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Payment Policy and Competition in the Medicare+Choice Program

HSRE Working Paper 11

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May 15, 2002

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This work was supported by the Centers for Medicare and Medicaid Services, contract No. HCFA-500-92-0014, and by a small grant from Abt Associates. An earlier version was presented at the Academy for Health Services Research and Health Policy, Atlanta Georgia, June 10, 2001.

PAYMENT POLICY AND COMPETITION IN THE MEDICARE+CHOICE PROGRAM

Abstract

Over the last two years, Medicare+Choice plans raised premiums and reduced benefits to an unprecedented degree, arguing that these were unavoidable consequences of inadequate payments. We investigate plan premium and benefit decisions, taking advantage of a natural experiment to separate the influences of payment rates, the intensity of interplan competition, and the underlying cost of providing coverage. We find that the effects of competition are comparable in importance to the effects of payment rates, confirming empirically that it is possible for the Medicare program to improve benefits without increasing spending or shifting additional costs to beneficiaries.

Introduction

The Medicare+Choice program (M+C) currently provides health insurance coverage to 5 million Medicare beneficiaries through privately operated managed care plans (CMS, 2002). In exchange for accepting some limits on utilization and choice of provider, M+C enrollees typically receive more extensive coverage than they would under traditional fee-for-service Medicare. Until recently, a substantial fraction of M+C enrollees received outpatient prescription drug coverage and paid either nothing or a small additional premium for their coverage. Starting in 2000, however, the program began to experience profound changes. Plans began to withdraw from a substantial number of markets, leaving enrollees to search for coverage elsewhere. In January 2001, over 150,000 Medicare beneficiaries previously enrolled in M+C were left with no M+C plans doing business in their counties (HCFA, 2000). In addition to the market withdrawals, plans began to increase premiums and reduce benefits in their remaining markets (Gold, 2001). Throughout this period, plans argued that changes in payment rates brought about by the Balanced Budget Act of 1997 and subsequent legislation combined with rapidly increasing costs to make these decisions unavoidable (AAHP, 2000; Fried and Zeigler, 2000).

In this article we investigate plan behavior with respect to premiums and benefits, ultimately separating the influences of payment rates, the intensity of interplan competition, and the underlying cost of providing health insurance coverage. The relative importance of each of these factors should help to determine the composition of an appropriate policy response to the recent turmoil in the M+C program. Furthermore, it is particularly important to develop a deeper understanding of plan behavior now, as

Congress and the Bush Administration consider alternative methods of providing outpatient prescription drug benefits to Medicare beneficiaries. The M+C program is seen (particularly by the administration) as a model of how Medicare benefits should be modernized in the future.

Although other studies have attempted to describe recent changes in premiums and benefits in M+C plans (Gold, 2001; MedPAC, 2000; GAO, 2000), this paper differs in two ways. First, we use multivariate methods over time to produce more precise results than previous work, and, second, we take advantage of the passage of the Benefits Improvement and Protection Act (BIPA) of 2000, which created a natural experiment.

Since BIPA passed in December of 2000, plans had already established their premium and benefit structures for 2001 in response to expected costs and the payment rates in force prior to BIPA. These levels were reported in the January 2001 Medicare Compare database. With the passage of BIPA, plans were permitted to change their premium and benefit levels to reflect the newly increased payment rates. The new levels were reported in the March 2001 Medicare Compare data. Since expected costs should not have changed substantially between January and March, a comparison of premiums and benefits from these two months should reveal effects of payment rate changes that are almost entirely free of the influence of unobserved intertemporal changes in cost. This is valuable because a principal obstacle to understanding the relationship between payments and plan behavior is the fact that the true costs of benefits are unobservable. It is impossible to evaluate whether payments are adequate to cover costs when costs are unobservable. Similarly, one cannot attribute observed changes in benefits or premiums to observed changes in payment rates or the intensity of competition if costs might be changing at the same time but cannot be observed. At a time when the cost of health care generally and prescription drugs in particular have been escalating rapidly, the fortuitous, last-minute change in payment rates brought about by BIPA allowed us to overcome these problems by revealing how plans responded to changes in payments implemented on an unusually compressed schedule.

Our principal finding is that the effects of competition are comparable in importance to the effects of payment rates. The finding that more intense competition increases benefits and reduces premiums, although predictable from a theoretical standpoint, empirically confirms that it is possible for the Medicare program to increase benefits without increasing spending or shifting additional costs to beneficiaries. Conversely, reduced competition would have the reverse effect. *Thus, the federal government has a strong institutional interest in safeguarding and promoting inter-plan competition in the Medicare+Choice program.*

Data

To measure benefits offered by risk plans, we obtained data from the Centers for Medicare and Medicaid Services' (CMS') Medicare Compare database. To measure urban/rural status, payment rates, and other county characteristics that might be associated with cost of coverage, we combined data from several standard sources, including the 2000 Area Resource File (ARF), CMS' State/County/Plan Files, and county-level average Principal In-Patient Diagnostic Cost Groups (PIP-DCG) risk scores calculated by CMS. In this section we provide details on the construction of the analytic files and the characteristics of the data.

File Construction

Four datasets were constructed, each representing a different point in time. Since M+C plans are generally permitted to change their benefits and premiums only once each year, we constructed three annual data sets, one each for January of 1999, 2000, and 2001. In addition, since plans were permitted to make a special set of changes in response to the passage of BIPA in December of 2000, we created a fourth data set for March 2001. Table 1 lists the record count for each data source used in the construction of each of these four data sets.

Table 2 describes the construction sequence for each of the four data sets, summarizing the number of matched records at each step. As Table 2 indicates, most of the plan-counties identified from Medicare Compare and the Service Area File were successfully merged with county characteristics, enrollment and payment rates, and risk scores.[1] In 2001, the match rate between service areas and the State/County/Plan (enrollment) data was lower than it had been because of the relatively large number of service area reductions and market withdrawals reflected in the State/County/Plan Files for that year.

Characteristics of the Data

The impact of these reductions and withdrawals can be seen in the first row of Table 3 that shows the percentage of Medicare beneficiaries who lived in a county with at least one M+C plan over time. The table includes descriptive results from the four data

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[1] In some cases (about 20% of plan-counties in 1999), plans offered more than one package of benefits in a county. Since the State/County/Plan files contain only one enrollment number for each plan in each county, some assignment rule was necessary. Following Gold (2001), we assigned each plan-county's enrollment to the package of benefits with the lowest premium and (in case of ties) the most generous drug benefits. Our qualitative findings, however, do not depend on this assignment (see note 4).

sets we constructed. The percentage of beneficiaries with access to a M+C plan declined slightly from 1999 to 2000 (68.3% to 68.1%), and then dropped in 2001 (to 62.6%). In general, the data also show that access to M+C plans was highest and declined the least in urban counties and counties with relatively high payment rates. Among plans remaining in the market, rows 2 through 5 of Table 3 show that the generosity of benefits followed a similar pattern. These rows contain the percentages of beneficiaries living in counties where the following benefits were offered: outpatient prescription drug coverage, drug coverage with an annual cap over $800, dental coverage, and coverage for eyeglasses. Data on premiums charged to M+C enrollees are displayed in the last two rows of Table 3, indicating that although the average nonzero premium did not change substantially between January 1999 and March 2001 (from $33.24 to $35.60), the percentage of enrollees in zero premium plans declined dramatically (from 62.3% to 17.2%). This decline was more evenly distributed than the changes in access to plans and benefits, affecting urban and high payment counties as well as rural and low payment ones.

Methods

The simple tabulation of benefits data by payment rate shown in Table 3 supports the plans' contention that higher payments are associated with more generous benefits, and, by extension, low payment growth might have been the cause of reduced benefits and increased premiums in 2000 and 2001. Although this is a simple and intuitively appealing argument, it is possible that other factors played a significant role. Among these factors may have been changes in the underlying cost of providing coverage and changes in the intensity of competition between plans.

To attempt to separate the influences of these potentially conflicting factors, we used a regression framework, limiting attention to data from January 2001 and March 2001. We focused the regression analysis on these two benefit periods because their close temporal proximity allowed us to minimize the potential influence of unobserved cost differences on plans' decisions regarding benefits and premiums. Similar models could be estimated using all four benefit periods, but without an accurate measure of underlying cost the results would suffer from omitted variable bias. Our model can be written as:

$$\text{benefit}_t^{p,c} = b_1 \text{payment}_t^c + b_2 \text{march}_t + b_3 \text{supply}_t^c + b_4 \text{demand}_t^c + b_5 \text{competition}_{t-1}^{p,c} + d^p + e_t^{p,c}$$

where t indexes the benefit period, p is a plan index, and c is a county index; $\text{benefit}_t^{p,c}$ denotes a particular continuous benefit or cost-sharing variable (we analyze seven such variables: premium greater than zero, premium amount, outpatient prescription drug benefit, generic copayment amount, brand-name copayment amount, dental benefit, and physician visit copayment amount); payment_t^c represents the government's base payment rate; march_t is an indicator of the benefit period (0 for January 2001 and 1 for March 2001); $\text{supply}_t^{p,c}$ is a vector of variables thought to affect plans' marginal costs; $\text{demand}_t^{p,c}$ is a similar vector thought to affect demand facing each plan; b_1, b_2, b_3, b_4, b_5 are coefficients to be estimated, d^p denotes a plan-level fixed effect, and $e_t^{p,c}$ is the residual.

Plan-level fixed effects were included in the specification because we suspected that benefit and premium decisions were not made independently at the county level, despite the fact that payment rates varied by county. There are two reasons why plan effects are likely to have been important: first, the administrative complexity of obtaining approval and managing different benefit and premium packages by county would have been burdensome, and, second, it would have been difficult for plans to explain to enrollees why premiums and benefits might be different across seemingly arbitrary county lines.

The vector of supply variables contained elements reflecting variation in input prices, bargaining power, capital intensity, and practice patterns. Permanent geographic variation in input prices was measured by historical per capita Medicare Part A spending (as in Wholey et al., 1995). Bargaining power is thought to vary with the number of physicians per capita (Wholey et al., 1993) and urban/adjacent/rural status (McBride, 1998). HMOs should have stronger bargaining positions in relatively urban counties with high numbers of physicians per capita because under these circumstances it is easier for plans to direct beneficiaries to preferred providers (because there are more providers to choose from and traveling distance is minimal). Plans' marginal costs should also vary with capital intensity, measured in our models by the per capita number of hospital beds in the county. Higher numbers of hospital beds per capita are thought to be associated with higher marginal costs because of the cost of maintaining additional beds (Gaynor and Anderson, 1995) and potentially as a reflection of regional practice patterns (Knickman and Foltz, 1985). Hospital utilization patterns also underlie the effects of PIP-DCG scores in our model because these risk scores rely on inpatient diagnoses and our specifications included historical Medicare Part A per capita spending to control for differences in input prices. Thus, when comparing two counties with the same input prices but different risk scores, the county with a higher risk score should have a practice pattern that relies more heavily on inpatient hospitalizations. Although this differs from the most common interpretation of the PIP-DCG risk score as a measure of average health status at the county level, it is appropriate in a model that also contains per capita Part A spending.

Plan decisions will also be affected by variations in the elasticity of demand for health insurance. Nyman (1998) argued that health insurance is valuable to its consumers primarily because it makes potentially needed procedures affordable. This motive, as well as the desire to avoid risk (Cutler and Zeckhauser, 2000), should vary with personal resources, so the demand vector in our models included per capita county income. In addition, markets where a high fraction of the population is over 65 years old may have more rapid exchanges of information among the elderly and therefore individual plans might face more elastic demand.

Although arguably another component of demand, we chose to highlight competition separately for clarity of presentation. The competition vector included the Herfindahl index (a measure of industry concentration),[2] and a variable reflecting the benefits offered by other plans in the county in the previous period. Higher industry concentration is expected to facilitate collusion, resulting in higher profits (Schmalensee, 1989) and therefore less generous benefits. The second competition variable depended on the model being estimated; for example, in a premium-level regression, it was the average premium charged in the county in 2000. Both this "other benefits" variable and the Herfindahl index were constructed using data from 2000, one year prior to the study period. We employed this time lag primarily because plans' benefit decisions would

[2] The Herfindahl index is defined to be the sum of squared market shares in a particular industry or market. In this case, we use the sum of squared market shares of M+C plans in each county.

have been made in the prior period and filed with CMS before going into effect. Additionally, this specification has the benefit of reducing any potential endogeneity that might have been introduced by including the contemporaneous versions of these variables. It should be noted that by including both the Herfindahl index and variables reflecting other plans' decisions in each model, we estimated the effect of industry concentration holding lagged competitors' decisions constant and *vice versa*.

There are several features of our benefits models that we wish to highlight. First, what we have just described is an ordinary least-squares (OLS) model for each continuous benefit or cost-sharing variable. We also modeled three discrete variables: premium greater than zero, prescription drug coverage, and dental coverage. A probit model was estimated for each of these binary variables. Second, for a model for a given benefit, we did not explicitly include the other benefits because they were simultaneously determined. Nevertheless, we recognized that plans' choices of benefit levels were related to each other and these influences were reflected in the residual terms.[3] Finally, for all models, observations were weighted by the number of enrollees in each plan-county-period so that smaller plans were given less weight and larger plans more weight.[4]

[3] The correlations between residuals could in principle be accounted for in a seemingly unrelated regression (SUR) framework. We have not implemented an SUR framework because the regressions run on different samples (as dictated by patterns of missing data) and have different functional forms (probit and OLS).

[4] Not only is it intuitively appropriate to give smaller plans less weight, this weighting also serves as a correction for possible heteroscedasticity in the OLS models. However, because of uncertainty in the assignment of enrollments (see note 1), we also estimated our models without weights, producing qualitatively similar results.

Results

When the results of all the benefit and premium models are reviewed together (Tables 4, 5, and 6), four broad similarities emerge. First, differences in payment rates continued to be strongly associated with variations in both premiums and benefits. As expected, higher payments corresponded to lower premiums and more generous benefits, and this relationship was statistically significant in all seven models. Second, at least one of the two variables intended to reflect the intensity of interplan competition was statistically significant in each of the seven models. These variables also had the expected effects, with lower competition corresponding to higher premiums and less generous benefits. Third, the county average PIP-DCG risk score had a statistically significant effect in four of the seven models, and the direction of the effect was consistent with our interpretation of the risk score as a measure of the hospital-intensity of practice patterns. The results indicate that plans in counties with higher average PIP-DCG risk scores were less likely to charge a premium and tended to charge lower copayments for physician visits and generic or brand name drugs. This relationship suggests that counties with high risk scores were counties with relatively high inpatient utilization rates and as such they constituted an attractive opportunity for managed care plans to profit if they could substitute other forms of care for inpatient care. Fourth, in four of the seven models, the coefficients on the $march_t$ variable indicated that premiums declined and benefits became more generous between January and March of 2001, even controlling for changes in payment rates. This suggests that plans may have judged the post-BIPA climate to be more promising, leading to renewed efforts to attract enrollees. It should also be noted that the historical per capita Medicare Part A spending variable was statistically significant in four of the seven models, having effects in the expected direction in each case.

Beyond these general findings, a more detailed examination of the individual models permits some comparisons of the importance of competition relative to payment rates. For monthly premiums, the intensity of competition appears to have had stronger effects than changes in payment rates. A 10-percentage-point increase in the Herfindahl index[5] would have increased the probability of charging a premium by 2.2%, while a 10% increase in payment rates (assuming a $500 rate[6]) would have reduced the probability by only 0.5% (Table 4).[7] Furthermore, an additional dollar in payment would have translated to $0.07 in lower premiums while a dollar decline in the lagged average premium of competitors would have corresponded to an own-premium decline of $0.32.

For outpatient prescription drug coverage, Table 5 shows that competition and payment rates had comparable effects on the probability of drug coverage. A 10-percentage-point increase in the Herfindahl index would have reduced the probability by 4.5% and a 10% increase in payment rates (assuming a $500 rate) would have increased the probability by 10%. The effects of competition and payment rates on drug copayments were similarly comparable. An additional dollar of payment would have led to reductions of $0.01 and $0.05 in generic and brand name copayments, respectively,

[5] The Herfindahl index is a nonlinear function of market shares, so a 10-percentage-point change could represent a variety of shifts in market shares. For example, a four-firm market that shifts from being evenly divided (25%, 25%, 25%, 25%) to (10%, 15%, 25%, 50%) would have a 9.5 point increase in the Herfindahl index. A similar increase would result in a two-firm market that shifts from (50%, 50%) to (72%, 28%).

[6] The weighted mean payment rate was $535 with a standard deviation of $70.

[7] These changes constitute 47% and 71% of a standard deviation for the Herfindahl index and the payment rate, respectively. Therefore, this comparison overstates the relative sensitivity of benefits and premiums to payment rates.

while a 10 percentage point increase in the Herfindahl index would have led to $0.24 higher generic copayments and a dollar increase in lagged competitors' copayments would have led to $0.27 more in own brand name copayments.

Competition and payment rates had similar effects on physician visit copayments and the probability of offering dental benefits (Table 6). An additional dollar in payment would have led to a $0.03 reduction in copayments while a 10 percentage point increase in the Herfindahl index or a one dollar increase in the lagged average copayment of competitors both would have led to about a $0.14 increase. The results from the dental coverage probit were perhaps the simplest and most striking. Both payment rates and the lagged decisions of competitors were statistically significant in the model, but while changes in payment rates had very small effects, the presence of another plan in the county that offered dental benefits in the previous period increased the probability of offering dental benefits by 54%.

Discussion and Conclusions

In this paper we took advantage of a natural experiment that occurred when Congress passed the Benefits Improvement and Protection Act in December of 2000. The passage of this law so late in the year resulted in two sets of benefit choices by plans in response to two sets of payment rates that were separated by only a few months. By choosing to focus attention on data from January and March of 2001, we minimized intertemporal variation in the cost of providing coverage to beneficiaries (a quantity that is notoriously hard to measure) while preserving BIPA-induced variation in payment rates, premiums, and benefits. Consequently, these data presented an unusual opportunity to study the premium and benefit decisions of plans in relation to the payment rates and levels of competition they face, without the potentially confounding influence of unobserved changes in cost.

We found that the data support the plans' contention that reduced payment rates led to higher premiums and less generous benefits (AAHP, 2000; Fried and Zeigler, 2000). However, we found that the level of interplan competition was also of substantial importance, and may have been more important for some benefits. This finding was robust in models for seven different dependent variables: premium greater than zero, premium amount, outpatient prescription drug benefit, generic copayment amount, brand-name copayment amount, dental benefit, and physician visit copayment amount. Competition had strong effects whether measured by the Herfindahl index (a measure of industry concentration) or by lagged variables reflecting the decisions of competitors (e.g., lagged average premium for competitors in the county).

These findings have a series of critical policy implications. Most importantly, they support the premise that intensified competition can deliver to beneficiaries more generous benefits at lower premiums with no additional cost to the government. Alternatively formulated, payment rate reductions can be offset by more intense interplan competition. These results indicate, for example, that the addition of one more plan to a market evenly divided among three existing plans would approximately offset the effects of a 10 percent reduction in payment rates. Of course, payment rates and the level of competition are related to each other and simply reducing payment rates also can be expected to reduce competition as marginal plans exit the marketplace. Nevertheless, there are numerous opportunities to intensify competition without changing payment rates. For example, Congress and CMS have intermittently considered setting M+C payment rates through competitive bidding, rather than through administered pricing. Under competitive bidding, rates across the country would be expected to correspond more closely to the cost of providing a basic package of benefits, encouraging plans to enter counties that currently have few Medicare HMOs (Dowd et. al., 2000). The results of this analysis indicate that such a change would be likely to result in better value for Medicare beneficiaries. In a less sweeping example, CMS is currently planning to allow M+C plans to offer premium rebates to beneficiaries, starting in 2003 (Feldman et. al., 2001). Since the current ban on rebates prevents zero premium plans from competing on the basis of price, this change would intensify competition and also can be expected to result in better value for Medicare beneficiaries.

Although our results are quite robust, at least two cautions apply. First, the strength of this analysis comes from its tight focus on a particular period in time, but this is also a weakness. The M+C program was in substantial turmoil during the first few months of 2001 and relationships observed during that time might not be as generalizable as they would be if a longer period of study could have been used. Second, the inability to directly observe the cost of providing coverage makes this type of analysis challenging, even under favorable conditions like those following the passage of BIPA. It will be difficult to confirm these results with future data without a method for observing and measuring this cost.

Acknowledgements

The authors wish to thank Ronald Deacon of CMS for his consistent support for this project. We also wish to thank Robert Coulam, Bryan Dowd, Roger Feldman, Alan White, Teresa Doksum, and seminar participants at the Academy for Health Services Research and Health Policy and the Centers for Medicare and Medicaid Services for their helpful comments. Louise Hadden and Carolyn Robinson contributed expert programming. All remaining errors are our own.

References

American Association of Health Plans. Survey shows inadequate payment and regulatory burden force more health plans out of the Medicare+Choice program. Peter D. Hart Research Associates, Inc. Washington D.C. (June, 2000).

Centers for Medicare and Medicaid Services (CMS). Medicare Managed Care Contract Report, February 1, 2002.

Cutler DM and Zeckhauser RJ. “The Anatomy of Health Insurance,” in *Handbook of Health Economics Volume 1A*, eds. AJ Culyer and JP Newhouse, Amsterdam: Elsevier (2000).

Dowd B, Coulam R, and Feldman R. “A Tale of Four Cities: Medicare Reform and Competitive Pricing,” *Health Affairs*, 19(5) (September/October, 2000) 9-29.

Feldman R, Dowd B, Coulam R, Nichols L, and Mutti A. “Premium Rebates and the Quiet Consensus on Market Reform for Medicare,” report submitted by Abt Associates Inc. to the Centers for Medicare and Medicaid Services in conjunction with contract number 500-92-0014 (August 30, 2001).

Fried BM and Ziegler J. *The Medicare+Choice Program: Is it Code Blue?* Health Insurance Association of America: Washington D.C. (June 8, 2000).

Gaynor M and Anderson GF. "Uncertain Demand, the Structure of Hospital Costs, and the Cost of Empty Hospital Beds," *Journal of Health Economics*, 14 (1995) 291-317.

Gold M. "Medicare+Choice: An Interim Report Card," *Health Affairs*, 20(4) (July/August, 2001) 120-138..

Health Care Financing Administration. "M+C Enrollees Affected by Non-Renewals and Service Area Reductions for 2001," Health Care Financing Administration: Baltimore, Maryland (July 21, 2000).

Knickman JR and Folz AM. "A Statistical Analysis of Reasons for East-West Differences in Hospital Use," *Inquiry,* 22(1) (Spring 1985) 45-58.

McBride TD. "Disparities in Access to Medicare Managed Care Plans and their Benefits," *Health Affairs* 17(6) (November/December 1998) 170-180.

Medicare Payment Advisory Commission (MedPAC). "Medicare+Choice: Trends Since the Balanced Budget Act," In *Report to the Congress: Medicare Payment Policy.* Washington D.C.: MedPAC, (March 2000) 117-126.

Nyman JA. "The Value of Health Insurance: The Access Motive," *Journal of Health Economics* 18 (April 1999) 141-152.

Schmalensee R. "Inter-Industry Studies of Structure and Performance," in *Handbook of Industrial Organization*, ed. R. Schmalensee and R. Willig, Amsterdam: Elsevier (1989).

U.S. General Accounting Office. *Medicare+Choice: Plan Withdrawals Indicate Difficulty of Providing Choice While Achieving Savings*, GAO/HEHS-00-183 (Washington: U.S. General Accounting Office, September 2000).

Wholey DR, Feldman R, and Christianson JB. "The Effect of Market Structure on HMO Premiums," *Journal of Health Economics* 14 (May 1995) 81-105.

Wholey DR, Christianson JB, and Sanchez S. "Professional Reorganization: The Effect of Physician and Corporate Interests on the Formation of Health Maintenance Organizations," *American Journal of Sociology* 99 (July 1993) 175-211.

Table 1

Record count by data source and year/month

Data Source	*Year/Month*			
	1999	*2000*	*2001/Jan*	*2001/Mar*
M+C Basic Plan Data	498 plans	319 plans	352 plans	359 plans
Service Area File[(a)]	2,755 plan-counties	2,543 plan-counties	1,801 plan-counties	1,825 plan-counties
State/County/Plan File[(b)]	33,352 plan-counties	33,913 plan-counties	28,650 plan-counties	28,650 plan-counties
Area Resource File	3,081 counties	3,081 counties	3,081 counties	3,081 counties
Risk Scores	3,249 counties	3,249 counties	3,249 counties	3,249 counties

(a) The 1999 and 2000 Service Area Files include all managed care plans while the 2001 files were pre-processed to include only M+C plans. Ultimately, only M+C plans are retained for analysis (see Table 2, step 7) so this represents a different sequencing of an otherwise equivalent data processing procedure.

(a) The State/County/Plan File includes plan-county data for all managed care plans and all counties in which they operate.

Table 2

Matched records in construction sequence by year/month

Step	*Year/Month*			
	1999	*2000*	*2001/Jan*	*2001/Mar*
1. Begin with M+C Basic Plan Data	498 plans	319 plans	352 plans	359 plans
2. Attach Service Area File data	2,745 matched plan-counties	2,031 matched plan-counties	1,785 matched plan-counties	1,809 matched plan-counties
3. Attach State/County/ Plan File data	2,741 matches	2,029 matches	1,161 matches	1,173 matches
4. Attach Area Resource File data	2,737 matches	2,027 matches	1,159 matches	1,171 matches
5. Attach Risk Scores	2,737 matches	2,027 matches	1,159 matches	1,171 matches
6. Append county dummy records	2,737 plan-county records and 3,132 county records	2,027 plan-county records and 3,132 county records	1,159 plan-county records and 3,129 county records	1,171 plan-county records and 3,129 county records
7. Drop non-M+C plans or plans with missing or zero enrollment	1,865 plan-county records and 3,132 county records	1,851 plan-county records and 3,132 county records	1,132 plan-county records and 3,129 county records	1,136 plan-county records and 3,129 county records

Table 3
Availability, Benefits, and Premiums in Medicare Risk Plans by Base Payment Rate and Urban/Rural Status, 1996, 1999, 2000 and 2001

			Payment rate in 1997 dollars			Urban/rural status		
							Nonurban	
	Year	All counties	$300 - $399	$400 - $499	$500 or more	Urban	Adjacent	Nonadjacent
Percent living in the county/GSA of any risk plan	Mar 01	62.6%	--	47.3%	96.7%	78.3%	20.5%	3.6%
	Jan 01	62.6%	--	47.2%	96.7%	78.2%	20.5%	3.6%
	2000	68.1%	27.3%	71.6%	97.1%	83.3%	30.6%	6.8%
	1999	68.3%	22.0%	64.0%	95.8%	83.4%	31.4%	7.5%
Percent in county where the following benefits are offered by any basic plan[1]								
Outpatient prescription drug coverage	Mar 01	45.3%	--	27.9%	84.0%	58.1%	7.4%	1.4%
	Jan 01	44.7%	--	27.3%	83.5%	57.4%	6.8%	1.4%
	2000	52.0%	13.3%	48.8%	89.7%	66.1%	12.2%	2.2%
	1999	61.2%	14.5%	52.7%	93.9%	75.8%	23.0%	5.3%
Outpatient prescription drug coverage over $800 per year	Mar 01	21.3%	--	10.4%	45.3%	27.5%	2.7%	0.0%
	Jan 01	21.1%	--	10.2%	45.3%	27.3%	2.5%	0.0%
	2000	44.0%	8.6%	40.6%	79.2%	56.4%	8.7%	0.2%
	1999	50.6%	5.4%	37.8%	87.8%	64.1%	13.0%	1.9%
Dental coverage	Mar 01	32.0%	--	17.6%	64.2%	41.3%	5.0%	0.1%
	Jan 01	29.5%	--	14.5%	62.8%	38.0%	4.3%	0.0%
	2000	32.6%	2.8%	23.4%	71.8%	42.4%	3.4%	0.1%
	1999	49.0%	6.8%	37.7%	82.9%	62.0%	12.8%	2.1%
Eye coverage, glasses	Mar 01	27.5%	--	13.7%	58.2%	35.7%	3.0%	0.0%
	Jan 01	27.5%	--	13.7%	58.2%	35.7%	3.0%	0.0%
	2000	56.3%	16.8%	54.7%	92.0%	70.2%	18.9%	4.4%
	1999	65.8%	16.4%	61.1%	95.5%	80.8%	27.4%	7.7%
Premiums charged to risk plan enrollees[2]								
Medicare risk-plan enrollees in zero- premium plans	Mar 01	17.2%	--	14.5%	19.3%	17.1%	22.9%	10.1%
	Jan 01	14.8%	--	14.4%	15.2%	14.7%	22.1%	10.1%
	2000	47.9%	30.1%	38.3%	58.2%	48.0%	45.2%	29.1%
	1999	62.3%	18.0%	46.5%	73.1%	63.3%	32.3%	18.0%
Average monthly premium	Mar 01	$35.60	--	$38.82	$31.75	$35.53	$35.60	$52.11
	Jan 01	$37.68	--	$42.62	$31.97	$37.54	$40.76	$53.23
	2000	$31.56	$42.17	$31.57	$23.90	$31.47	$29.57	$48.77
	1999	$33.24	$40.14	$32.57	$30.64	$32.72	$35.68	$50.35

(1) Following Gold (2001), in cases with multiple plan options, the basic plan was defined to be the option with the lowest premium (or most generous prescription drug benefit in case of ties).

(2) In cases with multiple plan options, enrollees were assigned to the basic plan.

Sources: 1999, 2000, January 2001 and March 2001 Medicare Compare database, Area Resource File, CMS Quarterly State/County File.

Table 4
Regression results:[1] monthly premium > $0 probit and monthly premium value regression

Variable	*Monthly Premium>0*		*Monthly Premium Value*
	Coefficient Value (Standard Error)	*Marginal Probability Effect*	*Coefficient Value (Standard Error)*
$payment_t^c$	-0.034*** (0.0044)	-0.7%[2]	-0.065*** (0.015)
$risk^c$	-24*** (6.5)	-49%[3]	-27 (16)
Adjacent	-2.2 (1.2)	-13%[4]	-6.9*** (2.2)
Rural	-[5]	-	-4.4 (5.3)
Income	0.00012*** (0.000032)	-0.0%[2]	-0.00017* (0.000082)
Proportion population 65+	-10** (3.9)	-22%[3]	-44** (15)
Herfindahl index	3.4** (1.1)	7%[3]	2.2 (3.2)
Beds per capita	-135 (102)	-2700%[2]	251 (225)
Part A per capita spending	0.0042*** (0.00065)	0.1%[2]	-0.0031 (0.0023)
Physicians per capita	-328* (714.8)	-6700%[2]	294 (351)
March	-0.51* (0.21)	-11%[2]	-1.2*** (0.27)
Lagged average premium	-[5]	-	0.32*** (0.054)
	N=403, Pseudo R^2 =0.63		N=1104, R^2 =0.82

(1) Although not listed here, the regression also includes a variable for each plan to control for plan-fixed effects as described in the text.
(2) Represents the change in probability due to a one unit increase in this independent variable.
(3) Represents the change in probability due to a 10 percentage point increase in this independent variable which ranges over [0,1].
(4) Represents the change in probability due to a change from 0 to 1 in this binary independent variable.
(5) Dropped due to collinearity.

* indicates significance at the 5% level.
** indicates significance at the 1% level.
*** indicates significance at the 0.1% level.

Table 5
Regression results:[1] outpatient prescription drugs probit and copayment regressions

Variable	*Outpatient prescription drug coverage*		generic copay	brand copay
	Coefficient Value (Standard Error)	*Marginal Probability Effect*	*Coefficient Value (Standard Error)*	*Coefficient Value (Standard Error)*
$payment_t^c$	0.013 (0.0072)	0.2%[2]	-0.012*** (0.0021)	-0.057*** (0.0036)
$risk^c$	-1.4 (12)	-1.8%[3]	-7.7* (3.2)	-36*** (5.9)
Adjacent	1.0 (2.0)	5.8%[4]	-0.33 (0.72)	0.58 (1.3)
Rural	-0.63 (12)	-12%[4]	-0.49 (14)	1.0 (25)
Income	0.00005 (0.00006)	0.0%[2]	0.00 (0.000013)	0.00 (0.000024)
Proportion population 65+	-8.5 (9.3)	-11%[3]	-4.7* (2.4)	4.9 (5.0)
Herfindahl index	-6.2*** (1.5)	-7.6%[3]	1.5* (0.71)	-1.7 (1.4)
Beds per capita	274 (206)	3400%[2]	23 (49)	96 (90)
Part A per capita spending	0.00067 (0.0012)	0.0%[2]	0.0011** (0.00037)	0.0067*** (0.00065)
Physicians per capita	-435 (258)	-5400%[2]	-6.0 (64)	-227 (122)
March	0.47 (0.29)	5.9%[2]	-0.023 (0.088)	-0.29 (0.17)
Lagged other Rx indicator or average copay	-2.1 (1.4)	-7.2%[4]	0.064 (0.057)	0.27*** (0.050)
	N=226 Pseudo R^2 =0.56		N=850 R^2 =0.93	N=769 R^2 =0.95

(1) Although not listed here, regressions also include a variable for each plan to control for plan-fixed effects as described in the text.
(2) Represents the change in probability due to a one unit increase in this independent variable.
(3) Represents the change in probability due to a 10 percentage point increase in this independent variable which ranges over [0,1].
(4) Represents the change in probability due to a change from 0 to 1 in this binary independent variable.

* indicates significance at the 5% level.
** indicates significance at the 1% level.
*** indicates significance at the 0.1% level.

Table 6
Regression results: (1) physician visit copayments and dental benefits

Variable	physician visit copay	dental coverage	
	Coefficient Value *(Standard Error)*	*Coefficient Value* *(Standard Error)*	*Marginal Probability Effect*
payment [c]	-0.026*** (0.0016)	0.12** (0.04)	0.0%(2)
risk [c]	-9.7*** (2.1)	42 (36)	0.0%(3)
Adjacent	-0.43 (0.28)	4.0 (7.4)	0.0%(4)
Rural	-0.30 (0.76)	-(5)	
Income	-0.000078*** (0.0000091)	0.00018 (0.00021)	0.0%(2)
Proportion population 65+	1.5 (1.6)	-17 (17)	-0.0%(3)
Herfindahl index	1.1** (0.39)	-2.5 (8.2)	-0.0%(3)
Beds per capita	-20 (31)	-145 (1466)	-0.02%(2)
Part A per capita spending	0.0018*** (0.00024)	-0.00098 (0.0028)	-0.0%(2)
Physicians per capita	152*** (44)	-1142 (1778)	-0.17%(2)
March	-0.19** (0.063)	2.3* (0.99)	0.0%(2)
Lagged average copay or other dental indicator	0.15*** (0.023)	6.3** (2.3)	57%(4)
	N=1636 R^2 =0.90	N=159 Pseudo R^2 =0.92	

(1) Although not listed here, the regression also includes a variable for each plan to control for plan-fixed effects as described in the text.
(2) Represents the change in probability due to a one unit increase in this independent variable.
(3) Represents the change in probability due to a 10 percentage point increase in this independent variable which ranges over [0,1].
(4) Represents the change in probability due to a change from 0 to 1 in this binary independent variable.
(5) Dropped due to collinearity.

* indicates significance at the 5% level.
** indicates significance at the 1% level.
*** indicates significance at the 0.1% level.