulation. The base model is $y_{it} = \alpha_i + .5x_{it} + e_{it}$ for t=1,2. The variables α_i, x_1 and x_2 are drawn from uniform distributions (with censoring) with a mean of .2, .035 and .2, respectively. About half of the x_1's are equal to zero. The e_{it} are drawn from a N(0,.1). The x's are correlated with α_i ($\rho = .5$). We performed 500 simulations each with 70 observations. Estimation of the coefficient on *x* using the approach of DSW resulted in a bias of approximately 50%.

[9] The PMSA with over 5% HMO penetration in 1981 include: Denver, CO; Portland, OR; Orange County, CA; Seattle, WA; New York, NY; Rochester, NY; Minneapolis, MN; Sacramento, CA; Los Angeles, CA; Miami, FL; Milwaukee, WI; Riverside Co., CA; San Diego, CA; Phoenix, AZ; Washington, DC and San Francisco, CA. DSW exclude Washington, DC, three California PMSAs, and Minneapolis from one of their analyses to test the robustness of their conclusions.

structure due to consolidation. The number of hospital competitors in a market can change due to consolidation (although a consolidation does not necessarily change the physical number of hospitals in a location) or because of entry or exit.

In related work, Baker and Brown (1999) examine the impact of managed care on the number of mammography providers. Their data is cross-sectional and they instrument for managed care penetration using the size distribution of employers in the locale. They find that managed care penetration reduces the number of mammography providers.

Methods

Our approach to estimate the impact of managed care on hospital consolidation is to relate the change in HMO penetration to one of several possible measures of the change in hospital concentration in a market area. Our empirical model is:

(1) $$H_{it} = \alpha_i + HMO_{it}\beta_t + x_{it}^{'}\gamma_t + e_{it}.$$

Hospital concentration in market *i* at time *t*, H_{it}, is a linear function of HMO penetration, HMO_{it}, market characteristics, x_{it}, a time-invariant unobserved (to the researcher) market fixed effect, α_i, and a market/time shock, e_{it}. We let β_t depend on time as changes in technology or unmeasured market conditions may affect the incentive for hospitals to consolidate in response to managed care.

It is possible and perhaps probable that the unobserved market fixed effect is correlated with HMO penetration. There are many possible reasons for this correlation. For example, locations in which hospitals have high α's may be unattractive for HMOs to enter because they have fewer hospitals with which to contract. If α is correlated with HMO penetration, OLS estimation of (1) will lead to biased estimates of β_t. This bias can be eliminated by taking differences of (1) across time.

Taking differences between period *t* and period *t-1* gives

$$(2) \qquad \Delta H_{it} = \Delta HMO_{it}\beta_t + HMO_{it-1}\beta^e + \Delta x_{it}^{'}\gamma_t + x_{it-1}^{'}\gamma^e + u_t$$

where Δ is the standard difference operator (i.e. $\Delta H_{it} \equiv H_{it} - H_{it-1}$), $\beta^e = \beta_t - \beta_{t-1}$, $\gamma^e = \gamma_t - \gamma_{t-1}$ and $u_{it} = e_{it} - e_{it-1}$. One may still be concerned that the error term in (2), u_t, may be correlated with our HMO variables. Our strategy is to correct this possible endogeneity by estimating the parameters of (2) with an instrumental variable approach. We discuss the instrument set later in this section.

We use a long difference of 10 years to define our change variable, with 1990 as the base year.[10] We chose 1990 as the starting year because that is the approximate beginning of the horizontal (within market) consolidation wave. The mean, population-weighted HHI in 1985 was .1900 and in 1990 it was .1903 – a very small difference over the five year period. However, the increase in average HHI between 1990 and 1991, the years of the smallest increase in our sample, was .0022, a rate of increase that is 55 times larger than the annual rate of increase over the preceding five years.

Measures of the Change in Concentration

The most common measure of competition used by economists is the Herfindahl-Hirschman Index (HHI). This measure serves as the basis for two of the three measures of the change in concentration in our analysis. The other measure of the change in hospital concentration that we study is the number of mergers between hospitals in the geographic area.

[10] We have performed our analysis on both a longer period, 1985 to 2000, and a shorter period, 1992-1998. The results from those analyses are qualitatively identical to those we present here. These results are available from the authors upon request.

The HHI is simply the sum of the squared market shares of all firms in a given market. That is, the HHI is defined as,

$$(3) \qquad HHI_{it} = \sum_{i=1}^{N} s_{it}^2$$

where s is the market share of the hospital system within the market, $s_{it} \equiv \frac{q_{it}}{\sum_{j=1}^{N} q_{jt}}$ and q is a measure of output/capacity while N is the number of market participants.

We made two important decisions in calculating the HHI. The first decision concerns the boundaries (both geographic and product) of the market. As we discuss in the data section, we use a Health Services Area (HSA) as the geographic boundary and short-term, non-governmental inpatient care as the product market. The second decision concerns the appropriate measure of output. Our measure of output is the total number of staffed beds. We tested the sensitivity of our results to these design decisions by repeating our analysis using HHI constructed from alternative measures of the geographic market (MSAs) and output (inpatient days) and found that our conclusions are robust to these alternatives.[11]

The standard HHI can change from period to period within a market for several reasons. The most obvious one is a merger or system formation between two market participants. However, this is not the only reason that the HHI can change. The HHI will change if the distribution of output/capacity changes or if there is entry into or exit from the market. If managed care does have an impact on the HHI through causing mergers we

[11] Results are available from the authors upon request.

would like to isolate the change in the HHI due to mergers. So, we developed a measure of the change in HHI due to consolidation (ΔMHHI).

We define ΔMHHI as follows,

$$(3) \qquad \Delta MHHI_{it,r} = \sum_{i=1}^{N_r} s(O_t)_{ir}^2 - \sum_{i=1}^{N_r} s_{ir}^2$$

where $s(O_t)_{ir}^2$ is the share based on the distribution of output/capacity from period *r* aggregated to the hospital using the ownership/system structure in period *t* (*r<t)*. If there is only one change in ownership/system structure in a location (denote the hospitals as *j* and *k*) during the period then the $\Delta MHHI_{it,r} = 2s_{jr}s_{kr}$. For the most part, ΔMHHI will be either 0 (if no changes in within-location ownership/system structures occur) or positive (if there is a consolidation in a location). If a de-merger occurs then ΔMHHI is negative.

Our third measure of the change in hospital concentration is simply an indicator taking the value of 1 if a merger occurred between hospitals in an HSA and 0 otherwise. In essence, this is simply an indicator of whether ΔMHHI is greater than zero. We include this variable in our set of dependent variables because ΔMHHI is a nonlinear function of market shares and the resulting distribution is highly skewed as well as censored. A high degree of skewness in the distribution of a left-hand side variable can lead to misleading inferences because of the presence of outliers – thus analyzing the parameter estimates from the indicator allows us to examine the robustness of our estimates from the ΔMHHI regression.

Instruments

The possibility that the error term in (2) may be correlated with our measures of HMO penetration caused us to seek out potential instruments for HMO penetration. There are two possible reasons for this correlation: omitted variable bias (primarily health status) and unobserved characteristics of hospital markets that may influence HMO penetration.

We use as instruments the number of potential HMO entrants in a market in 1990, the number of non-specialist physicians per capita in 1990, the percent of the population that is self-employed and the percent of the population employed in firms with over a 100 employees in 1990. These variables are plausibly unrelated to changes in hospital market structure and they predict HMO penetration.

The number of potential HMO entrants is simply the number of HMOs that are participating in a market within the state in 1990 but have yet to enter into the HSA in question. The rationale behind this measure is that an HMO must meet a number of state-level regulatory hurdles before it can enter a market, and if an HMO is already participating in a state, then its entry costs will be lower than if it tries to enter a market without any previous experience in that state. Our hypothesis is that more potential HMO entrants will increase actual entry in a market and that HMO penetration will be higher.

We hypothesize that HMOs will have more bargaining power and consequently lower costs of doing business in markets with higher numbers of primary care physicians per capita. Baker and Brown (1999) used the size distribution of employers to instrument for HMO penetration in their analysis of the impact of managed care on mammography providers.

It is well known that if the instrument set is a poor predictor of the endogenous variables conditional on all of the control variables, then small-sample bias can be very large (Stock and Staiger, 1997). Following the suggestion of Stock and Staiger (1997) and Bound, Jaeger and Baker (1995), we performed F-tests of the joint significance of the instrument set for the endogenous variables, ΔHMO Penetration and HMO Penetration. The F-statistics are 3.36 and 6.29 for the regressions with corresponding p-values of .003 and <.0001, respectively. Both of these F-statistics are significantly greater than 1—a value that Bound, Jaeger and Baker (1995) suggest should be "a cause for concern."[12]

Estimation Strategy

We have three different dependent variables – ΔHHI, ΔMHHI, and the merger indicator. Only ΔHHI has a distribution that is appropriate for using linear methods. The distribution of ΔMHHI is essentially censored with two-thirds of the observations equal to zero. Of course, the distribution of the merger indicator is discrete.

For each dependent variable we estimate two regressions – one treating the HMO penetration variables as exogenous and one instrumental variable regression. For ΔHHI we use linear regression with heteroskedasticity correction and instrumental variable estimation. As the distribution of ΔMHHI is essentially censored, we estimate the parameters using standard Tobit and instrumental Tobit as described by Newy (1989). Finally, we use Probit and instrumental variable Probit approaches to estimate the parameters from the merger indicator regression.

[12] The F-test suggests that the small sample bias will be less than .30. The presence of this bias will not overturn any of our conclusions.

Data

An observation in our analysis is a Health Services Area (HSA), defined by the National Center for Health Statistics (Makuc, et al., 1991) as one or more counties that are relatively self-contained with respect to the provision of routine hospital care. This definition is operationalized by using an algorithm that groups counties by minimizing travel by patients within the area to areas outside of the defined group. Thus, unlike Metropolitan Statistical Areas (MSAs), the HSA relies on patient flows instead of geo-political boundaries to identify hospital 'markets.'[13]

This definition is not ideal. The ideal unit of observation for our study would be an antitrust market. However, defining antitrust markets is notoriously difficult. It requires detailed analysis of each hospital under consideration, and there is still substantial disagreement within the economic and legal communities on the appropriate methodology to use (Capps et al., 2002; Werden, 1990). While it is important to recognize that HSAs are not antitrust markets, they are probably closer approximation to them than MSAs. Most economists believe that antitrust markets typically are smaller than MSAs, and HSAs, on average, are significantly smaller than MSAs.

The data in this study come from several sources. Data on hospitals come from the American Hospital Association (AHA) hospital survey. This commonly-used data set contains information on hospital location, size (beds and inpatient days) and system membership. The AHA survey also tracks mergers and acquisition of different hospitals in an appendix to the documentation. In most of the analysis, we use data from 1990 to

[13] A literature in antitrust analysis suggests that using patient flows to define markets can be misleading (Werden, 1990, and Capps et al., 2002). However, it is probably the case that using patient flow data to identify markets is more accurate than using existing geo-political boundaries.

2000 but some of our sensitivity analyses go back to the 1985 hospital survey. Each hospital is assigned to an HSA according to its address in the 1990 AHA Survey.

Our information on the structure of the managed care market comes from InterStudy. InterStudy collects information on the number of HMO enrollees by health plan. The HMO data come from InterStudy census data (InterStudy, 1985-1987; InterStudy, 1988-2001) and Group Health Association HMO Directories (Group Health Association of America, 1989-1992). The InterStudy Census includes HMO location, founding year, model type, not-for-profit status, federal qualification, and national affiliation. Enrollment by geographic area comes from multiple sources. From 1985 to 1988, enrollment data comes from the InterStudy Census and supplemental InterStudy reports on areas served by HMOs (Hartwell, et al. 1986). From 1988 to 1991, the list of counties an HMO operates in was obtained from Group Health Association HMO Directories (Group Health Association of America, 1989-1992). From 1992 to 1996, the list of counties where each HMO operates and enrollment by MSA was obtained from 1992 through 1996. Following 1996, InterStudy obtained county-level information on most HMOs in its annual survey and provided country enrollment measures. Where county enrollments were unavailable, enrollment for each HMO was pro-rated on the basis of county population to the each county served by the HMO.

There is a significant change in enrollment distribution between 1987 and 1988 because of the significant improvement in the listing of counties served. Because of the less complete listing of counties served in 1985 to 1987, MSA enrollment may be overstated. But, pro-rating HMO enrollment on the basis of a larger list of counties served may overstate rural country enrollment if HMOs predominantly developed in urban areas and

then expanded to adjacent rural areas. Since HMO development was more of an urban than rural phenomena in the mid- to late- 1980s, the 1985 enrollment measures probably reflect HMO development reasonably well.

County level market measures come from the Area Resource File (Bureau of Health Professions, 1999). State level wage data come from the Bureau of Labor Statistics Occupational Employment Statistics (Bureau of Labor Statistics, 2001). State HMO regulations come from reports compiled by Aspen Publishers (Aspen System Corporation 1985-1994; Aspen System Corporation 1988-1994; Aspen Systems Corporation 1996; Levy 1999).

Health System Areas (HSA) measures were constructed by first measuring HMO enrollment, Area Resource File characteristics, and wage data at the county level. Then, HSA-level measures were constructed as weighted averages of all counties in the HSA, with the weights being the proportions of the HSA's total HMO enrollment in each county. County data were aggregated into HSAs using a crosswalk between counties and HSAs.

We supplement the hospital and HMO information with data from the Area Resource File (ARF). For each year, the ARF has information on the number of primary care physicians, median per-capita income, unemployment rate, population of the HSA, and percent of the population over 65 years of age. We also acquired information on state hospital certificate of need (CON) regulations from the American Health Planning Association.

We limit our sample to HSAs with a population of more than 50,000 and more than one hospital 1990. We dropped the monopoly HSAs as those areas, by definition,

cannot experience a horizontal consolidation (unless there is entry and then merger/system formation – a very unlikely possibility).[14]

Table 1 lists the dependent and explanatory variables as well as the instruments in the analysis and the un-weighted means and standard deviations of these variables. Histograms of the ΔHHI and ΔMHHI are graphed in Figures 2a and 2b, respectively. The distribution of ΔHHI looks roughly symmetric while the distribution of ΔMHHI is essentially censored at zero.

Several patterns in Table 1 are noteworthy. As mentioned earlier, we see a large increase in hospital concentration and HMO penetration during the 1990s. Approximately 40% of all HSAs experienced a horizontal hospital merger. We also observe large declines in inpatient days per capita (20%) and the number of beds per capita (17%) over this 10-year period. In some ancillary analysis we have performed, the decline in inpatient days and beds per capita appears unrelated to HMO penetration.[15]

Results

Figure 3 graphs the number of *horizontal* hospital mergers, acquisitions and system expansions, total number (both horizontal and across HSAs) of hospital mergers, acquisitions and system expansions, and the number of hospital failures over time for our sample HSAs. Several features of the graph are of interest. First, in the early 1990s there was an increase in both the number of horizontal ownership changes and across-market ownership changes with the peak occurring in the mid-1990s. The height of the merger/system activity occurred two years before the peak of HMO penetration. Second,

[14] In order to check the robustness of our findings, we have also performed the analysis limiting the sample to those HSAs with less than 11 hospitals. Again, the qualitative results using this sample are identical to those we present here. These results are available from the authors upon request.
[15] These results are available from the authors upon request.

horizontal ownership changes account for approximately 40% of all ownership changes. Third, there is a modest inverse relationship between the number of hospital consolidations and the number of hospital failures.

Tables 2, 3 and 4 presents two sets of coefficient results for the three dependent variables--ΔHHI, ΔMHHI and the market consolidation indicator, respectively. For each dependent variable the first set of results presents the coefficient estimates treating HMO penetration as exogenous. The second set of results is the instrumental variable coefficient estimates. For each estimation method we present estimates from two specifications. The first set of coefficient estimates is from a specification that includes both the change in HMO penetration and the level of HMO penetration in 1990 as well as all of the control variables. The second set of estimates is from a specification that includes just the change in HMO penetration and all of the control variables. This specification tests the sensitivity of the results by imposing the restriction that β is time-invariant.

Table 2 presents the coefficient estimates The OLS estimated coefficient on ΔHMO Penetration is negative, small in magnitude and insignificant at traditional levels of confidence in both columns (1) and (2) of Table 2. That is, an increase in HMO penetration is associated with a *decrease* in measured hospital competition. In column (1) the coefficient on the level of HMO penetration in 1990 is small, positive and imprecisely estimated. The coefficient estimates are very similar between column (1) and (2) and the likelihood ratio and t-test do not reject the hypothesis that β is time invariant.

The IV estimates of the HMO penetration coefficients are also negative and larger in magnitude. The coefficient on ΔHMO Penetration in both columns (3) and (4) are just

significant at the 5% level. That is, there appears to be an *inverse* relationship between hospital concentration and HMO penetration. The coefficient on HMO Penetration in 1990 in column (3) is positive but insignificant. The IV estimates do not suggest that β varies over time.

Table 3 presents the coefficient estimates with ΔMHHI as the dependent variable. As ΔMHHI is censored, we estimate the parameters using Tobit and IV-Tobit procedures. Columns (1) and (2) present the maximum likelihood Tobit estimates. The coefficients on ΔHMO Penetration in both regressions are negative, small in magnitude and insignificant at traditional levels of confidence. The coefficient on HMO Penetration in 1990 is positive, small in magnitude and insignificant.

The IV estimates are presented in columns (3) and (4) of Table 3. The coefficient estimates on ΔHMO Penetration are again negative, large in magnitude and flirt with significance at the 5% level of confidence. That is, there is modest evidence that an increase in HMO penetration actually reduced consolidation activity. The coefficient on HMO Penetration in column (3) is positive and is imprecisely estimated.

Table 4 presents the coefficient results with the consolidation indicator as the dependent variable. The pattern of the coefficients is similar in the Probit and IV-Probit regressions. The coefficient on ΔHMO Penetration is negative and insignificant in all four regressions. The coefficient on HMO Penetration is positive and insignificant in columns (1) and (3). These results in combination with the results in Table 3 suggest that an increase in HMO penetration did not affect the likelihood of a consolidation but,

conditional on a consolidation occurring, reduced the average size of the consolidating parties.[16]

While the rise in HMO penetration does not seem to explain hospital consolidation, the presence of excess capacity is strongly correlated with hospital consolidation due to merger. The t-statistics on Hospitals Per Capita range from 2.60 to 4.39 in the ΔMHHI and hospital consolidation indicator regression. A one standard deviation increase in beds per capita (1.84) is associated with a non-trivial .048 increase in hospital concentration due to consolidation.

Several other variables are significantly associated with consolidation activity. The parameter estimates in all four regressions indicated that increases in the percent of the population in poverty and increases in the population level in 1990 are associated with increases in consolidation activity. Finally, an increase in the proportion of hospitals that are for-profit is associated with an increased likelihood of a consolidation.

Robustness

We estimated the parameters using different time frames (1990 to 1995, 1995 to 2000, and 1990 to 1998). The parameter estimates from these regressions lead to the same conclusions as those presented in Tables 2, 3 and 4. We also estimated the parameters on several different samples that were split by initial population size (split at 500,000 and 800,000) and HMO penetration (split at 5% and 10%) and again we found no significant relationship between HMO penetration and hospital consolidations in any of those samples. Finally, we also estimated the parameters weighting the observations by

[16] Two-stage estimation supports this conclusion.

the HSA population and the implications of the coefficients are not meaningfully different.

Explaining the Difference between Our Estimates and DSW

There are several possible reasons why our results differ from DSW. First, we difference both our left-hand and right-hand variables of interest while DSW do not difference their measure of HMO penetration. Second, DSW use a more inclusive measure of managed care but one that may have more measurement error. Third, they examine a different time frame, 1981 to 1994. Fourth, they use the MSA as the unit of analysis. Fifth, they limit their analysis to the largest 68 MSAs with population over 800,000.

In order to determine the source of our differing conclusions we first attempt to replicate their results with our data. We use the largest 68 MSAs and data from 1985 to 1994. We assumed that the final observed penetration level as a measure of penetration, assuming that initial penetration was zero. We use a similar, but not identical, set of control variables and instruments.[17]

Table A2 in the appendix presents the results of our attempts to replicate DSW. In column (1) we present the coefficient estimates of this effort. As in DSW, the coefficient on managed care penetration is positive and significant. In column (2) we estimate the same equation but replace the level of managed care penetration with the change in the managed care penetration. The coefficient on the managed care variable declines and becomes insignificant. However, the power of the instruments declines substantially raising the possibility of small sample bias. Nevertheless, inferences regarding the impact

[17] DSW use the percent of workforce self-employed and percentage of workforce employed in large firms in 1992 as instruments. We use the percentage of firms in each size category in 1992 and 1985 as instruments.

of managed care on hospital concentration are sensitive to the decision to difference the variable of interest.

In columns (3) and (4) of Table A2 we estimate the parameters as in columns (1) and (2) on the same sample, but using OLS. The parameter estimates in column (3) are very close to the estimates in column (1). However, the OLS estimates of the impact of the differenced managed care variable are essentially zero. In the last two columns of Table A2, we estimate the same specification as in the first two columns but use a more inclusive sample selection rule. We include all MSAs with a population greater than 500,000. Both coefficients on managed care population are insignificant in this sample.

Our estimates suggest that the results in DSW are sensitive to the decisions not to difference the managed care variable, the sample selection criteria and the time period they studied. While these findings are very suggestive, they are not conclusive as the data we use are not identical to DSW.

Discussion

In 1996, the two hospitals in Great Falls, MT, merged to form a "monopoly" hospital.[18] According to the State of Montana's Bureau of Business and Economic Research, "Managed care in Montana is virtually nonexistent, with recent data showing a miniscule number of people covered by health maintenance organizations."[19] That is, the anecdotal evidence from Great Falls suggests that managed care played no role in causing the large increase in hospital concentration. The parties to the merger argued that there was not enough demand for inpatient services to support two acute care hospitals in Great

[18] The two merging hospitals were Columbus Hospital and Montana Deaconess Medical Center.
[19] http://www.doj.state.mt.us/safety/greatfallshospital/decisionamended1996.pdf

Falls.[20] The data support this contention. In 1990, Great Falls had 7.56 beds per 1,000—the average in our sample is 3.85.

At the other end of the urban spectrum is Philadelphia. The Philadelphia metropolitan area (five counties in southeast Pennsylvania, 3 in New Jersey and 3 in Delaware) contained 65 hospitals, five of which were academic medical centers and three of which were children's hospitals. Philadelphia experienced a large growth in HMO penetration through the 1990s. In 1990, the HMO penetration was approximately 16%. By 2000, HMO penetration had risen to 53%. However, hospital HHI only increased a modest .023.

Furthermore, most of the increase in concentration in Philadelphia is unrelated to HMO penetration. The HHI increase is attributable, in large part, to the entry into the market by the Pittsburgh-based Allegheny Health Educational and Research Foundation (AHERF). AHERF acquired the poorest-performing academic medical center. AHERF's entry touched off a competitive rivalry among Philadelphia's academic medical centers to build bigger and bigger hospital systems. AHERF's acquisition of one of the three pediatric hospitals similarly touched off intense competitive strategies by hospitals to develop pediatric referral networks with community hospitals and physicians. (Burns, et al., 1997; Burns et al., 2000).

These two accounts of the evolution of hospital market structure from very different health care markets are consistent with our empirical findings. That is, the 'lore' that the hospital consolidation wave of the 1990s was caused by hospitals responding to the managed care revolution is not consistent with the data.

[20] See http://www.doj.state.mt.us/safety/greatfallshospital/decisionamended1996.pdf.

Our work raises an important and yet to be answered question: what really caused the horizontal hospital merger wave? Our results hint at a possible answer. The coefficient on Beds Per Capita is positive, large and significant in the all specifications in Table 3 and 4. This result is consistent with two possible and not mutually exclusive explanations. First, reduction of bed capacity is a public good (from the hospital's perspective) and consolidation may help hospitals internalize this externality. Second, hospitals located in markets in which there is higher bed capacity likely are in worse financial condition than their counterparts in under-bedded areas. Mergers may be the best alternative to closing the facility.

Conclusions

It is widely believed that rise of managed care caused the hospital consolidation wave of the 1990s. In this study we test this proposition using data on managed care penetration and hospital consolidation from 1990 to 2000. Our results suggest that the common wisdom is false—managed care penetration is not significantly related to hospital consolidation. This finding is robust to different specifications, time frames and sample selection criteria. This finding raises an obvious question: if managed care did not cause hospital consolidation—what did?

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Table 1

Summary Statistics
(standard deviations in parenthesis)

Variable	1990	2000	Δ or Percent Δ
Dependent Variables			
HHI	.31 (.075)	.37 (.16)	.068 (.11)
Merger HHI (Base = 1990)	.31 (.075)	.36 (.17)	.051 (.093)
Merger Indicator	.023 (.15)	.057 (.23)	.48 (.50)
Other Explanatory Variables			
HMO Penetration	5.4% (6.3%)	16.6% (12.4%)	11.1 (10.5)
Percent Elderly	13.4% (3.1%)	13.3% (2.8%)	-.18% (.84)
Population	413,894 (769,205)	466,792 (859,998)	10.6% (10.0%)
Median Per Capita Income	$15,956 ($3,197)	$24,101 ($5,203)	13.0% (6.5%)
Percent Population in Poverty	15.2% (6.24%)	13.6% (5.24%)	-1.65% (1.84%)
Inpatient Days per capita	.86 (.33)	.68 (.31)	-.17 (.20)
Beds per 1,000 population	3.85 (1.32)	3.17 (1.26)	-.65 (.80)
Percent FP Hospital in 1990		11.3% (17.8%)	
Hospital CON Regulation in 1990		54.8% (49.8)	
Land Area in 1990 (square miles)		5,928 (47,242)	
Instruments			
Potential HMO Entrants in 1990		12.7 (9.70)	
Primary Care Physicians per 1,000 population in 1990		.29 (.094)	
Percent of Establishments smaller than 5 employees in 1990		55.3% (3.5%)	
Percent of Establishments with between 5 and 9 employees in 1990		20.6% (1.2%)	
Percent Population Under 65 population Employed in Large Firms in 1990		7.5% (3.7%)	
N		583	

Table 2

Estimates of the Impact of HMO Penetration on Change in Hospital HHI
(standard errors in parentheses)

Variable	OLS		IV	
	(1)	(2)	(3)	(4)
ΔHMO Penetration	-.087*	-.088*	-1.26*	-.90*
	(.044)	(.044)	(.59)	(.45)
HMO Penetration in 1990	.023	---	.79	---
	(.062)		(.55)	
ΔPercent Elderly	.29	.42	.70	.13
	(.64)	(.63)	(.95)	(.76)
Δ Log Population	.051	.053	.18	.12
	(.049)	(.049)	(.088)	(.070)
ΔMedian Log Per Capita Income	.025	.021	-.21	-.11
	(.083)	(.085)	(.16)	(.13)
ΔPercent Population in Poverty	.0032	.0029	.012	.0080
	(.0038)	(.0038)	(.0076)	(.0056)
ΔUnemployment Rate	.00091	.00094	-.0082	-.0035
	(.0039)	(.0039)	(.0079)	(.0055)
Percent Elderly in 1990	.11	.10	.14	.0015
	(.17)	(.17)	(.26)	(.21)
Log Population in 1990	.17	.17	.16	.31
	(.080)	(.080)	(.082)	(.12)
(Log Population in 1990)^2	-.0071	-.0070	-.0051	-.012
	(.0031)	(.0031)	(.0033)	(.0049)
Median Log Per Capita Income in 1990	.00086	.00063	.030	.054
	(.044)	(.044)	(.081)	(.064)
Percent Population in Poverty in 1990	.00083	.00081	.0023	.000049
	(.0014)	(.0014)	(.0023)	(.0017)
Unemployment Rate in 1990	-.0054	-.0054	-.014	-.0088
	(.0033)	(.0033)	(.0068)	(.0047)
Beds per 1,000 population in 1990	.0070	.0070	.0077	.0042
	(.0041)	(.0041)	(.0063)	(.0051)
Percent FP Hospital in 1990	.035	.035	.0082	.012
	(.025)	(.025)	(.0041)	(.034)
Hospital CON Regulation in 1990	.0038	.0036	.042	.015
	(.0097)	(.0096)	(.025)	(.014)
Log Land Area in 1990 (square miles)	-.010	-.010	-.012	-.017*
	(.0054)	(.0054)	(.0096)	(.0075)
Hospital HHI in 1990	-.028	-.028	-.053*	-.042
	(.016)	(.016)	(.027)	(.022)
N	583	583	583	583
R^2	.048	.048	---	---
J statistic (p-value)	---	---	.56	.43

*Significant at 5% level of confidence
**Significant at 1% level of confidence.

Table 3

Estimates of the Impact of HMO Penetration on Change in Hospital Merger HHI
(standard errors in parentheses)

Variable	ΔMHHI Tobit Estimates		ΔMHHI IV Tobit Estimates	
	(1)	(2)	(3)	(4)
ΔHMO Penetration	-.074	-.082	-1.59*	-1.23
	(.083)	(.083)	(.79)	(.65)
HMO Penetration in 1990	.18	---	1.04	---
	(.12)		(.79)	
ΔPercent Elderly	-.65	-.81	-.66	-1.34
	(1.14)	(1.14)	(1.57)	(1.33)
Δ Log Population	-.089	-.097	-.065	.0023
	(.098)	(.098)	(.15)	(.12)
ΔMedian Log Per Capita Income	.060	.067	-.21	-.11
	(.14)	(.14)	(.23)	(.19)
ΔPercent Population in Poverty	.014*	.014*	.027*	.022*
	(.0070)	(.0070)	(.011)	(.0091)
ΔUnemployment Rate	-.0051	-.0046	-.016	-.010
	(.0075)	(.0075)	(.011)	(.0089)
Percent Elderly in 1990	-.0060	-.049	.0083	-.20
	(.32)	(.32)	(.44)	(.37)
Log Population in 1990	.060**	.059**	.86**	.73**
	(.016)	(.016)	(.24)	(.20)
(Log Population in 1990)^2	-.022**	-.021**	-.032**	-.032**
	(.0062)	(.0063)	(.010)	(.0098)
Median Log Per Capita Income in 1990	-.019	-.019	.056	.078
	(.084)	(.084)	(.12)	(.11)
Percent Population in Poverty in 1990	.0013	.00089	.0037	.014
	(.0026)	(.0026)	(.0039)	(.021)
Unemployment Rate in 1990	-.0065	-.0058	-.016	-.010
	(.0060)	(.0059)	(.010)	(.0070)
Beds per 1,000 population in 1990	.027**	.026**	.026**	.022**
	(.0072)	(.0071)	(.010)	(.0084)
Percent FP Hospital in 1990	.086	.0084	.050	.052
	(.047)	(.047)	(.064)	(.057)
Hospital CON Regulation in 1990	.0012	-.0042	-.074	.011
	(.017)	(.016)	(.14)	(.020)
Log Land Area in 1990 (square miles)	-.0060	-.0078	-.010	-.018
	(.0096)	(.0096)	(.015)	(.012)
Hospital HHI in 1990	-.024	-.022	-.063	-.044
	(.025)	(.025)	(.038)	(.031)
N	583	583	583	583
Log Likelihood	-41.30	-42.31	---	---

*Significant at 5% level of confidence
**Significant at 1% level of confidence.

Table 4

Estimates of the Impact of HMO Penetration on the Likelihood of a Horizontal Consolidation
(standard errors in parentheses)

Variable	Horizontal Consolidation Indicator Probit Estimates		Horizontal Consolidation Indicator IV Probit Estimates	
	(1)	(2)	(3)	(4)
ΔHMO Penetration	-.45	-.47	-3.26	-1.30
	(.63)	(.63)	(4.49)	(4.47)
HMO Penetration in 1990	.55	---	5.68	---
	(1.00)		(5.03)	
ΔPercent Elderly	-10.36	-10.86	-7.28	-11.15
	(8.27)	(8.23)	(9.87)	(8.93)
Δ Log Population	-.97	-1.00	-58	-92
	(.72)	(.73)	(.95)	(.87)
ΔMedian Log Per Capita Income	-.032	-.0034	-.71	-.12
	(1.21)	(1.13)	(1.46)	(1.32)
ΔPercent Population in Poverty	.11*	.11*	.14*	.11*
	(.051)	(.051)	(.070)	(.062)
ΔUnemployment Rate	-.082	-.080	-.11	-.083
	(.055)	(.055)	(.070)	(.063)
Percent Elderly in 1990	-.90	-1.06	.039	-1.15
	(2.50)	(2.49)	(2.90)	(2.60)
Log Population in 1990	3.73**	3.64**	4.51	3.77
	(1.50)	(1.51)	(1.88)	(1.72)
(Log Population in 1990)^2	-.12*	-.12*	-.16	-.12
	(.060)	(.061)	(.076)	(.069)
Median Log Per Capita Income in 1990	-.57	-.55	-.64	-.48
	(.66)	(.66)	(.82)	(.78)
Percent Population in Poverty in 1990	.014	-.0020	.012	-.0023
	(.021)	(.020)	(.025)	(.021)
Unemployment Rate in 1990	-.045	-.028	-.062	-.030
	(.046)	(.043)	(.058)	(.050)
Beds per 1,000 population in 1990	.25**	.25**	.27**	.25**
	(.057)	(.057)	(.063)	(.058)
Percent FP Hospital in 1990	.84*	.95**	.92*	.93*
	(.38)	(.35)	(.45)	(.40)
Hospital CON Regulation in 1990	-.057	-.072	.12	-.060
	(.13)	(.13)	(.22)	(.14)
Log Land Area in 1990 (square miles)	.038	.032	.070	.027
	(.078)	(.078)	(.099)	(.087)
Hospital HHI in 1990	-.45*	-.44*	-.57*	-.46*
	(.19)	(.19)	(.24)	(.21)
N	583	583	583	583
Log Likelihood	-296.62	-296.75	---	---

*Significant at 5% level of confidence
**Significant at 1% level of confidence

Figure 1

Mean, Population Weighted Hospital Concentration and HMO Penetration, 1990-2000

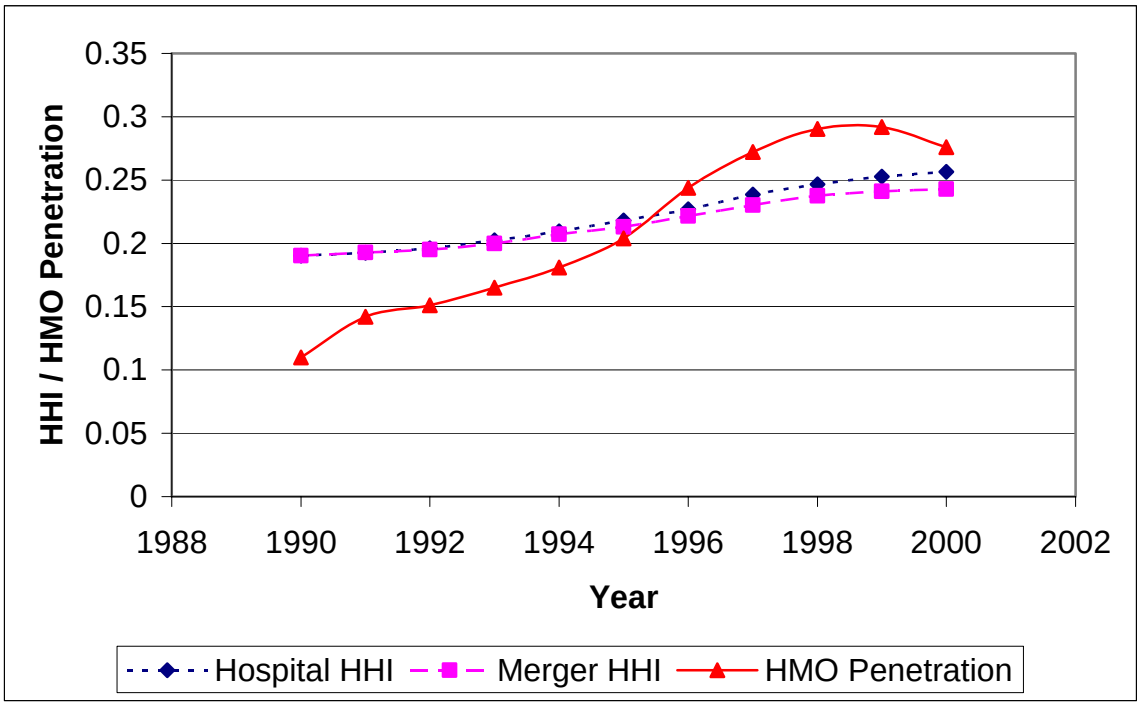


Figure 2a

Histogram of ΔHHI

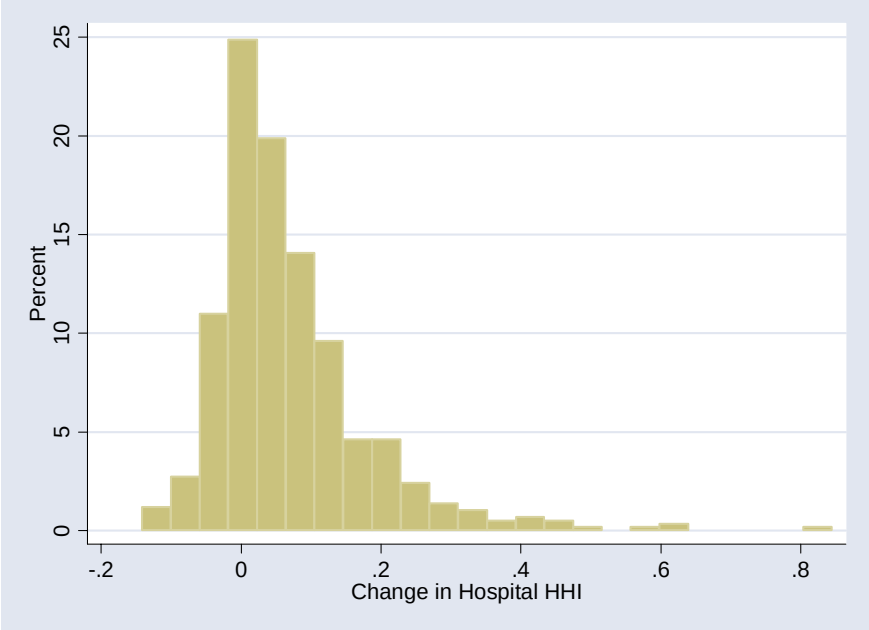


Figure 2b

Histogram of ΔMHHI

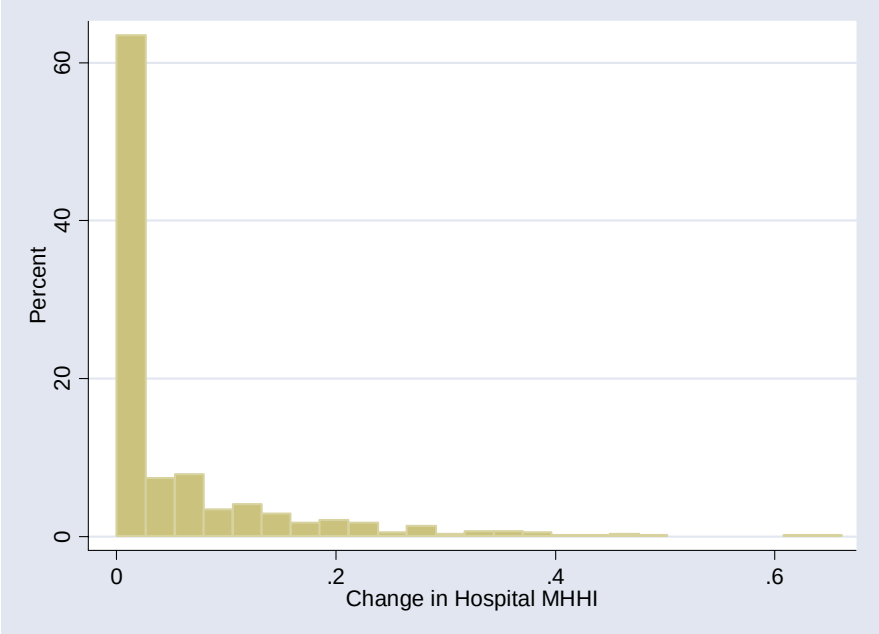
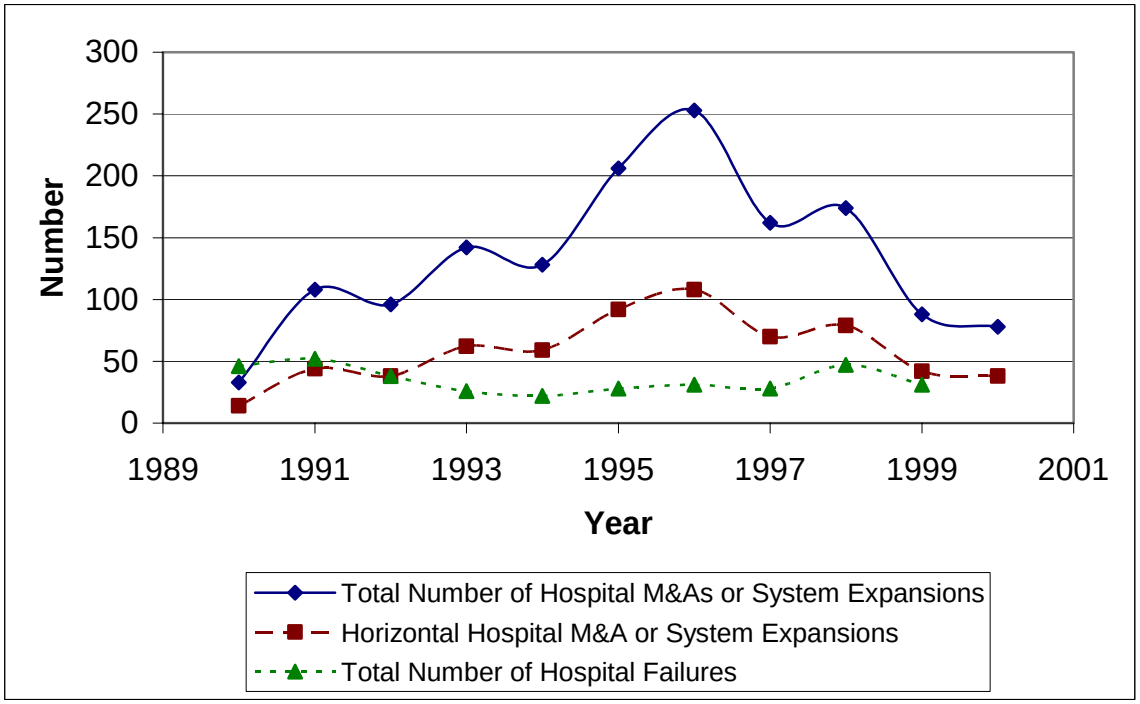


Figure 3

Number of Horizontal Hospital Mergers and Total Hospital Mergers



Appendix
Table A1
First Stage Regression Results
(standard errors in parentheses)

Variable	ΔHMO Penetration	HMO Penetration in 1990
Potential HMO Entry	.00043	.00074*
	(.00045)	(.00033)
Primary Care Physicians Per Capita	.10*	.055
	(.046)	(.031)
Percent of employers with 1-4 employees	-.018	-.28*
	(.21)	(.13)
Percent of employers with 5-9 employees	-.92	-.12
	(.49)	(.28)
Percent of employees in firms with over 250 employees	-.13	-.39**
	(.13)	(.084)
ΔPercent Elderly	-.31	-1.04**
	(.54)	(.31)
Δ Log Population	.076	-.011
	(.048)	(.029)
ΔMedian Log Per Capita Income	-.16*	.075
	(.065)	(.046)
ΔPercent Population in Poverty	.0062	-.0021
	(.0032)	(.0018)
ΔUnemployment Rate	-.0057	.0030
	(.0034)	(.0018)
Percent Elderly in 1990	-.21	-.29**
	(.15)	(.092)
Log Population in 1990	.16*	-.12*
	(.082)	(.048)
(Log Population in 1990)^2	-.0051	.0061**
	(.0032)	(.0020)
Median Log Per Capita Income in 1990	.056	.061
	(.042)	(.035)
Percent Population in Poverty in 1990	-.00021	-.0026**
	(.0012)	(.00080)
Unemployment Rate in 1990	-.0046	.0036*
	(.0027)	(.0018)
Beds per 1,000 population in 1990	-.0031	-.0036
	(.0030)	(.0018)
Percent FP Hospital in 1990	-.024	-.010
	(.025)	(.012)
Hospital CON Regulation in 1990	.019*	-.018**
	(.0088)	(.0054)
Log Land Area in 1990 (square miles)	-.012*	-.012**
	(.0053)	(.0039)
N	583	583
R^2	.31	.45
F-test of Excluding instruments	3.36	6.29
Partial R^2	.024	.054

Table A2

Replication of Dranove, Simon and White (2002)—Estimating the impact of HMO Penetration on Hospital Concentration 1985-1994.

	Dependent variable is ΔHospital HHI					
	IV Estimation, MSA Population > 800,000		OLS Estimation, MSA Population > 800,000		IV Estimation, MSA Population > 500,000	
	(1)	(5)	(6)	(2)	(3)	(4)
HMO Penetration	.10* (.045)	--	.099* (.033)	--	.060 (.051)	--
ΔHMO Penetration	--	-.071 (.048)	--	.0092 (.050)	--	.079 (.076)
Log Per Capita Income	.15 (.074)	.074 (.079)	.15 (.080)	.11 (.079)	.096 (.062)	.086 (.069)
Log Population	-.027 (.030)	-.045 (.036)	-.027 (.032)	-.035 (.033)	-.011* (.038)	-.0085 (.042)
Percent Elderly	-.030 (.039)	-.059 (.032)	-.030 (.040)	-.051 (.031)	-.037 (.020)	-.039* (.019)
N 1st stage F (p-value)	68 2.51 (.001)	68 1.81 (.067)	68 R^2=.20	68 R^2=.066	98 3.04 (.01)	98 1.58 (.11)

*Significant at the 5% level.