The Effect of Health Plan Characteristics on Medicare+Choice Enrollment

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Objective. To provide national estimates of the effect of out-of-pocket premiums and benefits on Medicare beneficiaries' choice among managed care health plans.

Data Sources/Study Setting. The data represent the population of all Medicare+Choice (M+C) plans offered to Medicare beneficiaries in the United States in 1999.

Study Design. The dependent variable is the log of the ratio of the market share of the j th health plan to the lowest cost plan in the beneficiary's county of residence. The explanatory variables are measures of premiums and benefits in the j th health plan relative to the premiums and benefits in the lowest cost plan.

Data Collection Methods. The data are from the 1999 Medicare Compare database, and M+C enrollment data from the Centers for Medicare and Medicaid Services (CMS).

Principal Findings. A \$10 increase in an M+C plan's out-of-pocket premium, relative to its competitors, is associated with a decrease of four percentage points in the j th plan's market share (i.e., from 25 to 21 percent), holding the premiums of competing plans constant.

Conclusions. Although our price elasticity estimates are low, the market share losses associated with small changes in a health plan's premium, relative to its competitors, may be sufficient to discipline premiums in a competitive market. Bidding behavior by plans in the Medicare Competitive Pricing Demonstration supports this conclusion.

Key Words. Medicare, Medicare+Choice, HMOs, health plan choice, price elasticity

The current Medicare program features a government-administered fee-for-service (FFS) health plan and private health plans under contract to the federal government. The latter part of the program is referred to as "Medicare+Choice," or M+C, and private health plans are referred to as M+C plans.¹ An ongoing controversy concerns the proper role of M+C plans in the Medicare program and the best way to contract with M+C plans. The best way to set the capitation rates paid by the federal government to M+C plans is an issue of particular interest.

The discussion over reform of the Medicare program has identified two broad approaches to paying M+C plans. We refer to the current approach as

“administered pricing.” Congress sets payment rates without regard to the M+C plans’ cost of providing any particular benefit package. The alternative payment approach is referred to as “competitive pricing.” The objective of competitive pricing is to reduce the government’s cost of Medicare benefits to the cost of a standard benefit package provided by an efficient health plan. Bids for the standard benefit package would be submitted by M+C plans, and perhaps by traditional FFS Medicare, as well. The bids would be collected in each market area (e.g., county or metropolitan statistical area), and the government’s “premium contribution” would be set at some reference point, such as the lowest bid, the median bid, or the enrollment weighted average bid.

The success of competitive pricing depends on Medicare beneficiaries being sufficiently sensitive to differences in out-of-pocket premiums across plans to discipline the health plan market. In fact, if the current administrative pricing system for setting payments to M+C plans was replaced with a competitive pricing system, the *only* factor disciplining prices in the market (and government expenditures, as well) would be the price sensitivity of demand for enrollment in M+C plans.

The existing studies of the price-elasticity of health plan choice among Medicare beneficiaries either are dated or confined to a small geographic area. The purpose of this study is to provide an estimate of the price elasticity of health plan choice among the elderly in the M+C program. We also present estimates of the effect of benefits on health plan choice.

BACKGROUND

Medicare is a federal health insurance program for people aged 65 or older, certain people with disabilities who are under 65, and people of any age who have permanent kidney failure. The program has two parts: Hospital Insurance (also called Part A) and Medical Insurance (also called Part B). Most people 65 or older are eligible for Part A based on their own or their spouse’s previous employment. Part A is financed by a payroll tax on current

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workers. Part B is a voluntary program, financed partly by a monthly premium (set at \$50 in 2001), but primarily from general federal tax revenues.

Medicare is the largest U.S. health insurance program, in terms of both enrollment and expenditures. In 1998, there were 38.8 million Medicare beneficiaries, an increase of 25 percent over the 1990 enrollment. The program spent an estimated \$230 billion on covered services in 1999, up from \$111 billion in 1990. Medicare spending has been growing rapidly, especially for hospital outpatient medical and surgical services under Part B and home health care under Parts A and B.

Medicare beneficiaries are legally entitled to a minimum level of insurance benefits, regardless of where they live. For example, Part A of the Medicare program pays 100 percent of the cost of hospital care (subject to an upper limit), after the beneficiary has paid a deductible (\$776 per admission in 2000). Since the early 1980s, Medicare beneficiaries have had a choice of two systems for delivering those benefits: traditional FFS Medicare and the M+C system of private health plans. Until 1997, the monthly payment for risk-bearing plans was tied directly to the cost of the entitlement benefit package in FFS Medicare in the beneficiary's county. The Balanced Budget Act of 1997 (BBA) (Public Law 105-33) altered the direct link between FFS costs and payment to M+C plans. Very high and very low-payment areas were compressed toward the national average. However, while year-to-year increase factors are now based on methods set forth in the 1997 BBA, rather than FFS costs per se,² the national variation in payments still reflect historic FFS costs to a great extent.

The current payment system has resulted in large disparities in payments to M+C plans because FFS Medicare costs vary widely among counties. In 2000, for example, M+C plans in Dade County, Florida (Miami), received \$794.02 per month for each "standard" beneficiary who enrolled, while plans in Hennepin County, Minnesota (Minneapolis), received only \$457.66 per month. Wide variance around the national average rate of \$504.96 per month occurs between other counties, as well.

Medicare+Choice plans can offer additional benefits beyond the entitlement, such as outpatient prescription drugs, and charge an out-of-pocket premium for such benefits. Plans in high-payment areas offer more benefit enhancements beyond the entitlement than plans in low-payment areas and are less likely to charge an out-of-pocket premium (McBride 1998; Pizer, Frakt, and Doksum 2001). The fact that generous benefits are often "free" to enrollees in high-payment areas implies that there is less variation in the cost of the entitlement benefit package among M+C plans nationwide than in the level of payment to plans.

Dowd and Feldman (1996) have suggested that competition among M+C plans over benefits may be “inefficient,” in the sense that some of the extra benefits offered by plans in high-payment areas are not worth as much to beneficiaries as they cost the health plan to produce. The theory of inefficient benefits has some empirical support. Annual payment increases to M+C plans in high payment areas have been fixed at 2 percent³ since the 1997 BBA legislation. In many cases, plans in those areas have reduced the level of benefits offered to enrollees in their zero-premium packages. Dental coverage, outpatient rehabilitation coverage, eyeglasses coverage, and routine eye exam coverage were the benefits with the largest cuts in high payment areas. Other benefit cuts included: coverage of chiropractic care, hearing exams, outpatient hospital visits, prescription drugs, laboratory tests, outpatient mental health, and group therapy mental health services. If beneficiaries were willing to pay at least as much as those benefits cost, plans should have held benefits constant and begun charging an out-of-pocket premium, as the real value of payments decreased year to year.⁴

The problem of inefficient benefits could be addressed by establishing a standard benefit package and having health plans compete with each other on the basis of price, rather than additional benefits.⁵ That approach underlies Dowd, Feldman, and Christianson’s (1996) proposal for competitive pricing in the Medicare program. It also was the model for a series of Medicare Competitive Pricing Demonstrations, two of which were first mandated, then subsequently blocked by Congress (Dowd, Coulam, and Feldman 2000).

In these demonstration projects, M+C plans were to bid on a standard benefit package that reflected the market norm benefit package among M+C plans at the time of the demonstration. Medicare+Choice plans were free to offer any *additional* benefits of their choosing at any premium the market would bear. The demonstration was designed to test whether beneficiary sensitivity to out-of-pocket premiums was sufficient to “discipline” the market and provide incentives for plans to submit low prices, thereby lowering the government’s cost. Currently, there is a lack of information about the responsiveness of enrollment to either out-of-pocket premiums or benefits, which makes it difficult to predict the outcome of competitive pricing arrangements.

PRIOR RESEARCH

Two studies have examined the price elasticity of health plan choice by Medicare enrollees. Feldman et al. (1993) analyzed data from 1989 that

reflected the early experience of the Medicare HMO program (now called Medicare+Choice). Feldman et al.'s primary interest was in estimating the demand equation for HMO enrollment. They found that demand was responsive to out-of-pocket premiums, with an elasticity of approximately -2 (a 1 percent increase in premiums would cause a 2 percent decrease in enrollment). They also were interested in the responsiveness of enrollment to the supplementary benefits offered by HMOs, but their measure of supplementary benefits was very rudimentary—a simple sum of 12 dummy variables representing different types of optional benefits (e.g., coverage of prescription drugs, eye examinations, and mental health services). Their estimates failed to show that the demand for HMO enrollment was responsive to this crude sum of supplementary benefits. This finding is counterintuitive, because beneficiaries should receive some utility from extra benefits, and M+C plans should not offer supplementary benefits that are not valued by beneficiaries.

Feldman et al. (1993) also estimated the relationship between the government payment rate, the out-of-pocket premium, and the level of supplementary benefits. They found that a \$1 increase in the payment rate was associated with a decrease in the supplementary premium of about \$0.26. This result suggests that payments to HMOs are more closely tied to their revenues than to their costs.

Feldman et al.'s study represented the early experience of the Medicare HMO program. In 1989, only 126 HMOs had signed risk-bearing contracts with the Centers for Medicare and Medicaid Services (CMS), and 29 of these had not yet started to enroll Medicare beneficiaries. Feldman's sample represented slightly more than one million Medicare HMO enrollees. By December 1999, M+C enrollment had grown to 6.3 million beneficiaries in 308 risk-bearing contracts.

The index of supplementary benefits in Feldman's study did not reflect either the HMO's cost of providing benefits or the value of different benefits to consumers. Subsequent research using beneficiary focus groups emphasized that by the late 1990s, prescription drug coverage was the most important optional benefit offered by M+C plans (Low and Associates 1999). In view of these suggestive findings, coding drug coverage with the same weight as all other optional benefits may have hidden the importance of prescription drugs as an incentive for beneficiaries to enroll in M+C plans.

Buchmueller (2000) examined the relationship between changes in out-of-pocket premiums from 1993 to 1994 and plan switching by University of California retirees. The UC retirees have a choice of health plans, including several HMOs and a fee-for-service plan that supplements basic Medicare.

Prior to 1994, none of the plans required an employee contribution. Because of a reduction in the University's contribution, the FFS plan and several HMOs began to charge monthly out-of-pocket premiums in 1994. The out-of-pocket premium for the FFS plan increased again in 1995 and 1996. Benefits and other plan features remained constant. The results from the HMO-only sample indicate an increase in switching from a baseline rate of 1.4 percent to 10.8 percent for a \$50 premium increase. Active University of California enrollees were much more likely to switch plans for smaller premium increases than retirees.

Buchmueller also used data from 1993–1996 to estimate the effect of price on the demand for FFS coverage versus any HMO. The HMOs in his data all offered similar benefits at zero or small premiums. The estimated price elasticity of demand for FFS was -0.16 . According to Buchmueller's estimates, a \$10 increase in monthly premium reduced the FFS market share by 1.3 percentage points. He found that switching nearly doubled from a baseline rate of 2.3 percent at a constant premium, to 4.4 percent when the relative premium of the FFS plan increased to \$50 per month.

Buchmueller's study reflects more recent market conditions than the study of Feldman et al., but it is limited to beneficiaries from one employer, in one part of the United States, and a limited number of health plans. Our study will provide estimates based on the full national population of M+C plans and their enrollees.

MODEL AND ESTIMATION METHODS

Our model of health plan choice assumes that beneficiaries choose an M+C health plan by comparing the expected utility from each plan in their "choice set" and picking the plan with the highest expected utility. Utility is derived from health plan characteristics such as breadth and depth of coverage and the premium. As in earlier studies by Feldman et al. (1989) and Dowd and Feldman (1994/95), the premium variable in this study is the premium paid out-of-pocket by beneficiaries.⁶

The beneficiary's "choice set" is defined as the M+C plans that have declared the beneficiary's county of residence to be within their official service area. Beneficiaries enrolled in health plans that do not serve their county of residence are excluded from the study.

Fee-for-service Medicare is excluded from the beneficiaries' choice sets in this study. Fee-for-service Medicare differs from private M+C plans in

important ways that we do not measure. For example, FFS Medicare is available nationally, and the beneficiary can obtain services from any Medicare-eligible provider who is willing to treat them. Proper analysis of these unobserved differences would require the use of a nested logit model in our demand analysis. The nested logit model is not implemented easily in aggregate data of the type used in this study. Atherly, Dowd, and Feldman (2001) have examined the nesting problem in a smaller sample from the Medicare Current Beneficiary Survey and found that approximately 90 percent of the price elasticity of M+C plan choice in those data is attributable to plan switching within the M+C sector, while 10 percent is attributable to switching between the M+C sector and FFS Medicare. Thus, it appears that limiting our analysis to the M+C sector captures the most important source of price elasticity. We also note that there is ongoing policy interest in Medicare reform that includes only M+C plans (Dowd, Coulam, and Feldman 2000).

Our choice model is a grouped-data version of McFadden's (1973) conditional logit model. The probability that the j th health plan is chosen from a set of M different health plans offered to beneficiaries is:

$$P_j = \frac{e^{X_j\beta + u_j}}{\sum_{m=1}^M e^{X_m\beta + u_m}} \quad (1)$$

The coefficients of interest in the demand equation (equation 1) represent the effects of the j th plan's own characteristics (e.g., its types and levels of coverage and its out-of-pocket premium) on the probability that a beneficiary will choose the j th plan. However, the coefficients are estimated by creating new variables that represent the j th plan's premium relative to the premium of a k th reference plan in each county. From equation (1), the ratio of P_j to P_k is:

$$\frac{P_j}{P_k} = e^{(X_j - X_k)\beta + (u_j - u_k)} \quad (2)$$

Notice that in equation 2, the choice set consisting of all the M+C plans serving the beneficiary's county differences away, leaving a pair-wise comparison between the j th plan and the reference plan. This pair-wise comparison becomes the unit of observation in our study. In theory, any plan could be the reference plan but, as in our previous work (Dowd and Feldman 1994/1995), we chose the lowest-priced plan as the reference plan, and designate this plan with the subscript L . If more than one plan charged the same lowest price, we chose the reference plan randomly from among the

lowest-priced plans. The effect of choosing the reference plan at random is discussed further in the results section.

By equating P_j to the proportion of a sample of beneficiaries in a county who choose the j th plan, P_j becomes the market share of the j th plan in the county (MS_j). Next, we calculate the ratio of MS_j to the share of the reference plan (MS_L) in each county. Allowing for stochastic error in the determination of the market share ratio, we arrive at the estimated equation:

$$\ln \frac{MS_j}{MS_L} = (X_j - X_L)\beta + u^* \quad (3)$$

where $u^* = \text{mean}(u_j - u_L)$.

We refer to the new variable $(X_j - X_L)$ as the *relative* out-of-pocket premium (RPREM). For example, if the out-of-pocket premium for the j th plan is \$30 per month and the out-of-pocket premium for the lowest-priced plan in that county is \$20 per month, the relative out-of-pocket premium for this observation would be \$10. The estimated coefficient of the relative premium in equation 3 is also the coefficient of the j th plan's own premium in equation 1. We will use this coefficient to calculate the premium elasticity of demand for the j th plan following estimation of equation 3.

The low-cost reference plans are excluded from the analysis because all of the explanatory variables, which are expressed relative to the low-cost plan, are zero for the reference plans. When the lowest premium is shared by two or more plans, RPREM is zero for *all* the lowest-priced plans. In fact, about three-quarters of the observations in our data have zero values of RPREM. This is because most M+C plans are found in counties with high government payments and as a result, do not charge an out-of-pocket premium. We investigate the sensitivity of our results to exclusion of plan-pairs with zero values of RPREM and report the findings in the results section.

DATA

Our analysis is based on counties, because the service areas of M+C plans, and thus the beneficiary's choice set, is generally defined at the county level. As shown in equation 3, the unit of observation in our analysis is a pair of observations at the "plan-county" level—the j th health plan's enrollment and benefits and premium relative to the (lowest cost) reference plan in the county.⁷ There were 319 M+C plans in the data. Thirty plans had no benefit data and were dropped. The service areas declared by the remaining 289 plans

produced 2,191 plan-county observations. The CMS does not report enrollment for plans with less than 10 enrollees, and there were 110 observations with zero enrollment in our data. Forty-one of those observations were coded zero, because the plan had announced that they would not be available in the county until the following year. The remaining 69 plan-counties with zero reported enrollment were assigned 10 enrollees because the HMO had enrollees in other counties in the metropolitan statistical area. After eliminating the plan-county pairs corresponding to the lowest cost plan in each county (because the right-hand side variables in equation 3 are zero), our final dataset consisted of 1,259 plan-county pairs.⁸

Data on benefits and premiums were obtained from the 1999 “Medicare Compare” section of the CMS “medicare.gov” web site. Data on enrollment came from the March 1999 quarterly enrollment data posted on the “cms.gov” web site within the “Market Penetration—Quarterly State/County/Plan Data Files” section.⁹

The benefit data from the web site were coded into meaningful categories (e.g., “covered at all” versus “amount of coverage”). Binary (0, 1) indicator variables were created for observations with missing variable values. Mammography coverage, pap smear coverage, and hepatitis vaccine coverage were highly correlated, so we created an additive index in which each type of coverage was scored 1 for 100 percent coverage, and 0 for less than 100 percent coverage. Details of other benefit coding algorithms are available from the authors. Some benefits were dropped because they did not have adequate variance to study their effects on enrollment. In all cases, we used the health plan’s lowest priced benefit package to characterize its “product.”

RESULTS

Table 1 shows the definitions of the variables used in the analysis and Table 2 shows their means and standard deviations. It is important to remember that all health plan characteristics (premiums and benefits) are expressed as differences in the value for the j th plan minus the value for the reference plan in that county. Thus, a positive value of an explanatory variable means that the j th (higher priced) plan had a larger value of the variable than the (lowest cost) reference plan. In general, we would expect positive means for the *relative* benefit variables, because more benefits typically will be associated with higher premiums, holding other factors constant. However, Table 2 shows that higher priced plans tend to have *higher* point-of-purchase cost sharing

Table 1: Variable Definitions

<i>Variable Name</i>	<i>Definition</i>
Log of the market share ratio	The log of $[MS_j/MS_k]$ where MS_j is the market share for the j th health plan in a county and MS_k is the market share for the reference health plan in a county.
Premium	Out-of-pocket premium
Chiropractor	= 1 if chiropractic services are covered beyond basic Medicare benefit and = 0 if chiropractic services are not covered beyond basic Medicare benefit
Dental	= 1 if some type of preventive dental exam coverage is provided and = 0 if no type of coverage (the beneficiary pays you pay 100%)
Eyeglasses	= 1 if there is some type of coverage for eyeglasses/contact lenses and = 0 if there is no coverage (the beneficiary pays you pay 100%)
Preventive Care	= the sum of mammography coverage, pap smear coverage, and hepatitis vaccine coverage where for each, 1 = 100% covered, 0 = not 100% covered
Laboratory	= 1 if laboratory services are 100% covered and = 0 if laboratory services are not 100% covered
Laboratory Missing	= 1 if "Laboratory" is missing and = 0 if "Laboratory" is not missing
Outpatient Hospital	= 1 if outpatient hospital visits are 100% covered and = 0 if outpatient hospital visit is not 100% covered
Outpatient Hospital Missing	= 1 if "Outpatient Hospital" is missing and = 0 if "Outpatient Hospital" is not missing
Physical Therapy	= 1 if occupational/physical therapy visits are 100% covered and = 0 if occupational/physical therapy visits are not 100% covered
Outpatient Rehab	= 1 if outpatient rehab visits are 100% covered and = 0 if outpatient rehab visits are not 100% covered
Care Outside U.S.	= 1 if services outside of the U.S. are covered and = 0 if services outside of the U.S. are not covered
Brand Name Copayment	Copayment for brand name prescription drugs
Brand Name Copayment Missing	= 1 if "Brand Name Copayment" is missing and = 0 otherwise

Generic Copayment	Copayment for generic prescription drugs
Generic Copayment Not Applicable	= 1 if "Brand Name Copayment" is not applicable (i.e., there was no prescription drug coverage) and = 0 otherwise
Generic Percentage	= 1 if a percentage, rather than a copayment is given for generic copayment, and = 0 otherwise
Drug > \$800	= 1 if the total annual prescription drug benefit is more than \$800 and = 0 if the total prescription drug benefit is \$800 or less
Physical Exam Copay	Copayment for a physical exam
Podiatry	= 1 if podiatry services are covered beyond basic Medicare benefit and = 0 if podiatry services are not covered beyond basic Medicare benefit
Physician Copayment	Copayment for an outpatient physician visit
No SNF Restriction	= 1 if you do not need to be in the hospital for 3 consecutive days before going to skilled nursing care and = 0 if you must be in the hospital for 3 consecutive days before going to skilled nursing care
XRAY	= 1 if x-rays are 100% covered and = 0 if x-rays are not 100% covered
Point of Service	= 1 if the M+C plan offers a point-of-service option
Staff Model	= 1 if the M+C plan is a staff model HMO
Group Model	= 1 if the M+C plan is a group model HMO (IPA model is the omitted category)
Herfindahl Index	The sum of the squared market shares for all M+C plans in the beneficiary's county of residence (based on March 1999 enrollment data)
M+C Risk Score	The average PIP-DCG risk score for all M+C enrollees in the beneficiary's county of residence (1999)
Percent White	The percent of the population in the beneficiary's county of residence that is white (1990 Census data)
Elderly Poverty	The percent of the elderly population in the beneficiary's county of residence that is below the federal poverty limit in the county (1989 Census data)

Note: In the regression equations, the variables chiropractic through Group Model are expressed as differences in the value for the j th plan minus the value for the reference plan.

Table 2: Variable Means and Standard Deviations ($N = 1,259$ Plan-counties)

<i>Variable</i>	<i>Mean</i>	<i>Standard Deviation</i>
Log of the market share ratio	.160	2.213
Premium	8.163	17.124
Chiropractor	-.012	.502
Dental	-.091	.579
Eyeglasses	-.003	.448
Preventive Care	-.006	1.202
Laboratory	.031	.388
Laboratory Missing	-.006	.149
Outpatient Hospital	-.0008	.570
Outpatient Hospital Missing	-.006	.152
Physical Therapy	-.0734	.492
Oupatient Rehab	-.044	.660
Care Outside U.S.	.048	.740
Brand Name Copayment	1.315	9.204
Brand Name Copayment Missing	.004	.254
Generic Copayment	.555	4.888
Generic Copyament Not Applicable	-.067	.420
Generic Percentage	.013	.1779
Drug > \$800	-.004	.511
Physical Exam Copay	1.450	4.040
Podiatry	-.029	.575
Physician Copayment	2.211	3.331
No SNF Restriction	.012	.381
XRAY	-.007	.353
Point of Service	-.104	.493
Staff Model	.048	.301
Group Model	.005	.637

Note: All variables (except log of the market share ratio) are expressed as differences in the value for the j th plan minus the value for the reference plan.

(coinsurance and copayments), and in some cases, higher priced plans cover fewer services. One possibility is that plans that tend to attract higher risk enrollees, due perhaps to intrinsic characteristics such as broader provider networks, attempt to offset their unfavorable selection by offering less generous coverage.

Model 2 in Table 3 shows the results for the demand equation. (We discuss Models 1 and 3 later.) As discussed earlier, when there was a tie for the lowest-priced plan in a county, we chose one of the lowest-priced plans randomly as the reference plan. Choosing a different reference plan has a slight effect on the estimated coefficients. We estimated the demand equation for five different samples defined by different random draws of the lowest cost

Table 3: Log of Relative Market Share Equation Dependent Variable: $\ln$ (market share of j th plan divided by market share of reference plan) Sample Size=1,259 (plan-counties t-statistics in parentheses)

<i>Variable</i>	<i>Model 1</i>	<i>Model 2</i>	<i>Model 3</i>
Constant	.3160** (4.558)	.2858** (3.547)	2.5301* (2.056)
Premium	-.0191** (-5.529)	-.0217** (-5.791)	-.0250** (-6.508)
Chiropractor		.0077 (0.054)	-.0124 (-.088)
Dental		.3362** (2.814)	.2968* (2.474)
Eyeglasses		.2950* (2.167)	.3073* (2.241)
Preventive Care		-.1964** (-2.610)	-.1910** (-2.580)
Laboratory		-.2516 (-.691)	-.2709 (-.737)
Laboratory Missing		.4391 (.490)	.5496 (.600)
Outpatient Hospital		-.5273** (-3.844)	-.5711** (-4.172)
Outpatient Hospital Missing		-.0788 (-.084)	-.3186 (-.330)
Physical Therapy		-.1793 (-1.281)	-.1919 (-1.378)
Outpatient Rehab		-.0291 (-.250)	-.0178 (-.154)
Care Outside U.S.		-.1644 (-1.717)	-.2028* (-2.117)
Brand Name Copayment		.0465** (4.062)	.0469 (4.109)
Brand Name Copayment Missing		.5616 (1.769)	.6640* (2.101)
Generic Copayment		.0411 (1.576)	.0379 (1.475)
Generic Copayment Not Applicable		1.3275** (4.441)	1.3455** (4.555)
Generic Percentage		1.6288** (3.400)	1.8749** (3.884)
Drug > \$800		.3886** (2.923)	.3618** (2.727)
Physical Exam Copay		-.0656** (-3.781)	-.0680** (-3.850)
Podiatry		.3382** (2.769)	.3694** (3.054)
Physician Copayment		.0414 (1.852)	.0452* (1.968)
No SNF Restriction		.3408 (1.922)	.3502* (1.982)
XRAY		.3692 (0.940)	.3939 (1.008)
Point of Service		-.3634* (-2.211)	-.3121 (-1.874)
Staff Model		.6928** (3.888)	.7281 (.023)

continued

Table 3: Continued

<i>Variable</i>	<i>Model 1</i>	<i>Model 2</i>	<i>Model 3</i>
Group Model		– .1536 (– 1.346)	– .1648 (– 1.449)
Herfindahl Index			.9047* (2.185)
M+C Risk Score			– 3.288* (– 2.552)
Percent White			.0041 (.800)
Elderly Poverty			– .0128 (– 1.033)
Adjusted R ²	0.021	0.119	0.129

**Prob $|\beta| = 0 < 0.01$

*Prob $|\beta| = 0 < 0.05$

Note: All standard errors are corrected for heteroskedasticity using White's heteroskedastic-consistent variance-covariance matrix.

plan, when there was a tie. The median value of the estimated premium coefficient, shown for Model 2 in Table 3, is -0.0217 . (The premium coefficients ranged from -0.0194 to -0.0259 across the five samples). The negative premium coefficient indicates that higher relative premiums are associated with lower market share, as expected.

Coverage of eyeglasses and outpatient prescription drugs more than \$800 are associated with significantly higher enrollment. Some coefficients have the wrong sign. More generous coverage of preventive care and hospital outpatient care are associated with significantly *lower* enrollment. Higher copayments for name-brand drugs are associated with significantly *higher* enrollment in the full sample.

We ran Model 2 on the subset of observations that had positive relative premiums ($n = 324$). The estimated relative premium coefficient in that subsample was -0.0257 , slightly larger in absolute value than the estimate based on the full sample.

The Premium Elasticity of Health Plan Choice

The out-of-pocket premium elasticity of health plan choice is defined as the percentage change in the j th M+C plan's market share (MS_j) associated with a one percent change in its own out-of-pocket premium ($PREM_j$). If the premium elasticity is small (i.e., close to zero in absolute value), a plan can raise its out-of-pocket premium without losing many enrollees. The formula for the

premium elasticity (ε) is:

$$\varepsilon = \frac{\partial MS_j}{\partial PREMIUM_j} \left[\frac{PREMIUM_j}{MS_j} \right] \quad (4)$$

The partial derivative of MS_j with respect to $PREMIUM_j$ is:

$$\beta_{RPREM} \times MS_j \times (1 - MS_j) \quad (5)$$

where β_{RPREM} is the coefficient of the j th plan's relative premium from equation 3.¹⁰ Thus the expression for the premium elasticity is:

$$\beta_{RPREM} \times (1 - MS_j) \times PREMIUM_j \quad (6)$$

The elasticity is different for different values of MS_j and $PREMIUM_j$.¹¹ Taken at the mean values of MS_j of 0.25 and $PREMIUM_j$ of \$7.95 (in the full sample), the elasticity estimate is $-0.0217 \times 0.75 \times 7.95 = -0.13$. This means that a one-percent increase in the plan's out-of-pocket premium is associated with a 0.13 percent (*not* percentage point) decrease in the plan's market share. A 10 percent increase in the out-of-pocket premium is associated with a 1.3 percent decrease in market share (i.e., from 25 percent to 24.7 percent). If we use the average out-of-pocket premium among plans charging a premium (\$40.14), the estimated premium elasticity is -0.65 . These estimates are smaller than those reported in the literature for under-65 employees (Scanlon, Chernew, and Lave 1997) and are consistent with Buchmueller's finding of low switching rates among UC retirees.

In most markets, profit maximizing health plans should set price in the elastic part of the demand curve, that is, the estimated elasticity should be less than -1 . However, in the M+C market, the plan's revenue is a combination of the government payment (A) and the plan's out-of-pocket premium ($PREMIUM$). The first-order condition for profit-maximization among premium-charging plans is:¹²

$$PREMIUM(1 + 1/\varepsilon) = (\partial C/\partial Q) - A \quad (7)$$

Where $\partial C/\partial Q$ is the plan's marginal cost. If the government payment exceeds the plan's marginal cost, the right-hand side of equation 8 will be negative, and assuming that premiums are positive (no cash rebates), ε must be less than -1 in absolute value. Thus, a low price elasticity among premium-charging plans is consistent with government payments that exceed the plans' marginal costs.

The Percentage-Point Change in Market Share

Price elasticities are subject to the choice of base premium, for example, the total premium versus the out-of-pocket premium, and are not defined for zero

premiums. These problems can be addressed by calculating the sensitivity of health plan choice to premiums as the *percentage-point* change in market share associated with a \$1 change in the premium (rather than the percentage change in market share for a 1-percent change in the premium). The equation for the percentage-point change in the j th plan's market share as its premium changes by \$1 is the partial derivative of MS_j with respect to $PREM_j$:

$$\beta_{RPREM} \times MS_j \times (1 - MS_j) \quad (8)$$

At the mean value of MS_j of 0.25, a \$1 increase in the j th plan's out-of-pocket premium is associated with a reduction of:

$$-0.0217 \times 0.25 \times 0.75 = -0.004 \quad (9)$$

in the j th plan's market share. Thus, we predict that a \$10 premium increase would produce a decrease of four percentage points in the j th plan's market share (i.e., from 25 percent to 21 percent).

The Probability of Switching

The effect of premiums on the probability of plan switching also can be derived from our model. The probability of plan switching associated with a change in the premium is the change in the plan's market share due to the change in premium, divided by the original market share:¹³

$$\partial_j \text{Probability of Switching} / \partial \text{Premium} = -(\partial MS_j / \partial \text{Premium}_j) / MS_j$$

Thus, the change in the probability of switching associated with a \$10 change in the out-of-pocket premium is $0.04/0.25 = 0.16$. This estimate is considerably larger than Buchmueller's estimate of 0.022 (Buchmueller 2000; Table 3, p. 961).

Potential Omitted Variables Bias

There are two potential sources of endogeneity among the right-hand side variables in our data—the use of aggregate market shares and omitted health plan characteristics. Analyses of *individual* consumers' health plan choices (e.g., Feldman et al. 1989) typically treat premiums and benefits as predetermined from the individual consumer's perspective. In this analysis, however, the dependent variable is constructed from market level enrollment data. Any omitted characteristics of beneficiaries that shift the demand curve for the j th health plan relative to the reference plan could alter the relative premium, benefits, and market share of the j th plan. Omitted variables bias could also arise simply because we are not able to observe every possible health plan characteristic that might be important to beneficiaries.

Omitted variables result in biased coefficients when they are correlated with the dependent variable and some of the explanatory variables in the regression equation. If our estimates are subject to omitted variables bias, the bias on our coefficients is likely to be positive. That is because omitted demand-shift variables and omitted plan characteristics are likely to be correlated positively with both enrollment in higher priced plans and observed plan characteristics, including premiums. We refer to explanatory variables that are subject to omitted variables bias as endogenous explanatory variables. The usual estimation strategy for potentially endogenous variables is to:

1. identify instruments (variables that are correlated with the endogenous variables, but not correlated with the error term in the equation of interest, that is, the demand equation);
2. test for endogeneity;
3. if endogeneity is present, estimate a two-stage least squares (2SLS) model, that is, a reduced form equation for out-of-pocket premium or benefits, followed by estimation of the demand equation which uses the predicted values from the premium and benefit equations as instruments.

There are two problems that make 2SLS difficult to implement in these data. The first problem, which affects only the benefit variables, is that the large number of benefit variables makes it impossible to find enough instruments to treat all the benefit variables as endogenous (or, by extension, even to test whether or not they are endogenous). If benefits and premiums were highly collinear, the omitted benefit variables could bias the estimated coefficients on premium and the observed benefits. We calculated the condition number for the vector of relative benefit variables and relative premium (Model 2 in Table 3). The condition number for the standardized variables¹⁴ was 8.06, considerably less than the threshold value for serious multicollinearity of 20 (Belsley, Kuh, and Welsch 1980). Thus, the benefits seem to vary independently of each other (and somewhat independently of premiums, as well).

The second problem is that the explanatory variables in our model are *differences* in plan characteristics between the j th plan and the reference plan. Generally, the demand equation is identified by variables that shift the firm's marginal cost curve. We observe several such variables, but only at the county level, not at the individual health plan level. These county-level variables should difference away in our conditional logit model, unless their coefficients

are different for individual health plans (as explained below). We observe no predetermined variables related to the marginal cost of individual health plans, and thus do not have sound theory to identify instruments for premiums. However, we use three alternative approaches to investigate the potential problem of omitted variables bias in our empirical analysis.

Our first approach, which addresses bias for the premium coefficient, is to estimate the demand equation with and without the benefit variables. Because we expect omitted benefits to be correlated positively with both premiums and enrollment, we would expect inclusion of more benefits to result in a more negative coefficient on the relative premium. Model 1 in Table 3 contains only relative premium and a constant. The coefficient on relative premium is $-.019$. Adding the benefit variables (Model 2) makes the relative premium coefficient more negative (-0.217), as expected. However, the difference in the estimated premium coefficient is quite small.

Our second approach addresses bias that might shift both relative premiums and benefits. We include four county-level variables that might shift demand for M+C plans. In order for these county-level variables to alter the demand curve for the j th health plan relative to the reference plan, the coefficients on the county-level variable must be different for the two plans. In other words, the specification of equation 3 must be:

$$\ln \frac{MS_{jm}}{MS_{km}} = (X_j - X_k)\beta + X_m(\beta_j - \beta_k) + u \quad (10)$$

where X_m is a vector of characteristics of the m th county.

The first county-level variable is the M+C Herfindahl index—the sum of the squared market shares of all M+C plans in the county. The second variable is average PIP-DCG risk score for M+C enrollees in the county. Higher scores indicate higher health risks (Ingber 2000). The third variable is the percent of the population that is white. White versus nonwhite disparities in HMO enrollment have been documented in previous studies (see Call et al. 1999, for a recent review). The fourth variable is the percent of the elderly with income below the federal poverty level. In high payment areas, the generous benefits offered by M+C plans at no additional out-of-pocket premium offer an attractive choice to low-income beneficiaries.

Several of these variables are significant, supporting the specification in equation 10, so we tested the sensitivity of our coefficients on relative premiums and benefits to inclusion and exclusion of those county-level demand factors. It is not clear exactly how the omitted demand factors should operate. However, it is likely that consumer income, for example, affects the

demand for the j th plan relative to the reference plan largely through demand for enhanced benefits offered by the j th plan. If that is the case, and some of those characteristics are unobserved, then exclusion of the demand factors will bias the premium coefficient toward zero. That is what we find. Neither the change in the relative premium coefficient nor the change in the benefit coefficients is large, however.

Our third approach to potential omitted variables affecting the relative premium coefficient is based on a “natural experiment.” In January 2000, Humana introduced a \$10 out-of-pocket premium in otherwise zero-premium M+C markets in Kansas City, Kansas and Corpus Christi, Texas.¹⁵ Humana’s market share fell 4.4 percentage points in Kansas City (from 41.8 to 37.4 percent) and 7 percentage points (from 77.7 to 70.7) in Corpus Christi.¹⁶ These results suggest that our estimates correspond closely to the experience of plans in the M+C market.

CONCLUSIONS

We have estimated the effect of premiums and benefits on health plan choice in the Medicare+Choice program. Our estimates generally are lower than estimates for the under-65 population. However, the market share losses associated with small changes in a health plan’s premium, relative to its competitors, are not trivial and may be large enough to discipline the premiums charged by M+C plans. Our estimates apply only to the counties that M+C plans have chosen to serve under the current payment system.

The importance of market share losses may depend on which beneficiaries are most likely to switch plans as out-of-pocket premiums rise or fall. For example, if the healthiest beneficiaries were the most price-elastic in their choice of plans, and if payments to plans were not adjusted accurately for risk,¹⁷ then the threat of losing low-cost beneficiaries might provide an additional incentive for plans to keep premiums low.

The estimated effects of benefits on enrollment are more difficult to assess. Coverage of some low-cost items such as dental and eye services seem to be very attractive to beneficiaries, while coverage of preventive care actually has a negative effect on enrollment. Several of the variables representing point-of-purchase cost sharing also had the wrong sign. Although we cannot provide further tests in our data, it is possible that some of the benefit variables may be correlated with unobserved characteristics of health plans. For example, plans with relatively little management of care may

employ higher point-of-purchase cost sharing to control moral hazard, and beneficiaries may take into account both factors when choosing a health plan.

There is further anecdotal evidence regarding M+C plans' response to a competitive market disciplined by out-of-pocket premiums. In June 1997, the CMS received bids from a number of M+C plans in the Denver area for a standard benefit package that approximated the market norm benefits offered by plans in Denver (i.e., one including an outpatient pharmaceutical benefit). The penalty for high bids was that high-priced plans would have to charge an out-of-pocket premium in an otherwise zero-premium market. If competition had forced the plans to spend all of their profits on free supplementary benefits, prior to the submission of bids, the bids should have come in close to the government's pre-bid payment rate. In fact, the bids for the market norm package were 4.6 to 16.7 percent below the government payment rate. The bids for the Medicare entitlement benefit package were 24 to 38 percent lower than the current payment rate. These results suggest three things. First, competition on the basis of benefits in zero-premium markets had failed to extract all surplus profits from the participating plans. Second, a competitive pricing payment system for M+C plans in 1997 could have resulted in substantial savings to the federal government with no reduction in benefits, because the plans were bidding on a market norm benefit package. Second, M+C plans *perceive* the price elasticity of health plan choice to be large enough to warrant substantial reductions in their prices when faced with the discipline of out-of-pocket premiums.

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NOTES

1. Prior to 1997, private health plans under full risk contracts were referred to as TEFRA-risk HMOs.
2. Under BBA, Medicare's monthly M+C capitation rate equals the *largest* of the following three amounts: (1) a blend of area-specific and national rates; (2) a minimum rate; and (3) a minimum percentage increase from the previous year's rate. High-payment areas are in the third category, because a minimum percentage increase (or even no increase) is greater than the minimum payment or the blended rate. The minimum percentage increase for 1998 and 1999 was 2 percent. In

addition, CMS began to phase in a new risk-adjustment system in 2000. Because it is widely believed that M+C plans selected favorable risks under the old risk-adjustment methodology (e.g., see Brown et al. 1993), the new system is expected to reduce the monthly capitation rate for many M+C plans. To avoid tightening M+C payments at a time of plan withdrawals and reaction against the BBA payment limits, Congress lengthened the transition to the new risk adjusters in the 1999 Balanced Budget Refinement Act. This legislation also provided for special payment increases for plans entering underserved, largely rural areas. Finally, in the Benefits Improvement and Protection Act of 2000, Congress eased the BBA payment limits in two major ways. First, Congress changed the minimum payment method, creating a new “urban” minimum payment of \$525 (for any payment area in a Metropolitan Statistical Area with a population of more than 250,000), along with an increase in the payment floor for all other areas to \$475. Second, Congress raised the BBA’s minimum percentage increase from 2 percent to 3 percent for 2001.

3. Except for a one-time “bonus” for some plans in 2001.
4. An HMO actuary suggested that the adjustments to benefits were driven by fears of unfavorable selection.
5. Establishing a standard benefit package on which M+C plans bid would not preclude M+C plans from offering separate supplementary benefits—that is, “add-ons” to the standard benefit—and charging an additional premium.
6. Throughout our empirical analyses, “out-of-pocket premium” refers to the difference between the plan’s total premium and the government’s premium contribution. For example, we do not have data on contributions to retiree’s out-of-pocket premiums by their, or their spouse’s, former employer.
7. An anonymous referee asked about the possibility that some M+C plans were operating under self-imposed caps on enrollment. In 1999, 7 of the 1,259 plan-county observations involved plans that had requested permission to impose an enrollment cap. Even if a cap is approved by CMS, plans do not always impose the cap. Plans also come back to CMS with proposals to change caps once they are in place.
8. Multiple observations of the same health plan’s premium and benefits result in a degree of autocorrelation in the data. However, our data consist of *pairs* of plans (the j th plan vis-à-vis the reference plan) and not all health plans declare the same counties in a metropolitan statistical area to be within their service area. Thus, the reference plan could be different for different counties within one plan’s market area. In fact, the plan might be the low-cost reference plan in one county, but not in another. Scanlon et al. (2002) were able to implement an empirical test for this type of autocorrelation and found that their results largely were unchanged.
9. We used the March 1999 enrollment data to avoid the problem of “anticipatory” enrollment and disenrollment that might occur if plans announced changes for 2000 during the summer of 1999.
10. We calculate the elasticity assuming that the premiums of all other plans are held constant. Thus, our estimates correspond to a change in the premium of a higher priced plan, holding the premium of the lowest priced plan constant.

11. We use the out-of-pocket premium for $PREM_j$ in our estimates of price elasticities.
12. See Feldman et al. (1993), p. 185.
13. The right-hand side of the expression has a negative sign because a decrease in market share corresponds to an increase in switching.
14. The standardized variables are the variables divided by their standard deviations.
15. One plan in Kansas City charged a \$110 out-of-pocket premium in 2000, but it had only 2 percent of the market.
16. Our estimates assume that other plan characteristics are held constant. In Corpus Christi, where Humana faced only one other competitor, Humana simultaneously increased its copayments for prescription drugs and inpatient and outpatient hospital services more than its competitor, but otherwise the benefits were similar.
17. Government payments to M+C plans are risk-adjusted, but the risk variables do not explain much of the variance in beneficiary costs. M+C plans are not allowed to vary their out-of-pocket premiums for enrollees in different risk categories. Thus, a substantial amount of the explainable variance in the risk of M+C plans' enrollees remains uncompensated in the current payment system.

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