Payment Policy and Competition in the Medicare+Choice Program

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Over the last 2 years, Medicare+Choice (M+C) 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 M+C Program currently provides health insurance coverage to 5 million Medicare beneficiaries through privately operated managed care plans (Centers for Medicare & Medicaid Services, 2002b). 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. However, the program has experienced profound changes as plans have withdrawn from a substantial number of markets, leaving enrollees to search for coverage elsewhere.[1] In January 2001, more than 150,000 Medicare beneficiaries previously enrolled in M+C were left with no M+C plans doing business in their counties (Health Care Financing Administration, 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 (American Association of Health Plans, 2000; Fried and Ziegler, 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

[1] The M+C Program enrolled 6.35 million beneficiaries in December 1999, 6.19 million in January 2000, and 5.6 million in January 2001 (Zarabozo, 2000; Pizer, Frakt, and Doksum, 2001).

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; Medicare Payment Advisory Commission, 2000; U.S. General Accounting Office, 2000), this article 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 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 2 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. (Although the compressed schedule permits a natural experiment with respect to cost, it also may have induced plans to translate less of the payment changes into benefit and premium adjustments than they would have otherwise. Researchers have found that more plans reported using the additional funds from BIPA to stabilize access to providers than to improve the benefit package (U.S. General Accounting Office, 2001; Centers for Medicare & Medicaid Services, 2002a). This included unprecedented deposits into stabilization funds in 2001, which were mostly withdrawn during 2002 (Zarabozo, 2002). If benefit and premium adjustments were muted by the schedule, we would expect our quantitative estimates to be affected accordingly, but qualitative relationships to stay approximately the same.)

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. We acknowledge that competition and spending are related by the fact that lower payments can be expected to induce plan exit, thereby undermining competition. Nevertheless, this research

Table 1

Data Sources and Record Counts, by Year and Month

	January			March
Source	1999	2000	2001	2001
Medicare+Choice Plan Data				
Number of Plans	498	319	352	359
Service Area File[1]				
Number of Plan-Counties	2,755	2,543	1,801	1,825
State/County/Plan File[2]				
Number of Plan-Counties	33,352	33,913	28,650	28,650
Area Resource File				
Number of Counties	3,081	3,081	3,081	3,081
Risk Scores				
Number of Counties	3,249	3,249	3,249	3,249

[1] Includes all managed care plans for 1999 and 2000, while the 2001 files were preprocessed to include only Medicare+Choice (M+C) plans. Ultimately, only M+C plans are retained for analysis (drop non-M+C plans or plans with missing or zero enrollment) so this represents a different sequencing of an otherwise equivalent data processing procedure.

[2] Includes plan-county data for all managed care plans and all counties in which they operate.

SOURCES: Centers for Medicare & Medicaid Services: Data from the Medicare Health Plan Compare database, Area Resource File, county-level principal inpatient diagnostic cost groups scores, and Quarterly State/County/Plan File: January 1999, 2000, 2001, and March 2001.

shows that the Federal Government has a strong institutional interest in safeguarding and promoting interplan competition in the M+C Program, independent of its policy on payment rates.

Data

To measure benefits offered by risk plans, we obtained data from CMS's Medicare Compare database (http://www.Medicare.gov/mphCompare/home.asp). 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's State/County/Plan Files, and county-level average principal inpatient 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 data sets 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 1999, 2000, and 2001. In addition, since plans were permitted to make a special set of changes in response to the passage of BIPA, we created a fourth data set for March 2001. Table 1 shows the record count for each data source used in the construction of each of these four data sets.

The construction sequence for each of the four data sets, summarizing the number of matched records at each step is described in Table 2. 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. In 2001, the match rate between service areas and the State/County/Plan (enrollment) data was

Table 2

Analytic File Construction Sequence: Matched Records, by Year and Month

Sequence	January 1999	January 2000	January 2001	March 2001
Start with M+C Basic Plan Data	498 Plans	319 Plans	352 Plans	359 Plans
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
Attach State/County/Plan File Data	2,741 Matches	2,029 Matches	1,161 Matches	1,173 Matches
Attach Area Resource File Data	2,737 Matches	2,027 Matches	1,159 Matches	1,171 Matches
Attach Risk Scores	2,737 Matches	2,027 Matches	1,159 Matches	1,171 Matches
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
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

NOTE: M+C is Medicare+Choice.

SOURCES: Centers for Medicare & Medicaid Services: Data from the Medicare Compare database, Area Resource File, and Quarterly State/County/File: January 1999, 2000, 2001, and March 2001.

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.

Enrollment data was used to weight observations in our empirical models. In some cases (approximately 20 percent 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. Since models estimated without enrollment weights produced similar results to the weighted versions, our qualitative findings do not depend on this assignment.

Data Characteristics

The impact of these reductions and withdrawals is shown in Table 3, which contains descriptive results from the four data 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 percent), and then dropped in 2001 (62.6 percent). In general, the data show that access to M+C plans was highest and declined the least in urban counties and counties with relatively high payment rates. Also shown are the percentages of beneficiaries living in counties where the following benefits were offered: outpatient prescription drug coverage, drug coverage with an annual cap more than $800, dental coverage, and coverage for eyeglasses. Data on premiums charged to M+C enrollees indicate that although the average non-zero premium did not change substantially between January 1999 and March 2001 (from $33.24 to $35.60), the percentage of enrollees having the option to enroll in zero premium plans declined dramatically (from 62.3 to 17.2 percent). 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.

Table 3

Descriptive Statistics for Medicare+Choice Plans, by Payment Rate and Urban and Rural Status: 1996, 1999, 2000, and 2001

Statistic	Month	Year	All Counties	Payment Rate in 1997 Dollars $300-$399	$400-$499	$500 or More	Urban and Rural Status Urban	Non-Urban Adjacent	Non-Adjacent
				Percent					
Living in the County	March	2001	62.6	NA	47.3	96.7	78.3	20.5	3.6
or GSA of Any Risk	January	2001	62.6	NA	47.2	96.7	78.2	20.5	3.6
Plan	January	2001	68.1	27.3	71.6	97.1	83.3	30.6	6.8
	January	1999	68.3	22.0	64.0	95.8	83.4	31.4	7.5
Basic Plan[1]									
Outpatient	March	2001	45.3	NA	27.9	84.0	58.1	74.0	1.4
Prescription Drug	January	2001	44.7	NA	27.3	83.5	57.4	68.0	1.4
Coverage	January	2000	52.0	13.3	48.8	89.7	66.1	12.2	2.2
	January	1999	61.2	14.5	52.7	93.9	75.8	23.0	5.3
Outpatient	March	2001	21.3	NA	10.4	45.3	27.5	27.0	0.0
Prescription Drug	January	2001	21.1	NA	10.2	45.3	27.3	25.0	0.0
Coverage Over $800	January	2000	44.0	8.6	40.6	79.2	56.4	87.0	0.2
per Year	January	1999	50.6	5.4	37.8	87.8	64.1	13.0	1.9
Dental Coverage	March	2001	32.0	NA	17.6	64.2	41.3	5.0	0.1
	January	2001	29.5	NA	14.5	62.8	38.0	43.0	0.0
	January	2000	32.6	2.8	23.4	71.8	42.4	34.0	0.1
	January	1999	49.0	6.8	37.7	82.9	62.0	12.8	2.1
Eye Coverage,	March	2001	27.5	NA	13.7	58.2	35.7	3.0	0.0
Glasses	January	2001	27.5	NA	13.7	58.2	35.7	3.0	0.0
	January	2000	56.3	16.8	54.7	92.0	70.2	18.9	4.4
	January	1999	65.8	16.4	61.1	95.5	80.8	27.4	7.7
Premiums[2]									
Medicare Risk-Plan	March	2001	17.2	NA	14.5	19.3	17.1	22.9	10.1
Enrollees in Zero-	January	2001	14.8	NA	14.4	15.2	14.7	22.1	10.1
Premium Plans	January	2000	47.9	30.1	38.3	58.2	48.0	45.2	29.1
	January	1999	62.3	18.0	46.5	73.1	63.3	32.3	18.0
				In Dollars					
Average Monthly	March	2001	$35.60	NA	$38.82	$31.75	$35.53	$35.60	$52.11
Premium	January	2001	37.68	NA	42.62	31.97	37.54	40.76	53.23
	January	2000	31.56	$42.17	31.57	23.90	31.47	29.57	48.77
	January	1999	33.24	40.14	32.57	30.64	32.72	35.68	50.35

[1] 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 (Gold, 2001).

[2] In cases with multiple plan options, risk-plan enrollees were assigned to the basic plan.

NOTES: GSA is general service area. NA is not applicable.

SOURCES: Centers for Medicare & Medicaid Services: Data from the Medicare Compare database, Area Resource File, and Quarterly State/County/Plan File: January 1999, 2000, 2001, and March 2001.

Methods

The simple tabulation of benefits data by payment rate (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 follows:

$$benefit_t^{p,c} = \beta_1 payment_t^c + \beta_2 march_t + \beta_3 supply_t^c$$
$$+ \beta_4 demand_t^c + \beta_5 competition_{t-1}^{p,c} + \delta^p + \varepsilon_t^{p,c}$$

where t indexes the benefit period, p is a plan index, and c is a county index; $benefit_t^{p,c}$ denotes a particular continuous benefit or cost-sharing variable (we analyze seven such variables: (1) premium more than zero, (2) premium amount, (3) outpatient prescription drug benefit, (4) generic copayment amount, (5) brand-name copayment amount, (6) dental benefit, and (7) physician visit copayment amount); $payment_t^c$ represents the Federal Government's base payment rate; $march_t$ is an indicator of the benefit period (0 for January 2001 and 1 for March 2001); $supply_t^{p,c}$ is a vector of variables thought to affect plans' marginal costs; $demand_t^{p,c}$ is a similar vector thought to affect demand facing each plan; $\beta_1, \beta_2, \beta_3, \beta_4, \beta_5$, are coefficients to be estimated, δ^p denotes a plan-level fixed effect,[2] and $\varepsilon_t^{p,c}$ 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: (1) the administrative complexity of obtaining approval and managing different benefit and premium packages by county would have been burdensome, and (2) it would have been difficult for plans to explain to enrollees why premiums and benefits might be different across seemingly arbitrary county lines.

[2] The term plan-level fixed effect refers to an unobservable variable that is constant across time, but varies between plans.

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 (Wholey, Christianson, and Sanchez, 1995). Bargaining power is thought to vary with the number of physicians per capita (Wholey, Feldman, and Christianson, 1993) and urban/adjacent/rural status (McBride, 1998). Health maintenance organizations (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 are constructed from demographics and inpatient diagnoses. The conventional interpretation of the PIP-DCG risk score is as a measure of average health status at the county level that can be used to make a prediction of total Medicare spending. Since our specifications included historical Medicare Part A per capita spending to control for differences in input prices, the remaining effect of the PIP-DCG risk score in our model comes from its reliance on inpatient hospital utilization. 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, it is appropriate in a model that also contains per capita Medicare Part A spending.

Plan decisions will also be affected by variations in the elasticity of demand for health insurance. Nyman (1999) 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 age 65 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[3] (Schmalensee, 1989) 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 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. (Endogeneity [also known as simultaneity] could arise here because plans' decisions regarding benefits could simultaneously influence each other. Therefore, a model that featured contemporaneous benefit decisions on both the left- and right-hand sides could produce biased results. Because current benefit decisions by one plan cannot influence past decisions by others, the lagged specification does not suffer from this problem.) 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: (1) premium more than zero, (2) prescription drug coverage, and (3) 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. (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

[3] The Herfindahl index (a measure of industry concentration) 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.

dictated by patterns of missing data] and have different functional forms [probit and OLS].) 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. (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. Note that, because of uncertainty in the assignment of enrollments previously discussed, 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 (Table 4), 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 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 (results not shown).

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 and changes in payment rates appear to have had effects of similar magnitude. A 10-percentage-point increase in the Herfindahl index[4] would have increased the probability of charging a premium by 7 percent, while a 10-percent increase in payment rates (assuming a $500 rate[5]) would have reduced the probability by 35 percent.[6] 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.

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[4] For example, a four-firm market that shifts from being evenly divided (25, 25, 25, and 25 percent) to (10, 15, 25, and 50 percent) would have a 9.5-point increase in the Herfindahl index.

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

[6] These changes constitute 47 and 71 percent 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.

Table 4
Effects of Payment and Competition Variables on Premiums and Benefits

Variable	Monthly Premium Greater than 0 — Coefficient Value	Monthly Premium Greater than 0 — Marginal Probability Effect Percent	Premium Dollars — Coefficient Value	Outpatient Drug Coverage — Coefficient Value	Outpatient Drug Coverage — Marginal Probability Effect Percent	Copay — Generic — Coefficient Value	Copay — Brand — Coefficient Value	Copay — Physician — Coefficient Value	Dental Coverage — Coefficient Value	Dental Coverage — Marginal Probability Effect Percent
payment	***-0.034 (0.0044)	[1]-0.7	***-0.065 (0.015)	0.013 (0.0072)	[1]0.2	***-0.012 (0.0021)	***-0.057 (0.0036)	***-0.026 (0.0016)	**0.12 (0.04)	[1]0.0
risk	***-24 (6.5)	[2]-49	-27 (16)	-1.4 (12)	[2]-1.8	*-7.7 (3.2)	***-36 (5.9)	***-9.7 (2.1)	42 (36)	[2]0.0
lagged	**3.4 (1.1)	[2]7	2.2 (3.2)	***-6.2 (1.5)	[2]-7.6	*1.5 (0.71)	-1.7 (1.4)	**1.1 (0.39)	-2.5 (8.2)	[2]-0.0
Herfindahl Index	*-0.51 (0.21)	—	***-1.2 (0.27)	0.47 (0.29)	—	-0.023 (0.088)	-0.29 (0.17)	**-0.19 (0.063)	*2.3 (0.99)	—
march		[3]-11	***0.32 (0.054)	-2.1 (1.4)	[3]5.9	0.064 (0.057)	***0.27 (0.050)	***0.15 (0.023)	**6.3 (2.3)	[3]0.0
lagged		—			[3]-7.2					[3]57.0
other [5]	[4]	—			—					—
Number of Observations	403		1,104	226		850	769	1,636	159	
Goodness of Fit	Pseudo R^2=0.63		R^2=0.82	Pseudo R^2=0.56		R^2=0.93	R^2=0.95	R^2=0.90	Pseudo R^2=0.92	

*** Significance at the 0.001 level.

** Significance at the 0.01 level.

* Significance at the 0.05 level.

[1] Represents the change in probability due to a 1-unit increase in this independent variable.

[2] Represents the change in probability due to a 10-percentage point increase in this independent variable which ranges over (0,1).

[3] Represents the change in probability due to a change from 0 to 1 in this binary independent variable.

[4] Dropped due to collinearity.

[5] Other refers to a different variable in each model. Respectively, these were: average premium, indicator of drug coverage, average copay, or indicator of dental coverage, all for other plans in the county, all lagged one period.

NOTES: Numbers in parentheses are standard errors. Complete regression results are available on request from the authors. Pseudo R^2=1-L1/L2, where L1 and L2 are the log likelihoods of a constant only model and the full model, respectively.

SOURCES: Centers for Medicare & Medicaid Services: Data from the Medicare Health Plan Compare database, Area Resource File, county-level principal in-patient diagnostic cost groups scores, and Quarterly State/County/Plan File: January 1999, 2000, 2001, and March 2001.

For outpatient prescription drug coverage, Table 4 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 7.6 percent and a 10-percent increase in payment rates (assuming a $500 rate) would have increased the probability by 10 percent. 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 $0.06 in generic and brand name copayments, respectively, while a 10 percentage point increase in the Herfindahl index would have led to $0.15 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. 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 and a $1.00 increase in the lagged average copayment of competitors would have led to increases of $0.11 and $0.15, respectively. The results from the dental coverage probit were the simplest and perhaps 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 57 percent.

DISCUSSION AND CONCLUSIONS

In this article we took advantage of a natural experiment that occurred when Congress passed BIPA in December 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 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 (American Association of Health Plans, 2000; Fried and Ziegler, 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 the seven different dependent variables previously mentioned. Competition had strong effects whether measured by the Herfindahl index 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 Federal 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, opportunities exist to intensify competition without changing average payment rates.

The most prominent of these is probably the frequently stalled effort to set M+C payment rates through competitive pricing rather than through administered pricing. Under recent variants of competitive pricing proposals, plans would submit bids to provide a standard benefit package. Based on the information in these bids, CMS would set the payment rate, and base premiums for each plan would be determined by the difference between the rate and the bid. Plans would be free to offer supplemental benefits and charge additional premium if they so chose. This process would be expected to produce rates that need not differ in the aggregate from current rates, but would correspond more closely at the county level to the cost of providing a basic package of benefits, thereby encouraging plans to enter counties that currently have few Medicare HMOs. CMS has repeatedly attempted to demonstrate competitive pricing in sites where payment rates are relatively high (and optional benefits are relatively rich), but each time has met with successful opposition spearheaded by local plans and members of Congress (Dowd, Coulam, and Feldman, 2000). Despite this difficult history, the concept remains attractive enough to have been included in some of the major Medicare reform bills before Congress during 2002 (Johnson, 2002).

In a less sweeping example, BIPA authorized M+C plans to offer premium rebates to beneficiaries, starting in 2003. This will intensify competition because the current ban on rebates prevents zero premium plans from competing on the basis of price, forcing them to compete by offering optional benefits instead. Premium rebates would constitute a new, highly visible dimension for competition between these plans. Under the new law, plans will be permitted to elect payment reductions of up to 125 percent of the Medicare Part B premium, with 80 percent of any reduction (up to 100 percent of the Part B premium) distributed to the enrollee and the remainder to the Federal Government. Unfortunately, the fact that only 80 percent of the benefit flows to enrollees may make these products relatively unattractive, thus defeating the purpose of the rebate program (Feldman et. al., 2001). The results presented in this article indicate that either competitive pricing or premium rebates would be likely to result in better value for Medicare beneficiaries and the Federal Government.

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 experiencing unprecedented plan withdrawals and benefit changes 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.

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