Managed Care and Medicare Expenditures

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I. Introduction

The 1990s saw a dramatic increase in the percentage of Medicare enrollees who joined an HMO. By 2001, approximately 15% of Medicare enrollees had opted out of the traditional fee-for-service (FFS) program and had joined a Medicare managed care organization. The rise of a meaningful managed care sector in Medicare has broad policy implications for the program. A large Medicare managed care sector may affect both the financial health of the program (the focus of this paper) and the physical health of Medicare enrollees.[1]

Conceptually, the aggregate impact of managed care on Medicare expenditures is comprised of two components—a spillover effect and a selection effect. Spillover effects refer to changes in the care delivered to fee-for-service enrollees that arise due to changes in HMO enrollment among Medicare beneficiaries, holding the health status of FFS enrollees constant. Spillover effects may arise from either aggregate HMO enrollment or Medicare specific HMO enrollment. There are several reasons why we might expect spillovers. For example if physicians tend to practice similarly for all patients, more managed care enrollment may alter practice patterns for FFS patients. Alternatively, managed care enrollment may influence the availability of services by influencing aspects of market structure such as the number of hospitals, or beds, or available services (Chernew 1995). Such changes could influence practice patterns for all individuals in the market.

Selection effects refer to the impact of non-random selection of beneficiaries into Medicare HMOs. If beneficiaries who are healthier systematically enroll in Medicare HMOs, the costs for those remaining in FFS will rise because the FFS population will be, on average, sicker. This

[1] In late 2000, the Health Care Financing Administration (HCFA), now called the Center for Medicare and Medicaid Services (CMS), convened a technical review panel to examine the assumptions used by the Office of the Actuaries to assess the financial health of the Medicare Trust Funds. The panel concluded that these assumptions were in need of revision. One specific area of concern was forecasting the impact of Medicare managed care on Medicare costs.

will occur even if care for any given FFS patient is not affected by managed care penetration (i.e., even in the absence of spillover effects).

In this paper we assess the aggregate impact of Medicare HMO enrollment on FFS Medicare expenditure and utilization and decompose it into spillover and selection effects. Our basic approach is to regress spending by FFS Medicare beneficiaries on the share of Medicare beneficiaries in their county enrolled HMO plans. The aggregate effect is estimated using a model with only basic demographics as covariates, whereas the spillover effect is estimated using a model with a much larger set, including extensive controls for health status.

Because HMO market share is potentially endogenous, we use variation in Medicare payment policy as an instrument for HMO penetration. This approach has been used successfully in other contexts (Town and Liu (2002); Gowrisankaran and Town (2004)). Our identification comes from longitudinal variation in payment rates over our study period (1994 – 2001) and reflects in large part reforms instituted in the balanced Budget Act of 1997 (BBA) and idiosyncrasies in Medicare payment rules.

We find evidence of substantial spillover in a sample of fee-for-service Medicare beneficiaries. In particular, in instrumental variables models we find that a one percentage point increase in county-level Medicare HMO penetration is associated with between a one and two percent reduction in individual-level annual Medicare spending. These estimates are considerably larger in magnitude than corresponding least squares estimates. To better understand the source and validity of our findings, we also estimate models which examine the impact of Medicare HMO penetration on various categories of health care utilization. We find that increases in county-level Medicare HMO enrollment reduce both inpatient and outpatient events, with the largest effects coming on the intensive utilization margins. These estimates are consistent with our main finding that increased Medicare HMO penetration reduces spending by FFS beneficiaries.

In the following section, we provide background on the progression of Medicare managed care and its relation to our work. Section III presents our empirical strategy, which relies on county-level payment rates as instruments for county-level Medicare HMO penetration. Section IV describes our data, including the construction of our key variables and a detailed description of the samples we analyze. Section V presents our estimates and discusses their relevance. Section VI concludes the paper.

II. Background

In 1982, Congress passed the Tax Equity and Fiscal Responsibility Act (TEFRA), which mandated the provision of managed care plan options to Medicare beneficiaries. Under the statute, the Health Care Financing Administration was directed to contract with HMOs to provide a managed care option to Medicare enrollees. Under M+C Medicare enrollees can forgo the traditional Medicare insurance program and enroll in a qualified HMO. The HMO agrees to provide health insurance that covers all Medicare-covered services (Parts A and B) for the enrollee in exchange for a per-capita fee from CMS. In addition, HMOs may offer benefits beyond those available to fee-for-services Medicare beneficiaries. The rationale underlying TEFRA is that HMOs may be more efficient at providing care thereby reducing the federal Medicare expenditures. Beginning in the early 1990s and extending into the late 1990s, there was a surge in the share of Medicare beneficiaries that took advantage of this option.

HMO enrollment may be beneficial for beneficiaries and the Medicare program if Medicare HMOs provide care more efficiently than the traditional FFS Medicare system. More efficient care can be manifest through lower costs of care, higher quality, or through broader benefit coverage. If savings exist from the HMOs, Medicare ultimately may save money (depending on how payment rates are set) and/or beneficiaries could receive enhanced benefits because of competition among plans.

An important lever that Medicare has to influence beneficiary participation in HMOs is payment policy. Our first aim is to understand the relationship between payments and enrollment. Several studies suggest that payment rates affect HMO participation in the Medicare program (Cawley, Chernew et al. 2002; Town and Liu 2002). However, none of these studies directly measures the impact of payment changes on aggregate HMO enrollment.

In addition to estimating the impact of payments on enrollment, it is important for forecasting and policy purposes to understand the fiscal impact of Medicare HMO enrollment on the program. Medicare HMO enrollment has both direct and indirect impacts on the Medicare program. The direct fiscal impact of a Medicare beneficiary choosing to enroll in an HMO depends on Medicare's payment rate to HMOs, relative to what the individual would have spent had they remained in the traditional FFS Medicare program. Because payment rates for Medicare HMOs were historically tied to the local costs of care for enrollees in the traditional Medicare program, and because HMOs attracted a healthier population than average, analysts have felt that growing enrollment in HMOs increased the costs of the Medicare program. Any cost savings obtained by HMOs were either captured by the HMOs or competed away as they offer richer benefit packages to beneficiaries. Though the Balanced Budget Act of 1997 (BBA) changed payment rules, analysis by MedPAC suggests that spending by Medicare for HMO participants was 4% higher than for demographically similar beneficiaries in traditional Medicare (MedPAC March 2002).

Yet, the preceding calculation does not adjust for any market spillover effects. If there are spillover effects from Medicare HMO practice patterns to practice patterns in traditional FFS Medicare, efficiencies from HMO practice may reduce the costs of caring for individuals who do not enroll in Medicare HMOs. These savings may offset the direct effect of Medicare HMO enrollment.

Several strands of work suggest that spillover may be important. Most directly is the relatively small body of work examining the impact of managed care enrollment on Medicare costs (Baker and Corts (1996); McClellan and Baker (2001); Bundorf et al. (2001); Baker (1997); Baker and Shankarkumar (1997); Cutler and Sheiner (1997)). Also relevant is the somewhat larger literature examining the impact of overall HMO activity on the market as a whole (Wickizer and Feldstein (1995); Hill and Wolfe (1997); Gaskin and Hadley (1997); Robinson (1991); and Robinson (1996); Robinson and Luft (1988); Melnick and Zwanziger (1995)). Overall, this research has provided strong support for the general proposition that markets are connected and thus we would expect activities in the Medicare HMO market to influence Medicare FFS costs.[2]

Although the overall body of literature supports the existence of spillover effects, the literature explicitly examining the impact of Medicare HMO enrollment on Medicare FFS expenditures is relatively small and most of the relevant studies use relatively similar methods. Overwhelmingly, the research cited above treats HMO penetration as exogenous. Omitted area level variables may be influence the magnitude of the findings and, as the review article pointed out, some of the findings are implausible. Baker (1997) is an exception, reporting instrumental

[2] Note also that a series of studies by Zwanziger, Melnick and colleagues reach a similar qualitative conclusion using a somewhat different approach, emphasizing importance of selective contracting on costs, without explicitly controlling for managed care penetration (Zwanziger and Melnick (1988); Melnick and Zwanziger (1988); Melnick et al. (1989a); Melnick et al. (1989b); Zwanziger et al. (1994)).

variable (IV) results, though he does not distinguish between spillover and selection effects, which might account for the concave relationship between overall HMO enrollment and FFS Medicare expenditures that he finds.

None of the existing work has directly examined the impact of payment rate changes on spillover. Assessing the impact of payment rates on FFS Medicare expenditures requires measuring the impact of payment on enrollment, the extent to which the change in enrollments are systematically related to health status, and the extent to which changes in enrollment will affect practice patterns, holding health status constant. Our approach allows us to examine these issues more directly than the existing literature because our identification strategy is based on payment changes. This may be particularly important if payment changes affect not only the penetration of HMOs, but also the characteristics of the HMOs in the market.

III. Methods

Using a sample of individuals enrolled in Medicare FFS, we estimate models of the form:

$$(1)\ Log\ Expenditure_{ict} = \alpha_c + \gamma_t + \beta\ MC_{ct} + \lambda\ X_{it} + \varepsilon_{ict}$$

where *i* indexes the individual, *c* indexes county of residence and *t* indexes year of interview. *Expenditure* represents total annual enrollee spending on all covered services. In later specifications, we replace spending with measures of health care utilization (e.g. inpatient days, outpatient hospital visits, physician visits, etc.) in an attempt to better understand our spending estimates. *MC* represents the fraction of Medicare beneficiaries enrolled in an HMO. In this approach we are including county fixed effects, (α_c) and thus identifying the impact of managed care on spending through changes in the penetration of managed care in a given location over time. In addition, we also include a vector of year effects (γ_t) to account for trends that are common to all counties in our sample. The vector X_i represents individual covariates that will affect demand (e.g. age and sex). Equation (2) is designed to measure an aggregate effect, including selection and spillover, effects. Thus the X vector will include age and gender only, because CMS can easily adjust payment rates for these demographic characteristics. Other correlates of demand are intentionally omitted for now.

In this specification β measures the gross impact of M+C enrollment on utilization. That is, the estimated β embodies both the impact of changes in M+C enrollment on resource utilization for a given individual and the impact of changes in M+C enrollment on the characteristics of the population in the fee-for-service sector. Below we discuss our strategy for decomposing these effects.

This specification is similar to the existing literature examining the impact of HMO enrollment on Medicare spending. Unlike the existing literature, we use individual level data. Although not as relevant in this analysis, this will become important later when we expand the model to isolate selection and spillover effects.

Since the disturbance term in equation (1) is likely correlated with *MC* penetration, Ordinary Least Squares estimation may generate a biased estimate of β . Assuming that HMOs enter areas with rising FFS spending, we would expect the bias in the OLS estimates to be positive. We correct for this bias using an instrumental variable (IV) approach. In this setting, the county-level payment rates are the instruments used to identify the HMO effect. [3] We expect the IV estimates to be more negative than the OLS estimates.

[3] Other potential instruments could be based on firm size in an area. Baker (1997) advocates the use of the size distribution of business establishment in a given area as an instrument for commercial HMO penetration. The interaction between the Medicare payment rate and the size distribution of firms is a good potential instrument set for commercial HMO penetration.

Variation in payment rates comes from two sources. First, prior to the BBA, Medicare based its payment to HMOs on the per capita costs for FFS enrollees in the county. This may seem to suggest that payment rates would be a poor instrument for HMO penetration in our model because of their apparent relationship with FFS spending, the Medicare payment rates at time t were based on average FFS spending rates between periods t-8 and t-3.[4] The validity of payment as an instrument will depend on the autocorrelation of FFS spending over time. To explore the potential for using payment as an instrument, we estimated a first-order autoregression of the residuals from a regression of log spending on all of the exogenous variables, including the instruments.[5] The autocorrelation parameter appears to be sufficiently small to allow this to be a useful source of identifying variation.[6]

The other source of payment variation is the BBA of 1997 (and its subsequent refinements), which changes the formula for paying HMOs to reduce the reliance on local FFS costs. The BBA fundamentally modified Medicare's payment methodology.[7] While the changes in the payment formula are rather technical, for our purposes the important feature of the new payment formula is that now updates to the county payments are divorced from changes in the Medicare FFS experience in the county.[8] In most counties the post-BBA payment formula lead to a substantial decrease in CMS payment rates over what the HMOs would have received prior to the BBA. It is estimated that the BBA methodology lowers the payments to HMOs by an average of 6%.[9] In

[4] More specifically, these are five year averages, starting eight years prior to time t.
[5] This required collapsing the residuals to county-year cells, so the residuals used in the autoregression are averaged over all individuals in a given county in a particular year.
[6] In particular, the parameter ranged from 0.04 to 0.07 and was not statistically different from zero at conventional levels of significance.
[7] With the passage of the BBA, the Medicare HMO program also received a new name: "Medicare+Choice."
[8] The county rates in 1998 and onward are based on the 1997 county rate book and not on the experience of fee-for-services enrollees in each county in the previous years. Those county rates are set equal to the maximum of three rates: blended input price—an adjusted national rate and an area-specific rate; a floor payment designed to increase the rates in low-paying counties; and minimum rate increases of 2% per year.
[9] Source: Congressional Budget Office (1999).

addition to reducing the level of CMS payments, the BBA diminished the variance of the payments across the counties.

While the impact of the BBA on the payment rate is likely unrelated to the error term in (2) implying that the instruments are valid, the payment rate may still be a weak instrument. We test the strength of our instrument set via a standard F-test. As will be seen, all F-tests strongly reject the hypothesis that our instruments are unrelated to county-level Medicare HMO enrollment rates.[10] The validity of payment as an instrument also requires payment changes to be unrelated to existing trends on spending across counties. The counties that experienced relatively generous or stingy growth payments due to the BBA might differ.

To examine this possibility we divided the counties into those whose payment growth was slowed by the BBA and those whose spending growth was accelerated by the BBA. The results from this exercise are presented in Table 1. This was based on the ratio of payment growth in each county after the BBA relative to before the BBA. Prior to the BBA, counties which were generously treated by the BBA (had above the median relative payment growth) had roughly the same percent growth in FFS expenditure (9.1%) as those that were less generously treated by the BBA (10.3% FFS spending growth). This suggests spending trends prior to the BBA were similar across counties that later were differentially treated by the BBA. After the BBA, and consistent with our results reported below, the counties whose payment growth was slowed by the BBA had higher percentage FFS spending growth (25.7%) relative to those counties whose payment growth was accelerated by the BBA (14.5% FFS spending growth). One important issue is spillover from HMO penetration in the non-Medicare market. Because we use an IV approach, we can interpret the findings as spillovers only from Medicare HMO penetration. Any

[10] A standard rule of thumb is that this F-statistic be greater than ten. All of our F-statistics are greater than thirty. In addition, we report partial R^2 for each first stage regression.

confounding effects of system wide HMO penetration will not affect the important coefficients in the base specification.

To identify selection and spillover effects we estimate equation (1), adding controls for health status and other demographic variables related to utilization. Specifically, the estimated effect of managed care penetration from equation (2), which included only age and sex as covariates, is 'unconditional'. It measures only how expenditure in the FFS Medicare sector varies with HMO penetration and thus may include both spillover and selection effects. The better we are able to control for individual health status and other correlates of demand, the more we isolate the spillover HMO effect from the selection effect. In order to do this, we estimate the following:

$$\text{(2)} \qquad Log\ Expenditure_{ict} = \alpha_c + \gamma_t + \beta\ MC_{ct} + \lambda\ X_{it} + \varphi\ X^*_{it} + \varepsilon_{ict}$$

where X^* represents the additional array of health status measures and other variables correlated with demand.

Because we are controlling for health status in this model (as well as other covariates that may be related to utilization), the estimated managed care effect, β, is largely the spillover effect. The difference between the coefficient estimates from this model and from the more parsimonious one is an estimate of the magnitude of the selection effect.

IV. Data

We use data from the annual Cost and Use files of the Medicare Current Beneficiary Survey (MCBS) for the years 1994 to 2001, inclusive. The MCBS is a nationally representative survey of Medicare beneficiaries that gathers information on respondents via a rotating panel. While the sampling frame includes elderly and disabled beneficiaries, we limit our analysis to individuals aged sixty-five and older. In addition, we exclude individuals who completed facility interviews.[11] Since we examine potential spillovers associated with Medicare managed care, we also exclude beneficiaries enrolled in HMOs when surveyed. In what follows, we describe key variables and the construction of our analysis samples.

The MCBS contains detailed information on respondent demographics (e.g., income, race, living arrangements), health status (e.g., self-reported health status, past experience with a wide variety of diseases and disorders) as well as health care utilization and expenditure. With respect to the latter, respondents are linked with Medicare claims data to ensure the accuracy of individual spending measures. The MCBS staff uses this information to construct each respondent's total annual Medicare expenditure, which is our main dependent variable. To better understand our findings, we also examine the impact of Medicare HMO penetration on selected categories of health care utilization including inpatient events, outpatient events, medical provider events and office visits.[12] Another set of key variables includes county-level estimates of HMO enrollment and the county-specific payment rates CMS uses to compensate manage care companies. Both are available from CMS. We merge this county-level information to MCBS data using geographic identifiers available in the restricted-use version.

We analyze two samples. The first eliminates the relatively few individuals with zero annual expenditure and the second includes them, assigning these individuals an expenditure of one dollar since we model log expenditure in our spending models. As mentioned, we eliminate institutionalized individuals and those under sixty-five years old which results in a sample of 76,974 individuals. Eliminating those enrolled in HMOs reduces this figure to 64,321 individuals

[11] These individuals could not complete the interview on their own and required a proxy to do so. They amount to about ten percent of sample members on an annual basis.
[12] Medical provider events include doctor visits, surgical and laboratory services, or purchases of medical equipment and supplies.

and missing covariate information further reduces the sample size to 62,790. The corresponding sample size without zero expenditure individuals is 61,208, which represents a conditional loss of about 2.5 percent of cases.

Since the MCBS contains several counties with few individuals, we restrict our analysis to individuals in counties that contribute at least fifteen observations over the eight years of data we examine.[13] This restriction reduces the sample that includes zero expenditure individuals from 62,790 to 60,067 and the sample that excludes these individuals from 61,208 to 58,569.[14] Both figures represent over ninety-five percent of their unrestricted sample sizes.[15] Table 1 presents selected means and standard deviations for these four samples. In particular, the first two columns represent samples we use to generate regression estimates, while the last two columns represent samples that include all counties regardless of the number of observations they contribute. Comparing the first and third columns as well as the second and fourth ones, it is apparent that there are no substantial differences associated with our restriction. However, as expected, there are differences in average expenditure between samples that do and do not contain zero expenditure individuals.

Table 1 indicates that there is substantial variance in medical expenditures across individuals. The variance is close to twice the mean expenditure. Importantly, we know from the research on local area variation that much of the difference in expenditures across individuals is do to geography. There is also substantial variation in the "exposure" to the HMO penetration

[13] Counties contributing fewer than fifteen observations contribute an average of less than four observations over the eight years in question or less than one half of an observation per year, on average. In the near future, we will attempt to better understand the nature of these individuals. For example, they may be movers followed by MCBS or perhaps they are related to what MCBS calls "zip fragments", which are zip code areas that overlap county boundaries. If so, such individuals may be assigned to nearby counties already in the sample.

[14] With the exception of models that check the robustness of our main findings, all estimates are generated from these two samples.

[15] In particular, 60,067/62,790≈0.957 and 58,569/61,208≈0.957

rate and the Medicare payment rate to HMOs. This variation is present both in the cross-section and over time within a county.

V. Results

In Table 3 we present the OLS estimates of the impact of HMO penetration on Medicare FFS expenditures. In all samples and specifications the estimated coefficients are small and insignificant. Also, the coefficients estimates are very similar across the different specifications implying that the OLS estimates do not suggest important selection effects of managed care. The coefficients on age and age squared are significant and imply a concave relationship between age and FFS expenditures.

Table 4 presents the IV estimates. In all samples and specifications coefficient on HMO penetration is negative, large in magnitude and significantly different from zero at the 1% level. Focusing on the coefficient estimate in the third column (this corresponds to the specifications with a limited set of control variables), the estimate implies that a 1 percentage increase in Medicare HMO penetration would reduce expected FFS expenditures by roughly 1.6%. Column (4) presents the coefficient estimates with the full set of demographic and risk controls. The coefficient estimates are modestly smaller in magnitude. The estimates imply that that a 1 percentage increase in Medicare HMO penetration would reduce expected FFS expenditures by roughly 1.3%. Over our sample period the mean HMO penetration increased approximately 8 percentage points. Our estimates imply that the rise of managed care in Medicare reduced FFS expenditures by approximately 10% per FFS enrollee.

Thus, the implications of the results in Tables 3 and 4 are three-fold. First, increasing Medicare HMO enrollment reduces FFS expenditures. Second, there is little evidence of a significant risk selection suggesting that the mechanism through which managed care affects FFS spending is primarily through spillover. Third, the difference between the OLS and IV estimates suggests that managed care penetration is higher in areas that experience positive expenditure shocks.

Of course, the reliability of our estimates is only as good as the validity of our instruments. In the bottom half of Table 4 we present some evidence on this issue. The instruments explain a significant amount of the variation in HMO penetration controlling for county fixed-effects and the other right hand side variables. The coefficients on the payment rate and squared payment rate are both large in magnitude and significant at the 1% level in all specifications. Estimates from the first stage model, suggest that a 10% increase in HMO payments leads to a 1 percentage point increase in Medicare HMO penetration. The partial R^2 is .14 and the F-test that the coefficients on the instruments are all zero is over 33 in all specifications. Thus, there is no evidence that our estimates suffer from a weak instrument problem. The Hausman test of the consistency of the OLS estimates resoundingly rejects, while the Hansen test of the over identifying restrictions does not reject in any specification. Finally, the Hahn-Hausman (2002) test suggests that standard (first-order) asymptotics hold and, as such, supports the validity of the preceding tests. Therefore, failure to reject this test, in conjunction with the above results of several IV diagnostics, is consistent with the validity of our instruments.[16] In sum, the test statistics uniformly support the validity of the instruments. When combined with the diagnostic results of minimal autocorrelation in FFS spending growth and similar spending growth prior to

[16] In general, the Hahn-Hausman test compares the coefficient estimates on endogenous right-hand side variables from "forward" and "reverse" regressions (where the latter switches the endogenous right-hand side variable and the dependent variable) since the forward estimate should equal the reciprocal of the reverse estimate under standard asymptotics. See Hahn and Hausman (2002) for more details. As mentioned in the notes to Table 4, our forward coefficient estimates are nearly identical to the reciprocal of the corresponding reverse estimates in all cases.

the BBA in counties treated more and less generously by the BBA, we believe we these are reasonable instruments.

In Table 5 we present some robustness checks of our findings. Medicare managed care grew extremely rapidly in both California and Florida during the 1990s. Both of these states may be anomalous for other reasons unrelated to the impacts of managed care on FFS expenditures and those anomalies may bias our inferences. The results in Table 4 indicate that our conclusions are not driven by the Florida and California experience. The coefficient estimate from column (4) (the sample that drops both California and Florida) is -.0098 and it is statistically significant at the 5% level. Since the implied effect is only slightly smaller relative to the full sample, it continues to represent a practically significant impact of Medicare HMO penetration on FFS expenditures.

In order to isolate where managed care spillover is impacting medical expenditures we estimate IV specifications of the impact of HMO penetration on the number of inpatient, outpatient, medical provider and office visit events. These results are presented in Table 6 and Table 7. The distribution of the dependent variable is discrete and skewed thus we present the results using three different specifications of the dependent variable. The three specifications are indicators whether any event occurred, the number of events, and the number of events conditional on the number being greater than zero. The results indicate that managed care spillovers appear to primarily affect the frequency outpatient and medical provider events conditional upon there being an event. That is, the impact of managed care is on the intensive margin of medical care usage for outpatient and medical provider events. Again, the magnitude of the managed care spillover is large. Conditional on having an outpatient event, a one percentage point increase in HMO penetration reduces the expect number of visits by .08 or

1.4% when evaluated at the mean of the dependent variable. There is also some evidence that managed care penetration impacts inpatient events but the significance of the coefficient is above the 5% level in only one of the three specifications making definitive conclusions regarding the impact of managed care on inpatient use difficult.

In sum, our results indicate that there are substantial spillovers from Medicare managed care to the FFS program. Furthermore, the impacts of managed care spillovers appear to be on intensive margins of outpatient and medical event use.

VI. Conclusions

Quantifying the impact of managed care enrollment on Medicare spending is an important policy exercise. For this reason we must know the extent to which managed care enrollment influences Medicare FFS spending. This paper suggests that the effects are reasonably large. Using IV models that adjust for the endogeneity of HMO penetration changes across counties, we estimate that a one percentage point increase in county-level Medicare HMO penetration is associated with between a one and two percent reduction in individual-level annual Medicare spending. The findings are robust to several sensitivity checks and a number of diagnostic exercises suggest that the instruments are reasonable. The results are considerably stronger than OLS estimates, which is consistent with the notion that HMOs systematically enter areas with rapid cost growth, presumably because there is more opportunity for them to profit if the costs of the competing FFS are rising rapidly.

The findings should be interpreted as applying to the range of managed care penetration influenced by payment policy. Given the substantial magnitude of these findings, we suspect that larger changes in managed care would translate into somewhat smaller effects. Yet given these estimated effects, policy makers might be well advised to undertake policies to encourage greater HMO penetration because some of the costs of greater payments to plans would be offset by savings in the FFS system, and likely the health care system overall.

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Table 1. Pre and post-BBA growth in mean FFS expenditure related to payment growth.

	Ratio of payment growth post-BBA to payment growth pre-BBA	
	Below Median	Above Median
% growth in mean FFS expenditure (1994-1996)	9.1%	10.3%
% growth in mean FFS expenditure (1998-2001)	25.7%	14.5%

Notes: Payment growth pre-BBA is the percent increase in payment from 1994 to 1996 and payment growth post-BBA is the percent increase in payment from 1998 to 2001. The ratio of these two percentages is computed for each county and the resulting distribution is divided into two groups—those counties below the median of this ratio and those counties above it. These two groups are represented in the columns above. Conceptually, counties below the median are those whose payment rates are slowing, while those above the median are counties whose payment rates are accelerating over time.

Table 2. Selected sample means and standard deviations.

	With county restrictions		Without county restrictions	
	(1)	(2)	(3)	(4)
	Without zeroes (N=58,569)	With zeroes (N=60,067)	Without zeroes (N=61,208)	With zeroes (N=62,790)
Total annual expenditures	$8,031.84 (14,572.83)	$7,832.54 (14,444.37)	$8,066.77 (14,585.48)	$7,863.55 (14,455.93)
Fraction zero expenditure	-----	0.025 (0.156)	-----	0.025 (0.156)
Medicare HMO penetration	9.053 (12.334)	9.035 (12.324)	8.914 (12.245)	8.893 (12.234)
Payment rate	$451.11 (103.99)	$451.10 (104.01)	$449.45 (104.29)	$449.36 (104.37)
Age	76.70 (7.49)	76.63 (7.50)	76.68 (7.51)	76.61 (7.52)
Female	0.586 (0.493)	0.583 (0.493)	0.585 (0.493)	0.582 (0.493)
Health--Excellent	0.150 (0.357)	0.154 (0.361)	0.148 (0.355)	0.152 (0.359)
Health--Very good	0.274 (0.446)	0.275 (0.447)	0.273 (0.446)	0.274 (0.446)
Health--Good	0.322 (0.467)	0.320 (0.467)	0.322 (0.467)	0.320 (0.466)
Health--Fair	0.181 (0.385)	0.179 (0.383)	0.181 (0.385)	0.179 (0.383)
Health--Poor	0.073 (0.260)	0.072 (0.258)	0.074 (0.261)	0.072 (0.259)

Notes: All four samples include only beneficiaries enrolled in traditional fee-for-services Medicare (i.e., they exclude beneficiaries enrolled in an HMO). The first two columns represent the samples we use to generate regression estimates. The second two columns relax our restriction that a county contribute at least fifteen observations over the eight years of data in question. Standard deviations are in parentheses.

Table 3—Estimated effect of Medicare HMO penetration on log annual spending—OLS

	Without Zeroes		With Zeroes	
	(1)	(2)	(1)	(2)
HMO penetration	-0.00220	-0.00260	-0.00202	-0.00225
	(1.15)	(1.72)	(1.09)	(1.54)
	{9.1%}	{9.1%}	{9.0%}	{9.0%}
Age	0.19560	0.08708	0.18527	0.08599
	(9.04)	(4.66)	(8.86)	(4.73)
Age squared	-0.00106	-0.00044	-0.00100	-0.00044
	(7.69)	(3.70)	(7.49)	(3.76)
Female	-0.00966	0.02823	-0.01005	0.02827
	(0.57)	(1.71)	(0.60)	(1.77)
R^2	0.06	0.25	0.45	0.57
N	58,569	58,569	60,067	60,067

Notes: Dependent variable is log of annual spending. Models "without zeroes" drop individuals with zero expenditure, while models "with zeroes" assign one dollar of spending to these individuals and include a dummy variable indicating if an individual has zero expenditure. In addition to county and year fixed effects, model (1) includes controls for age, age squared and gender. Model (2) adds controls for race, income, household size, marital status, general health status, sixteen disease indicators, smoking status, and body mass index. Absolute values of t-ratios in parentheses and mean Medicare HMO penetration in curly brackets. Standard errors adjusted for clustering at the county-level.

Table 4—Estimated effect of Medicare HMO penetration on log annual spending—IV estimates

	Without Zeroes		With Zeroes	
	(1)	(2)	(1)	(2)
Structural equation estimates				
HMO penetration	-0.01649 (3.10) {9.1%}	-0.01377 (3.19) {9.1%}	-0.01595 (3.05) {9.0%}	-0.01322 (3.11) {9.0%}
Hahn-Hausman test	-0.20	-0.39	-0.17	-0.41
Overidentification test	0.31	1.02	0.27	1.07
Hausman test	16.59	12.75	16.65	12.94
First stage estimates				
Payment rate	-0.00137 (7.76)	-0.00137 (7.83)	-0.00137 (7.81)	-0.00137 (7.82)
Payment rate squared	0.0000167 (6.06)	0.0000167 (6.11)	0.0000167 (6.09)	0.0000167 (6.09)
F-statistic	33.84	34.35	34.09	34.22
Partial R^2	0.140	0.140	0.140	0.140
N	58,569	58,569	60,067	60,067

Notes: Dependent variable is log of annual spending. Models "without zeroes" drop individuals with zero expenditure, while models "with zeroes" assign one dollar of spending to these individuals and include a dummy variable indicating if an individual has zero expenditure. In addition to county and year fixed effects, model (1) includes controls for age, age squared and gender. Model (2) adds controls for race, income, household size, marital status, general health status, sixteen disease indicators, smoking status, and body mass index. Absolute values of t-ratios in parentheses and mean Medicare HMO penetration in curly brackets. Overidentification test statistic Hausman test statistic are distributed $\chi^2(1)$ and the Hahn-Hausman test statistic is a t-ratio. With respect to the Hahn-Hausman test, the reciprocals of the "reverse" regression coefficients are, moving from left to right, -0.01676 (t=3.17), -0.01461 (t=3.40), -0.01617 (t=3.11) and -0.01409 (t=3.33). Standard errors adjusted for clustering at the county-level.

Table 5—Estimated effect of Medicare HMO penetration on log annual spending—IV estimates without California and Florida counties.

	Without Zeroes		With Zeroes	
	(1)	(2)	(1)	(2)
Structural equation estimates				
All counties	-0.01649 (3.10) [58,569]	-0.01377 (3.19) [58,569]	-0.01834 (2.53) [60,067]	-0.01419 (2.52) [60,067]
Without CA counties	-0.01445 (2.91) [55,010]	-0.01300 (3.19) [55,010]	-0.01399 (2.86) [56,401]	-0.01248 (3.11) [56,401]
Without FL counties	-0.01404 (2.74) [55,820]	-0.01113 (2.67) [55,820]	-0.01343 (2.69) [57,261]	-0.01061 (2.59) [57,261]
Without CA and FL counties	-0.01179 (2.53) [52,261]	-0.01033 (2.65) [52,261]	-0.01275 (2.48) [53,595]	-0.00984 (2.57) [53,595]

Notes: Dependent variable is log of annual spending. Models "without zeroes" drop individuals with zero expenditure, while models "with zeroes" assign one dollar of spending to these individuals and include a dummy variable indicating if an individual has zero expenditure. In addition to county and year fixed effects, model (1) includes controls for age, age squared and gender. Model (2) adds controls for race, income, household size, marital status, general health status, sixteen disease indicators, smoking status, and body mass index. Absolute values of t-ratios in parentheses and sample sizes in brackets. Standard errors adjusted for clustering at the county-level.

Table 6—Estimated effect of Medicare HMO penetration on health care utilization—IV estimates (without zeroes, full covariates)

Dependent variable	Any	Number	Number>0
Inpatient events	-0.00172	-0.00552	-0.01158
	(1.85)	(2.14)	(1.39)
	{0.220}	{0.390}	{1.709}
	[58,569]	[58,569]	[12,877]
Outpatient events	-0.00178	-0.07490	-0.08201
	(0.88)	(2.56)	(2.48)
	{0.709}	{3.939}	{5.536}
	[58,569]	[58,569]	[41,530]
Med. provider events	-0.00070	-0.24472	-0.23862
	(1.79)	(2.21)	(2.22)
	{0.980}	{24.987}	{25.494}
	[58,569]	[58,569]	[57,410]
Office visits	-0.00049	0.00688	0.00886
	(0.37)	(0.28)	(0.38)
	{0.851}	{5.976}	{7.051}
	[58,569]	[58,569]	[49,848]

Notes: This table presents estimates from twelve separate IV regressions. Columns represent the model specification and rows represent the dependent variable in question. In particular, the column labeled "any" presents models where the dependent variable equals one if the individual has experienced an event indicated by a given row, the column labeled "number" presents models where the dependent variable is the number of relevant events, and the column labeled "number>0" present models where the sample is restricted to individuals with strictly positive events. In addition to county and year fixed effects, models include controls for age, age-squared, gender, race, income, household size, marital status, general health status, sixteen disease indicators, smoking status and body mass index. The sample includes all counties which contribute at least fifteen observations over eight years of data, 1994-2001, inclusive. Absolute values of t-ratios in parentheses, dependent variable mean in curly brackets, and sample sizes in square brackets. Standard errors adjusted for clustering at the county level.

Table 7 —Estimated effect of Medicare HMO penetration on health care utilization—IV estimates (with zeroes, full covariates)

Dependent variable	Any	Number	Number>0
Inpatient events	-0.00165	-0.00538	-0.01158
	(1.78)	(2.12)	(1.39)
	{0.214}	{0.380}	{1.709}
	[60,067]	[60,067]	[12,878]
Outpatient events	-0.00176	-0.07312	-0.08212
	(0.90)	(2.56)	(2.48)
	{0.691}	{3.842}	{5.536}
	[60,067]	[60,067]	[41,534]
Med. provider events	-0.00070	-0.23573	-0.23865
	(1.89)	(2.17)	(2.22)
	{0.956}	{24.367}	{25.493}
	[60,067]	[60,067]	[57,412]
Office visits	-0.00050	0.00688	0.00886
	(0.39)	(0.28)	(0.38)
	{0.830}	{5.828}	{7.051}
	[60,067]	[60,067]	[49,848]

Notes: This table presents estimates from twelve separate IV regressions. Columns represent the model specification and rows represent the dependent variable in question. In particular, the column labeled "any" presents models where the dependent variable equals one if the individual has experienced an event indicated by a given row, the column labeled "number" presents models where the dependent variable is the number of relevant events, and the column labeled "number>0" present models where the sample is restricted to individuals with strictly positive events. In addition to county and year fixed effects, models include controls for age, age-squared, gender, race, income, household size, marital status, general health status, sixteen disease indicators, smoking status, body mass index and a dummy variable that equals one if an individual has zero expenditure. The sample includes all counties which contribute at least fifteen observations over eight years of data, 1994-2001, inclusive. Absolute values of t-ratios in parentheses, dependent variable mean in curly brackets, and sample sizes in square brackets. Standard errors adjusted for clustering at the county level.