NBER WORKING PAPER SERIES

# HMOS AND FEE-FOR-SERVICE HEALTH CARE EXPENDITURES: EVIDENCE FROM MEDICARE

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Working Paper 5360

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
November 1995

I wish to thank Joel Cantor, Victor Fuchs, Alan Krueger, Jean LeMasurier, Jennifer Messersmith, Uwe Reinhardt, and Dave Skellan for helpful comments and discussions. Dave Skellan and Clifton Maze of the Office of the Actuary at HCFA provided comments and invaluable assistance with the data. This paper is part of NBER's research program in Health Care. Any opinions expressed are those of the author and not those of the National Bureau of Economic Research.

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# HMOS AND FEE-FOR-SERVICE HEALTH CARE EXPENDITURES: EVIDENCE FROM MEDICARE

## ABSTRACT

Increasing levels of HMO activity may influence health expenditures in other sectors of the market. Medicare provides FFS coverage to the majority of its beneficiaries and may thus provide a way of examining these so-called spillover effects. This paper examines 1986-1990 Medicare FFS expenditures at the county- and MSA-levels, coupled with county- and MSA-level measures of HMO market share. Fixed-effects and IV estimates of the relationship between market share and expenditures are presented. All of the models imply that FFS expenditures are concave in market share and that expenditures are decreasing in market share for market shares above about 18%. Many of the estimates suggest that expenditures become decreasing in market share at much lower levels (between 0% and 10%). Fixed-effects estimates imply that increases in market share from 20 to 30 percent would be associated with expenditure reductions of 3.4% - 6.6% in Part A expenditures and 2.5% - 5.6% in Part B expenditures. IV estimates imply larger responses. The results are consistent with the hypothesis that managed care can affect non-managed-care expenditures.

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## I. Introduction

Rising levels of managed care activity in the United States are likely to influence health care provision and costs in a number of ways. One important possibility is that managed care affects utilization patterns and expenditures in the non-managed-care sector of the market. Understanding whether such so-called "spillover" effects occur, and, if so, how significant they are, is important for understanding the implications of the ongoing shift toward managed care and for evaluating policies that would further encourage growth in managed care. This paper examines the relationship between HMO market share and expenditures for the care of Medicare beneficiaries who have traditional, fee-for-service (FFS), Medicare coverage. Understanding the effects of HMOs on Medicare expenditures may contribute to current policy discussions about Medicare financing. Moreover, since Medicare FFS expenditures are likely to be correlated with non-Medicare FFS expenditures, the ability of managed care to influence Medicare FFS expenditures may be evidence that managed care can also have broader effects.

Managed care may influence non-managed-care expenditures and utilization in a number of ways. For example, managed care organizations may compete with non-managed-care providers or insurers, influence levels of technology availability, influence the organizational structure of the overall health care market, or contribute to the spread of conservative practice patterns among non-managed-care providers. Supporters of managed care argue that, through these and other mechanisms, managed care can lower FFS health care costs. While several studies have examined the effects of managed care on the FFS sector, the evidence remains inconclusive--some studies find evidence for the existence of spillovers and other do not. However, most studies of spillovers have been confined to examining hospital expenditures and utilization, or have examined data from small sets of markets.

Medicare offers an excellent opportunity to study spillover effects since nationwide data on both ambulatory care and hospital FFS expenditures are available for multiple years. Medicare is the U.S. government health insurance program for the elderly and disabled[1] and is overseen by the Health Care Financing Administration (HCFA). Broadly, Medicare provides two types of coverage to its beneficiaries: traditional FFS coverage and HMO coverage. Most beneficiaries have traditional coverage, which is provided in two parts. Part A covers inpatient hospital care, limited nursing home services, and some home health services, subject to deductibles and coinsurance. Part B covers payments for physicians, services and supplies ordered by physicians, outpatient hospital visits, rural health clinic visits, durable medical equipment (e.g. wheelchairs), and some other services, also subject to a deductible and coinsurance.

This study examines nationwide panel data on county- and MSA-level Medicare FFS expenditures for both Part A and Part B care between 1986 and 1990. These data are matched to measures of Medicare HMO market share at the county and MSA level. Some analyses also use county-level measures of system-wide HMO market share available for 1990. Since OLS estimates of the effect of market share on expenditures may be biased by unobserved or omitted variables that are correlated with both market share and expenditures, and by simultaneous determination of the level of expenditures and HMO activity, fixed-effects models are initially used.

Results from quadratic fixed-effects models that use Medicare HMO market share as the key independent variable suggest that both hospital and ambulatory care expenditures are concave in

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[1]Medicare also provides coverage for individuals with End Stage Renal Disease (ESRD) and some poor populations. Since ESRD patients make up less than 1% of Medicare enrollees and tend to have distinct health care needs, they are not included in this analysis. In some states, the poor are included under OBRA 1989, in which Congress authorized state governments to pay the Medicare hospital insurance premium on behalf of Medicaid eligible individuals who are not otherwise entitled to Medicare. HCFA estimated that about 220,000 individuals would enroll under this provision in 1992 (Palsbo, 1992).

market share, and are decreasing for all positive values of market share. The estimated coefficients indicate that an increase in market share from 10 to 20 percent would be associated with a 4.5% decrease in Part A expenditures and a 4.1% decrease in Part B expenditures. Increasing market share from 20 to 30 percent would be associated with decreases in Part A and Part B expenditures of 6.6% and 5.6%, respectively.

The fixed effects estimates rely on variation in HMO activity within counties or MSAs to identify the effect of HMOs. Since there is also substantial variation across counties and MSAs, IV estimates that exploit this variation are presented. Firm size and the proportion of workers who are considered white collar are used as instruments. As with the fixed-effects models, the results indicate that expenditures are concave functions of market share. Here, however, expenditures reach their peak at about 15% market share. Above these levels, the estimates suggest larger expenditure responses to changes in market share than the fixed effects models. Increasing market share from 20 to 30 percent would be associated with declines of 25%-38% in Part A expenditures and 24%-45% in Part B expenditures. IV estimates obtained using 1990 measures of overall, rather than Medicare, market share produce similar estimates.

The remainder of this paper proceeds as follows. The next section outlines the ways in which HMOs may be expected to influence expenditures and reviews relevant literature. Section 3 describes the data used. Section 4 describes the estimation strategy. Section 5 presents estimation results, and section 6 concludes.

## II. HMOs and Health Care Expenditures

Managed care may influence FFS expenditures in a number of ways. First, in the overall health care system, managed care networks may compete with other health care providers or insurers

for the business of employers or other purchasers of health insurance in their market area. If non-managed-care providers or insurers earn excess profits in the absence of managed care, increasing the level of competition could enhance the level of market discipline and lead to lower prices and, thus, expenditures. Competition could also force FFS providers to change the ways in which they provide care, or prompt insurers to expand utilization review and other oversight efforts, both of which could lead to expenditure reductions. It should be noted that competition could also be associated with increasing expenditures. For example, competition from managed care plans may prompt other providers to compete on the basis of quality or technology. Alternately, if managed care pulls patients away from FFS physicians, they may respond by inducing demand from or increasing amounts charged to non-managed-care patients.[2]

Within Medicare, the ability of competition from managed care to directly influence expenditures is limited since Medicare does not compete for the business of the elderly or disabled. Nor does Medicare operate under the same incentives that face for-profit insurance companies and health care providers. In addition, the rules that govern reimbursement limit the amount of price variation that can take place . HCFA pays providers for Part A services using the Prospective Payment System (PPS), which HCFA began phasing in during 1983. Under the PPS, hospitals are paid a fixed rate, set by HCFA, for the treatment of most patients. The rates are based on the diagnosis of the patient and the treatments received, adjusted for geographic cost differences. Some price variation may be evident in Part B expenditures, however. During the time period under study here, HCFA reimbursed physicians for Part B services on an FFS basis, determining the amount a

[2]The willingness and ability of physicians to do this is a subject of debate. Some evidence, e.g. Mitchell et al., 1989, Cromwell and Mitchell, 1986, indicates that physicians can induce demand and may do so in response to reductions in demand or prices.

billing physician would be reimbursed for each service based on the usual, customary, and reasonable charges for similar services performed by the billing physician and other physicians in his or her area.[3]

Some variation in Medicare expenditures may be indirectly induced by competition from HMOs in the non-Medicare market. Most importantly, system-wide changes in utilization patterns may be reflected in Medicare expenditures. Most health care providers and most hospitals care for both Medicare and non-Medicare patients. If FFS providers are influenced by competition to adopt more conservative practice styles for their non-Medicare patients, Medicare patients may also be treated more conservatively. Since Part B reimbursement is based on the number of procedures performed, variation in utilization will be reflected in Part B expenditures. Part A expenditures will also reflect variation in the number of hospitalizations and, to some degree, variation in the procedures performed once patients are in the hospital.[4] Some indirect price variation may also occur. Changes in hospital costs that result from competition with HMOs may be incorporated into the PPS payment rates. The usual, customary, and reasonable charges paid for Part B services may also reflect the overall price structure of physician's practices, which may be influenced by the level of HMO competition.[5]

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[3]Beginning in the late 1980s, HCFA began to introduce the Medicare Fee Schedule, under which reimbursement amounts reflect the resources used for each procedure. To begin to move toward the Fee Schedule reimbursement amounts, the Omnibus Budget Reconciliation Acts of 1987 and 1989 reduced fees for several procedures thought to be overpriced, effective in the following year. HCFA began phasing in the entire fee schedule in 1991.

[4]McClellan (1995) shows that substantial portions of the variance in hospital reimbursements under the PPS can be explained by variation in procedure codes and outlier payments that may reflect variation in utilization patterns.

[5]This effect could go either way. Increases in competition from HMOs could lead FFS physicians to reduce costs or take other steps to reduce prices, which may lead to reductions in Medicare prices. Alternately, faced with declining prices in the non-Medicare portion of their practice, they may increase

In addition to competition, managed care may also influence expenditures in the FFS sector by influencing the environment in which FFS providers practice. For example, managed care may change the incentives associated with the purchase of high-cost medical technologies, leading, perhaps, areas with high levels of managed care to have low levels of technology availability, and vice versa. If the practice patterns of FFS providers are influenced by the system-wide level of technology availability, then FFS utilization and expenditures could be affected. Similar effects could occur if HMOs change the structure of the hospital market (e.g., Chernew, 1995), the distribution of generalist and specialist physicians, or other market characteristics. Effects of this type could easily affect both Medicare and non-Medicare expenditures.

Managed care could also affect expenditures if, as some models of physician learning (e.g. Phelps, 1992) suggest, physicians tend to adopt the practice patterns of other physicians around them. If this occurs, then increases in the number of managed care physicians practicing in a given area may result in faster promulgation of conservative practice techniques. A related possibility is that IPA-affiliated FFS physicians who retain some FFS patients may adopt more conservative practice styles throughout their practices, reducing expenditures for their FFS patients.

As with other empirical examinations of spillovers, the purpose of this paper is to examine the net effect of managed care. Some of the mechanisms described above may associate increases in managed care with increases in expenditures and some may do the opposite. Since aggregate data on FFS expenditures and HMO activity are used, it will not be possible to determine the individual contributions of each mechanism.

Evidence on the extent to which managed care is able to affect utilization and expenditures

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prices charged to Medicare patients.

in the non-Medicare market is not conclusive. A number of studies suggest that managed care lowers FFS expenditures (e.g. Goldberg and Greenberg, 1979; Frank and Welch, 1985; Dowd, 1986; McLaughlin, 1987; Noether, 1988; Robinson, 1991; Chernew, 1995). On the other hand, several studies failed to find evidence of spillover effects (e.g. Feldman et al., 1986; Luft et al., 1986; McLaughlin, 1988; Rossiter, 1989). However, data availability has forced most of these studies to focus on narrow sets of markets, which raises questions of generalizability, or on aggregate measures of hospital expenditures, which include expenditures for both HMO and FFS patients and may not reflect ambulatory care expenditures.

Two studies have examined Medicare data for evidence of spillovers, but neither fully exploited its potential. Welch (1994) examined the effect of HMO market share on overall Medicare expenditures per beneficiary between 1984 and 1987 in 295 Metropolitan Statistical Areas (MSAs) using a partial adjustment model. He found a weakly significant negative relationship between market share and expenditures. There are a number of methodological issues that limit the strength of this finding. Most importantly, the study focused on aggregate expenditures, which include both FFS and HMO expenditures, making it difficult to identify spillovers with certainty. The model also assumed a linear relationship between market share and expenditures, and did not account for the possibility of bias induced by simultaneous determination of expenditures and HMO market share. Finally, it was confined to urban areas. Although urban counties tend to have the most significant managed care presence, estimates of the effects of managed care may be improved by utilizing the experience of additional non-urban counties.

Clement et al. (1992) constructed and analyzed data on the expenditures and utilization of a sample of Medicare FFS beneficiaries between 1985 and 1988 in 108 metropolitan areas. Using a linear fixed-effects model, they estimated that increases of 10 percentage points in HMO market

share are associated with 5% decreases in Medicare FFS expenditures, although the results were sensitive to specification. In the end, they rejected this result as implausibly large. This study was also confined to a small set of urban areas.

In this paper, I explore the relationship between HMO market share and expenditures made for the care of Medicare beneficiaries who have traditional FFS coverage. Nationwide, county-level data on only FFS expenditures are examined,[6] flexible functional forms are used, and explicit attempts are made to account for simultaneity. Compared to the two previous studies of Medicare expenditures, the focus of this paper is also broader. There, the focus was on spillovers that may occur entirely within Medicare. For example, increases in the number of HMOs participating in Medicare may draw patients away from Medicare FFS physicians who may, in response, attempt to become more cost-effective. However, as Clement et al. (1992) themselves note, the potential for internal factors to have a strong effect on Medicare expenditures is limited since incentives facing Medicare FFS physicians are weak. I am concerned here with the effect of system-wide HMO activity on expenditures. In this broader context, it is more plausible that HMOs may influence FFS expenditures. A practical consequence of this shift in focus is that, while Welch (1994) and Clement et al. (1992) used only measures of Medicare risk HMO market share, I use measures of market share that capture HMO activity more broadly.

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[6]This data is expected to be an improvement over the data used by Welch (1994) and Clement et al. (1992). Welch used data on aggregate expenditures, while this data includes only data on FFS expenditures. The data used by Clement et al. were based on the records from samples of patients drawn from each county. The data used here are based on the HCFA database, which contains data on the expenditures of all Medicare beneficiaries in each county.

## III. Data

Medicare Expenditures

Data on Medicare expenditures and overall enrollment by county for 1986-1990 were obtained from HCFA, Office of the Actuary. Initially, data on all 3,141 counties in the United States (including Puerto Rico) were obtained. Counties in Puerto Rico were excluded since not all of the variables used in the analyses were available for them. Counties with fewer than 50 beneficiaries in any of the years were also excluded.[7] The analysis sample included 3,073 counties per year (N=15,365).

The expenditure data include only expenditures made on behalf of FFS beneficiaries--payments to HMOs and other providers for the care of HMO-enrolled patients are not included.[8] In addition, expenditures that are not covered by Medicare, such as copayments, deductibles, payments made for non-Medicare-covered services, and services covered by Medigap insurance,[9] are not

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[7]In addition, to facilitate the addition of demographic and other data, several adjustments were made for compatibility with the Area Resource File county definitions: most of the Virginia independent cities were merged with their surrounding counties (cities remaining separate are: Alexandria, Chesapeake, Hampton, Newport News, Norfolk/Portsmouth, and Virginia Beach); Cibola, NM and Valencia, NM were merged; LaPaz, AZ and Yuma, AZ were merged, and counties in Alaska were aggregated to form a single unit.

[8]In some cases, it is difficult to determine whether payments made are for the care of FFS or HMO patients. For example, payments for hospitalization of beneficiaries enrolled in some HMOs may be sent directly to hospitals and would then be counted as FFS expenditures. Since this occurs in only a small number of cases, it is not expected to affect the findings. To further ensure that this did not affect the conclusions, I re-estimated the models shown below using data from 233 MSAs collected by Mathematica Policy Research as part of the evaluation of the risk contracting program, which are not likely to contain such errors, and the findings were consistent with those reported.

[9]In addition to Medicare coverage, most elderly patients also have some form of supplementary "Medigap" insurance. In 1991, almost 75% of the elderly had such coverage (Chulis et al., 1993). These policies typically provide benefits not covered by Medicare such as prescription drugs and stop-loss limits on large out-of-pocket expenditures.

included. Nonetheless, the included expenditures are expected to account for the majority of spending by Medicare FFS beneficiaries.

To construct the county-level measures of spending, HCFA apportions expenditures for each beneficiary to his or her county of residence,[10] regardless of where the expenditures were incurred. To construct yearly totals, HCFA assigns expenditures to the calendar year in which the related claim was processed by the HCFA intermediary or carrier.[11]

The PPS for hospital reimbursement was phased-in between 1983 and 1990. During the transition, hospitals were paid varying blends of institutional, regional, and national rates. To ensure that this did not affect the findings, the data for each year have been adjusted to reflect what payments in that year would have been under the fully-phased in 1990 PPS (see Palsbo, 1991, for further detail on the adjustment process).

Mean nominal Part A and Part B expenditures per beneficiary are presented in Table 1. In 1986, mean Part A and Part B expenditures per beneficiary were $1,571 and $831, respectively. By 1990, these amounts had risen to $1,920 for Part A and $1,233 for Part B, increases of 22% and 48%, respectively. In contrast, the CPI rose by 19% over this time period.

Market Share Measures

The aim of this work is to examine the relationship between managed care and FFS expenditures. Although the arguments advanced above apply to managed care broadly, data

[10]County of residence is determined using the zip code on each beneficiary's social security check as of July 1 each year.

[11]There are often lags in processing claims and, therefore, some claims filed for services rendered near the end of a calendar year may be assigned to the following year. Since lags are probably similar from year to year, they are not expected to bias the findings.

availability constrains me to examine only HMO activity. If levels of HMO activity are correlated with levels of non-HMO managed care activity, then HMO measures will adequately capture the effects of managed care. This paper focuses on HMO market share as a measure of HMO activity. Measures of other aspects of HMO activity are not available at sufficiently geographically-detailed levels for the complete set of years examined here.

Two market shares measures are used here. The first is Medicare HMO market share. As an alternative to traditional Medicare FFS coverage, HCFA has taken steps to make HMO coverage available to most Medicare beneficiaries. To do this, HCFA contracts with willing HMOs to provide care for beneficiaries.[12] Under federal law, HMOs contracting with HCFA must draw least 50% of their membership from the non-Medicare population. Medicare beneficiaries who live in areas served by contracting HMOs may elect HMO coverage. Enrollees typically pay a small monthly premium, and receive their care through the normal HMO channels. Because contracting organizations must serve the overall market, and because enrollment patterns among the elderly are likely to be correlated, albeit imperfectly, with enrollment patterns in the non-Medicare market, Medicare market share is likely to be correlated with overall market share.

County-level data on the number of Medicare beneficiaries enrolled in HMOs for 1985-1990 were obtained from HCFA, Office of the Actuary.[13] Market shares are simply the ratio of the

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[12]Almost all contracting is done using either "risk contracts," under which organizations are paid on a capitated basis, or "cost contracts" under which organizations are retrospectively reimbursed for the cost of providing care to their Medicare enrollees. Contracting organizations must accept all beneficiaries who desire to enroll. The organizations are required to provide all Medicare covered services (an exception is the "Health Care Prepayment Plan" contract, under which enrollees receive ambulatory services through the HMO, but receive hospital services under the traditional Part A program), and often provide some supplementary services as well.

[13]To construct as broad a measure of HMO activity as possible, enrollment in Medicare HMOs operating under both risk and cost contracts, and enrollment in HMOs operating as Health Care Prepayment Plans, are included here.

number of HMO enrollees to the number of beneficiaries in each county.[14] The top portion of Table 2 presents summary statistics describing the distribution of Medicare HMO market shares by year. Between 1985 and 1990, mean Medicare HMO market share grew from 4.4% to 6.4%, although it declined somewhat from 1987 to 1989. The distribution of county Medicare market shares is highly skewed, as evidenced by the fact that the median market shares are all well below the means. Market shares in the sample range from 0 to 64 percent. Table 3 summarizes the observed distributions of market share levels in each of the years examined and 1985-1990 changes. Unweighted distributions and distributions weighted by county Medicare enrollment are shown.

Since Medicare market shares are available at the county level, they permit the identification of variation across individual market areas. They are also available for the complete set of years being examined. Their principal disadvantage is the likelihood that they measure overall HMO market share with error. Since there are few alternate measures of market share for comparison, it is difficult to assess the extent of the error. To obtain a rough idea, I compared estimates of state-level Medicare market share, constructed by aggregating the Medicare data described above to the state level, to state-level measures of overall market share constructed using Interstudy data (Interstudy, 1986-1991).[15] Baker (1995) notes that state-level estimates based on the Interstudy data appear to measure overall market share quite accurately. The correlation between the two sets of state-level

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[14]Because of complexities in the ways in which HCFA contracts with HMOs, some beneficiaries are counted as members of HMOs for their Part B care, but not for their Part A care. Part A market shares are used here. Part A and Part B market shares are highly correlated. For example, the correlation between 1990 Part A and Part B market share is 0.999.

[15]Interstudy provides data on the total enrollment and market share of all HMOs operating in the U.S. To construct the state estimates, I apportioned the enrollment of each HMO to the state in which its headquarters were located. The market share measure used is the ratio of the estimated number of enrollees in each state to state population. Some difficulties were encountered with Washington, DC, which contains the headquarters of several HMOs that serve Maryland and Virginia, and so data from Washington DC are not included in the comparisons presented here.

estimates, weighted by state population, is 0.788. Additional information can be obtained by comparing the 1990 Medicare HMO market shares to estimates of county-level HMO market shares in the overall (Medicare and non-Medicare) market that are available for 1990 (described further below). Although the mean Medicare and overall market shares differ by 8 percentage points (14.6% overall versus 6.4% in Medicare) the correlation coefficient for the two series, weighted by county Medicare enrollment, was 0.640. While these comparisons are limited in strength since the state-level market shares and 1990 overall county market shares to which the Medicare market shares are compared are themselves estimates, the results do provide some reason to believe that Medicare market share is a usable, if noisy, proxy for overall market share.

The second market share measure used is a set of county-level estimates of overall (Medicare and non-Medicare) HMO market share in 1990. These estimates were constructed using data from the Group Health Association of America, which surveys all HMOs in the U.S. each year (GHAA, 1991). Among the data elements obtained are total enrollment, a list of the counties served by the HMO, and the location of the HMO headquarters. To construct estimates of county-level enrollment, the enrollment of each HMO was distributed among the counties in its service area based on county population and the distance from HMO headquarters. County-level enrollment estimates were computed by summing over all HMOs serving the county. County market shares are the ratio of county HMO enrollment to county population. Additional details of the process by which the data were constructed can be found in Appendix A; full details are found in Baker (1995). Overall county-level market shares range from 0 to 57%; summary statistics are shown in Table 2.

The overall county-level HMO market share data are expected to accurately represent overall HMO market shares, and are also able to capture small area variation in market share. However, because of limitations on the GHAA data available before 1990, reliable estimates are not available

for earlier years.[16]

Other Data

Additional data used in the analysis came from a number of sources. Many of the covariates were obtained from the 1992 Area Resource File (ARF), which is compiled by the Bureau of Health Professions, Office of Data Analysis and Management. The ARF summarizes county-level information from a variety of sources including the American Medical Association, the American Hospital Association, the Census Bureau, the National Center for Health Statistics, and HCFA. Variables not obtained from the ARF include per capita income for 1990 from Slater and Hall (1993), county population data by sex, race, and age, from the Census Bureau intercensal estimates tape, and firm size data from the County Business Patterns files for 1985-1990.

## IV. Estimation Issues and Strategy

This section briefly develops the framework in which estimation is attempted. Aggregate expenditures in market $i$ are assumed to be the sum of expenditures for each of the $j$, $j=1, \ldots, J$, types of health care services performed in the market:

$$E_i = \sum_j p_{i,j} q_{i,j} \qquad (1)$$

where $E$ denotes expenditures, $p$ denotes price, and $q$ denotes quantity.

[16]As an additional measure of HMO market share I experimented with the use of overall state-level HMO market share measures constructed using Interstudy date. State-level estimates are available for all of the years studied and are likely to accurately reflect state-level market share. However, they cannot capture within-state variation, which is substantial in some cases. The results of specifications that use these variables were similar to those obtained using the two measures described above.

Both price and quantity may be influenced by HMO activity ($H$), as discussed above, and possibly other characteristics as well. Thus, I write:

$$p_{i,j} = f(H_i, X_{i,j}^P), \tag{2}$$

where $H_i$ represents the level of HMO activity and $X_{i,j}^P$ is a vector of variables influencing price. In the context of Medicare, price variation is expected to be limited, but some indirect price effects may be possible. Similarly, the quantity of services provided is expected to be a function of HMO activity and other factors:

$$q_{i,j} = f(H_i, X_{i,j}^Q), \tag{3}$$

where $X_{i,j}^Q$ is a vector of variables influencing quantity. Variables in $X_{i,j}^Q$ may include things like the health status of the population, treatment preferences of providers, preferences of patients, and local market conditions. HMOs may influence quantities through competition, by influencing the availability of health care technologies, through learning effects, and through IPA effects.

Combining (1), (2), and (3) gives

$$E_i = f(H_i, X_{i,1}^P, \ldots, X_{i,J}^P, X_{i,1}^Q, \ldots, X_{i,J}^Q). \tag{4}$$

As a practical matter, market-level data on the potential influences of prices and quantities of individual procedures are not available. Hence I consider models of the form:

$$E_i = f(H_i, X_i^P, X_i^Q). \tag{5}$$

With the procedure subscripts ($j$) dropped, $X_i^P$ and $X_i^Q$ denote vectors of market-level variables that may influence aggregate prices and quantities.

There are a number of issues important for estimating equation (5). First, since variables are defined at the market level, market areas must be identified. As others have noted (e.g. Garnick et al, 1987; Chernew, 1995) selection of a unit of analysis for which data are available and which represents a plausible market for health care services is an important but difficult issue in this context. Because of data availability, most of the analyses presented here are conducted at the county-level. Since counties may be too small in some cases, some models are estimated using MSA-level data.

A second issue is the possibility that there are unobservable variables that are correlated with both market share and expenditures. For example, preferences of patients and providers for conservative care might increase HMO market share and decrease expenditures. Unobserved elements of the health status of the population may also influence both market share and expenditures. Estimates of the effect of HMOs that ignore variables like these will be biased and inconsistent.

A third issue for estimation is the possibility that HMO market share and expenditures are simultaneously determined. Forward-looking HMOs may consider current and expected expenditure levels when deciding whether to enter or expand operations in a market. HMOs that can effectively reduce utilization may be most successful in markets where FFS expenditures are high. A number of studies (e.g. Porell and Wallack, 1990; Welch, 1984; Goldberg and Greenberg, 1981) have concluded that HMO market share is positively related to health care costs and utilization. If increases in FFS expenditures cause increases in HMO market share, then estimates of the effect of HMOs on expenditures that do not account for simultaneity will understate any expenditure-reducing effect of HMOs.

As an initial attempt to deal with omitted variables and simultaneity, models that include

fixed-effects for counties (or MSAs) and years are examined below. If the omitted variables are constant over time within areas, or are constant within years across areas, then the inclusion of fixed effects will remove the bias. The use of fixed-effects models is also expected to control for simultaneity bias to a large degree. By relying on changes over time to identify the effect of HMOs, the bias induced by high expenditure levels causing high HMO market share levels will be removed. However, some bias could remain if changes in market share are prompted by expected future changes in expenditure levels.[17]

The fixed-effects models rely on variation within areas over time to identify they effect of HMO market share. These estimates do not exploit a considerable amount of variation across areas. To attempt to exploit the across variation, a set of models that excludes the county or MSA fixed-effects is also estimated. To control for omitted variables and simultaneity, IV techniques are used. If the instruments are uncorrelated with the omitted variables and uncorrelated with the expenditures, then these estimates will also be unbiased. The specific instruments are discussed below.

A final issue for estimation is the possibility of selection bias. Many studies have found that HMOs and other managed care organizations receive a favorable selection of beneficiaries (e.g. Hellinger 1995, Hill and Brown, 1990, Riley et al., 1989, Luft and Miller, 1988, Brown, 1988, Hellinger 1987, and Wilensky and Rossiter, 1986). Compared to FFS beneficiaries, HMO enrollees tend to have lower levels of health care utilization in prior years and have lower mortality rates, after adjusting for demographic characteristics. Given the evidence, selection is expected to bias

[17]If instrumental variables that could reliably identify changes in market share within counties over time were available, it would be possible to investigate the extent of the remaining simultaneity bias. However, such variables are not available.

estimates of the effect of HMOs on FFS expenditures in a positive direction. Since data on the characteristics of Medicare FFS and HMO beneficiaries are not available at the county level, I am not able to control for the effects of selection bias. Nonetheless, I proceed with the expectation that the expenditure-reducing effects of HMOs that are identified are likely to be conservative estimates of the true effect of HMOs.

## V. Estimation

### Fixed effects models

I begin the estimation by specifying fixed-effects regression models of the form:

$$\log(E_{i,t}) = \beta_0 + \beta_1 H_{i,t-1} + \beta_2 H_{i,t-1}^2 + \beta_3 X_{i,t} + \beta_4 C_i + \beta_5 Y_t + \epsilon_{i,t} \qquad (6)$$

where *log(E)* represents the natural logarithm of FFS expenditures per beneficiary,[18] $H$ represents HMO market share, $X$ is a vector of covariates expected to influence expenditures, $C$ is a set of county-specific intercepts,[19] and $Y$ is a set of year-specific intercepts. The errors, $\epsilon_{i,t}$, are assumed to be independently and identically distributed normal random variables.[20] Subscript $i$ denotes county $i$

[18]Log expenditures are used since both visual inspection and Box-Cox analysis indicated that a logarithmic specification is superior to a linear specification.

[19]Hausman's (1978) technique was used to test the hypothesis that the county-fixed effects do not belong in the model. Under the null hypothesis that there are no county-specific components in the error term, OLS estimation without fixed-effects is consistent and efficient and inclusion of fixed effects generates a consistent but inefficient estimator. If there are county-specific components, fixed effects models remain consistent but OLS models do not. Using these two estimators, the $\chi^2[9]$ test statistics are 2,926.87 for Part A and 25,477.16 for Part B, which reject the null. A similar test can be used for the null hypothesis that there are county-specific components in the error terms, but that they are uncorrelated with market share or other independent variables. In this case, the $\chi^2[9]$ test statistics are 1,818.81 for Part A and 5,733.65 for Part B.

[20]This implies that there is no autocorrelation in $\epsilon$, which can be tested. As pointed out by Nickell (1981) and Solon (1984), the standard autocorrelation test statistics are biased in panel data. Following

and subscript $t$ denotes year $t$.

Since the effect of HMO activity may vary with the level of activity, I experimented with a variety on non-linear models. Initially, linear and quadratic splines with knots at 5, 15 and 25 percent market share were estimated. Since the results of these models were consistent with the results of a simple quadratic model, the latter is used here. Figures 1 and 2 plot results from the linear and quadratic splines, along with the quadratic model.

Lagged market share is used as the key dependent variable since some effects of HMOs may take time to manifest themselves. I also experimented with the use of current and twice-lagged market share. Since market share measures are highly correlated over time,[21] these alternate measures produced estimates very similar to those shown.

Equation (6) was estimated separately for Part A and Part B expenditures since differences between the content and reimbursement of ambulatory and hospital care may cause the effect of HMOs to vary. To correct for possible heteroskedasticity arising from variation in enrollments across counties, weighted least squares regression was used, with county Medicare Part A enrollment as the weight.

Table 4 presents means and standard deviations of the variables used in the models. The control variables include per capita income, the proportions of the population age 65-74, 75-84, and over age 85, and the proportions of the over-65 population that are female, black, and "other race."

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Solon (1984), I tested for autocorrelation using coefficients from autoregressions of errors from a first differenced version of equation (6). For the most part, the results were consistent with the hypothesis of no autocorrelation. However, in one case, the results suggested that $\epsilon$ may follow an AR(1) process with an autocorrelation coefficient of about 0.1. To be sure that this did not bias my results, I estimated a quasi-differenced version of equation (6) and the results were consistent with those reported.

[21]For example, for Medicare market share (weighted by county Medicare Part A enrollment), corr($H_{i,t}$ , $H_{i,t-1}$) = 0.976, and corr($H_{i,t}$ , $H_{i,t-2}$) = 0.946.

The inclusion of county- and year-specific intercepts will capture the effects of additional county- and year-specific unobserved or omitted variables.

Coefficients from estimation of equation (6) for Part A (column 1) and Part B (column 2) expenditures using Medicare HMO market share as the key independent variable are shown in Table 5. The estimated relationship between market share and expenditures is downward sloping for all positive values of market share for both Part A and Part B expenditures. Note that the coefficients have been scaled to represent the effect of a 10 percentage point change in market share (e.g. moving from 5% to 15% market share). Figure 3 plots the relationships.

If expenditures are lognormally distributed, the ratio of expenditures at market share *H2* to expenditures at market share *H1* can be estimated using:

$$\begin{aligned} E_{H2}/E_{H1} &= \exp(\hat{\log} E_{H2} - \hat{\log} E_{H1}) \\ &= \exp(\hat{\beta}_1(H2 - H1) + \hat{\beta}_2(H2^2 - H1^2)). \end{aligned} \tag{7}$$

where $\hat{\log} E_{H1}$ and $\hat{\log} E_{H2}$ are the predicted values of log(expenditures) at market shares *H1* and *H2*, respectively.

To better envision the implied relationship and illustrate the magnitude of the coefficients, the bottom portion of Table 5 shows the market share level at which expenditures reach a maximum, $H_{max}$ (constrained to be non-negative), and estimates of the proportional change in expenditures associated with moving from 10 to 20 and from 20 to 30 percent market share. Estimates are not computed for market shares above 30% since there are relatively few data points in that range.

The estimates presented in Table 5 imply that moving from 10 to 20 percent market share would be associated with a 4.5% decrease in Part A expenditures and a 4.1% decrease in Part B expenditures. Moving from 20 to 30 percent market share would be associated with expenditure

reductions of 6.6% and 5.6% in Part A and Part B expenditures, respectively.

To examine the robustness of these results, I examined two alternate specifications of equation (6). The estimated market share coefficients are shown in Table 6.

First, since many counties had Medicare HMO market shares between 0 and 1%, I re-estimated equation (6) using only data from counties in which market share exceeded 1% in all of the years examined. The results are presented in columns 1 and 3 of Table 6. The relationships implied by these coefficients are similar to those obtained from the models that used all of the counties, although the curves are somewhat less steep at high levels of market share.

The preceding models were estimated using county-level measures of HMO market share. A difficulty with the use of county-level measures is the possibility that counties are too small to adequately represent market areas. Since MSAs may better represent market areas, I re-estimated equation (6) using data on expenditures per beneficiary and market share aggregated to the MSA level.[22] Results are shown in columns 2 and 4 of Table 6. The estimated relationships remain concave, although they are increasing in market share for low positive levels of market share. Part A and Part B expenditures rise with market share until market share reaches 3% and 9%, respectively, and then decline. The slope of the relationship between market share and expenditures also is less steep, particularly at high market share levels.

Instrumental Variables Estimation

To exploit variation across areas in producing the estimates, IV models that do not include the county or MSA fixed effects are estimated. The IV estimates are produced using a two-step

[22]Data on 322 MSAs in each year are used, for a total sample size of 1610. Counties not in MSAs are excluded.

procedure in which equations of the following forms are estimated (Bowden and Turkington, 1990):

$$H_{i,t} = \alpha_0 + \alpha_1 X_{i,t} + \alpha_2 I_{i,t} + \alpha_3 Y_t + \eta_{i,t} \quad (8)$$

$$\log(E_{i,t}) = \delta_0 + \delta_1 \hat{H}_{i,t-1} + \delta_2 \hat{H}^2_{i,t-1} + \delta_3 X_{i,t} + \delta_4 Y_t + \psi_{i,t} \quad (9)$$

In the first equation, market share is regressed on the set of exogenous variables, $X$, and excluded instruments, $I$. the predicted values of $H$ and their squares are then used as instruments for market share in the second step.[23] The set of control variables used here includes the population demographic characteristics used in the fixed-effects models, along with dummy variables to control for urban and rural areas. Since these estimates incorporate cross-area variation in market share and expenditures, I also experimented with the use of additional variables designed to control for cross-sectional variation in the health status of the population. These variables included the proportion of the population with disabilities, the proportion of the population in long-term care facilities, and other measures of income and wealth (e.g. housing characteristics and social security payments). Use of these additional variables did not affect the market share coefficients and they are not included in the specifications presented here.

Two instruments are used in equation (8): firm size, measured as the average number of workers per firm in each county or MSA, and the proportion of workers in each county or MSA who are considered white-collar. Since large firms and white collar firms are more likely to offer HMOs to their employees, areas with large and white collar firms are expected to be favorable climates for

[23]An alternate procedure is to instrument directly for the linear and quadratic terms using 2-stage least-squares. Estimation using this technique produced estimates consistent with those reported.

HMOs.[24]

The instruments are expected to remove bias from omitted variables and simultaneity. In this case, the most obvious omitted variables are patient and provider preferences about care and the health status of the population. Firm size and the proportion of workers who are white collar are unlikely to be correlated with provider preferences, but may be associated with population income and wealth levels. Income and wealth are, in turn, likely to be associated with health status. To attempt to prevent this from biasing the results, the model controls directly for per capita income. Further, as noted above, models that included additional variables to control for other health status characteristics yielded coefficients similar to those reported.

Firm size and the proportion of workers who are white collar are also unlikely to have a strong direct relationship to expenditures. One possible difficulty is that large firms may have significant power in the health care purchasing market and may have taken independent steps to help control costs. This could cause the IV estimates to overstate a cost-reducing effect of HMOs. Re-estimation of IV models using only the percent of workers who are white collar yielded results that were consistent with those reported.

An additional advantage of the IV estimates may be the removal of errors-in-variables bias. If the instruments are uncorrelated with the measurement error in the market share estimates, then IV will purge the model of this bias as well (Hausman et al. 1991).

Equations (8) and (9) were first estimated using Medicare market share and pooled data from 1986-1990.[25] Following this, they were estimated using the 1990 overall market share measures.

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[24]I also experimented with using the presence of state laws favorable to HMOs as instruments. These variables produced results very similar to those reported.

[25]An assumption implicit in pooling the data from all of the years is that the relationship between market share and expenditures did not vary strongly by year. Estimation of equations (8) and (9) for

Table 7 presents the results from estimation of equation (8); column 1 shows the case in which Medicare HMO market share is the dependent variable and column 2 shows the case in which overall HMO market share is the dependent variable. In both equations, the instruments perform as expected--increases in average firm size and increases in the proportion of workers that are white collar are both associated with increases in market share. Following the suggestion of Bound, Jaeger, and Baker (1993), I computed the F-statistic for the joint significance of the excluded instruments. In both cases, the instruments are highly significant.

Table 8 presents the market share coefficients from estimation of the second stage using Medicare HMO market share. (For comparison, OLS regression coefficients are shown in Table 9.) As before, the market share coefficients in all of the specifications are statistically significant and imply that expenditures are concave functions of market share. IV estimates were first computed using the entire set of counties (columns 1 and 4). Part A and Part B expenditures reach maxima at 16% and 18% market share, respectively. Compared to the fixed effects results, the relationships between market share and expenditures are much more sharply sloped. The estimates imply that increasing market share from 10 to 20 percent would be associated with a 6.3% increase in Part A expenditures, and that moving from 20 to 30 percent would be associated with a 38.4% decrease. For Part B expenditures, moving from 10 to 20 and 20 to 30 percent would be associated with a 22.3% increase and a 32.0% decrease, respectively. The implied relationships are plotted in Figure 4. As a check on the validity of the instruments, I estimated equations (8) and (9) using each instrument separately. For both instruments, the results obtained were very similar to those shown.

Equations (8) and (9) were also estimated using only data from counties that had at least 1%

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each year separately yielded coefficients that are broadly consistent with the coefficients presented below.

market share in all of the years examined (columns 2 and 5). For both Part A and Part B, expenditures reach a maximum at similar levels, but the relationships between market share and expenditures are less steep than the estimates obtained using all counties. Increases in market share from 20 to 30 percent would be associated with decreases of 24.8% in Part A expenditure and 24.0% in Part B expenditures. Estimating equations (8) and (9) at the MSA-level (columns 3 and 6) implied a somewhat shallower concave relationships for Part A but not for Part B.[26] Here, increasing market share from 20 to 30 percent would be associated with declines of 25.8% and 44.8% in Part A and Part B expenditures.

Estimates using 1990 county market share estimates

Alternate estimates of HMO market share, that both capture within-state variation and reflect overall market share better than Medicare market share, are available for 1990. As a check on the results presented above, I estimated equations (8) and (9) using these estimates as the market share measure.[27] Results from specifications that include data from all counties are consistent with the findings reported earlier (Table 10). In all cases, expenditures are concave in market share. The estimated curves reach maxima at about 17% market share and imply that moving from 10 to 20 percent market share would be associated with 7.8% and 17.4% increases in Part A and B expenditures. Moving from 20 to 30 percent would be associated with 27.9% and 44.9% declines in Part A and B expenditures. The implied relationships between market share and expenditures are shown in Figure 5. The models were re-estimated using data from counties with at least 1% market

[26]To construct the IV estimates at the MSA level, the county-level instruments were aggregated to the MSA level.

[27]Because only 1990 overall HMO market share estimates are available, current market share, rather than lagged market share, is entered into the equation.

share, and at the MSA level, and the results were similar to those reported.

## VI. Discussion

This paper examined the relationship between HMO market share and Medicare FFS expenditures. Fixed-effects and IV models were estimated using Medicare HMO market share and overall HMO market share as measures of HMO activity. In all cases, the estimated relationship between market share and expenditures is concave. All of the estimates suggest that expenditures are decreasing in market share for market shares above 18%. Many of the estimates suggest that expenditures are decreasing in market share at much lower levels of market share.

Results from fixed-effects models that measured HMO market share using Medicare HMO market share suggest that expenditures reach a maximum at market shares between 0 and 10% and are decreasing thereafter. Estimates imply that increases in market share from 20 to 30 percent would be associated with expenditure reductions of 3.4% - 6.6% in Part A expenditures and 2.5% - 5.6% in Part B expenditures. IV estimates of the effect of market share imply much larger responses. IV estimates that use Medicare market share suggest that expenditures reach maxima between 14% and 18%. Moving from 20 to 30 percent market share would be associated with decreases of 25% - 38% in Part A expenditures and 24% - 45% in Part B expenditures. IV estimates that use county HMO market share measures from 1990 suggest that an increase from 20 to 30 percent would be associated with 27.9% and 44.9% decrease in Part A and B expenditures. Because there are relatively few counties with market share levels over 30%, it is difficult to draw conclusions about the effects of HMOs above this level.

There are at least two possible reasons that the results presented here may understate the true effect of HMOs on expenditures. First, they do not account for selection bias. Data needed to

control for selection bias were not available, but, since the existing literature consistently argues that HMOs receive a favorable selection of patients, it is likely that selection biases these results in a positive direction. Cost shifting may also have influenced the results. If high levels of managed care reduce the revenues of non-Medicare FFS providers or insurers, they may attempt to shift costs onto Medicare. This could result in increased utilization and, during the time period examined here, increases in Part B prices, possibly causing the results to understate the effect that HMOs could have on expenditures if there were no cost shifting.

Although these difficulties, and the wide range in the size of the estimated effects, make it hard to draw strong conclusions about the exact amount of savings that would be associated with changes in market share, the results suggest that HMOs are able to reduce non-HMO health care costs in Medicare, at least above certain levels of market share. This also suggests that HMOs can reduce costs in the non-Medicare sector of the market since Medicare HMO market share and Medicare FFS expenditures are likely to be correlated with their non-Medicare counterparts. In fact, it is possible that the results presented here understate the effect that HMOs can have in the non-Medicare market since the structure of Medicare limits the ability of HMOs to compete with FFS providers, whereas stronger competition is possible in the non-Medicare market. In addition, HCFA's payment mechanisms limit the amount of price variation that HMOs can induce in Medicare Part A prices, but more price variation is possible outside of Medicare.

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Figure 1: Spline and quadratic estimates for Part A expenditures

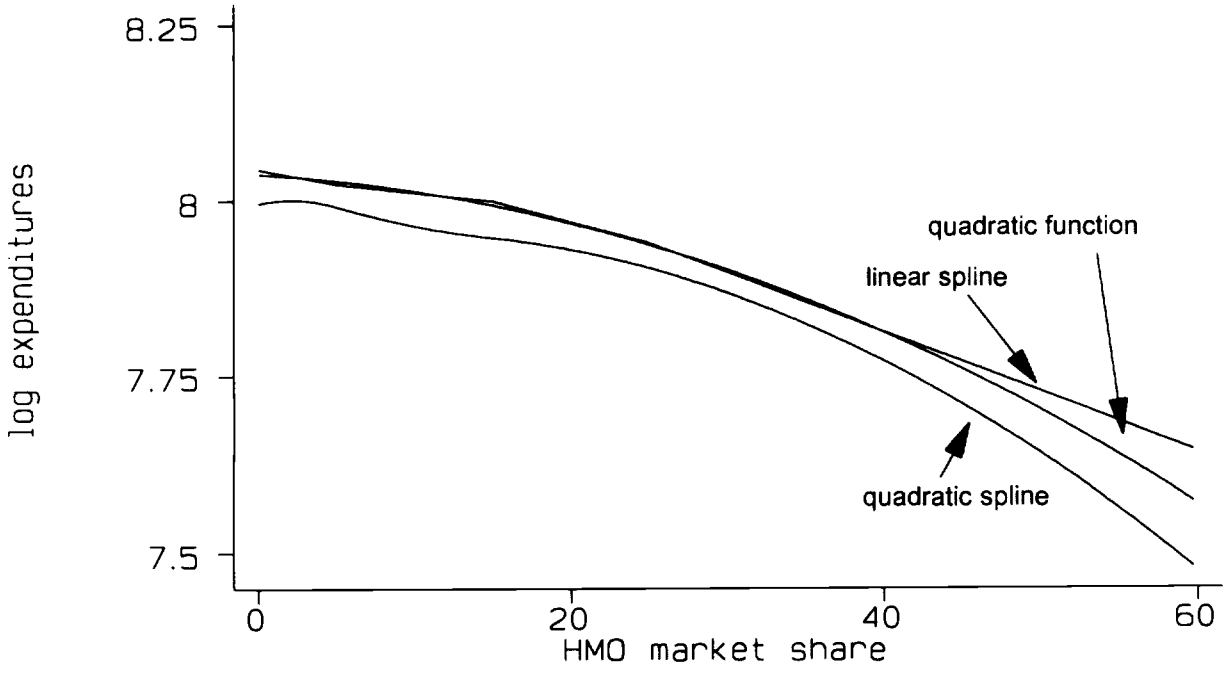


Figure 2: Spline and quadratic estimates for Part B expenditures

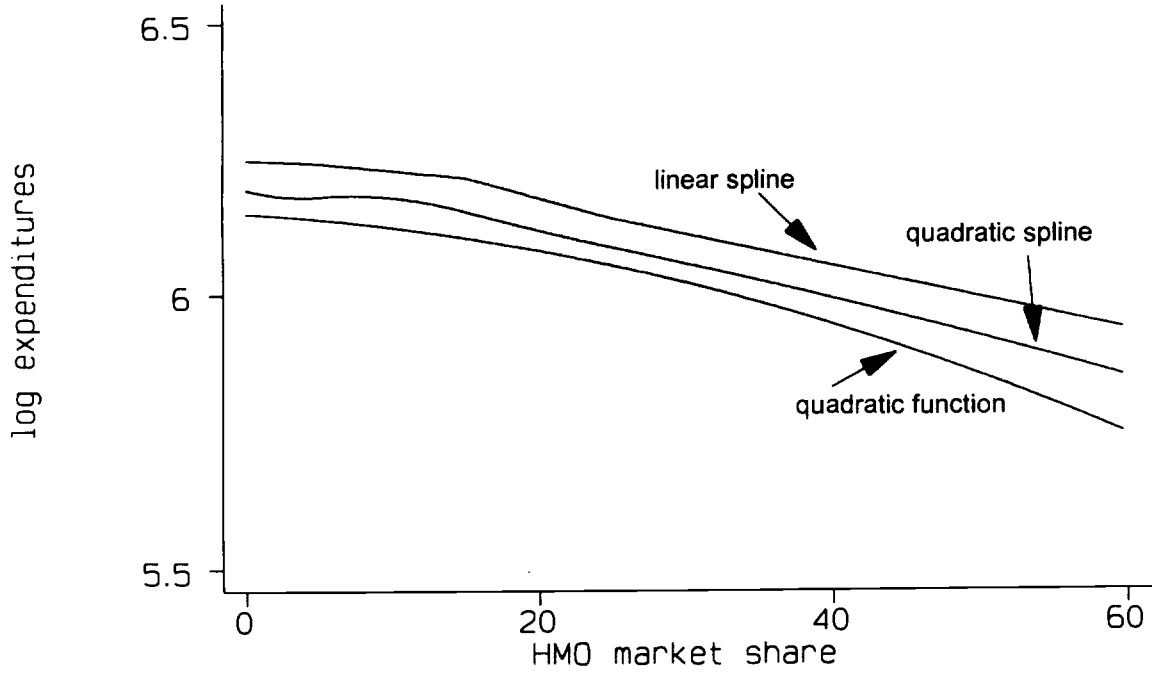

Figure 3: Fixed-effects estimates using Medicare market share

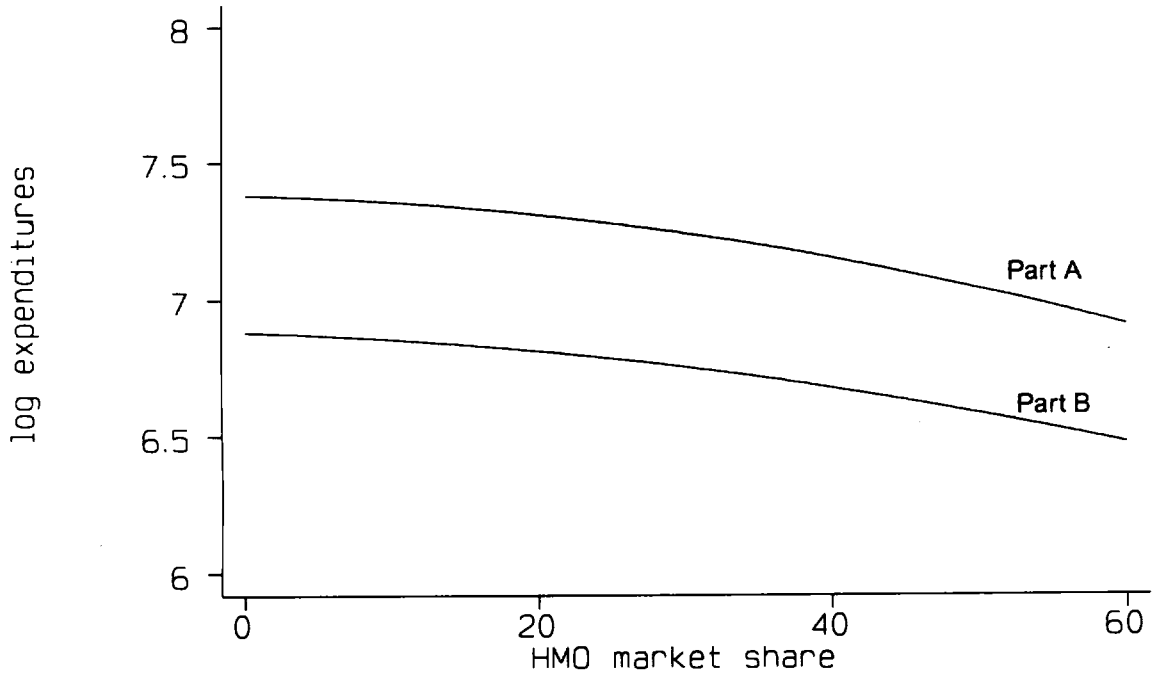


Figure 4: IV estimates using Medicare market share

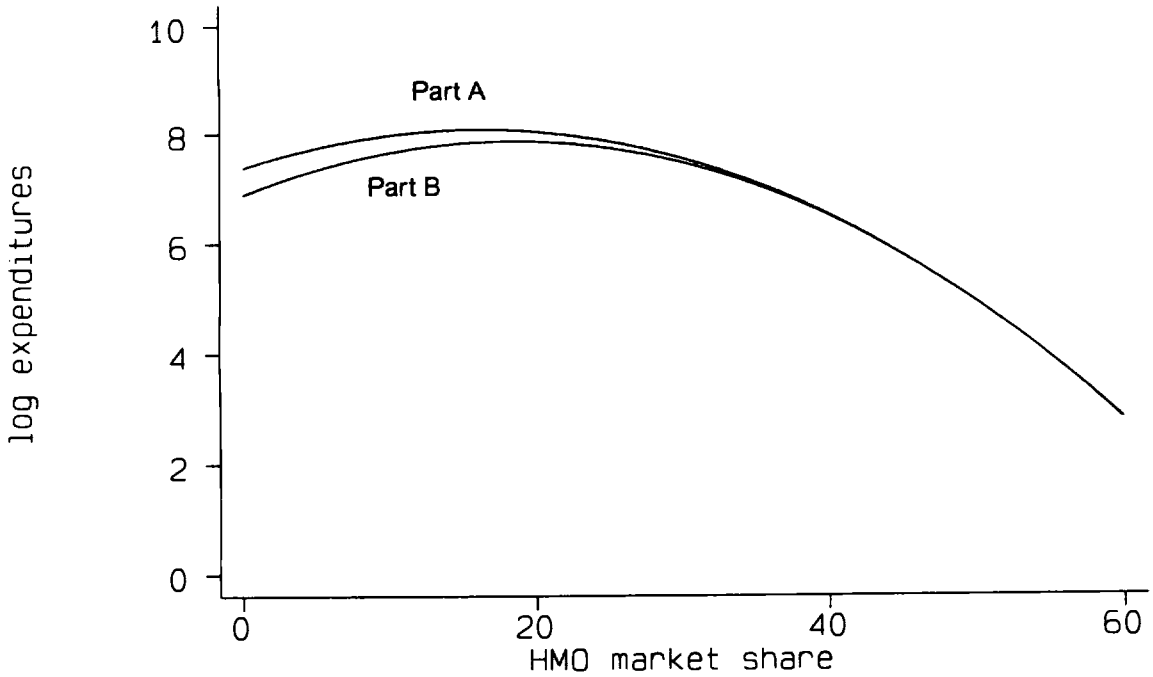

Figure 5: IV estimates using overall market share

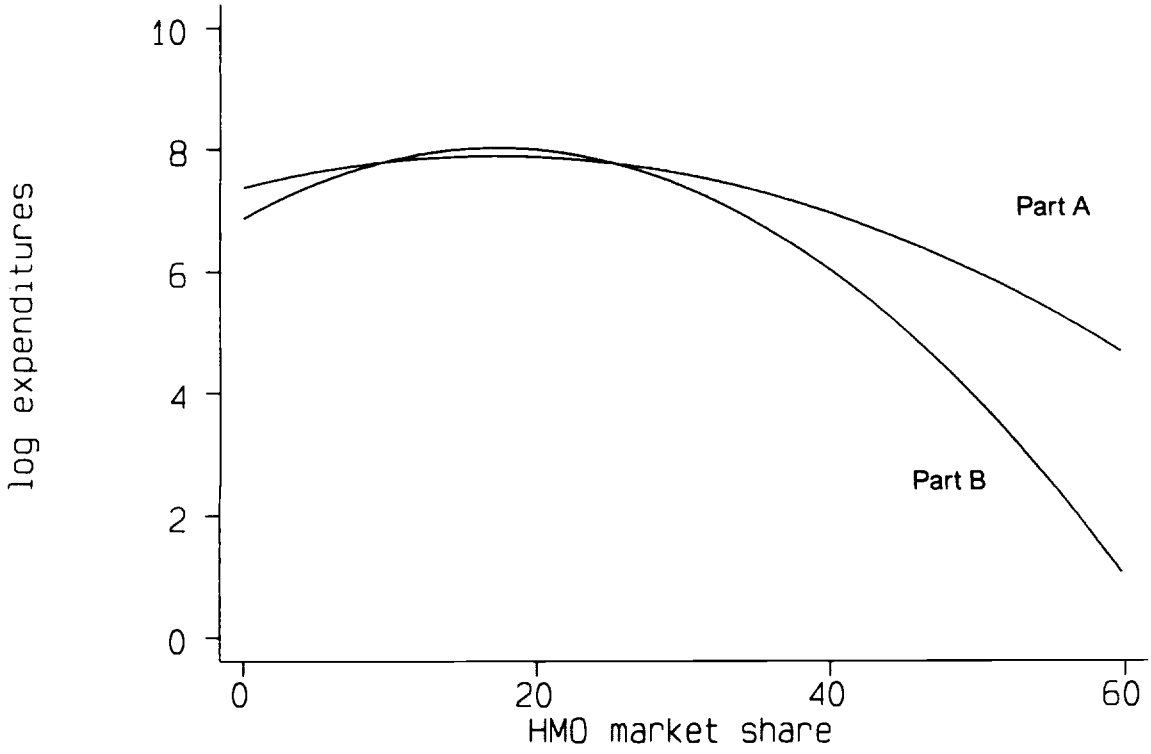

**Table 1: Mean Nominal Medicare FFS Expenditure per Beneficiary, 1986-1990**

| Year | Part A | % Change | Part B | % Change |
|---|---|---|---|---|
| 1986 | $1570.89 | | $831.20 | |
| | (372.80) | --- | (270.25) | --- |
| 1987 | $1660.07 | | $949.61 | |
| | (413.36) | 5.68% | (301.96) | 14.25% |
| 1988 | $1654.95 | | $956.65 | |
| | (418.13) | -0.31% | (280.30) | 0.74% |
| 1989 | $1772.78 | | $1115.77 | |
| | (190.71) | 7.12% | (354.88) | 16.63% |
| 1990 | $1919.99 | | $1233.11 | |
| | (517.87) | 8.30% | (346.12) | 10.52% |

Note: Standard errors in parentheses. N = 3,073 counties per year. Results are weighted by county Medicare enrollment.

**Table 2: Market Share Measures**

| Year | Mean | Standard Deviation | 25th pctl | Median | 75th pctl |
|---|---|---|---|---|---|
| Medicare HMO Market Share | | | | | |
| 1985 | 4.419 | 6.973 | 0.161 | 1.045 | 6.020 |
| 1986 | 5.710 | 8.392 | 0.241 | 1.601 | 8.243 |
| 1987 | 6.162 | 8.961 | 0.261 | 1.874 | 8.571 |
| 1988 | 6.101 | 8.872 | 0.247 | 1.642 | 8.693 |
| 1989 | 6.003 | 8.999 | 0.220 | 1.561 | 7.584 |
| 1990 | 6.369 | 9.500 | 0.225 | 1.604 | 8.648 |
| Overall County HMO Market Share | | | | | |
| 1990 | 14.632 | 12.118 | 4.656 | 13.065 | 21.078 |

Note: Sample size is 3,073 counties per year. Medicare market share measures are weighted by Medicare enrollment and overall county market shares are weighted by county population.

**Table 3: Distribution of Counties by Medicare HMO Market Share Levels and Changes**

| | | HMO Market Share | | | | | |
|---|---|---|---|---|---|---|---|
| Year | | Less than 1% | 1-5% | 5-15% | 15-25% | 25-40% | Over 40% |
| 1985 | Unweighted N | 2425 | 390 | 159 | 57 | 31 | 11 |
| | Unweighted % | 78.91 | 12.69 | 5.17 | 1.85 | 1.01 | 0.36 |
| | Weighted % | 49.65 | 23.38 | 16.16 | 9.32 | 0.96 | 0.52 |
| 1986 | Unweighted N | 2291 | 447 | 211 | 59 | 42 | 23 |
| | Unweighted % | 74.55 | 14.55 | 6.87 | 1.92 | 1.37 | 0.75 |
| | Weighted % | 43.92 | 23.34 | 18.41 | 10.70 | 2.79 | 0.83 |
| 1987 | Unweighted N | 2234 | 470 | 234 | 64 | 45 | 26 |
| | Unweighted % | 72.70 | 15.29 | 7.61 | 2.08 | 1.46 | 0.85 |
| | Weighted % | 41.24 | 23.44 | 20.60 | 10.78 | 2.84 | 1.10 |
| 1988 | Unweighted N | 2280 | 450 | 221 | 67 | 36 | 19 |
| | Unweighted % | 74.19 | 14.64 | 7.19 | 2.18 | 1.17 | 0.62 |
| | Weighted % | 42.28 | 23.75 | 18.41 | 11.57 | 3.02 | 0.98 |
| 1989 | Unweighted N | 2349 | 426 | 187 | 63 | 37 | 11 |
| | Unweighted % | 76.44 | 13.86 | 6.09 | 2.05 | 1.20 | 0.36 |
| | Weighted % | 43.49 | 23.97 | 16.81 | 9.22 | 5.55 | 0.97 |
| 1990 | Unweighted N | 2338 | 424 | 201 | 60 | 40 | 10 |
| | Unweighted % | 76.08 | 13.80 | 6.54 | 1.95 | 1.30 | 0.33 |
| | Weighted % | 43.93 | 21.47 | 18.50 | 7.30 | 7.89 | 0.92 |

| | | 1985-1990 change in market share | | | | | |
|---|---|---|---|---|---|---|---|
| | | Less than -5 | -5 to -1 | -1 to 0 | 0 to 1 | 1 to 5 | Over 5 |
| 1985-90 | Unweighted N | 41 | 132 | 923 | 1675 | 199 | 103 |
| | Unweighted % | 1.33 | 4.30 | 30.04 | 54.51 | 6.48 | 3.35 |
| | Weighted % | 1.21 | 4.51 | 19.54 | 39.80 | 22.17 | 12.76 |

Note: N = 3,073 counties per year. Weighted market share percentages are weighted by Medicare enrollment. Weighted 1985-1990 changes are weighted by 1990 Medicare enrollment. Distributions may not sum to 1 due to rounding.

**Table 4: Means and Standard Deviations of Selected Variables**

| Variable | Mean | Standard Deviation |
|---|---|---|
| Log (Part A Expenditures per Beneficiary) | 7.417 | 0.252 |
| Log (Part B Expenditures per Beneficiary) | 6.877 | 0.316 |
| Lagged Medicare HMO Market Share | 5.691 | 8.518 |
| (Lagged Medicare HMO Market Share)$^2$ | 104.947 | 276.693 |
| County- level Overall HMO Market Share | 13.491 | 11.768 |
| (County-level Overall HMO Market Share)$^2$ | 320.497 | 482.147 |
| Percent Population Age 65 - 74 | 7.774 | 2.273 |
| Percent Population Age 75 - 84 | 4.301 | 1.480 |
| Percent Population Over Age 84 | 1.294 | 0.468 |
| Percent Elderly Population Female | 59.866 | 2.347 |
| Percent Elderly Population Black | 8.087 | 10.161 |
| Percent Elderly Population "Other Race" | 1.834 | 5.202 |
| Per capita Income | 16.441 | 4.410 |

Note: N=15,365 (3,073 counties per year) for all variables except county-level overall market share (N=3,073). Results are weighted by county Medicare enrollment.

**Table 5: Fixed-Effects Results Using Medicare HMO Market Share**

| Variables | Part A | Part B |
|---|---|---|
| Lagged Medicare HMO Market Share /10 | -0.013<br>(0.007) | -0.018<br>(0.006) |
| (Lagged Medicare HMO Market Share /10)$^2$ | -0.011<br>(0.002) | -0.008<br>(0.001) |
| % Population 65-74 | -0.035<br>(0.065) | 0.172<br>(0.050) |
| % Population 75-84 | -0.241<br>(0.119) | -0.552<br>(0.092) |
| % Population Over 84 | 0.056<br>(0.257) | 3.006<br>(0.198) |
| % Elderly Population Female | -0.179<br>(0.053) | 0.070<br>(0.041) |
| % Elderly Population Black | 0.380<br>(0.025) | -0.177<br>(0.020) |
| % Elderly Population Other Race | -0.028<br>(0.008) | -0.101<br>(0.006) |
| Per capita Income | 1.354<br>(0.122) | -0.003<br>(0.094) |
| Year = 1987 | 0.051<br>(0.003) | 0.139<br>(0.002) |
| Year = 1988 | 0.037<br>(0.003) | 0.153<br>(0.003) |
| Year = 1989 | 0.088<br>(0.005) | 0.299<br>(0.004) |
| Year = 1990 | 0.159<br>(0.006) | 0.407<br>(0.005) |

continued

**Table 5: Fixed-Effects Results Using Medicare HMO Market Share**
continued

| Variables | Part A | Part B |
|---|---|---|
| Intercept | 8.038 | 6.152 |
| | (0.328) | (0.252) |
| N | 15365 | 15365 |
| $R^2$ | 0.9175 | 0.9692 |
| F[df] (market share) | 118.82 | 159.86 |
| | [2,12279] | [2,12279] |
| $H_{max}$ | 0 | 0 |
| $E_{20}/E_{10}$ -1 | -0.045 | -0.041 |
| $E_{30}/E_{20}$ -1 | -0.066 | -0.056 |

Note: Standard errors in parentheses. The dependent variable is log (expenditures per beneficiary). The F-statistic shown tests the hypotheses that the linear and quadratic market share coefficients are jointly 0; $p<.0001$ in both cases. Regressions are weighted by county Medicare enrollment.

**Table 6: Fixed Effects Regression Results from Alternate Specifications**

| | Part A | | Part B | |
|---|---|---|---|---|
| | 1 | 2 | 3 | 4 |
| Variables | Counties with >1% Mkt. Shr. | MSA-Level | Counties with >1% Mkt. Shr. | MSA-Level |
| Lagged Medicare HMO Mkt. Shr./10 | -0.030 | 0.005 | -0.009 | 0.015 |
| | (0.017) | (0.020) | (0.012) | (0.015) |
| (Lagged Medicare HMO Mkt. Shr./10)$^2$ | -0.007 | -0.008 | -0.007 | -0.008 |
| | (0.004) | (0.005) | (0.002) | (0.004) |
| N | 2795 | 1610 | 2795 | 1610 |
| $R^2$ | 0.9141 | 0.9058 | 0.9735 | 0.9694 |
| F[df] (market share) | 37.97 | 4.28 | 36.60 | 4.44 |
| | [2,2223] | [2,1294] | [2,2223] | [2,1249] |
| P(F) | 0.0000 | 0.0140 | 0.0000 | 0.0119 |
| $H_{max}$ | 0 | 3.125 | 0 | 9.375 |
| $E_{20}/E_{10}$ - 1 | -0.050 | -0.019 | -0.030 | -0.009 |
| $E_{30}/E_{20}$ - 1 | -0.063 | -0.034 | -0.043 | -0.025 |

Note: Standard errors in parentheses. The dependent variable is log (expenditures per beneficiary). All models also include controls for area demographics, year dummies, and county dummies. The F-statistic shown tests the hypothesis that the linear and quadratic market share terms are jointly 0. Regressions are weighted by county Medicare enrollment.

**Table 7: First Stage Regression Coefficients**

| | Dependent variable | |
|---|---|---|
| | 1<br>Medicare HMO<br>Market Share | 2<br>Overall HMO<br>Market Share |
| % Workers White Collar , 1990 | 0.011<br>(0.001) | 0.021<br>(0.003) |
| Firm Size | 0.013<br>(0.002) | 0.066<br>(0.006) |
| % Population 65 - 74 | -1.965<br>(0.093) | -1.363<br>(0.229) |
| % Population 75 - 84 | 3.848<br>(0.202) | -0.226<br>(0.486) |
| % Population Over 84 | -2.387<br>(0.366) | 5.376<br>(0.895) |
| % Elderly Population Female | -0.569<br>(0.043) | -0.714<br>(0.115) |
| % Elderly Population Black | -0.092<br>(0.007) | -0.079<br>(0.019) |
| % Elderly Population Other Race | 0.432<br>(0.013) | 0.421<br>(0.030) |
| Per capita Income | 1.619<br>(0.246) | 3.178<br>(0.566) |
| Metropolitan County | 0.299<br>(0.023) | 0.377<br>(0.061) |
| Rural County | -0.067<br>(0.023) | -0.417<br>(0.061) |

continued

**Table 7: First Stage Regression Coefficients**
continued

| | Dependent variable | |
|---|---|---|
| | 1<br>Medicare HMO<br>Market Share | 2<br>Overall HMO<br>Market Share |
| Year = 1987 | 0.098<br>(0.019) | -- |
| Year = 1988 | 0.103<br>(0.020) | -- |
| Year = 1989 | 0.057<br>(0.020) | -- |
| Year = 1990 | -0.024<br>(0.021) | -- |
| Intercept | 2.834<br>(0.248) | 3.072<br>(0.657) |
| N | 15365 | 3073 |
| $R^2$ | 0.2357 | 0.4451 |
| F[df] (instruments) | 50.59<br>[2,15349] | 82.66<br>[2,3061] |

Note: Standard errors in parentheses. The dependent variable is lagged HMO market share. The F-statistic tests the hypothesis that the coefficients on the instrumental variables are jointly 0; $p<.0001$ in both cases. Regressions are weighted by county Medicare enrollment (column 1) and county population (column 2).

**Table 8: IV Regression Results**

| | Part A | | | Part B | | |
|---|---|---|---|---|---|---|
| | 1 | 2 | 3 | 4 | 5 | 6 |
| Variables | All Counties | Counties with >1% Mkt. Shr. | MSA-Level | All Counties | Counties with >1% Mkt. Shr. | MSA-Level |
| Lagged Medicare HMO Mkt. Shr./10 | 0.880 | 0.515 | 0.596 | 1.080 | 0.546 | 0.840 |
| | (0.043) | (0.071) | (0.055) | (0.044) | (0.071) | (0.083) |
| (Lagged Medicare HMO Mkt. Shr./10)$^2$ | -0.273 | -0.160 | -0.179 | -0.293 | -0.164 | -0.287 |
| | (0.021) | (0.016) | (0.034) | (0.021) | (0.016) | (0.051) |
| N | 15365 | 2795 | 1610 | 15365 | 2795 | 1610 |
| F[df] (market share) | 213.57 | 52.12 | 83.94 | 326.45 | 54.82 | 66.83 |
| | [2,15346] | [2,2776] | [2,1593] | [2,15346] | [2,2776] | [2,1593] |
| F[df] (instruments) | 50.59 | 12.04 | 14.77 | 50.59 | 12.04 | 14.77 |
| | [2,15349] | [2,2779] | [2,1596] | [2,15349] | [2,2779] | [2,1596] |
| $H_{max}$ | 16.117 | 16.094 | 16.648 | 18.430 | 16.646 | 14.634 |
| $E_{20}/E_{10}$ -1 | 0.063 | 0.036 | 0.061 | 0.223 | 0.055 | -0.021 |
| $E_{30}/E_{20}$ -1 | -0.384 | -0.248 | -0.258 | -0.320 | -0.240 | -0.448 |

Note: Standard errors in parentheses. The dependent variable is log (expenditures per beneficiary). Regressions also contain controls for area population demographics, MSA status, year dummies, and an intercept. The F-statistics shown test (1) the hypotheses that the coefficients on the linear and quadratic market share terms are jointly 0, and (2) the hypothesis that the coefficients on the instruments in the first stage are jointly 0; $p<.0001$ in all cases. Regressions are weighted by county Medicare enrollment.

**Table 9: OLS Regression Results**

| Variables | Part A | | | Part B | | |
|---|---|---|---|---|---|---|
| | All Counties | Counties with >1% Mkt. Shr. | MSA-Level | All Counties | Counties with >1% Mkt. Shr. | MSA Level |
| Lagged Medicare HMO Mkt. Shr./10 | 0.249 | 0.191 | 0.262 | 0.350 | 0.313 | 0.359 |
| | (0.004) | (0.009) | (0.012) | (0.005) | (0.012) | (0.015) |
| (Lagged Medicare HMO Mkt. Shr./10)$^2$ | -0.047 | -0.032 | -0.050 | -0.075 | -0.065 | -0.078 |
| | (0.001) | (0.002) | (0.004) | (0.001) | (0.003) | (0.004) |
| N | 15365 | 2795 | 1610 | 15365 | 2795 | 1610 |
| $R^2$ | 0.527 | 0.563 | 0.5656 | 0.602 | 0.542 | 0.6044 |
| F[df] (market share) | 2330.24 [2,15351] | 305.99 [2,2781] | 312.05 [2,1596] | 2860.10 [2,15351] | 369.70 [2,2781] | 322.02 [2,1596] |
| $H_{max}$ | 26.489 | 29.844 | 26.200 | 23.333 | 24.077 | 23.013 |
| $E_{20}/E_{10}$ - 1 | 0.114 | 0.100 | 0.119 | 0.133 | 0.125 | 0.133 |
| $E_{30}/E_{20}$ - 1 | 0.014 | 0.031 | 0.012 | -0.025 | -0.012 | -0.031 |

Note: Standard errors in parentheses. The dependent variable is log(expenditures per beneficiary). All model also include controls for area demographics, year dummies, and an intercept. The F-statistics shown tests the hypotheses that the linear and quadratic market share coefficient are jointly 0. Regressions are weighted by county Medicare enrollment.

**Table 10: IV Estimates Using Overall County HMO Measures**

| | Part A | Part B |
|---|---|---|
| Overall HMO Market Share | 0.606 | 1.329 |
| | (0.206) | (0.379) |
| (Overall HMO Market Share/10)$^2$ | -0.177 | -0.385 |
| | (0.059) | (0.108) |
| N | 3073 | 3073 |
| F[2,3058] (market share) | 4.57 | 6.33 |
| P(F) | 0.0104 | 0.0018 |
| F[2, 3061] (instruments) | 82.66 | 82.66 |
| P(F) | 0.0000 | 0.0000 |
| $H_{max}$ | 17.119 | 17.260 |
| $E_{20}/E_{10}$ -1 | 0.078 | 0.174 |
| $E_{30}/E_{20}$ -1 | -0.279 | -0.449 |

Note: Standard errors in parentheses. The dependent variable is log (expenditures per beneficiary). The models also contain controls for area population demographics, MSA status, and an intercept. The first F-statistic shown tests the hypotheses that the linear and quadratic market share terms are jointly 0. The second tests the hypotheses that the first-stage instruments are jointly 0. Regressions are weighted by county population.

Appendix A: Construction of County-level HMO market share estimates.

The process by which the estimates were constructed is described briefly here; full details are presented in Baker (1995). Conceptually, construction took place in three steps. First, for each HMO in the United States, its total enrollment and service area, specified by county, were obtained. Second, the enrollment of each HMO was distributed among the counties in its service area. Finally, the total number of enrollees in each county was computed by summing county enrollments over all of the HMOs serving the county. Using the total number of HMO enrollees in each county, HMO market share was computed as the ratio of enrollees to total population.

The primary source of information on HMO enrollments and service areas is the *National Directory of HMOs*, published annually by the Group Health Association of America (GHAA). Each year the GHAA conducts a mail survey, with telephone follow up, of all known HMOs in the country and, among other things, asks their total enrollment and their service area. The results of the survey are published in the annual *Directories*. To construct estimates of 1990 county market share, the 1991 *Directory*, which lists enrollment and service area for each of the 567 HMOs in the mainland U.S., Alaska, and Hawaii as of December 31, 1990, was used.

All but one of the HMOs in the directory indicated their enrollment. In the missing case, data from the 1992 *Directory* was used. Most HMOs (459 of 567) also indicated the counties that they served. However, 108 HMOs (19%), did not provide a clear definition of their market area in terms of counties. For these HMOs, market areas were determined by reference to subsequent *Directories* and/or telephone contact.

The next step was to distribute the enrollment of each HMO among the counties in its service area. Initially, this was done by simply distributing enrollment proportionally to county population. In addition, since HMO enrollment may be concentrated near HMO headquarters or since HMOs may locate their headquarters in areas where their enrollment is concentrated, estimates that incorporate both county population and distance from HMO headquarters were constructed. The correlation between estimates produced by the two methods is approximately 0.97. Estimates that incorporate both population and distance are used in subsequent analyses.

Once enrollments had been distributed over service areas, the total number of enrollees in each county was computed by summing over the set of HMOs serving that county. Using the set of county enrollment estimates, market share estimates were computed as the proportion of the population enrolled in HMOs.

Since the county service areas on which the series are based are quite accurate, it is likely that the series themselves are also quite accurate. Comparisons with limited HMO market share data available from independent sources also suggest that the new estimates are reasonable measures of HMO activity levels.