HMO Growth and the Geographical Redistribution of Generalist and Specialist Physicians, 1987–1997

José J. Escarce, Daniel Polsky, Gregory D. Wozniak, and Phillip R. Kletke

Objective. To assess the impact of the growth in HMO penetration in different metropolitan areas on the change in the number of generalists, specialists, and total physicians, and on the change in the proportion of physicians who are generalists.
Data Sources/Study Setting. The American Medical Association Physician Masterfile, to obtain the number of patient care generalists and specialists in 1987 and in 1997 who were practicing in each of 316 metropolitan areas in the United States. Additional data for each metropolitan area were obtained from a variety of sources, and included HMO penetration in 1986 and 1996.
Study Design. We estimated multivariate regression models in which the change in the number of physicians between 1987 and 1997 was a function of HMO penetration in 1986, the change in HMO penetration between 1986 and 1996, population characteristics and physician fees in 1986, and the change in population characteristics and fees between 1986 and 1996. Each model was estimated using ordinary least squares (OLS) and two-stage least squares (TSLS).
Principal Findings. HMO penetration did not affect the number of generalist physicians or hospital-based specialists, but faster HMO growth led to smaller increases in the numbers of medical/surgical specialists and total physicians. Faster HMO growth also led to larger increases in the proportion of physicians who were generalists. Our best estimate is that an increase in HMO penetration of .10 between 1986 and 1996 reduced the rate of increase in medical/surgical specialists by 10.3 percent and reduced the rate of increase in total physicians by 7.2 percent.
Conclusions. The findings of this study support the notion that HMOs reduce the demand for physician services, particularly for specialists' services. The findings also imply that, during the past decade, there has been a redistribution of physicians—especially medical/surgical specialists—from metropolitan areas with high HMO penetration to low-penetration areas.
Key Words. Managed care, health maintenance organizations, physician location, physician supply

The rapid growth of managed care has altered the patterns of utilization of physician services and, consequently, the practice and employment options available to U.S. physicians. Managed care reduces medical care costs through its effects on the demand for medical care and on the prices of medical services. The stricter forms of managed care arrangements, especially health maintenance organizations (HMOs), influence demand through a variety of complex mechanism for delivering care, paying providers, and monitoring provider performance. Typical mechanisms that apply to physician services include selective contracting, use of generalists as "gatekeepers" who must approve referrals to specialists, sharing of financial risk with physicians, physician profiling, clinical practice guidelines, and utilization review (Gold, Hurley, Lake, et al. 1995; Hurley et al. 1996; Remler, Donelan, Blendon, et al. 1997). Through these methods HMOs reduce the demand for specialists' services relative to the demand for generalists' services (e.g., Miller and Luft 1994; Weiner 1994; Council on Graduate Medical Education [COGME] 1995). HMOs also make physician services markets more price competitive, leading to lower physician fees (Baker 1995).

HMO growth has not been uniform across the United States. Although HMOs have gained enrollees rapidly in certain parts of the country, other areas remain minimally affected by managed care. The variation in HMO growth appears to have spawned wide geographic disparities in earning opportunities for generalists and specialists. For example, Hadley and Mitchell (1997) found that higher HMO penetration reduced the number of patients seen per week by both generalists and specialists, although the reduction in patient visits was twice as great for specialists as for generalists. Hadley and Mitchell (1996) also found that higher HMO penetration did not affect the incomes of generalists, but reduced specialists' incomes. Simon, Dranove, and White (1998) found that growth in HMO penetration increased generalists' incomes, did not affect the incomes of medical and surgical specialists, and reduced the incomes of radiologists and anesthesiologists.

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Address correspondence to José J. Escarce, M.D., Ph.D., Senior Natural Scientist, RAND, 1700 Main Street, M-28, Santa Monica, CA 90401. Daniel Polsky, Ph.D. is Research Assistant Professor, Division of General Internal Medicine, Dept. of Medicine, University of Pennsylvania, and the Leonard Davis Institute of Health Economics, University of Pennsylvania. Gregory D. Wozniak, Ph.D. is Senior Economist, and Phillip R. Kletke, Ph.D. is Senior Social Scientist, Center for Health Policy Research, American Medical Association. This article, submitted to *Health Services Research* on May 3, 1999, was revised and accepted for publication on August 26, 1999.

Geographic disparities in earning opportunities engendered by HMOs may influence decisions by physicians regarding their geographical location and professional activities. In a recent study, we found that HMO penetration had sizable effects on the practice location choices of new physicians, that is, of physicians who were completing their graduate medical education and beginning to practice for the first time (Escarce, Polsky, Wozniak, et al. 1998). Specifically, we found that both generalists and specialists who completed graduate medical education in 1994 were less likely to locate their first practice in a market area with high HMO penetration than in a low-penetration area, although the effect was much larger for specialists. By contrast, in another study we found that higher HMO penetration did not affect the probabilities that established physicians in early or mid-career relocated their practices or left patient care altogether to go into, for example, administration, teaching, or research (Polsky et al. Forthcoming). HMO penetration could affect older physicians' retirement decisions, although this has not been studied.

It is possible that the cumulative impact of these labor market decisions by individual physicians is to change the aggregate distribution of physicians across market areas with differing rates of growth in HMO penetration. To address this question, this article examines the effect of HMO growth between 1986 and 1996 on the geographical distribution of generalist and specialist physicians across metropolitan areas in the United States. In particular, we assess the impact of the growth in HMO penetration in different metropolitan areas on the change in the number of generalists, specialists, and total physicians, and on the change in the proportion of physicians who are generalists.

CONCEPTUAL FRAMEWORK

The conceptual framework underlying the empirical analyses in this study assumes that HMOs influence the demand for physician services through a variety of complex mechanisms for delivering care, paying providers, and monitoring provider performance. As discussed earlier, many of these mechanisms reduce the demand for specialists' services relative to generalists' services. The demand for the services of generalists and specialists in a particular market area is assumed to be reflected in the number of full-time equivalent (FTE) generalist and specialist physicians, respectively, practicing in the market.[1] Therefore, the change over time in the number of generalist and specialist physicians in a market area may be modeled as a function of HMO penetration and additional factors that affect the demand for physician services. These additional factors include sociodemographic characteristics of the population in the market area and the out-of-pocket price of physician services.

Of course, HMOs do not randomly choose which market areas to enter or in which markets to expand their presence. Rather, HMO penetration is determined by the demand for managed care and by insurers' costs of contracting with and monitoring health care providers, including physicians, in different markets (Dranove, Simon, and White 1998). Because HMOs are widely believed to generate cost savings, demand for HMOs may be greatest and HMOs may gain market share most readily in market areas where health care costs are high or rising rapidly.[2] Similarly, HMOs may gain market share most readily in markets where the number of physicians is high or growing quickly. Therefore, it is likely that HMO penetration is endogenous in a model of the change over time in the number of physicians as a function of HMO penetration. The empirical analyses in the study, described in detail below, were designed to address the potential for endogeneity bias in quantifying the effects of HMO penetration.

EMPIRICAL MODEL

We estimated multivariate regression models to examine the change between 1987 and 1997 in the number of FTE generalist and specialist physicians practicing in 316 Metropolitan Statistical Areas in the United States. (Non-metropolitan areas were excluded from the study.) The purpose of the analysis was to determine how the level and rate of growth in HMO penetration influenced the changes in the number of FTE physicians. Under the conceptual framework described in the preceding section, changes in the numbers of physicians are assumed to reflect changes in the demand for physician services. Therefore, the regression models controlled for other market-level variables that may influence demand. The general form of the regression models was:

$$\Delta MD_{87-97} = \alpha + \beta_0 HMO_{86} + \beta_1 \Delta HMO_{86-96} + \gamma_0 X_{86} + \gamma_1 \Delta X_{86-96} + e$$

where ΔMD_{87-97} is the change in the number of FTE physicians between 1987 and 1997 (or in the proportion of physicians who were generalists); HMO_{86} is the level of HMO penetration in 1986; HMO_{86-96} is the change in HMO penetration between 1986 and 1996; X_{86} is a vector of additional variables that affect the demand for physician services, measured in 1986; ΔX_{86-96} is the change in these variables between 1986 and 1996;[3] e is an error disturbance; and the Greek symbols are regression coefficients to be estimated.[4]

Dependent Variables

We estimated five different regression models with the following dependent variables: (1) the change in the logarithm of the number of FTE generalist physicians between 1987 and 1997, (2) the change in the logarithm of the number of FTE medical/surgical specialists, (3) the change in the logarithm of the number of FTE hospital-based specialists, (4) the change in the logarithm of the total number of physicians, and (5) the change in the proportion of physicians who were generalists.

The numbers of FTE physicians were derived from the American Medical Association (AMA) Physician Masterfile, which provides current and historical data for each allopathic physician in the United States, including the physician's main professional activity (e.g., patient care, administration, research, teaching, etc.), the physician's primary specialty and secondary specialty (if any), and the physician's state and county (Randolph, Seidman, and Pasko 1996). We used the 1987 and 1997 year-end data to identify the nonfederal physicians in each specialty who had completed graduate medical education (i.e., internship, residency, and fellowship) and whose main professional activity was patient care. To obtain the number of FTE physicians in each specialty in each metropolitan area, we counted physicians who reported only one specialty as 1.0 FTE in the reported specialty, and physicians who reported two specialties as 0.6 FTE in the primary specialty and 0.4 FTE in the secondary specialty.[5] To develop the dependent variables for the models, general internists, general pediatricians, family physicians, and general practitioners were categorized as generalists; internal medicine subspecialists, pediatric subspecialists, dermatologists, neurologists, general surgeons, obstetricians/gynecologists, surgical subspecialists, psychiatrists, occupational medicine physicians, rehabilitation physicians, and physicians in several smaller specialties were categorized as medical/surgical specialists; and radiologists, pathologists, anesthesiologists, and emergency physicians were categorized as hospital-based specialists.

Explanatory Variables

The key explanatory variables in the regression models were HMO penetration in 1986, defined as the proportion of the population in the metropolitan area who were enrolled in HMOs, and the change in HMO penetration between 1986 and 1996. The 1986 penetration data were developed in two steps. First, county-level estimates of the number of HMO enrollees were obtained using data on total enrollment and service areas for all HMOs in the United States (Group Health Association of America [GHAA] 1988). Each HMO's enrollment was allocated to the counties in its service area using an algorithm that accounted for county population and for the concentration of HMO enrollees (Wholey, Feldman, and Christianson 1995). Second, county-level estimates were aggregated to the level of metropolitan areas. The 1996 penetration data were obtained from InterStudy (1998).[6] The HMO enrollment totals used to calculate HMO penetration included enrollment in point-of-service HMOs.

The additional explanatory variables in the regression models included the following characteristics of the population in each metropolitan area, measured as both the level in 1986 and the change from 1986 to 1996: the logarithm of total population, the proportion of the population who were younger than 20 years old, the proportion of the population who were older than 64 years old, the proportion of the population who were white, per capita income, and a wealth index that reflects the proportion of income from investments.[7] The additional explanatory variables also included the following population characteristics, measured in 1989: the proportion of the population who were women, the proportion of the adult population who graduated from high school, the proportion of the adult population who graduated from college, and the proportion of the population who were poor. (Data were unavailable to assess changes over time in these variables.) The additional explanatory variables also included a geographic index of Medicare physician fees in 1984 and the change in the fee index between 1984 and 1996.[8] Finally, the models for the change in the number of physicians between 1987 and 1997 also included the logarithm of the number of physicians in 1987 as an explanatory variable.

The population characteristics are intended to capture differences across metropolitan areas in the demand for physician services. In addition, certain of the population characteristics, such as the proportion of elderly persons, per capita income, the wealth index, and the poverty rate, tend to proxy for the extent of insurance coverage. The combination of these characteristics and the physician fee indexes would be expected to capture some of the variation across metropolitan areas in the out-of-pocket price of physician services.

Total population, the proportion of the population who were women, the proportion of the adult population who graduated from high school and from college, the poverty rate, and the index of Medicare physician fees in 1984 were obtained from the Area Resource File. The population age distribution, the proportion of the population who were white, per capita income, and the wealth index were obtained from Woods and Poole Economics (1995). The Medicare fee index in 1996 was obtained from the *Federal Register* (Health Care Financing Administration 1995).

ESTIMATION METHOD

The regression models were estimated using two alternative approaches: ordinary least squares (OLS) and two-stage least squares (TSLS). TSLS estimation was used to address the possibility that HMO penetration is endogenous, because OLS estimation results in inconsistent coefficient estimates when explanatory variables are endogenous (Greene 1993).[9] The level of HMO penetration in 1986 and the change in HMO penetration between 1986 and 1996 would be endogenous, as we suspect, if unobservable market area characteristics exist that are correlated with both HMO penetration and the rate of increase in the number of physicians. For instance, as discussed earlier, HMOs may gain market share more readily in markets where there are many physicians or where the number of physicians is growing rapidly. All other explanatory variables were considered exogenous.

To implement TSLS estimation, the HMO penetration variables in each regression model were first regressed on a set of instrumental variables that consisted of all of the exogenous explanatory variables in the model plus additional (identifying) variables suggested by prior studies of the determinants of managed care penetration (e.g., Dranove, Simon, and White 1998). The identifying variables included the distribution of firm size in the metropolitan area in 1988, the proportion of total employment in the manufacturing and in the service sectors in 1988, and the proportion of generalist physicians who were in solo practice in 1985. Although these variables were expected to be correlated with HMO penetration, they were not expected to have a direct effect on the change in the number of physicians.[10] The predicted values of the HMO penetration variables from these first-stage equations were then used as explanatory variables in estimating the main regression models (also called the second-stage equations).

Each regression model was estimated using all 316 metropolitan areas. Standard errors were estimated using heteroscedasticity-consistent covariance matrix estimators (White 1980, 1982).[11]

RESULTS

Descriptive Data

The total number of physicians in the 316 metropolitan areas whose main professional activity was patient care increased by 35.9 percent between 1987 and 1997 (Table 1). The increase in generalists was 35.4 percent; in medical/surgical specialists, 33.6 percent; and in hospital-based specialists, 44.6 percent. The proportion of physicians who were generalists was .34 in both 1987 and 1997. HMO penetration in the 316 metropolitan areas averaged .14 in 1986 and .28 in 1996. The distribution of HMO penetration across metropolitan areas is shown in Figure 1. Table 2 presents descriptive statistics for the explanatory variables used in the regression models.

Ordinary Least Squares Estimates

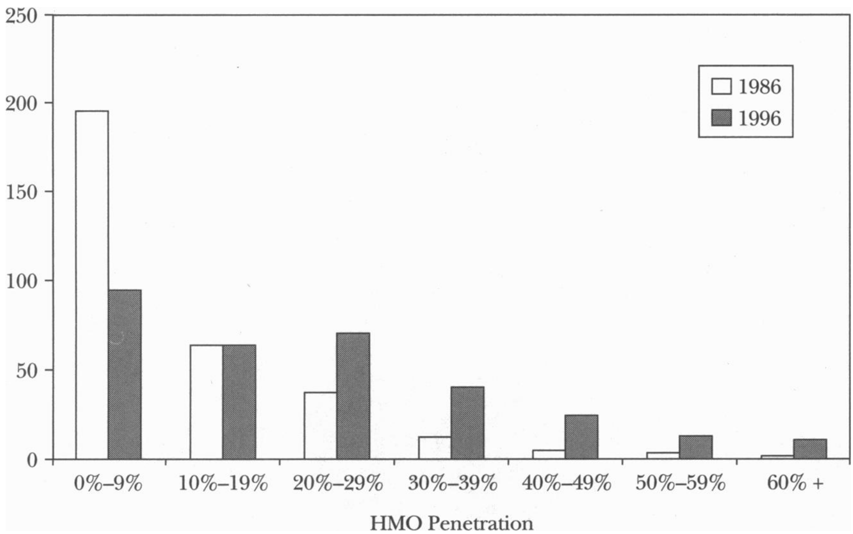
Table 3 reports the results of OLS estimation of the regression models. Neither HMO penetration in 1986 nor the change in HMO penetration between 1986 and 1996 had a statistically significant effect on the change in the number of generalists or hospital-based specialists. By contrast, the change in HMO penetration between 1986 and 1996 had a significant negative influence on the change in the numbers of medical/surgical specialists and total physicians. Specifically, an increase in HMO penetration of .10 reduced the rate of increase in medical/surgical specialists by 1.1 percent, and reduced the rate of increase in total physicians by 0.6 percent. Both the level of HMO penetration in 1986 and the change in penetration between 1986 and 1996 had statistically significant positive effects on the proportion of physicians who were generalists.

Other results of OLS estimation generally were consistent with theoretical expectations. The rates of population and income growth had statistically significant positive effects on the change in the numbers of generalists, medical/surgical specialists, hospital-based specialists, and total physicians. These findings are consistent with the expected effects of population and income on the demand for physician services. In particular, the crucial role of population growth in attracting physicians has been documented previously (e.g., Ernst and Yett 1985). Conversely, the level of Medicare physician fees in 1984 and the change in fees between 1984 and 1986 had significant negative effects on the change in the numbers of medical/surgical specialists, hospital-based specialists, and total physicians, consistent with higher fees reducing the demand for the services of these physicians. Estimated coefficients of other explanatory variables were less often statistically significant.

Table 1: Full-time Equivalent Physicians in 316 Metropolitan Areas, 1987 and 1997

	1987	*1997*	*% Change*
Generalists	108,466	146,817	35.4
Medical/Surgical Specialists	157,871	210,895	33.6
Hospital-based specialists	50,143	72,520	44.6
Total physicians	316,480	430,232	35.9
Proportion of physicians who were generalists	.34	.34	—

Figure 1: Distribution of HMO Penetration across Metropolitan Areas, 1986 and 1996



Two-Stage Least Squares Estimates

Tables 4 and 5 report the results of TSLS estimation. Table 4 reports the coefficient estimates for the first-stage equations, in which HMO penetration in 1986 and the change in penetration between 1986 and 1996 were each regressed on the sociodemographic characteristics of the population in each metropolitan area, the Medicare physician fee indexes, and the additional identifying variables described earlier.[12]

Consistent with the findings of Dranove, Simon, and White (1998), a higher proportion of employment in large firms and a lower proportion of

Table 2: Descriptive Statistics for Explanatory Variables Used in Regression Models

Variable	*Mean*	*s.d.*
HMO penetration in 1986*	.093	.115
Δ in HMO penetration, 1986–1996*	.114	.132
Logarithm of population, 1986	12.665	1.022
Δ in log of population, 1986–1996	.102	1.05
Proportion of population < 20 years, 1986	.297	.029
Δ in proportion of population < 20 years, 1986–1996	−.005	.007
Proportion of population > 64 years, 1986	.118	.033
Δ in proportion of population > 64 years, 1986–1996	.010	.005
Proportion of whites in population, 1986	.875	.100
Δ in proportion of whites, 1986–1996	−.016	.013
Logarithm of per capita income, 1986	9.507	.178
Δ in log of per capita income, 1986–1996	.466	.060
Wealth index, 1986	.946	.159
Δ in wealth index, 1986–1996	.005	.040
Proportion of women in population, 1989	.513	.013
Proportion of adult high school graduates, 1989	.763	.067
Proportion of adult college graduates, 1989	.198	.063
Proportion of population poor, 1989	.130	.050
Medicare fee index, 1984	.912	.147
Δ in Medicare fee index, 1984–1996	.041	.124

*Means differ from those reported in the text because the means in the text are weighted by the population of each metropolitan area.

generalists in solo practice were associated with higher HMO penetration in 1986. Interestingly, however, a higher proportion of employment in large firms was associated with slower growth in HMO penetration between 1986 and 1996, whereas a higher proportion of employment in the service sector was associated with faster growth in HMO penetration.[13] The partial F-statistics for the identifying variables were $F = 5.38$ ($p < .001$) in the model for HMO penetration in 1986, and $F = 3.07$ ($p = .006$) in the model for the change in HMO penetration between 1986 and 1996 (Table 4). These partial F-statistics support the use of the identifying variables as predictors of HMO penetration.

Table 5 reports the results of TSLS estimation of the second-stage equations. As before, neither HMO penetration in 1986 nor the change in penetration between 1986 and 1996 had a statistically significant effect on the change in the number of generalists or hospital-based specialists. However, the change in HMO penetration between 1986 and 1996 had a significant

Table 3: Ordinary Least Squares Estimates (t-statistics in parentheses)

Explanatory Variable	*Change in Generalists*	*Change in Medical/ Surgical Specialists*	*Change in Hospital-based Specialists*	*Change in Total Physicians*	*Change in Generalist Proportion*
HMO penetration in 1986	.049	−.071	−.065	−.042	.021*
	(0.84)	(1.03)	(0.87)	(0.78)	(1.65)
Δ in HMO penetration, 1986–1996	.036	−.107***	−.047	−.057*	.026**
	(0.82)	(2.58)	(0.76)	(1.69)	(2.56)
Logarithm of population, 1986	.080**	.006	.047	−.020	.001
	(2.25)	(0.24)	(1.47)	(0.81)	(0.28)
Δ in log of population, 1986–1996	.763***	.678***	.634***	.713***	0.24
	(11.09)	(10.22)	(6.16)	(12.08)	(1.44)
Proportion of population < 20 years, 1986	−.682	−1.100***	−1.622**	−.989***	.133
	(1.41)	(2.76)	(2.48)	(2.59)	(1.41)
Δ in proportion of population < 20 years, 1986–1996	−1.528	−1.671	−.984	−1.657	.002
	(1.10)	(1.38)	(0.46)	(1.42)	(0.01)
Proportion of population > 64 years, 1986	−.106	−.507	−1.212**	−.529*	.124
	(0.28)	(1.41)	(2.07)	(1.67)	(1.44)
Δ in proportion of population > 64 years, 1986–1996	.367	−2.683*	−4.460*	−1.894	.829**
	(0.25)	(1.83)	(1.89)	(1.56)	(2.10)
Proportion of whites in population, 1986	−.134	−.019	−.102	−.065	−.023
	(1.49)	(0.25)	(0.88)	(0.94)	(1.26)
Δ in proportion of whites, 1986–1996	−.165	.372	.614	.173	−.149
	(0.32)	(0.67)	(0.72)	(0.37)	(1.20)
Logarithm of per capita income, 1986	.008	.019	.004	−.023	−.014
	(0.05)	(0.13)	(0.02)	(0.17)	(0.41)
Δ in log of per capita income, 1986–1996	.317*	.579***	.656**	.510***	−.059
	(1.71)	(3.55)	(2.47)	(3.63)	(1.49)
Wealth index, 1986	−.157	−.084	−.040	−.086	−.012
	(0.96)	(0.59)	(0.16)	(0.64)	(0.36)
Δ in wealth index, 1986–1996	.317	.069	.136	.124	.040
	(1.31)	(0.33)	(0.50)	(0.70)	(0.83)
Proportion of women in population, 1989	.289	−.579	−1.006	−.680	.164
	(0.44)	(1.00)	(0.80)	(1.39)	(1.07)
Proportion of adult high school graduates, 1989	.007	−.022	.019	−.027	.008
	(0.04)	(0.14)	(0.07)	(0.19)	(0.19)
Proportion of adult college graduates, 1989	.569**	.267	.307	.264	.045
	(2.42)	(1.19)	(1.01)	(1.31)	(0.93)
Proportion of population poor, 1989	.146	−.051	.036	.015	.036
	(0.53)	(0.22)	(0.10)	(0.07)	(0.65)

Continued

Table 3: Continued

Explanatory Variable	*Change in Generalists*	*Change in Medical/ Surgical Specialists*	*Change in Hospital-based Specialists*	*Change in Total Physicians*	*Change in Generalist Proportion*
Medicare fee index, 1984	−.019 (0.12)	−.408*** (2.60)	−.729*** (3.01)	−.323** (2.50)	.110*** (3.19)
Δ in Medicare fee index, 1984–1986	.010 (0.07)	−.336** (2.07)	−.630*** (2.59)	−.241* (1.87)	.103*** (3.02)
Logarithm of no. of generalists, 1987	−.073** (2.01)	—	—	—	—
Logarithm of no. of medical/ surgical specialists, 1987	—	−.003 (0.12)	—	—	—
Logarithm of no. of hospital-based specialists, 1987	—	—	−.031 (1.03)	—	—
Logarithm of total physicians, 1987	—	—	—	.027 (1.11)	—
Intercept	−.453 (0.27)	.902 (0.70)	1.458 (0.62)	1.451 (1.19)	−.099 (0.33)
F-statistic	16.01	16.75	10.41	24.23	2.74
R-square	0.50	0.52	0.40	0.60	0.14

$^{*}p < .10$; $^{**}p < .05$; $^{***}p < .01$.

negative influence on the change in the numbers of medical/surgical specialists and total physicians. The estimated effects were considerably larger than was suggested by the OLS regression (see Table 3). Thus, an increase in HMO penetration of .10 reduced the rate of increase in medical/surgical specialists by 10.3 percent and reduced the rate of increase in total physicians by 7.2 percent. Faster growth in HMO penetration also led to a larger increase in the proportion of physicians who were generalists.

Other results of TSLS estimation paralleled the OLS results. In particular, the rates of population and income growth had consistently positive and significant effects on the change in the numbers of physicians of all types.

Which Estimates Are Preferable?

As shown in Tables 3 and 5, OLS and TSLS estimates of the influence of HMO penetration on the change, between 1987 and 1997, in the numbers of physicians in different metropolitan areas yield the same qualitative conclusions. Nonetheless, the TSLS coefficient estimates imply much larger effects

Table 4: Two-stage Least Squares Estimates, First-stage Equations (t-statistics in parentheses)

	Dependent Variable	
Explanatory Variable	*HMO Penetration in 1986*	*Change in HMO Penetration 1986–1996*
Logarithm of population, 1986	.025***	.014
	(3.42)	(1.19)
Δ in log of population, 1986–1996	.074	−.041
	(1.20)	(0.54)
Proportion of population < 20 years, 1986	.244	.218
	(0.71)	(0.43)
Δ in proportion of population < 20 years, 1986–1996	−1.661	2.276
	(1.52)	(1.59)
Proportion of population > 64 years, 1986	−.032	.222
	(0.11)	(0.47)
Δ in proportion of population > 64 years, 1986–1996	−.062	−3.070
	(0.05)	(1.52)
Proportion of whites in population, 1986	.211***	.015
	(3.39)	(0.18)
Δ in proportion of whites, 1986–1996	−1.755***	−.135
	(3.22)	(0.16)
Logarithm of per capita income, 1986	−.101	.271
	(0.78)	(1.63)
Δ in log of per capita income, 1986–1996	−.010	−.010
	(0.07)	(0.50)
Wealth index, 1986	.089	−.652***
	(0.66)	(3.89)
Δ in wealth index, 1986–1996	−.001	.176
	(0.01)	(0.49)
Proportion of women in population, 1989	.319	−.206
	(0.45)	(0.24)
Proportion of adult high school graduates, 1989	.540***	−.445
	(3.10)	(1.60)
Proportion of adult college graduates, 1989	−.220	.646**
	(1.10)	(2.20)
Proportion of population poor, 1989	.014	−.817**
	(0.05)	(2.25)
Medicare fee index, 1984	.345**	.499**
	(2.09)	(2.19)

Continued

Table 4: Continued

	Dependent Variable	
Explanatory Variable	*HMO Penetration in 1986*	*Change in HMO Penetration 1986–1996*
Δ in Medicare fee index, 1984–1996	.284*	.536**
	(1.79)	(2.35)
Proportion of generalists in solo practice, 1985†	−.246***	.097
	(4.97)	(1.62)
Proportion employment in firms with 20–99 employees, 1988†	1.105**	−.558
	(1.97)	(0.83)
Proportion employment in firms with 100–499 employees, 1988†	.663**	−.613
	(2.40)	(1.63)
Proportion employment in firms with > 500 employees, 1988†	.774**	−.610*
	(2.50)	(1.70)
Proportion of employment in manufacturing sector, 1988†	−.197	.277*
	(1.64)	(1.72)
Proportion of employment in service sector, 1988†	−.104	.752***
	(0.47)	(3.08)
Intercept	−.966	−1.974
	(0.87)	(1.35)
F-statistic	12.03	3.44
R-square	0.41	0.17
Partial F-statistic for identifying variables (df = 6,291)	5.38	3.07

*$p < .10$; **$p < .05$; ***$p < .01$.
†Identifying variable.

of HMO penetration on the numbers of medical/surgical specialists and total physicians. Therefore, we conducted tests to assess the relative merits of the OLS and TSLS estimates.

TSLS estimates are consistent only if the identifying variables in the first-stage equation are uncorrelated with the error disturbance in the second-stage equation (Greene 1993). As discussed earlier, these variables would not be expected to have a direct effect on the change in the number of physicians in a metropolitan area. Nonetheless, variables such as the proportion of employment in the service sector and in manufacturing could be correlated with the error disturbances in the second-stage equations if, for example, differences exist in the extent of insurance coverage across sectors and the influence of

Table 5: Two-stage Least Squares Estimates, Second-stage Equations (t-statistics in parentheses)

Explanatory Variable	*Change in Generalists*	*Change in Medical/ Surgical Specialists*	*Change in Hospital-based Specialists*	*Change in Total Physicians*	*Change in Generalist Proportion*
HMO penetration in 1986	.156	−.557	.051	−.457	.053
	(0.50)	(1.43)	(0.17)	(1.44)	(0.91)
Δ in HMO penetration, 1986–1996	−.012	−1.086**	−.197	−.752*	.141*
	(0.03)	(2.20)	(0.61)	(1.90)	(1.91)
Logarithm of population, 1986	.082	−.038	.045	−.059	−.002
	(1.53)	(0.81)	(1.31)	(1.26)	(0.59)
Δ in log of population, 1986–1996	.751***	.722***	.619***	.749***	.023
	(9.51)	(6.61)	(5.99)	(8.70)	(1.16)
Proportion of population < 20 years, 1986	−.710	−.571	−1.617**	−.577	.087
	(1.32)	(0.81)	(2.47)	(1.03)	(0.73)
Δ in proportion of population < 20 years, 1986–1996	−1.304	−.502	−.560	−.938	−.186
	(0.91)	(0.28)	(0.26)	(0.64)	(0.55)
Proportion of population > 64 years, 1986	−.071	−.252	−1.136**	−.343	.081
	(0.18)	(0.42)	(1.97)	(0.74)	(0.76)
Δ in proportion of population > 64 years, 1986–1996	.337	−5.316*	−4.706*	−3.851*	1.127**
	(0.20)	(1.69)	(1.89)	(1.65)	(2.14)
Proportion of whites in population, 1986	−.149	.167	−.112	.068	−.037
	(1.28)	(1.01)	(0.79)	(0.54)	(1.40)
Δ in proportion of whites, 1986–1996	.077	−.574	.896	−.634	−.097
	(0.10)	(0.45)	(0.85)	(0.65)	(0.46)
Logarithm of per capita income, 1986	.019	.134	.030	.053	−.035
	(0.10)	(0.62)	(0.10)	(0.29)	(0.81)
Δ in log of per capita income, 1986–1996	.312*	.599**	.651**	.532**	−.058
	(1.69)	(2.19)	(2.41)	(2.50)	(1.22)
Wealth index, 1986	−.173	−.534	−.099	−.406	.043
	(0.68)	(1.58)	(0.32)	(1.48)	(0.74)
Δ in wealth index, 1986–1996	.311	.124	.132	.164	.028
	(1.28)	(0.29)	(0.47)	(0.52)	(0.39)
Proportion of women in population, 1989	.247	−.378	−1.056	−.566	.107
	(0.37)	(0.37)	(0.83)	(0.74)	(0.63)
Proportion of adult high school graduates, 1989	−.067	.005	−.124	.023	.016
	(0.31)	(0.02)	(0.40)	(0.10)	(0.26)
Proportion of adult college graduates, 1989	.602**	.542	.380	.459	−.008
	(2.40)	(1.40)	(1.23)	(1.49)	(0.13)
Proportion of population poor, 1989	.135	−.744	.033	−.489	.106
	(0.38)	(1.38)	(0.08)	(1.16)	(1.41)

Continued

Table 5: Continued

Explanatory Variable	*Change in Generalists*	*Change in Medical/ Surgical Specialists*	*Change in Hospital-based Specialists*	*Change in Total Physicians*	*Change in Generalist Proportion*
Medicare fee index, 1984	.008 (0.03)	.360 (0.75)	−.627* (1.67)	.232 (0.61)	.022 (0.29)
Δ in Medicare fee index, 1984–1996	.042 (0.12)	.438 (0.92)	−.521 (1.40)	.321 (0.84)	.015 (0.21)
Logarithm of no. of generalists, 1987	−.078 (1.20)	—	—	—	—
Logarithm of no. of medical/ surgical specialists, 1987	—	.067 (1.22)	—	—	—
Logarithm of no. of hospital-based specialists, 1987	—	—	−.031 (0.98)	—	—
Logarithm of total physicians, 1987	—	—	—	.087 (1.62)	—
Intercept	−.470 (0.27)	−.597 (0.26)	1.336 (0.52)	.412 (0.22)	.197 (0.45)
F-statistic	15.90	7.02	10.20	12.01	1.91

*$p < .10$; **$p < .05$; ***$p < .01$.

such differences on the demand for physician services is not captured by other explanatory variables in the second-stage equations. We evaluated this correlation using a regression test of overidentifying restrictions, in which the residuals from each second-stage equation are regressed on the explanatory variables in the corresponding first-stage equation, including the identifying variables (Greene 1993). The null hypothesis of zero correlation was rejected for the model for the change in the number of generalists ($p = .018$), but the null hypothesis could not be rejected at the 10 percent level of significance for any of the other models. The identifying variables performed especially well in the models for the change in the number of medical/surgical specialists and for the change in the total number of physicians. The regressions of the residuals from these second-stage equations had $R^2 = .0020$ ($\chi^2 = .63$, $p = .96$) and $R^2 = .0024$ ($\chi^2 = .75$, $p = .94$), respectively, which indicates that no evidence existed to support even a weak correlation between the identifying variables and the error disturbances.

In addition, we used the specification test developed by Spencer and Berk (1981) to test the null hypothesis that HMO penetration in 1986 and the change in penetration between 1986 and 1996 were exogenous in each model that we estimated.[14] The null hypothesis of exogeneity was rejected in the models for the change in the number of medical/surgical specialists ($p = .008$) and for the change in the total number of physicians ($p = .063$). However, the null hypothesis of exogeneity could not be rejected at the 10 percent level for any of the other models.

Taken together, these tests suggest that the TSLS estimates are preferable in both the case of the model for the change in the number of medical/surgical specialists and the model for the change in the total number of physicians. By contrast, the OLS estimates are preferable in the case of the models for the change in the number of hospital-based physicians and for the change in the proportion of physicians who were generalists. Although TSLS yields similar coefficient estimates for the models, the OLS estimates are more efficient. Finally, the test of overidentifying restrictions raises concerns about using TSLS to estimate the model for the change in the number of generalists; consequently, the OLS estimates are preferable.

Of note, although TSLS estimates are consistent, they may retain some degree of bias in small samples. Recent work has shown that the bias of TSLS estimates is in the same direction as the bias of OLS estimates, and that the reciprocal of the partial F-statistic for the identifying variables in the first-stage equation is an approximate measure of the size of the TSLS bias relative to the OLS bias (Bound, Jaeger, and Baker 1995; Staiger and Stock 1997). The modest partial F-statistics for the identifying variables in this study (see Table 4) suggest that even our TSLS estimates may understate the damping effect of HMO growth on the rate of increase in medical/surgical specialists and total physicians.

Simulations

We developed simulations to provide a more intuitive illustration of the effects of HMO growth on the change in the numbers of medical/surgical specialists and total physicians. Using the TSLS estimates, the simulations predicted the change in the number of physicians between 1987 and 1997 in each of seven hypothetical metropolitan areas constructed to be alike in every respect except that the growth in HMO penetration varied from zero to .30 (in steps of .05)—the typical range of values observed in the study data.

To predict the change in the number of medical/surgical specialists, we assumed that each metropolitan area had 100 medical/surgical specialists in 1987. Absent an effect of HMO growth, each metropolitan area would have

134 medical/surgical specialists in 1997 (see Table 1). In contrast, predictions that account for the impact of HMO growth reveal that the metropolitan area that experienced no change in HMO penetration would have 157 medical/surgical specialists in 1997, whereas the metropolitan area that experienced an increase in HMO penetration of .30 would have 113 medical/surgical specialists in 1997.

To predict the change in the number of total physicians, we assumed that each metropolitan area had 100 total physicians in 1987. Absent an influence of HMO penetration, each metropolitan area would have 136 total physicians in 1997 (see Table 1). In contrast, predictions that account for the effect of HMO growth indicate that the metropolitan area that experienced no growth in HMO penetration would have 152 total physicians in 1997, while the metropolitan area that experienced an increase in HMO penetration of .30 would have 121 total physicians in 1997.

DISCUSSION

This study examined the impact of HMO penetration on the change between 1987 and 1997 in the number of physicians practicing in different metropolitan areas. The study found that faster growth in HMO penetration during the study period led to smaller increases in the numbers of medical/surgical specialists and total physicians, whereas HMO penetration did not affect the change in the number of generalists. Not surprisingly, faster HMO growth also led to larger increases in the proportion of physicians who were generalists. These results support the notion that HMOs reduce the overall demand for physician services, and that the impact of HMOs in curbing the demand for specialists' services is especially pronounced. The results also imply that a redistribution of physicians has taken place during the past decade—particularly among medical/surgical specialists—from metropolitan areas with high HMO penetration to low-penetration areas. Effects on the number of physicians may be an important mechanism for "spillover" effects of HMOs on patients with traditional indemnity insurance (Baker 1997).

The results of this study are consistent with previous research on the impact of HMO penetration on individual physicians' labor market decisions. In particular, Escarce, Polsky, Wozniak, et al. (1998) found that specialists who were completing graduate medical education were much less likely to locate their first practice in a market area with high HMO penetration than in a low-penetration area. Generalists who were completing graduate medical education also were less likely to locate their first practice in a high-penetration area, but the effect was smaller for generalists than for specialists. However, Polsky et al. (Forthcoming) found that higher HMO penetration did not affect the probabilities that established medical/surgical specialists relocated their practices or left patient care altogether.

Our comparison of OLS and TSLS estimates of the impact of HMO penetration on the increase in physicians is revealing. We found evidence that the growth in HMO penetration was endogenous to the changes in the numbers of medical/surgical specialists and total physicians, which suggests that there are unobservable market area characteristics correlated with both HMO growth and the rate of increase in these physicians. More specifically, the finding that the estimated effects of HMO penetration were larger under TSLS estimation than OLS estimation implies that HMOs gained market share most readily in those metropolitan areas where, for reasons that could not be observed in the data, the number of medical/surgical specialists was growing the most rapidly. Possibly in these areas health care costs were high or rising quickly, and consequently, the demand for managed care was substantial (e.g., Goldberg and Greenberg 1981; Welch 1984; Porell and Wallack 1990).

Our analyses may not have accounted for all of the characteristics of managed care markets that affect the rate of change in the numbers of different types of physicians. Because of a lack of data, our regression models did not include measures of the predominant type of HMO (e.g., staff, group, network, or independent practice association) or the penetration of preferred provider organizations (PPOs) in each metropolitan area. Recent evidence suggests that alternative HMO types differ little with regard to the use of systems for controlling utilization or plan performance (Gold, Hurley, Lake, et al. 1995; Hurley et al. 1996; Miller and Luft 1994). Also, PPOs probably have much weaker effects than HMOs on physician behavior because the former are much less likely to use payment and monitoring systems that influence the demand for physician services (Gold, Hurley, Lake, et al. 1995; Hurley et al. 1996); nonetheless, because PPOs discount fees, they could affect physician incomes and hence the attractiveness of different market areas to physicians. If levels of HMO activity are correlated with levels of non-HMO managed care activity, however, then HMO penetration will capture the effects of a broader range of managed care (Baker 1997).

Our study is limited by the exclusion of nonmetropolitan areas. If our findings are generalizable to these areas, however, HMO growth in metropolitan areas during the past decade may have led to larger increases in the availability of physicians in nonmetropolitan areas than otherwise would have occurred. Additional research is needed to clarify whether the migration of specialists to small towns and rural areas has accelerated in recent years.

This study provides new and compelling evidence of the effects of HMOs on the demand for physician services. The finding that HMOs are leading to a redistribution of physicians from metropolitan areas with high HMO penetration to low-penetration areas complements prior studies of utilization (e.g., Miller and Luft 1994), physician labor supply (Hadley and Mitchell 1997), and physician incomes (Hadley and Mitchell 1996; Simon, Dranove, and White 1998). Of course, any redistributive effects of HMOs can persist only so long as wide geographic differences in HMO penetration remain. Continued growth in HMOs and other strict managed care arrangements may result eventually in marked reductions in physicians' hours of work and incomes, and in lower economic returns for entering the medical profession and for specializing.

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NOTES

1. Most studies treat the number of physicians in a market area as a measure of supply. While this is obviously correct, it is also true that the number of FTE physicians in a market area must reflect demand, at least in urban markets where there is unlikely to be unmet demand. Formally, we can consider a system of two equations and one identity: $Q_s = f(S)$, $Q_d = g(D)$, and $Q_s = Q_d$, where Q_s is the quantity of physicians' services supplied, Q_d is the quantity demanded, S is the vector of supply variables, D is a vector of demand variables, and f and g are functions. The conceptual framework and empirical analyses in this study adopt a demand perspective.
2. For example, Goldberg and Greenberg (1981), Welch (1984), and Porell and Wallack (1990) concluded that the level of and rate of growth in HMO penetration are positively correlated with health care costs.
3. Whenever data for 1986 and 1996 were unavailable, we used the nearest years for which data were available.

4. If physician services markets were in equilibrium, only the change variables would be relevant in explaining the change in the number of physicians. However, given the rapidity of the growth in managed care and in other recent changes in health care financing and delivery, the numbers of physicians in different markets may not have reached equilibrium. The 1986 levels of the variables are included to account for this possibility.
5. These were the FTE weights used by Newhouse et al. (1982) in their landmark studies of physician practice location. Internal medicine physicians are much more likely than other physicians to report more than one specialty in the Masterfile, and combinations of general internal medicine and a subspecialty are especially common. Shea, Kletke, Wozniak, et al. (1999) show that internal medicine physicians' specialties reported in the Masterfile reflect the clinical content of their practices.
6. Prior to the mid-1990s, the InterStudy census and the Group Health Association of America directories provided information on the headquarters location, total enrollment, and service area for each HMO, but no information on enrollment by place of enrollee residence was given. Therefore, methods to allocate enrollment to the counties and metropolitan areas where enrollees resided were necessary (e.g., Wholey, Feldman, and Christianson 1995; Baker 1997). The 1996 InterStudy census, by contrast, reports data on HMO penetration by enrollee place of residence.
7. The wealth index is a weighted average of income per capita in the metropolitan area divided by U.S. income per capita; the proportion of income from dividends, interest, and rent divided by the U.S. proportion; and the U.S. proportion of income from transfers divided by the proportion in the metropolitan area (Woods and Poole Economics 1995).
8. The 1984 index was a Laspeyres index of Medicare-allowed charges under the now defunct "current, prevailing, and reasonable" pricing system. The 1996 index was based on the geographic adjustment factors used in the resource-based Medicare Fee Schedule.
9. Consistency refers to the properties of estimation methods in large samples. Consistent estimation methods yield unbiased regression coefficient estimates in large samples, even though they may yield biased estimates in small samples. In contrast, inconsistent estimation methods yield biased coefficient estimates even in large samples.
10. The firm size distribution and sector employment variables were expected to be correlated with the demand for managed care, whereas the proportion of generalists in solo practice was expected to be correlated with managed care organizations' costs of contracting with and monitoring physicians (Dranove, Simon, and White 1998). Data on the distribution of firm size and employment in the manufacturing and service sectors were obtained from the U.S. Census Bureau (1996). The proportion of generalists in solo practice was obtained from the AMA Physician Masterfile.
11. In preliminary analyses, we used Chow tests to assess whether it was appropriate to estimate separate models for large (>500,000 population) and for small

metropolitan areas. The Chow tests did not reject the null hypothesis that the large and small metropolitan areas could be pooled. However, Goldfeld-Quandt tests found that the variation in metropolitan area population was a source of heteroscedasticity. The null hypothesis of homoscedasticity was rejected at $p < .001$ for every model.

12. In the interest of space, Table 4 presents the first-stage equations used to estimate the model for the proportion of physicians who were generalists. The first-stage equations used to estimate the models for the change in the numbers of physicians have one additional explanatory variable each that corresponds to the number of physicians in 1987 (see the section on Empirical Model). The results for these first-stage equations were similar to those presented in Table 4.
13. A possible explanation for these findings is that markets with a large proportion of employment in large firms and with many generalists in group practice achieved high levels of HMO penetration early, but that HMO enrollment in those markets subsequently grew more slowly as other markets began to catch up. An anonymous reviewer suggested that large firms not in the service sector may be more unionized and hence that HMO penetration growth may be slowed because of the negotiated nature of benefits.
14. The Spencer and Berk test is a single-equation version of the well-known Hausman (1978) test.

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