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Physician response to Medicare fee reductions: changes in the volume of coronary artery bypass graft (CABG) surgeries in the Medicare and private sectors

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Abstract

The demand inducement hypothesis predicts that physicians will respond to reductions in their income by increasing the volume of their services when the income effect is strong and negative. I test for such inducement in the market for coronary artery bypass grafting (CABG), using a longitudinal panel of physicians in New York and Washington states. The results show that physicians whose incomes were reduced the most by Medicare fee cuts performed higher volumes of CABGs, and they did so in both the Medicare and private markets. © 1998 Elsevier Science B.V. All rights reserved.

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1. Introduction

The subject of demand inducement, both its theoretical underpinnings and empirical evidence, has generated heated debate for decades (Evans, 1974; Rein-

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hardt, 1985; Fuchs, 1986; Pauly, 1994). The debate has been rekindled by Medicare's recent change in payment method for physician services from the cost-based Customary, Prevailing and Reasonable (CPR) system to the Medicare Fee Schedule (MFS), leading in part to the adoption of the Volume Performance Standard. [1] Although voluminous empirical studies exist, most do not constitute direct tests for demand inducement and many are plagued with methodological problems, making the interpretation of results ambiguous (Holahan and Scanlon, 1979; Rice, 1983; Gabel and Rice, 1985; Feldman and Sloan, 1988; Mitchell et al., 1989; Rice and Labelle, 1989; Wedig et al., 1989; Hurley et al., 1990; Christensen, 1992; Escarce, 1993a,b; Ginsburg and Hogan, 1993; Physician Payment Review Commission, 1993). This paper attempts to provide a direct test for demand inducement by explicitly estimating the impact of the income effects of fee changes, the impact which the theory predicts as the critical factor in determining physician response to fee changes (McGuire and Pauly, 1991). It also adds to the existing empirical literature by incorporating the multiple-payer feature of the US health care system.

Using the 'overpriced procedures' provisions under the Omnibus Budget Reconciliation Act of 1987 (OBRA87) as a natural experiment, and focusing on thoracic surgeons and a set of heart revascularization surgical procedures—coronary artery bypass graft (CABG) surgery (one to six vessels), this paper addresses the following questions: (1) Do thoracic surgeons respond to Medicare fee cuts by increasing Medicare volume in order to compensate for income loss? (2) Is there evidence of spillover of the response to the private sector? (3) Do higher private-to-Medicare price ratios matter in determining the amount of spillover? [2] The data used for this study consists of a panel of physicians from New York and Washington states in 1987 and 1989, constructed from state inpatient discharge abstracts, the Part B Medicare Annual Data (BMAD) Prevailing Files, and the Surgical Prevailing Charge System from the Health Insurance Association of America (HIAA).

The results show evidence of a negative income effect that leads to increased volume as Medicare prices are cut, especially among the more intensive CABG procedures (CABG with more vessels). There is also a spillover of the negative income effect into the private sector, leading to increased private volume after Medicare fees are reduced. However, private-to-Medicare price ratios are not significant in determining the amount of spillover.

The paper is organized as follows. Section 2 briefly describes the institutional background. Section 3 describes the theoretical model. Section 4 discusses empiri-

[1] The Volume Performance Standard system is a mechanism to adjust Medicare fee schedule updates based on how annual increases in actual expenditures compare with previously determined 'performance standard' rates of increase.

[2] The behavioral response of 'cost-shifting' is analyzed in a separate paper (Yip, 1995).

cal estimation methodologies. Section 5 describes the data. Section 6 presents the results. Section 7 discusses policy implications and future research.

2. Institutional background

Prior to the adoption of the MFS, reimbursement for physician services was based on historical charges, using the CPR system. Under this system, the 'allowed' charge that a physician received was the lowest of (1) the 'actual' charge by the physician; (2) the physician's 'customary' charge for the service (defined as the physician's median charge for that service during the previous year); (3) the 'unadjusted prevailing' charge in the locality [3] (defined as the 75th percentile of the distribution of customary charges for all physicians in the locality); or (4) the 'adjusted prevailing' charge (defined as the prevailing charge in June 1973, updated by the Medicare Economic Index [MEI]).

Under OBRA89, the CPR has been gradually replaced since 1992 by the MFS, a fee schedule based on the Resource Based Relative Value Scale (Hsiao et al., 1988a,b). Under the MFS, relative fees for evaluation and management services increase, while those for surgical and diagnostic procedures decrease. Such fee changes can add up to profound changes in physician income. For example, Medicare income was projected to increase by 34% for general practitioners, but decrease by 26% for thoracic surgeons (Levy et al., 1990).

Before the implementation of the MFS, several laws were passed as part of OBRA87, OBRA89, and OBRA90, which reduced the prevailing charges for selected 'overpriced' procedures. In particular, OBRA87 reduced the prevailing charges for 12 procedure groups, effective April 1, 1988. Among these were CABG surgeries (CABG1–CABG6: one to six vessels). The prevailing charges for these procedures were reduced by 2%, and then further reduced between 0% and 15% by a sliding scale amount, depending on the deviation of the 1987 (locality) prevailing charges from the national average. [4] These changes implied that for the same procedure, physicians in Medicare localities where the prevailing charges were higher than the national average (e.g., New York City) experienced a greater reduction than those in localities where the prevailing charges were closer to or below the national average (e.g., most rural areas).

[3] A Medicare locality is a geographic area defined for the administration of Medicare fee policies.

[4] Prevailing charges above 150% of the national average were reduced by 15% on top of the 2%, while prevailing charges less than 85% of the national average received no further reduction. Prevailing charges between 85% and 150% of the national average were reduced by three-thirteenths of one percent for each percent of the prevailing charge exceeding 85% of the national average.

3. Basic model and predictions

The model used for motivating the empirical estimation is the McGuire and Pauly (1991) model of physician response to fee changes in a multiple-payer market. This model differs from previous models of physician behavior in two major aspects. First, rather than assuming the existence or non-existence of target income, it is a general model of physician behavior with a subjective inducement cost that encompasses both benchmark cases of profit maximization and target income. Second, it incorporates multiple payers, an important feature of the US health care system.

Formally, the physician supplies services to both the Medicare (m) and private (p) sectors, and maximizes utility ($U(Y, L, I_m, I_p)$) by 'deciding how much inducement to undertake' [5] in each market, subject to constraints on time ($24 = L + t_m I_m + t_p I_p$) and income ($Y = P_m Q_m(I_m) + P_p Q_p(I_p)$). Y is income; L is leisure; I_m and I_p are inducement in the Medicare and private markets, respectively; t_m, t_p are the time costs per unit of inducement effort; P_m, P_p are the net prices paid by Medicare and private payers, respectively; and Q is volume. Utility increases in income (Y) and leisure (L) but decreases in inducement (I). [6] Inducement is a choice variable, and utility maximization yields the optimal amount of Medicare and private inducement ($I_m *$, $I_p *$). Since quantity is increasing in inducement, the supply functions for Medicare and private services can be derived:

$$Q_m = Q_m(P_m, P_p, t_m, t_p) \tag{1}$$

$$Q_p = Q_p(P_m, P_p, t_m, t_p) \tag{2}$$

3.1. Income and spillover effects

The model shows that the strength of the income effect is the key determinant of a physician's volume response to a fee cut. At the one extreme, when the income effect is very strong and negative, the physician pursues a target income, increasing volume to completely compensate for the income lost from the price cut. At the other extreme, when the income effect is small, the physician behaves like a profit-maximizing firm, reducing the volume of services when prices are reduced. In between these two polar cases, when the income effect is strong but not quite dominant, the physician exhibits a negatively sloped supply curve:

[5] McGuire and Pauly (1991), p. 390.

[6] Physicians are assumed to suffer disutility from inducement due to their 'internal conscience.'

quantity increases partially, but not enough to fully offset the price cut. To the extent that the private market provides another avenue for recouping income losses, the model also predicts that a strong and negative income effect will lead to an increase in private volume after a Medicare price cut (the spillover effect). Measuring the sign and magnitude of the income effect therefore provides guidance regarding the size and direction of the potential volume response to a contemplated price cut.

For thoracic surgeons, since revenue from CABGs usually constitutes a considerable share of total practice revenue, one would expect the income effect to dominate the substitution effect. On the other hand, given the time cost and the risk involved, CABG is a relatively costly procedure to induce. Moreover, thoracic surgeons have to rely on cardiologists for patient referrals, and therefore may not have ultimate control over the number of patients. However, thoracic surgeons and cardiologists do not necessarily have different or conflicting incentives, since any patient undergoing a CABG would bring an expected stream of income to the cardiologist as well. Without empirical evidence, it is therefore difficult to predict a priori whether thoracic surgeons would increase volume of services to compensate for income lost due to Medicare price reductions.

3.2. The relative price effect

The model also shows that in the context of multiple payers, there is a second substitution effect beyond the standard substitution effect between labor and leisure. The existence of the private sector provides another avenue to recoup income losses. The extent to which physicians substitute private for Medicare inducement depends on the relative costs and returns of inducement in the private and the Medicare sectors, such as the relative ease of inducement, the sensitivity of demand to inducement, and the relative payment for services in each market. Among these factors, the relative payment of private payers to Medicare is of special interest because it is an observable market variable that has policy importance. Since inducement is costly, as private sector payment increases, the relative net return to inducement in the private sector also increases, making it more attractive to induce in the private sector. I will call this differential inducement effect due to differences in payment rates of the private and Medicare sectors the ‘relative price’ effect here after. [7]

4. Empirical estimation

To analyze physician response in volume, a first-difference model is used. This model purges physician and locality fixed effects that can potentially be correlated

[7] For readers who are not familiar with this theoretical model, please see McGuire and Pauly (1991).

with Medicare price changes and hence bias the estimates of Medicare price changes on volume. I also construct a BITE variable to measure the income effect of a price change. Linearizing Eqs. (1) and (2) in Section 3,

$$Q^{\mathrm{m}}_{ijt} = \beta^{\mathrm{m}} P^{\mathrm{m}}_{tj} + \gamma^{\mathrm{m}} P^{\mathrm{p}}_{itj} + X_{jt}\delta^{\mathrm{m}} + \alpha^{\mathrm{m}}_i + \mu^{\mathrm{m}}_j + \nu^{\mathrm{m}}_t + \varepsilon^{\mathrm{m}}_{ijt} \tag{3}$$

$$Q^{\mathrm{p}}_{ijt} = \beta^{\mathrm{p}} P^{\mathrm{m}}_{tj} + \gamma^{\mathrm{p}} P^{\mathrm{p}}_{itj} + X_{jt}\delta^{\mathrm{p}} + \alpha^{\mathrm{p}}_i + \mu^{\mathrm{p}}_j + \nu^{\mathrm{p}}_t + \varepsilon^{\mathrm{p}}_{ijt} \tag{4}$$

Q^{m}_{ijt} is the number of times the procedure (for example, CABG3) is performed on Medicare patients by physician i in Medicare locality j in year t. P^{m}_{tj} is the Medicare prevailing charge for the overpriced procedure in locality j in year t; P^{p}_{itj} is the private price that physician i in locality j receives in year t; X_{jt} is a vector of time-varying exogenous demand shifters at the locality level; α_i is a physician- specific effect; μ_j is a Medicare locality fixed effect; ν_t is a time effect; and $\varepsilon^{\mathrm{m}}_{ijt}$ is a random error. Medicare is indexed by ‘m’ and the private sector by ‘p’.

4.1. Income and spillover effects

In this specification, the β’s incorporate both the income and substitution effects of a Medicare price change. As described in the theoretical model section, whether or not physicians respond to a price cut by increasing volume depends on the strength of the income effect. To separate the income effect from the substitution effect, a BITE variable is constructed to measure the impact on volume of price change as it changes the income of the physician’s practice. [8] Eqs. (3) and (4) are amended to:

$$Q^{\mathrm{m}}_{ijt} = \Psi^{\mathrm{m}} P^{\mathrm{m}}_{tj} + \pi^{\mathrm{m}}\mathrm{BITE}_{ti} + \gamma^{\mathrm{m}} P^{\mathrm{p}}_{itj} + X_{jt}\delta^{\mathrm{m}} + \alpha^{\mathrm{m}}_i + \mu^{\mathrm{m}}_j + \nu^{\mathrm{m}}_t + \varepsilon^{\mathrm{m}}_{ijt} \tag{5}$$

$$Q^{\mathrm{p}}_{ijt} = \Psi^{\mathrm{p}} P^{\mathrm{m}}_{tj} + \pi^{\mathrm{p}}\mathrm{BITE}_{ti} + \gamma^{\mathrm{p}} P^{\mathrm{p}}_{itj} + X_{jt}\delta^{\mathrm{p}} + \alpha^{\mathrm{p}}_i + \mu^{\mathrm{p}}_j + \nu^{\mathrm{p}}_t + \varepsilon^{\mathrm{p}}_{ijt} \tag{6}$$

The ψ’s capture the substitution effect and the π’s capture the income effect of a price change, respectively. Conceptually, the size of the income effect varies with the share of the physician’s practice income derived from Medicare overpriced procedures. Under OBRA 87, overpriced procedures for thoracic surgeons included CABGs (i.e., CABG1–CABG4 [9]) and pacemaker insertions. In other words, faced with the same per-unit price change, a physician with a larger share

[8] BITE parallels the ‘Bite’ variable used in the Prospective Payment System (PPS) literature for measuring the financial pressure of PPS on hospitals (Staiger and Gaumer, 1990; Gruber, 1994). Conceptually, the ‘Bite’ variable is Medicare’s PPS payment to a hospital weighted by the hospital’s share of costs covered by Medicare in the year prior to PPS.

[9] As explained in the Section 5, due to different coding methods, CABG4–CABG6 cannot be distinguished, so they are all aggregated as one procedure: CABG4.

of Medicare overpriced procedure income experiences a larger income effect than one with a small share of Medicare overpriced income. Formally,

$$\mathrm{BITE} = \frac{P_j^{\mathrm{m1}} Q_i^{\mathrm{m1}} + P_j^{\mathrm{m2}} Q_i^{\mathrm{m2}} + P_j^{\mathrm{m3}} Q_i^{\mathrm{m3}} + P_j^{m4} Q_i^{\mathrm{m4}} + P_j^{\mathrm{m5}} Q_i^{\mathrm{m5}}}{\sum_{n=1}^{N} Q_i^n}$$

where P_j^{m1} = Medicare price for CABG1, and similarly for CABG2–CABG4; P_j^{m5} = Medicare price for pacemaker insertions; Q_i^{m1} = mean Medicare CABG1 volume in 1987 and 1989, similarly for CABG2–CABG4 and pacemaker insertions; Q_i^n = mean total quantity of each service (overpriced and non-overpriced, and for all payers) that the surgeon performed in 1987 and 1989. Differencing this model between periods $t-1$ and t yields the following:

$$\Delta Q_{ij}^{\mathrm{m}} = \Delta \nu_t^{\mathrm{m}} + \Psi^{\mathrm{m}} \Delta P_j^{\mathrm{m}} + \pi^{\mathrm{m}} \Delta \mathrm{BITE}_i + \gamma^{\mathrm{m}} \Delta P_j^{\mathrm{p}} + \Delta X_{jt} \delta^{\mathrm{m}} + \Delta \varepsilon_{ij}^{\mathrm{m}} \qquad (7)$$

$$\Delta Q_{ij}^{\mathrm{p}} = \Delta \nu_t^{\mathrm{p}} + \Psi^{\mathrm{p}} \Delta P_j^{\mathrm{m}} + \pi^{\mathrm{p}} \Delta \mathrm{BITE}_i + \gamma^{\mathrm{p}} \Delta P_j^{\mathrm{p}} + \Delta X_{jt} \delta^{\mathrm{p}} + \Delta \varepsilon_{ij}^{\mathrm{p}} \qquad (8)$$

Using OBRA87, which took effect in 1988, the corresponding $t-1$ and t periods are 1987 and 1989. By first differencing, the unobserved physician and locality fixed effects that may be correlated with the price variable are eliminated. [10] ΔBITE is the first difference version of BITE and is equal to the weighted sum of Medicare price changes faced by the physician for all CABGs and pacemaker insertions, with the weights equal to the shares of the respective Medicare procedures in the physician's total practice volume. That is, [11]

$$\Delta\mathrm{BITE} = \frac{\Delta P_j^{\mathrm{m1}} Q_i^{\mathrm{m1}} + \Delta P_j^{\mathrm{m2}} Q_i^{\mathrm{m2}} + \Delta P_j^{\mathrm{m3}} Q_i^{\mathrm{m3}} + \Delta P_j^{\mathrm{m4}} Q_i^{\mathrm{m4}} + \Delta P_j^{\mathrm{m5}} Q_i^{\mathrm{m5}}}{\sum_{n=1}^{N} Q_i^n}$$

Theoretically, there is no justification for preferring the base-, terminal-, or mean-year volumes as weights. However, econometrically, since the dependent variable is the difference between the terminal- and base-year physician volumes, using either the base- or terminal-year volumes as weights can lead to potential bias of ΔBITE as a result of mean reversion. In particular, if the errors are

[10] For example, a physician with high motivation is likely to have high volume and quality. High quality, however, is correlated with high price. Then, measuring the effect of price on volume without controlling for the unobserved fixed effect (motivation) leads to biased estimates of the price variable. First differencing removes all time-invariant fixed effects, resulting in unbiased estimates of key independent variables.

[11] This is analogous to the Slutsky decomposition of the income and substitution effects of a price change, where ΔBITE is the change in gross income normalized by total practice volume.

identically and independently distributed, using the base-year volumes as weights leads to negative bias, while using the terminal-year volumes as weights leads to positive bias. However, using the mean-year volumes will result in unbiased estimates. [12]

ΔBITE, however, is measured with error. To the extent that there may be non-exogenous price changes of other non-overpriced procedures, and by other payers, ΔBITE may understate or overstate the true income effect. Also, ΔBITE should ideally be normalized by total practice revenue rather than total practice volume. [13] If there are wide variations in the prices of the different services that the thoracic surgeon supplies, then total practice volume may understate or overstate total practice income. Consequently, the *t*-statistic is biased towards zero.

I assume that in Eqs. (5) and (6), the physician-specific private price (P_{itj}^{p}) is made up of two components, the individual physician's markup $(P_{\alpha i}^{\mathrm{p}})$ above the market private price (P_{tj}^{p}), that is, $P_{itj}^{\mathrm{p}} = P_{tj}^{\mathrm{p}} + P_{\alpha i}^{\mathrm{p}}$. I also assume that $P_{\alpha i}^{\mathrm{p}}$ is constant over time. First differencing implies that $\Delta P_{ij} = \Delta P_j^{\mathrm{p}}$, which is assumed to be exogenous, at least to the individual physician. [14] $\Delta \nu_t$ captures any trend and technology effects on volume. If the income effect is significant in determining the volume increase in response to a Medicare price cut, the π's will be negative and significant in both Eqs. (7) and (8). [15]

4.2. The relative price effect

To capture the heterogenous effects of relative payments on spillover of inducement, interactive variables between ΔP_j^{m} and $(P_j^{\mathrm{p}}/P_j^{\mathrm{m}})$ and between

[12] In the case where serial correlation between 1987 and 1989 exists, I assume that the correlation takes the specific form in which the correlation is captured in the physician fixed effect (α_i^{m}); hence by first differencing, the $\Delta \varepsilon_{ij}^{\mathrm{m}}$ only contains random noise and is not correlated with the α_i^{m} inherent in the ΔBITE variable. This assumption, however, cannot be tested with the data available.

[13] Data on volume for this study is based on hospital discharge files, and data on Medicare and private prices is based on Medicare Part B and Health Insurance Association of America's (HIAA) data, respectively. While procedures are identified by ICD-9 codes in hospital discharge data, they are identified by CPT-4 in Medicare Part B and HIAA data. For a considerable number of procedures, there is no one-to-one correspondence between ICD-9 and CPT-4 codes; therefore, income cannot be calculated using these two data sources. Furthermore, physician identifiers used in the hospital discharge files cannot be linked to other physician surveys which may contain data on physician income or practice revenue.

[14] Data on private prices are only available at aggregates of three-digit zip level, therefore this assumption cannot be tested for more formally.

[15] If the different CABG procedures are close substitutes, the coefficient estimates of ΔBITE could be negative as a result of cross price effects even if there were no income effect.

ΔBITE and $(P_j^{\mathrm{p}}/P_j^{\mathrm{m}})$ are included.

$$\Delta Q_{ij}^{\mathrm{m}} = \Delta\nu_t^{\mathrm{m}} + \Psi^{\mathrm{m}}\Delta P_j^{\mathrm{m}} + \pi^{\mathrm{m}}\Delta \mathrm{BITE}_i + \theta^{\mathrm{m}}\Delta\, BITE_i \times \left(P_j^{\mathrm{p}}/P_j^{\mathrm{m}}\right)_t + \varphi^{\mathrm{m}}\Delta P_j^{\mathrm{m}} \times \left(P_j^{\mathrm{p}}/P_j^{\mathrm{m}}\right)_t + \gamma_t^{\mathrm{m}}\Delta P_j^{\mathrm{p}} + \Delta X_{jt}\delta^{\mathrm{m}} + \Delta\varepsilon_{ij}^{\mathrm{m}} \tag{9}$$

$$\Delta Q_{ij}^{\mathrm{p}} = \Delta\nu_t^{\mathrm{p}} + \Psi^{\mathrm{p}}\Delta P_j^{\mathrm{m}} + \pi^{\mathrm{p}}\Delta \mathrm{BITE}_i + \theta^{\mathrm{p}}\Delta \mathrm{BITE}_i\left(P_j^{\mathrm{p}}/P_j^{\mathrm{m}}\right)_t + \varphi^{\mathrm{p}}\Delta P_j^{\mathrm{m}}\left(P_j^{\mathrm{p}}/P_j^{\mathrm{m}}\right)_t + \gamma_t^{\mathrm{p}}\Delta P_j^{\mathrm{p}} + \Delta X_{jt}\delta^{\mathrm{p}} + \Delta\varepsilon_{ij}^{\mathrm{p}} \tag{10}$$

A higher relative price ratio moderates the income effect in the Medicare sector, but augments the income effect in the private sector, as returns to inducement efforts are higher in the private than in the Medicare sector. θ^{m} is therefore hypothesized to be positive, while θ^{p} is hypothesized to be negative. Independent of the income effect, a higher relative payment in the private sector should also lead the physician to substitute more private for Medicare procedures; hence, both φ^{m} and φ^{p} are hypothesized to be negative.

There are altogether eight equations, one for each of CABG1–CABG4, and for the Medicare and private sectors separately. Since the unit of observation is the individual physician, who is likely supplying all or most of the four types of CABG procedures in each payer market, the error terms in each equation would be correlated. Seemingly unrelated regression (SUR) is therefore used for estimation in order to account for correlations across equations to produce more efficient estimates.

5. Data and variables

Since no standard medical survey or claims records serve the purpose of this analysis, a data set was assembled from various sources to provide information on the volume and price of both Medicare and private sector services for a panel of uniquely identified physicians. The primary data source was state inpatient discharge abstracts from New York and Washington. These two states [16] were chosen because the data sets contain physician identifiers that can be linked between 1987 and 1989 at the time of the study. Based on physician identifiers, procedure codes, and payer variables, measures of Medicare and private volume for each CABG procedure were created, with volume defined as the number of times the procedure was performed by the physician in a year. All patients who

[16] These two states consist of 12 Medicare localities.

report the primary payer as commercial insurance or Blue Shield are counted as private patients. [17] The independent variables include the following.

Medicare prices. Medicare prevailing charges were computed from HCFA's BMAD Prevailing Files. [18] These locality-level [19] Medicare prices were merged with physician volume data by county of physician practice location. [20] Prevailing charges (the ceiling of allowed charges) were used as the Medicare price measure. Though the actual allowed charges more accurately reflect physician reimbursements, they are endogenous. [21] Since changes in prevailing charges are legislated, they are relatively exogenous. [22]

Private prices. Private prices are based on the Health Insurance Association of America's (HIAA) Surgical Healthcare Prevailing Charges System. These are three-digit, zip-code-level private list prices based on claims submitted to private insurance companies, Blue Shield, and the self-insured. [23]

Other covariates. Data on county-level factors hypothesized to affect demand was obtained from the Area Resource File. Since a first difference model is used, only factors that changed between 1987 and 1989 were included. A dummy variable indicating counties in an MSA with a population of more than 1 million

[17] HMOs' and other types of health care service contractors' patients are excluded from private volume counts because these payers usually have very different payment arrangements with providers and hence will elicit different behavioral responses.

[18] HCFA and other researchers familiar with the Medicare BMAD Prevailing Charge File claim that these data are 'unreliable'. Prevailing charge data were also obtained from Medicare carriers directly, to cross check with the data from the BMAD Prevailing charge files. For New York, prevailing charge data are only available from carrier 00801 (which covers localities 80101–80104), and these data do not significantly deviate from the BMAD data. In Washington, carrier prevailing charges tend to be higher than the BMAD charges. Carrier prevailing charges were used for analysis because changes in the prevailing charges between 1987 and 1989 obtained from the carrier approximated more closely the changes according to the price reduction rule under OBRA87. For New York, BMAD data are used.

[19] Physician-level allowed charges are available from the BMAD Provider and Beneficiary Files; however, there exist no common identifiers between these files and the inpatient discharge abstracts such that physicians can be linked.

[20] Provider files containing zip codes of physician practice were obtained from the New York State Department of Health and Healthcare Business Services of Washington.

[21] The allowed charge is the lowest of the actual, customary, or prevailing charge; see Section 2.

[22] Based on BMAD Beneficiary files, the percentage of claims constrained by prevailing charges for CABGs was approximately 75% in 1987 and 90% in 1989 for New York; and 65% in 1987 and 80% in 1989 for Washington. Under the CPR system, physicians could choose to accept assignment on a claim-by-claim basis. If the physician accepted the assignment, he/she agreed to accept the Medicare allowed charge as full payment. Otherwise, he/she could 'balance bill' by charging patients an amount above the Medicare allowed charge. Information on assignment is not available from the data set I used. To the extent that physicians can charge above the prevailing charges, the effect of price controls based on prevailing charges can be mitigated by balance billing. Based on BMAD Beneficiary files, the percentage of assigned claims for CABGs was approximately 75% in 1987 and 85% in 1989 for New York, and 65% in 1987 and 80% in 1989 for Washington.

[23] The data does not contain any information in discounts.

(Large MSA) is also included. Since Medicare localities that experience steep prevailing charge cuts are mostly concentrated in urban areas, the Medicare price measure may capture urban/rural differences in utilization trends. By first differencing, unobserved time-invariant urban/rural factors will be removed. If there still remain time-variant factors that are correlated with the Medicare fee change, estimates of the Medicare price change will be biased. The urban dummy is included to capture these differences in urban/rural trend effects. Physician density (Phy/Pop) in an area is included to control for differences in utilization associated with physician supply that may be correlated with Medicare fee cuts. Furthermore, to adjust for differences in cost of production, the change in the health care wage index (Wage) from HCFA is included. In the private equation, an additional variable, the change in the percentage of the population enrolled in HMOs (HMO/Pop) is also included. If the growth of managed care during the study period occurred preferentially in areas with relatively small Medicare fee reductions, an increase in private volume in areas with large fee cuts may be misinterpreted as evidence of physician's behavioral response to fee cuts.

The goal of the sample is to include all thoracic surgeons in New York and Washington. The data, however, do not provide information on physician specialty. All physicians who performed any CABGs in either 1987 or 1989 constitute a sample of 439 physicians. Around 50% of these physicians were excluded because they performed fewer than 10 CABGs (irrespective of payer and vessels) for 1987 and 1989 combined. Also excluded were all physicians for whom more than 60% [24] of procedures were diagnostic and radiology procedures (such as catheterization and ultrasound of the heart) or procedures that are commonly performed by cardiologists (e.g., percutaneous transluminal coronary angioplasty). These exclusions resulted in a sample of 232 physicians. [25]

6. Results

Tables 1 and 2 show summary statistics on prices and volume for both the Medicare and private sectors. Between 1987 and 1989, while Medicare prices fell as a result of OBRA87, both total Medicare and private volumes increased. Table 3a presents summary statistics on the changes in physician volume of CABGs (the dependent variables of interest) and the ΔBITE variable. Table 3b provides descriptive statistics on physicians' practice characteristics in 1987. On average, CABGs (irrespective of payer type) account for 41% of total practice volume, with a payer mix of 51% for Medicare, 36% for private, and 13% for health services

[24] I also tried 50% and 70% cutoffs or the entire sample; the results remain robust.

[25] The total number of thoracic surgeons in New York and Washington was 219 in 1989 according to the 1992 Area Resources File.

Table 1
Summary statistics for medicare prevailing charges and private prices

Procedure	Medicare prices		Private price		Private to medicare ratio	
	Changes	Percentage	Changes	Percentage	1987	1989
CABG1	−US$289.2[a] (278.9)[b]	−6.6 (6.1)	US$789.4 (602.2)	19.1 (15.7)	1.04 (0.15)	1.33 (0.13)
CABG2	−506.5 (422.3)	−9.0 (7.2)	357.0 (477.7)	8.4 (12.6)	1.06 (0.10)	1.29 (0.16)
CABG3	−569.4 (532.0)	−9.5 (8.5)	518.1 (424.4)	11.2 (16.1)	1.08 (0.12)	1.34 (0.17)
CABG4	−557.3 (370.9)	−10.5 (6.8)	632.7 (562.4)	10.1 (9.8)	1.18 (0.12)	1.46 (0.12)

[a] Sample mean weighted by the number of physicians in the Medicare locality.
[b] Standard deviations.
CABG4: prices for CABG4 are weighted averages of prices for CABG4, CABG5, and CABG6, where the weights are locality-level quantities.

contractors, Medicaid and self pay patients. Medicare CABG therefore accounted for approximately 21% of thoracic surgeons' practice volume in this sample. [26]

6.1. The income effect

The results from the SUR estimation of Eqs. (7) and (8) are presented in the first column of each block of Table 4A and B. They support the hypothesis that there are negative and significant income effects from fee changes, and that there is spillover of the income effects into the private sector. The estimated coefficients of ΔBITE are negative across all equations, and except for Medicare CABG1, all are statistically significant. These estimated coefficients translate into income elasticities of considerable magnitude (Table 5‘). Using relative volumes as weights, the average income effect across all CABG procedures and across both payers is approximately -0.74. On the other hand, the substitution effects are relatively smaller, and all of them are imprecisely estimated. The lack of statistical significance of the own price variable can be partly due to multicollinearity between the price and ΔBITE variables. The Breush–Pagan test of independence rejects the null hypothesis that there is no correlation across the different equations at the 99% confidence level ($\chi^2(28) = 351.13$).

6.1.1. Growth of other technology

Note that the absolute magnitudes of the income effects are in general bigger for CABG3 and CABG4 than for CABG1 and CABG2. This can partly be explained by the increased use of percutaneous transluminal coronary angioplasty (PTCA) during the 1980s as a substitute for single vessel, and to a smaller extent double vessel, bypass procedures. During that decade, the primary success rate of PTCA reached almost 85% (although the restenosis rate—recurrence of the

[26] Based on information from a national survey of physicians by Schneider and Rosenbach (1991) and the BMAD Procedure Files, Mitchell and Cromwell (1995) estimated that the national average of Medicare CABG as a share of thoracic surgeons' practice revenue is 21.6%.

Table 2
Total volume of each procedure by payer

Procedure	Medicare			Private		
	1987	1989	Percentage change	1987	1989	Percentage change
CABG1	524 (3.0)	694 (4.0)	32.4	556 (3.3)	594 (3.5)	6.8
CABG2	1352 (7.2)	1589 (8.7)	17.5	1116 (6.3)	1210 (6.8)	8.4
CABG3	2438 (12.2)	2909 (14.5)	19.3	1936 (10.0)	2103 (11.0)	8.6
CABG4	2761 (13.5)	3043 (15.1)	10.2	2567 (13.0)	2736 (13.9)	6.6

Volume is defined as the number of procedures in a year.
CABG4 includes volume for CABG4, CABG5, and CABG6.
Mean physician-level volume in parentheses.

disease at the site of the angioplasty—was also quite high at around 25%). For New York state, the total volume of PTCA grew almost 30% from 6403 in 1987 to 8515 in 1989 (New York State Department of Health, 1990). As the use of PTCA

Table 3
(a) Summary statistics of dependent variables and BITE

	Mean	S.D.	Min	Max
Change in physician-level Medicare volume				
CABG1	0.98	3.31	−6	11
CABG2	1.30	6.25	−18	33
CABG3	2.35	9.27	−31	53
CABG4	1.40	11.36	−42	47
Change in physician-level private volume				
CABG1	0.22	3.63	−16	13
CABG2	0.52	6.62	−34	32
CABG3	0.88	8.66	−31	33
CABG4	0.86	10.65	−39	49
ΔBITE	−US$107.1	US$219.1	−US$1923.2	US$1260.9

(b) Summary statistics of service and payer mix of physician practice in 1987

	Mean	S.D.
Total CABG as a share of total practice volume	0.409	0.261
Medicare CABG as a share of total practice volume	0.213	0.146
Medicare CABG as a share of total CABG volume	0.514	0.159
Private CABG as a share of total CABG volume	0.360	0.145

CABG volume includes volume counts for CABG1–CABG4.
Total CABG volume includes volume counts for all payers.
Total practice volume include volume counts of CABGs and other procedures in the physician practice.

Table 4

(A) SUR results of Eqs. (7) and (9)—Medicare

	CABG1		CABG2		CABG3		CABG4	
	Eq. (7)	Eq. (9)	Eq. (7)	Eq. (9)	Eq. (7)	Eq. (9)	Eq. (7)	Eq. (9)
ΔP^{m}	0.0007	0.0024	0.0001	−0.0162	−0.0015	−0.0161	−0.0010	0.0567**
	(0.0010)	(0.0080)	(0.0022)	(0.0136)	(0.0019)	(0.0168)	(0.0039)	(0.0237)
ΔBITE	−0.0020	−0.0188	−0.0092***	−0.0188	−0.0165***	0.0132	−0.0261***	−0.0078
	(0.0013)	(0.0141)	(0.0024)	(0.0219)	(0.0034)	(0.0319)	(0.0044)	(0.0416)
ΔBITE(P^{p}/P^{m})	–	0.0124	–	0.0074	–	−0.0207	–	−0.0118
		(0.0104)		(0.0170)		(0.0221)		(0.0293)
$\Delta P^{m}(P^{p}/P^{m})$	–	−0.0010	–	0.0121	–	0.0098	–	−0.0416**
		(0.0060)		(0.0103)		(0.0110)		(0.0172)
ΔP^{p}	0.0006	0.0008	0.0015	0.0018	0.0027**	0.0029**	0.0008	−0.0015
	(0.0005)	(0.0006)	(0.0010)	(0.0011)	(0.0012)	(0.0013)	(0.0015)	(0.0017)
Large MSA	−0.8034	−1.0224	−2.7969	−2.5900	−6.9233***	−7.5448***	−7.6512	−6.3971
	(0.7211)	(0.8205)	(1.8237)	(1.8379)	(2.0257)	(2.3357)	(4.3876)	(4.4578)
Δ(Phy/Pop)	−0.0139	−0.0024	−0.0272	0.0807	−0.0975	−0.0934	−0.2549*	−0.6602***
	(0.0420)	(0.0432)	(0.0741)	(0.1005)	(0.0971)	(0.1058)	(0.1375)	(0.1996)
Δ(Wage)	−32.4855*	−33.8624*	−1.9422	6.4780	12.9610	22.2006	−47.7943	−125.7596**
	(17.7764)	(18.3369)	(34.4275)	(35.0584)	(48.6312)	(53.9188)	(59.4809)	(65.1956)
Constant	0.7658	0.6255	1.2178	0.0768	2.9499*	3.1135*	3.8069	5.8556***
	(0.5539)	(0.5805)	(1.0159)	(1.2476)	(1.5540)	(1.5962)	(2.1415)	(2.2331)
Coeff_[ΔBITE][a]		−0.0023		−0.0095***		−0.0136***		−0.0234***
		(0.0014)		(0.0026)		(0.0046)		(0.0049)
Coeff_[ΔP^{m}][a]		0.0011		−0.0009		−0.0034		0.0017
		(0.0014)		(0.0017)		(0.0029)		(0.0039)

(B) SUR results of Eqs. (8) and (10)—Private

	CABG1		CABG2		CABG3		CABG4	
	Eq. (8)	Eq. (10)	Eq. (8)	Eq. (10)	Eq. (8)	Eq. (10)	Eq. (8)	Eq. (10)
ΔP^{m}	−0.0007	−0.0013	−0.0019	−0.0265	−0.0023	−0.0190	−0.0019	0.0212
	(0.0012)	(0.0092)	(0.0026)	(0.0199)	(0.0022)	(0.0196)	(0.0038)	(0.0239)
ΔBITE	−0.0033**	0.0190	−0.0123***	−0.0127	−0.0149***	0.0200	−0.0230***	0.0288
	(0.0014)	(0.0150)	(0.0025)	(0.0229)	(0.0033)	(0.0308)	(0.0042)	(0.0393)
ΔBITE(P^{p}/P^{m})	–	−0.0165	–	0.0000	–	−0.0243	–	−0.0360
		(0.0111)		(0.0178)		(0.0214)		(0.0277)
$\Delta P^{m}(P^{p}/P^{m})$	–	0.0002	–	0.0177	–	0.0108	–	−0.0159
		(0.0067)		(0.0140)		(0.0122)		(0.0171)
ΔP^{p}	0.0012**	0.0007	0.0015	0.0020	0.0038**	0.0040**	0.0012	−0.0003
	(0.0005)	(0.0008)	(0.0011)	(0.0012)	(0.0017)	(0.0014)	(0.0015)	(0.0016)
Large MSA	−1.3741*	−0.9719	−3.6791*	−3.9239	−3.7267*	−4.7270*	−3.7267*	−6.0778
	(0.8312)	(1.0268)	(2.0232)	(2.1004)	(2.1241)	(2.8494)	(2.1241)	(4.2558)
Δ(HMO/Pop)	−7.0188	−8.9298*	−6.7718	−15.4242	−24.4274**	−30.3185**	−23.5543*	−22.6630**
	(4.7744)	(5.3482)	(9.3410)	(13.4527)	(12.3125)	(15.1399)	(13.4038)	(14.4475)
Δ(Phy/Pop)	−0.0650	−0.0792*	−0.1348	−0.0337	−0.2330**	−0.2439**	−0.2591*	−0.4772**
	(0.0447)	(0.0457)	(0.0844)	(0.1070)	(0.0987)	(0.1033)	(0.1343)	(0.1901)
Δ(Wage)	−55.8732**	−55.2936**	−59.5889	−59.4005	−60.6130	−60.1740	−38.3175	−82.4791
	(19.2181)	(19.5244)	(37.1617)	(37.7185)	(51.1077)	(53.2471)	(57.4806)	(62.4863)
Constant	0.1320	0.4245	0.8035	0.0048	1.0907	1.6108	4.5071*	5.7146**
	(0.6725)	(0.7579)	(1.2383)	(1.3166)	(1.6542)	(1.7018)	(2.3030)	(2.3355)
Coeff_[ΔBITE]		−0.0029***		−0.0126*		−0.0115**		−0.0188***
		(0.0010)		(0.0687)		(0.0043)		(0.0049)
Coeff_[ΔP^{m}]		−0.0010		−0.0042		−0.0050		0.0002
		(0.0014)		(0.0033)		(0.0044)		(0.0039)

[a] Coeff_ [ΔBITE] and Coeff_ [ΔP^{m}] are the marginal effects of ΔBITE and ΔP^{m} on the dependent variable of interest, respectively, taking into account the first order effect and the interactive effect with (P^{p}/P^{m}), evaluated at the mean of (P^{p}/P^{m}).

Standard errors in parentheses. *Statistically significant at 10% level, **statistically significant at 5% level, ***statistically significant at 1% level, based on two-tailed test.

Table 5
Estimated elasticities

	CABG1	CABG2	CABG3	CABG4
Medicare				
Income Elasticity	−0.108 (0.066)	−0.457*** (0.125	−.710*** (0.240)	−1.257*** (0.263)
Private				
Income Elasticity	−0.122*** (0.042)	−0.554*** (0.114)	−0.552** (0.206)	−0.929*** (0.242)

became the norm of the profession for diagnoses that are candidates of both CABG and PTCA, cardiologists opted for PTCA, and thoracic surgeons lost referrals for candidate patients for CABG1, making inducement in this procedure more difficult. The growth of PTCA can possibly bias the coefficient estimates of ΔBITE if it is correlated with Medicare CABG fee cuts. That is, if PTCA grew faster in areas with smaller fee cuts, a larger volume growth in areas with greater fee cuts may be misinterpreted as evidence of physician behavioral response to fee cuts. To test for this possibility, Eqs. (7) and (8) are re-estimated, including the change in locality-level PTCA volume as an independent variable. The coefficient estimates for the PTCA variables are both economically and statistically insignificant, and the estimates for ΔBITE remain robust.

6.1.2. Practice size effect

In the market for physician services, one would expect growth in demand to be distributed proportionally to physician practice volume (due to better quality and/or referral). To make sure that the coefficient estimates of ΔBITE are not capturing this practice size effect, Eqs. (7) and (8) are re-estimated, including total base year CABG volume (regardless of number of vessels or payer type) of the physician as an independent variable. The coefficient estimates for ΔBITE remain robust, and there is no clear pattern of relationship between practice size and volume growth. These results indicate that the income effect detected is not an artifact of the practice size effect. [27]

6.1.3. Exclusion of other procedures

Under OBRA87, a number of other overpriced procedures also experienced Medicare fee reductions. How much confidence can one attach to the detection of income effects for CABGs as evidence of income effects for physicians in general? In other words, how can one be sure that the findings for CABG are not a matter of chance? To answer these questions, I apply the same empirical test to all the other overpriced procedures to the extent that my data allow them to be identified. I then employ the sequentially rejective Bonferroni test for hypothesis testing (Holm, 1979), which has a prescribed level of significance protection

[27] Results are available from the author.

against Type I error—error of rejecting the null hypothesis when it is true, in the context of multiple procedures testing. Among the 13 [28] procedures examined, the null hypothesis that there were no income effects was rejected for CABG2–CABG4, bronchoscopy with biopsy, and upper gastrointestinal endoscopy with biopsy for the Medicare sector; and CABG2–CABG4, upper gastrointestinal endoscopy with biopsy, and knee arthroscopy for the private sector (Yip, 1995). The CABG results therefore survive the rather conservative Bonferroni test for multiple procedure testing.

6.2. The relative price effect

There does not seem to be evidence that the payment of private payers relative to that of Medicare is important in determining the amount of spillover. None of the coefficient estimates of the interactive variables, $\Delta\mathrm{BITE}(P_j^{\mathrm{p}}/P_j^{\mathrm{m}})_t$, are significant (Table 4A and B). Surprisingly, independent of the income effect, a higher private price relative to Medicare price also does not lead physicians to substitute private for Medicare procedures. Except for Medicare CABG4, none of the coefficient estimates for $\Delta P_j^{\mathrm{m}}(P_j^{\mathrm{p}}/P_j^{\mathrm{m}})_t$ is significant. This probably suggests that Medicare and private procedures are not perfectly substitutable at the physician practice level, or perhaps there is insufficient demand in the private sector. *F*-test results indicate that both interactive variables are jointly insignificant for all equations. Another possible explanation for failing to detect any ‘relative price’ effect may be measurement error in the private-to-Medicare price ratio. The data on private prices are listed prices; the data does not provide any information on discounts. Table 1 shows that on average private prices are 30% above those of Medicare, but this difference may be overstated if discounts exist. Also, private prices are based on aggregate data which may not reflect what physicians really receive. Alternatively, the lack of statistical significance can be a result of high multicollinearity among the main and interaction effects.

6.3. Other covariates

Results for other covariates are largely consistent with expectations. With the recent growth in managed care, it is perhaps worth highlighting the general negative coefficient estimates for the (HMO/Pop) variable, which indicate that areas with greater HMO penetration are associated with slower growth in private volume. When growth of HMOs is not controlled for, the coefficient estimates for the BITE variables are consistently less negative in all private equations, suggesting that HMO penetration is correlated with Medicare fee cuts, and that areas with greater fee cuts are associated with higher HMO penetration. Increased HMO

[28] They include CABG1-4, broncoscopy (with and without biopsy), upper gastrointestinal endoscopy (with and without biopsy), hip replacement, knee replacement, knee arthroscopy, knee arthroplasty, and transurethral prostatectomy (TURP).

penetration may exert competitive pressure on physicians, thus limiting their ability to induce demand. To test for this hypothesis, an interactive variable between ΔBITE and Δ(HMO/Pop) was attempted in estimation. Although the estimated coefficients are statistically insignificant, they all exhibit positive coefficients, lending support to the hypothesis that market competition may have some constraining effect on demand inducement.

6.4. Is there any treatment pattern change?

To the extent that increased intensity leads to increased income, the McGuire and Pauly model of physician behavior makes no distinction between a volume and an intensity response in the event of strong income effects. At the physician practice level, intensity response leads to a greater mix of more intensive procedures (for example, CABG3 and CABG4). The larger income elasticities detected for CABG3 and CABG4 are indicative of such response behavior. At the patient level, increased intensity can take the form of more ancillary services per bypass procedure, a decision probably in the hand of the cardiologist, and the additional income generated thereby would mostly accrue to the cardiologist rather than the thoracic surgeon. Intensity can also take the form of a more intensive bypass procedure (i.e., bypassing more vessels). While one must not infer that the medical provider is purely driven by financial incentives in making clinical decisions, there are circumstances of clinical uncertainty in which individual judgement is necessary. It is under those circumstances that financial incentives would play a role in clinical decision making. For example, before any bypass procedure, an angiogram is performed to provide the surgeon with a road map of sites of vessel occlusion. However, there is often ambiguity about whether certain lesions seen on the angiogram should be grafted, or about how severe the lesions are (e.g., 50% vs. 70%). The relevant question is, under these circumstances, would surgeons be more likely to be driven by financial incentives to bypass the additional vessel?

Since the price changes for all types of CABGs are highly correlated, it is not feasible to construct a ‘price’ change measure for intensity. Empirically, I use a difference-in-difference approach to test for whether there is any differential change in treatment intensity between areas with different fee cuts. I define Medicare localities that had prevailing charges above the national average in 1987 (and hence experienced bigger price cuts under OBRA87) as ‘binding’, and localities that had prevailing charges below the national average (and hence received minor price changes) as ‘non-binding’. [29] The question is whether patients in ‘Binding’ localities are more likely to be treated with more intensive

[29] Several other cutoffs for defining the ‘bindingness’ of a locality are also attempted; results are robust to these different definitions.

procedures than those in 'non-binding' localities, after OBRA87 went into effect. That is,

$$\left[(\text{Intensity})_{\text{B},t} - (\text{Intensity})_{\text{B},t-1}\right] > \left[(\text{Intensity})_{\text{NB},t} - (\text{Intensity})_{\text{NB},t-1}\right]$$

To operationalize the testing of this hypothesis, I use an ordered probit regression. Intensity is parameterized such that CABG4 is considered a more intensive treatment than CABG3, CABG3 more intensive than CABG2, and so on, given the diagnosis and characteristics of patients.

$$\Pr\begin{pmatrix}\text{CABG4}\\ \text{CABG3}\\ \text{CABG2}\\ \text{CABG1}\end{pmatrix} = \beta_1 \text{Bind} + \beta_2 \text{Yr89} + \beta_3 \text{BindYr89} + \beta_4 \text{BindYr89} * \text{BITE}$$

$$+ \beta_5 \begin{pmatrix}\text{Patient Age}\\ \text{Patient Sex}\\ \text{Diagnosis}\end{pmatrix} + \beta_6 \begin{pmatrix}\text{Large MSA}\\ \text{Wage}\\ \text{Phy/Pop}\\ \text{HMO/Pop}\end{pmatrix} + \epsilon$$

'Bind' is a dummy variable that equals one if the patient is in a 'Binding' locality. Yr89 is a dummy variable that equals one for 1989 and zero for 1987. BindYr89 is an interactive variable between Bind and Year89. The dependent variable is an index variable, for which CABG4 takes the highest value. β_1 captures the difference in treatment intensity between the binding and non-binding areas. β_2 captures the difference between 1987 and 1989. β_3 captures the difference in change in intensity between the two years between binding and non-binding areas. If intensity increases significantly in the binding areas between 1987 and 1989, then β_3 will be positive and significant. Parallel to the volume equations, β_4 measures whether patients treated by physicians with a bigger income effect are more likely to be given more intensive procedures after the Medicare fee cuts in 1988.

The results (Table 6) show that after the fee cut, patients in 'Binding' localities are more likely to be treated with more intensive CABG procedures than patients in 'Non-binding' localities are, controlling for patient characteristics. For Medicare patients, this intensity effect increases with the physician's share of practice that is affected by the exogenous price reduction.

To further understand where the increase in intensity occurs (i.e., from CABG1 to CABG2 or from CABG3 to CABG4, etc.), I performed three separate probit regressions for CABG1 vs. CABG2, CABG2 vs. CABG3, and CABG3 vs. CABG4. The results show that most of the increase in intensity is driven by increased probability of performing CABG4+ as opposed to CABG3. This has important policy implications in terms of cost containment. It implies that physicians faced with similar percentage reductions of Medicare prices for each kind of

Table 6
Treatment pattern change (Ordered probit regression: difference-in-difference analysis)

	Medicare ($N=14270$)		Private ($N=11915$)	
Bind	0.1921*** (0.0433)	−0.2036*** (0.0437)	0.0890* (0.0467)	0.0916* (0.0471)
Yr89	−0.1609*** (0.0325)	−0.1603*** (0.0325)	−0.0700* (0.0376)	−0.0699* (0.0376)
Bind*Yr89	0.1669*** (0.0405)	0.0835 (0.0591)	0.1469*** (0.0460)	0.1263** (0.0654)
Bind*Yr89* BITE	–	0.000047* (0.000024)	–	0.000012 (0.000027)
Age	0.0000 (0.0015)	0.0000 (0.0015)	0.0093*** (0.0013)	0.0093*** (0.0013)
Female	−0.2325*** (0.0197)	−0.2331*** (0.0198)	−0.2513*** (0.0266)	−0.2512*** (0.0266)
Large MSA	−0.0581* (0.0316)	−0.0653* (0.0318)	−0.1502*** (0.0340)	−0.1519*** (0.0342)
Wage	−1.222*** (0.1459)	−1.2525*** (0.1467)	−0.3819** (0.1739)	−0.3876** (0.1744)
Phy/Pop	0.0003*** (0.0001)	0.0003* (0.0001)	0.0005*** (0.0002)	0.0005*** (0.0002)
HMO/Pop	–	–	−0.3275*** (0.0780)	−0.3300*** (0.0800)
AMI	0.9248*** (0.0963)	0.9248*** (0.0963)	0.6427*** (0.1463)	0.6424*** (0.1463)
Unstable Angina	1.1952*** (0.0802)	1.2723*** (0.0727)	0.8397*** (0.1298)	0.8395*** (0.1298)
Stable Angina	1.2721*** (0.0728)	1.2720*** (0.0728)	1.0678*** (0.1187)	1.0674*** (0.1187)
Chest Pain	1.4055*** (0.1364)	1.4091*** (0.1364)	1.1800*** (0.1644)	1.1787* (0.1645)
_cut1	−1.7813 (0.1828)	−1.8174 (0.1837)	−0.5167 (0.2097)	−0.5239 (0.2105)
_cut2	−0.9448 (0.1828)	−0.9801 (0.1837)	0.2744 (0.2097)	0.2672 (0.2105)
_cut3	0.0034 (0.1828)	−0.0385 (0.1837)	1.1372 (0.2097)	1.1300 (0.2105)

CABG will concentrate increase-in-volume responses in more intensive procedures, which are also the more expensive procedures.

Estimates on other covariates are consistent with expectations. Within the under-65 private sector, older patients are more likely to undergo more intensive procedures, probably because older patients are more likely to have multiple vessel occlusion. However, among Medicare patients, older patients are not likely to have more intensive procedures because of the higher risks involved. Conditioning on positive treatment, females are less likely to receive more intensive treatment. This may reflect that women have less severe disease after controlling for age. This may also reflect treatment bias against females for coronary diseases, as has been documented elsewhere (Ayanian and Epstein, 1991). It is interesting to note that patients in areas with higher HMO penetration are less likely to undergo the more intensive procedures, controlling for patient characteristics and diagnosis. While this is not conclusive evidence on the impact of competition on treatment pattern, it at least suggests that there may be a relationship between competition/managed care and treatment pattern that may warrant future research.

7. Discussion

There has been a long-standing debate in the physician service literature on both the existence and extent of physician demand inducement. McGuire and Pauly (1991) develop a general model of physician behavior which focuses on the extent of demand inducement, and examine the conditions under which inducement is strong enough to lead to a negative volume response to price changes. Their model shows that the strength of the income effect is the critical factor in determining the magnitude of demand inducement, suggesting that empirical research should focus on estimating the income effect, which is precisely the objective of this paper. This paper further adds to the existing empirical literature of demand inducement by incorporating the multiple-payer feature of the US health care market. The empirical results support the hypothesis that physicians compensate for income losses due to public price regulations by increasing supply in both the Medicare and private sectors. By assuming the existence of demand inducement, and focusing on estimating the magnitude of demand inducement, both the McGuire and Pauly model and this empirical study fall short of answering the debate on whether demand inducement exists or not. However, given asymmetric information between providers and patients, it may be argued that physicians would most likely induce; investigating how much physicians induce, and the impact of inducement, are perhaps more fruitful than investigating whether inducement exists or not per se. Empirically, there are also good reasons to believe that the CABG market is demand constrained, so that the results can be interpreted as demand inducement rather than a neoclassical backward bending supply curve. First, excess supply of surgical specialties is well documented, with thoracic surgeons at the top of the list (Rutkow, 1988; Leape, 1989). Second, even though elderly patients are relatively well insured by Medicare, private patients may still face high costs of use (in the form of cost sharing or lack of coverage.)

From a policy point of view, these results raise concerns regarding the extent to which price regulation, such as the Medicare Fee Schedule (MFS) can be relied on to control costs. Recognizing physicians' ability to mitigate income changes through their influence on volume and intensity of service supply, HCFA incorporated an asymmetric 50% 'volume offset' in the calculation of the fee schedule. That is, HCFA assumed that physicians experiencing a Medicare fee reduction would increase volume and intensity to recoup half of the revenue loss, while physicians receiving fee increases would not offset the revenue increase by reducing services. Results from this study suggest that thoracic surgeons recoup about 70% of their revenue loss by increasing volume and intensity of CABGs, making the MFS a less effective tool for controlling costs. In view of this, HCFA's incorporation of a 'volume offset' in its fee calculations seems to be justified. The question remains whether 50% is a reasonable assumption and whether the asymmetric treatment of cuts and increases in reasonable. Existing empirical evidence of income effects on physician labor supply is sparse and inconclusive (Sloan, 1974, 1975; Hurdle and Pope, 1989; Lee and Mroz, 1991; Rizzo and Blumenthal, 1994). Using a sample of young, self-employed physicians, Rizzo and Blumenthal found an income elasticity of -0.27 for male physicians, smaller than the income elasticities documented in the present study. This can be partly explained by differences in specialties, geographic locations, and years of practice experience. Future research may explore the possibility of differential volumes offset based on these characteristics to allow for a 'fairer' adjustment of the Medicare Fee Schedule based on these dimensions.

Evidence on the spillover of the income effects of Medicare fee changes into the private sector shows that analyses of fee policies that focus on a single sector are incomplete and can result in misleading conclusions about the efficacy of the policy. Since the private sector provides an alternative avenue for physicians to recoup income losses, controlling Medicare program costs (assuming that this is successful) may not result in significant control of economy-wide health care costs. However, demand inducement is limited by its cost and relative ease of inducement, which in turn depend on market conditions (McGuire and Pauly, 1991; Dranove, 1988). With increasing competition among payers and providers in the private market, rapid growth of managed care, widespread adoption of utilization review and care management, and increased information available for consumers on quality of care, will the spillover effect be reduced? Or will competition provide an overall constraint on physicians' inducement behavior? These are empirical questions that warrant future research taking into account the organizational structure and financial incentives of the growing managed care sector.

Several limitations of this study are worth mentioning. First, with only two states, results may not be easily generalized to the rest of the nation. With only 12 localities, variation in the Medicare price variable is limited, reducing the precision of estimation. Second, a controversial aspect of the volume offset is its assumption of asymmetric physician response (i.e., physicians experiencing a price increase will not reduce volume.) Given that OBRA87 only legislated price reductions but not price increases, this study cannot address this aspect of the volume offset. Third, this analysis is restricted to thoracic surgeons and a set of overpriced procedures, CABGs. There may be spillover of volume increases to non-overpriced procedures that are substitutes for, or complements of, CABGs. [30] Analysis of data such as the BMAD Beneficiary File (which has comprehensive records of all services billed for an episode of disease) can shed light on whether physicians respond to Medicare fee cuts by changing the service mix for patients and its impacts on cost control and patient outcomes.

Despite these limitations, this paper is the first attempt to decompose the income and substitution effects of a price change to offer a direct test of demand inducement. The results show that the income effect of a price change is an important predictor of volume response. An ideal measure of the income effect is the share of a physician's total practice income (including all services and payers) that is derived from Medicare overpriced procedures, for which the BITE variable constructed in this study is only an approximation. Since the income elasticity provides information on physicians' volume response behavior, future research should focus on this measure.

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[30] For example, physicians can increase the number of hospital visits or other diagnostic tests, which are services more malleable to inducement.

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