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MEDICARE'S INFLUENCE ON PRIVATE PHYSICIAN PAYMENTS

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ABSTRACT

We demonstrate Medicare's influence on private insurers' payments for physicians' services. Using a large administrative change in payments for surgical versus medical care, we find that private prices follow Medicare's lead. A $1 change in Medicare's fees moved private prices by $1.16. A second set of Medicare payment changes, which generated area-specific reimbursement shocks, had a similar effect on private sector prices. Medicare's influence is strongest in areas with concentrated insurers, small physician groups, and competitive physician markets. The public sector's influences on system-wide resource allocation and costs extend well beyond the share of health expenditures it finances directly.

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The United States spends 3.5 percent of GDP, or more than \$573 billion annually, on physician care and similar medical services.[1] The workings of the markets and public programs that allocate this care thus have substantial welfare implications (Chandra, Jena and Skinner 2011). When prices signal relationships between production costs and consumers' willingness to pay, they steer markets towards efficient outcomes. But most medical services are purchased through insurance, which purposefully severs consumers from the price mechanism (Gaynor, Haas-Wilson and Vogt 2000, Baicker and Goldman 2011).[2]

We ask how physicians and private insurers determine the prices that insurers pay on their beneficiaries' behalf. Using two sources of administratively induced price variation, we show that the fee schedules set by Medicare, the federal insurer of the elderly and disabled, exert a predominant influence over private insurers' payment rates. We show that changes in Medicare's prices move both relative payments across services and average payments across geographic areas. Further, we find that Medicare's influence varies significantly across markets. Consistent with intuition, theory, and past work (Dunn and Shapiro 2012, Dafny, Duggan and Ramanarayanan 2012), more concentration among insurers and less concentration among physicians are associated with lower reimbursement rates in the cross-section. In the context of our natural experiments, we find that these same two conditions predict stronger linkages between Medicare and private payments.

Policy makers increasingly recognize the relevance of provider payments to the health system's performance (Cutler 2011). Our results highlight Medicare's centrality as a large, public player. Specifically, they imply that private payment arrangements amplify Medicare's capacity to steer resources across both physician specialties and geographic areas. Resulting

[1]This figure comes from the "Physician and Clinical Services" line of the National Health Expenditure Data for 2012 (Centers for Medicare and Medicaid Services (CMS) 2013), inflated to 2013 dollars using the CPI-U. The non-physician part of this category includes freestanding outpatient clinics and some laboratories.

[2]Researchers who focus on the contributions of technology to health expenditure growth also acknowledge the key role of insurance in isolating consumers from the costs of this new technology (Weisbrod 1991, Chandra and Skinner 2011).

resource reallocations can shape the system's long run evolution through their effects on incentives for entry and innovation.

Our results also contribute to a longstanding question in the public and academic debates over hospital pricing. Both researchers and policymakers have considered the extent to which health care providers engage in *cost-shifting* (Dranove 1988).[3] The cost-shifting hypothesis suggests a negative relationship between public and private prices for health care services, especially in the face of government payment cuts.[4] This theory draws heavily on the idea of revenue or income targets, which may be necessary to cover large fixed costs. It is frequently invoked to mitigate concerns about public sector payment cuts.[5] We show that prices for physician services exhibit the opposite relationship, which we term *cost-following*. Our evidence thus advises against evaluating physician payment policies under the presumption that private markets will offset changes in public insurers' payments. Instead, the private sector amplifies these changes.

Our empirical analysis draws on details of how Medicare sets physicians' fees. In the absence of exogenous shocks to Medicare's payments, we would be concerned that private and Medicare prices would naturally covary because of changes in productivity or demand. We use two overhauls of Medicare's administrative payment mechanisms to overcome this concern. First, we exploit a sharp reduction to Medicare's payments for surgical procedures relative to other medical services that occurred in 1998. Because the price change varied substantially in dollar terms across services, our identification comes from payment changes

[3]For overviews of the extensive cost-shifting literature (including Cutler 1998, Kessler 2007, Wu 2010, Robinson 2011), see Frakt (2011, 2013). Foster (1985) and Dranove (1988) highlight that cost-shifting behavior will tend to be inconsistent with profit maximization, making it more plausible in the hospital context than among the physician groups we study. Recent work in the hospital setting finds evidence against cost shifting from price shocks (White 2013, White and Wu 2013).

[4]Note that the empirical work in some papers, including Dranove, Garthwaite and Ody (2013), interprets the concept of cost-shifting differently than we do here. Dranove et al. (2013) find private price reductions in response to a wealth shock, rather than a public sector pricing shock.

[5]Frakt (2011) and Dranove et al. (2013) present numerous examples of this phenomenon, ranging from ProPAC (1992) through the recent Supreme Court decision in *National Federation of Independent Business v. Sebelius* (2012).

both between medical and surgical services and within each of those two categories.

To study these changes, we construct a novel link between databases of Medicare and private sector claims. We use these claims to construct a rich panel of public and private prices that vary across years, states, and 2,194 individual medical services.[6] Using these data, we estimate that a \$1 decrease in Medicare's payment for a surgical service led on average to a \$1.16 decline in private payments for that service. Private prices moved in sync with Medicare's relative payments; the response emerged in full during the year following the implementation of Medicare's administrative change. The private prices show no pre-trends prior to the Medicare payment changes, supporting the view that these shocks were exogenous for our purposes.

We next study a set of across-the-board payment changes that varied across geographic areas. We again find a positive relationship between Medicare's reimbursements and private prices. Over the medium to long run, we estimate that a \$1 decrease in Medicare's fees led to a \$1 decrease in private payments. Private sector responses to these broad-based rate changes appear to unfold over several years.

We present a stylized model of physician price-setting to elucidate these results. In our model, the doctor can choose at the margin between treating Medicare patients at a fixed, exogenous reimbursement rate, or treating those with private insurance, who exhibit downward-sloping demand. The physician's optimal pricing is a markup over her outside option, which is determined by Medicare's payment rate. The markup depends in turn on how elastic demand is for that doctor. The model thus predicts that variation in the demand elasticity, such as that associated with provider or insurer competition, should influence how Medicare's rates affect private prices.

Motivated by this model, we next analyze heterogeneity in Medicare's effect on private

[6]The services we consider are defined quite precisely. A 20-minute office visit is distinct from a 30-minute office visit, and coronary artery bypass grafts (CABG) are counted differently depending on the number of grafts and whether arterial grafts are used in addition to venous grafts.

prices across geographic and specialty markets. Medicare's influence is particularly strong in areas with relatively concentrated insurance markets and with relatively competitive provider groups, both of which are likely to imply more elastic demand for the physician. These relationships must be interpreted with the usual caution afforded to analyses of treatment effect heterogeneity, as our variation in market characteristics is not quasi-experimental. Nonetheless, these relationships are robust to flexible controls for the relationship between the Medicare price change and various regional economic and demographic characteristics. The heterogeneity in private responses to Medicare is consistent with a variable mark-ups version of our physician pricing model.

To better understand the magnitude of Medicare's effects on private prices, and to explore our findings at a more practical level, we turn to the institutional details of physician-insurer contracting. Industry participants report that insurers make simple, take-it-or-leave-it offers to small provider groups.[7] These offers involve fixed fee schedules based in large part on Medicare's payment menu, perhaps with a constant markup. Linking physician payments to Medicare's relative values involves a tradeoff. On the one hand, such contracts incorporate any inefficiencies embedded within Medicare's menu. On the other hand, capitalizing on Medicare's familiar pricing schedule reduces the complexity of contract negotiation and bill processing, which are substantial in this setting (Cutler and Ly 2011). When insurers interact with small physician groups, the latter consideration is relatively likely to outweigh the former. Empirically, we find that high prevalence of small physician groups predicts relatively strong transmission of Medicare's relative prices.

Our results show that Medicare strongly influences both aggregate health expenditures and the relative valuations placed on physicians' services. Effects on average prices imply

[7]In Appendix A we present industry participants' own descriptions of these negotiations. Anecdotally, Medicare's relative values are prevalent in private insurers' default physician payment menus (see Appendix A). This is not limited to traditional insurance arrangements and includes some health maintenance organizations (see, *e.g.*, Blue Cross and Blue Shield of Texas 2010)

that Medicare has substantial effects on health sector inflation and thus, to a non-trivial degree, overall price inflation (Clemens, Gottlieb and Shapiro 2014). Furthermore, the payment spillovers we estimate have additional implications for the health system's long-run performance and the political economy of payment reform.

The connection between private payments and Medicare's relative valuations is particularly relevant to health system performance. A hefty literature has demonstrated how medical care prices influence future entry (Dezee et al. 2011), investment decisions (Acemoglu and Finkelstein 2008, Clemens and Gottlieb 2014), and innovation.[8] It is beyond this paper's scope to determine optimal absolute or relative price levels. Instead, we analyze how private reimbursements are set in practice.

When insurers rely directly on Medicare's pricing menu, they amplify both the costs of Medicare's inefficiencies and the potential gains from payment reform. Newhouse (2002), Cutler (2011) and many others note that Medicare's reimbursements are unlikely to be efficient.[9] For example, absent an implausibly large and sudden shift in patient needs or medical technology, Medicare's relative payments for surgical and non-surgical services cannot have been optimal both before and after the change we analyze. More generally, prices purposefully set on an average cost basis can only be optimal by chance.[10] Despite these inefficiencies, our evidence shows that Medicare's relative values influence a wide swath of private sector payment contracts. We provide documentary evidence that this includes payments within

[8]Acemoglu and Linn (2004), Finkelstein (2004), Blume-Kohout and Sood (2013), Budish, Roin and Williams (2013), and Clemens (2013) show that innovation responds to potential market sizes, which affect the return to practice more generally.

Beyond entry, investment and innovation, other aspects of health care supply—such as drug choice (Jacobson, Earle, Price and Newhouse 2010, Alpert, Duggan and Hellerstein 2013), imaging and relatively elective procedures (Clemens and Gottlieb 2014)—also respond to payment rules. Glied and Graff Zivin (2002) show that changes in physician's financial incentives when treating one set of patients affect the doctors' behavior throughout their practice.

[9]We think of optimal pricing as setting the reimbursement rate equal to the health benefit of a marginal treatment, as valued by society.

[10]The appropriate average cost is hard even to define conceptually in the presence of fixed costs, since it depends on the quantity of care, which itself responds to payment rates.

health maintenance organizations (HMOs) and other network-based plans.

The co-movement of public and private prices unambiguously worsens the returns to specialties experiencing Medicare payment cuts. Between its direct effects ($2.6 billion) and private sector spillovers ($7.6 billion), we estimate that the surgical-medical pricing change reallocated $10.2 billion annually across providers and types of services. The specific context of surgical versus nonsurgical payments remains relevant today. Similar debates over the value of preventive care relative to curative therapy, and of interventional relative to non-intensive treatments, are lively in both academic and policy spheres.[11]

Our analysis of changes to Medicare's geographic adjustments also has substantial implications for the incomes of urban and rural practices. Absent private sector spillovers, the geographic payment overhaul would have redistributed $282 million annually from urban to rural physicians. Accounting for spillovers brings the total to $1 billion per year. This is more than three times as large as the formal subsidies spent under federal legislation to support rural hospitals through the Critical Access Hospital program.

The paper proceeds as follows. In section 1, we present institutional background on Medicare payments and describe our first empirical strategy. We estimate the effects of this change in section 2. We then examine our second set of Medicare payment shocks, which affected Medicare's payments across the board to certain geographic areas, in section 3. This section also illustrates our results' implications for the extent to which Medicare shapes sector-wide resource allocation. Section 4 presents a stylized model of physician payments in order to understand the underlying economics, and section 5 shows that private sector pricing is broadly consistent with the model. Motivated by the model, section 6 demonstrates heterogeneity in Medicare's effects across markets depending on the degrees of competition across physician groups and insurers. We consider the institutional details

[11] On prevention, see for example Cohen, Neumann and Weinstein (2008), Cohen and Neumann (2009) and Baicker, Cutler and Song (2010). On intensity of care, see for example Fuchs (2004), Skinner, Staiger and Fisher (2006), Chandra and Staiger (2007), and Doyle (2011).

underlying physician-insurer contracting and explore their implications in section 7. Section 8 concludes.

1 Estimating the Effects of Changes in Medicare's Reimbursement Rates

To estimate Medicare's influence on private sector pricing, we exploit two large, administrative changes in its reimbursement rates. While these payments are set according to administrative rules, these rules can be changed by Congress or Medicare's administrators in Washington (the Centers for Medicare and Medicaid Services, or CMS). Such changes deliver variation in payment rates that may be independent of patient demand and technological change or other supply-side market pressures. Our identification strategies rely on two changes that significantly altered Medicare's payments. In this section we describe our first source of identification, namely a change in Medicare's payments for surgical services relative to non-surgical services. In section 3 we explain our second source of identification, which is an overhaul of Medicare's system of geographic adjustments. Before discussing the details of these changes, we briefly describe the institutions that determine prices in public and private markets for health care services.

1.1 How Are Private Medical Payments Set?

Public and private payments for health care services are set through very different mechanisms. In the physician setting we study, public rates are set through an administrative apparatus mandated to set payments according to the resource costs of providing care. In the world of private health insurance, payment rates are set on markets with varying degrees of competition (Dafny et al. 2012).

U.S. private sector health care prices are largely unregulated.[12] Rather than being set

[12]Some exceptions apply to this statement. For instance, all hospital payment rates in Maryland are set

according to measured resource utilization, as in Medicare, they are agreed upon through negotiations between insurance carriers and the providers with whom they contract.[13] Negotiated prices are often unknown to final consumers and can vary substantially, for ostensibly similar services, across both providers and insurers (Dunn and Shapiro 2012).

Existing research sheds light on some of health care prices' determinants. Cutler, McClellan and Newhouse (2000) find significant differences between the prices negotiated by HMOs and traditional health insurance plans. Price variation also stems from producer heterogeneity, with more attractive hospitals commanding higher prices (Ho 2009, Moriya, Vogt and Gaynor 2010, Gowrisankaran, Nevo and Town 2013, Lewis and Pflum forthcoming). Robust insurance-market competition increases payments to physicians and hospitals (Town and Vistnes 2001, Dafny 2005, Dafny et al. 2012), while competition among provider networks reduces them (Dunn and Shapiro 2012). Most closely linked to our setting, Showalter (1997) finds a positive cross-sectional relationship between state Medicaid fees and private insurers' physician payments.

1.2 A Large Shock to the Relative Prices of Outpatient Services

Since 1992, Medicare has paid physicians and other outpatient providers through a system of centrally administered prices, based on a national fee schedule. This fee schedule, known as the Resource-Based Relative Value Scale (RBRVS), assigns relative values to more than 10,000 distinct billing codes according to the resources CMS believes the services to require. Medicare scales these relative valuations by multipliers called "Conversion Factors" (CF). Our first natural experiment involves a large, administrative change in the CFs associated with surgical and non-surgical services. Because input costs vary across areas, the fee

by a state government board.

[13]When serving self-pay patients (generally meaning the uninsured), prices are simply set by the provider as in traditional markets for goods and services, and consumers can choose which firm receives their business. In these transaction, however, the threat of personal bankruptcy filings leads to substantial price renegotiations after treatment has taken place (Mahoney 2012).

schedule further adjusts payments to partially offset such differences. Our second natural experiment exploits an administrative change in this system of geographic adjustments.

For service j, supplied by a provider in payment area i, the provider's fee is approximately:

$$\text{Reimbursement}_{i,j,t} = \text{Conversion Factor (CF)}_{t,c(j)} \times \text{Relative Value Units (RVU)}_{j,t} \times \text{Geographic Adjustment Factor (GAF)}_{i,t}. \quad (1)$$

The Conversion Factor is a national adjustment factor, updated annually and generally identical across broad categories of services, $c(j)$. In the early 1990s, wrangling over payments across specialties led to the institution of separate CFs for surgical procedures and other services. Surgeons argued that slower growth in the use of procedures relative to other medical services should be rewarded. Congress implemented this plan, and CMS first distinguished between the CFs for surgery, primary care, and other services in 1993.[14]

From 1993 to 1995, payments for surgical procedures grew relative to payments for other services. 1995 to 1997 marked a period of relative stability, with an average bonus of 15.5 percent for surgical RVUs relative to primary care and other non-surgical RVUs. Because the Conversion Factors are set nationally, the surgeons' relative price bonus was constant across geographic regions. These unequal payments for equal RVUs spawned political discontent among non-surgeons. In 1998, this 15.5 percent bonus was eliminated through a budgetarily neutral merger of the CFs.[15] This administrative change generated large and sudden changes to Medicare payments. We take advantage of these changes to identify spillovers from Medicare reimbursement rates to private insurance payments.

Figure 1 shows how the surgical and non-surgical Conversion Factors evolved over this time period. As the graph shows, the two CFs evolved in parallel from 1995 through 1997. In 1998, they were merged together. The CF merger thus resulted in large changes to

[14]We owe our knowledge of this political history to Newhouse (2002).

[15]62 *Federal Register* 59048, 59102 (1997).

relative payments for broad categories of services that were not associated with changes in beneficiary demographics or other determinants of demand. Within these broad categories, it also generated variation in the dollar value of payment shocks across services.

To take two illustrative examples, consider Medicare's payments for a coronary artery bypass grafts (CABG) and a cardiac stress tests with nuclear imaging (SPECT). In 1997, Medicare's fee for CABG averaged $1,428.[16] In 1998, the average fee fell to $1,283, or by just over 10 percent. In 1997, SPECT generated an average Medicare fee of $475. Because SPECT is a test rather than a surgical procedure, its fee rose to $513 (an increase of 8 percent).[17]

1.3 Estimation Strategy

This section summarizes our strategy for estimating the effects of Medicare's administrative price changes on private sector reimbursement rates. We harness a rich linkage between databases of Medicare and private sector claims, which allows us to construct a novel panel of public and private price data that vary across years, areas, and individual service codes. The data set is uniquely suited for assessing how Medicare's menu influences average market-by-service payments from private insurers.[18]

We use the aforementioned CF shock to construct an instrument for Medicare's payments. As detailed below, our baseline approach expresses these shocks in level terms, giving our estimates an interpretation similar to a pass-through coefficient. The levels specification is suggested by practitioner characterizations of physicians' contracts (e.g. Gesme and Wiseman 2010, Mertz 2004), which regularly emphasize a linear benchmarking relative to Medicare's rates, for instance according to

[16]This example refers specifically to CPT code 33533.

[17]Specifically, CPT code 78465.

[18]The private sector claims do not longitudinally follow the payments associated with identifiable physician-insurer pairs. Our analysis harnesses some of the least aggregated moments of these data that can be assembled into a panel.

$$P^{\text{Private}} = b \cdot P^{\text{Medicare}} + \text{other factors}.$$

If this characterization of physician contracts is correct, the parameter of interest, b, must be estimated using the levels of Medicare payments as opposed to logs.[19] Nevertheless we also present a log-log analogue of the framework outlined below. Results from the level and log specifications are very similar.

Stage 0: Compute the Instrument: Predicted Price Change

Using Medicare payment data from 1995 to 1997, we compute each service's average price, $\overline{P^{\text{Medicare}}}_{j,\text{pre}}$, prior to the policy change. We then construct a variable that captures the price change implied by the 1998 CF merger.

$$\text{PredChg}_j^{\text{Medicare}} = \overline{P^{\text{Medicare}}}_{j,\text{pre}} \cdot \begin{pmatrix} -0.104 \cdot \text{Surgical}_j \\ +0.05 \cdot \text{Non-Surgical}_j \end{pmatrix} \tag{2}$$

where the factors -0.104 and 0.05 are the average changes in the nominal Conversion Factors for surgical and non-surgical services, respectively.

Stage 1: First Stage

We then use this predicted price change, $\text{PredChg}_j^{\text{Medicare}}$, as an instrument for the actual Medicare reimbursement rate. Specifically, we run the following first stage regression:

$$\begin{aligned} P_{j,s,t}^{\text{Medicare}} = \pi \cdot \text{PredChg}_j^{\text{Medicare}} \cdot \text{Post1998}_t + X_{j,s,t}\psi + \mu_j \mathbb{1}_j + \mu_s \mathbb{1}_s + \mu_t \mathbb{1}_t \\ + \mu_{j,s} \mathbb{1}_j \cdot \mathbb{1}_s + \mu_{t,s} \mathbb{1}_t \cdot \mathbb{1}_s + e_{j,s,t} \end{aligned} \tag{3}$$

[19] We are grateful to Michael Dickstein and Neale Mahoney for making this point.

at the service (j), by state (s), by year (t) level.[20] We weight observations by the number of times the service was performed in 1997.[21] Because there could be persistent idiosyncratic changes in the market for a service, and the payment changes vary at that same level, we cluster standard errors by service codes.[22]

Equation (3) represents a linear formulation of Medicare prices with respect to the predicted policy-driven shock. We expect to estimate a coefficient of $\hat{\pi} = 1$ in the absence of measurement error and correlated reimbursement changes. We control for service, state, and year fixed effects as well as full sets of service-by-state ($\mathbb{1}_j \cdot \mathbb{1}_s$) and state-by-year ($\mathbb{1}_s \cdot \mathbb{1}_t$) effects.

The most important elements of the vector of additional controls ($X_{j,s,t}$) are indicators that capture major payment changes for relevant services. Specifically, our first stage most cleanly tracks the policy change of interest when we control separately for major mid-1990s payment changes associated with cataract surgery.[23] We further include controls, defined in section 1.4, for the types of insurance plans associated with the services in our data.

Stage 2: Second Stage

The Medicare price predicted in equation (3), $\widehat{P_{j,s,t}^{\text{Medicare}}}$, then serves as an instrument for actual Medicare prices in the following second stage equation:

[20]While we construct $P_{j,s,t}^{\text{Medicare}}$ at the service-by-year level, we use service-by-state-by-year observations to maintain consistency through subsequent analysis of heterogeneity in Medicare's effects across services and states. Appendix D.1 shows that our results remain similar when using national level observations.

[21]With the regression estimated in levels, weighting by the 1997 service count accounts appropriately for the surgical payment shock's budgetary neutrality. When we run specifications on log prices, we weight each service by its service count times its assigned number of relative value units.

[22]Our results are robust to using larger clusters, such as 1-, 2-, or 3-digit Betos codes, and to clustering at the state level. See footnote 54 below.

[23]Cataract surgery has long been viewed as a procedure provided in excess and, in an effort to reduce its usage, was subjected to significant payment reductions in the years leading up to the 1998 price shock on which we focus. With cataract surgery accounting for a non-trivial fraction of Medicare's payments for surgical services, we find that "dummying out" these earlier payment reductions allows us to cleanly track the natural experiment of interest. Alternative specifications, including those that either do nothing to account for the cataract-surgery reductions or that drop cataract surgery from the sample, generate similar estimates of the effect of Medicare payment changes on private sector prices. In the first stage, however, these alternative specifications have inferior ability to track the 15 percent reduction in the surgical CF.

$$P^{\text{Private}}_{j,s,t} = \beta \cdot \widehat{P^{\text{Medicare}}_{j,s,t}} + X_{j,s,t}\phi + \nu_j \mathbb{1}_j + \nu_s \mathbb{1}_s + \nu_t \mathbb{1}_t$$
$$+ \nu_{j,s}\mathbb{1}_j \cdot \mathbb{1}_s + \nu_{t,s}\mathbb{1}_t \cdot \mathbb{1}_s + \varepsilon_{j,s,t} \tag{4}$$

Our use of the predicted Medicare prices as an instrument is valid under the following assumptions. First, the predicted change $\text{PredChg}_{j,s,t}$ must be reflected in the actual Medicare prices in the first stage equation (3). Second, the shock used to generate these predicted prices must be conditionally independent of other sources of change in private sector payment rates, as captured by the error term $\varepsilon_{j,s,t}$. These include technology shocks, demand shocks, and other changes in market conditions. We use the large, one-time nature of the payment shocks to investigate the potential relevance of threats to identification as carefully as possible. Most importantly, we check for the presence of pre-existing trends in both Medicare and private payments by graphically presenting parametric event study estimates from the following two equations:

$$P^{\text{Medicare}}_{j,t} = \sum_{t \neq 1997} \gamma_t \cdot \mathbb{1}_t \cdot \text{PredChg}^{\text{Medicare}}_j + X_{j,s,t}\psi + \mu_j \mathbb{1}_j + \mu_s \mathbb{1}_s + \mu_t \mathbb{1}_t$$
$$+ \mu_{j,s}\mathbb{1}_j \cdot \mathbb{1}_s + \mu_{t,s}\mathbb{1}_t \cdot \mathbb{1}_s + u_{j,s,t} \tag{5}$$
$$P^{\text{Private}}_{j,s,t} = \sum_{t \neq 1997} \delta_t \cdot \mathbb{1}_t \cdot \text{PredChg}^{\text{Medicare}}_j + X_{j,s,t}\alpha + \upsilon_j \mathbb{1}_j + \upsilon_s \mathbb{1}_s + \upsilon_t \mathbb{1}_t$$
$$+ \upsilon_{j,s}\mathbb{1}_j \cdot \mathbb{1}_s + \upsilon_{t,s}\mathbb{1}_t \cdot \mathbb{1}_s + v_{j,s,t}. \tag{6}$$

If pre-existing trends in either public or private payments are correlated with $\text{PredChg}^{\text{Medicare}}_j$, they will be apparent in the estimates of δ_t and γ_t for years prior to 1997. Estimates for 1998 and beyond will trace out the dynamic effects of Medicare's payment shocks. For Medicare itself, the post-1997 coefficients in equation (5) should hew closely to 1.

1.4 Health Care Price Data

We study the public sector's influence on private sector health care prices by linking health insurance claims data across the two environments. In both settings, providers request reimbursement by submitting claims to the relevant third-party payer. For Medicare claims, we use a 5 percent random sample of the Medicare Part B beneficiary population for each year from 1995 through 2002. Part B, formally known as Supplementary Medical Insurance, is the part of Medicare that covers physician services and outpatient care. The data contain service-by-service reports of the relevant care purchased by Medicare for these beneficiaries. For pricing purposes, they include the Health Care Procedure Coding System (HCPCS) code for each service along with Medicare's payment (the "allowed charge"). We construct a measure of Medicare's payment rates by aggregating claims to compute the average allowed charge for a service at the code-by-year level.

We measure private sector prices similarly, using private insurance claims data from the ThompsonReuters MarketScan database (also known as "MedStat"). Private insurers use procedure codes that overlap substantially with the HCPCS system. MarketScan obtains these codes, along with service-level payment rates and additional information, from large self-insured employers' claims records. The data are thus sufficient to allow us to estimate how the service-specific payments negotiated between insurers and providers vary across space and over time. We aggregate these claims and compute the average allowed charge at the code-by-state-by-year level.

Our baseline estimation sample includes 2,194 individual HCPCS codes that satisfy two criteria. First, they must be linked across the Medicare and MarketScan databases. Second, we require that our panel be balanced in the following sense: a state-by-service pair is only included in the sample if it appears in each year from 1995 through 2002. Appendix C provides further detail on these primary data sources and the comprehensiveness of our merge procedure. Summary statistics describing Medicare and private sector prices across services and states are shown in Table 1, separately for surgical and non-surgical services. In this sample, the average surgery price is $239 in Medicare and $374, or nearly 60 percent higher, in the private market. The average non-surgical service is reimbursed $114 in Medicare and $125 in the private sector. Both the public and private price data represent in excess of 100 million underlying services.

We observe private sector prices from a range of insurance plan types. In 1996, 38 percent of Medstat service claims came from Major Medical or Comprehensive Insurance (CI) plans, 52 percent from less generous Preferred Provider Organization (PPO) plans, and 10 percent from even more restrictive Point of Service (POS) plans. By 2006, 8 percent of Medstat service claims came from CI plans, 59 percent from PPO plans, 12 percent from POS plans and roughly 27 percent from other less generous plans including Health Maintenance Organizations (HMO) and Consumer-Driven Health Plans (CDHP). The data thus reflect a national trend away from comprehensive coverage towards forms of coverage designed to control costs.

To ensure that our results are not affected by changes in the composition of plans included in our data, we construct a variable to measure the types of plans covered. While changes in this composition are largely captured by the state-by-year fixed effects that we include throughout the analysis, one might still be concerned that different types of plans tend to procure different services for their beneficiaries. To control for this, we construct a variable called "Plan Type Payment Generosity" by first regressing payments on plan type indicators. Using the resulting coefficients and changes in the plan type composition, we generate predicted payments that we aggregate to the state-by-year-by-service level. We also construct a control for plan generosity based on patient cost sharing. This variable, "Service Specific Cost Sharing," is constructed at the state-by-year-by-service level by dividing out-of-pocket payments by the total payments made to providers for the service. Because changes in plan types unfolded relatively smoothly over time, we would expect any associated concerns to reveal themselves in the event study estimates of equation (6).

Despite the changes in coverage types that we observe, data from the Community Tracking Study reveal that these changes in plan design had little impact on insurers' methods for paying physician groups. Between 1995 and 2004, the fraction of physicians' revenues associated with capitated, as opposed to fee-for-service, payments *declined* from 16 percent to 13 percent (CSHSC 1999, 56; 2006, 4-29). This reflects the fact that many managed care arrangements ultimately pay physicians through relatively traditional fee for service arrangements. While our estimates will be internally valid to plans for which payments are uniformly fee-for-service, Medicare's influence extends to alternative arrangements. For example, many insurers' provider newsletters make explicit their use of Medicare's relative values within HMO plans (*e.g.* Blue Cross and Blue Shield of Texas 2010, Anthem Blue Cross and Blue Shield 2012a). These are non-trivial applications, as the HMO Blue Texas plan advertised having 38,000 physicians in its provider network as of 2009 (Blue Cross and Blue Shield of Texas 2014).[24]

2 Empirical Effect of Medicare Prices on Private Prices

Panel A of Figure 2 illustrates the raw correlation between public and private prices. Public and private payments are tightly related in the cross-section, with Medicare paying roughly 40 percent less than private insurers for identical services. Despite substantial variation in private payments both across and within geographic markets, average Medicare payments predict 89 percent of the variation across services in the average private payment. Changes in public and private prices over time are also tightly related, as illustrated in Panel

[24]While the details of physician payment within HMOs have not been systematically studied, additional sources confirm that their payments regularly take on fee for service structures (Marton, Yelowitz and Talbert 2014). Because the Medstat data do not report reliable per-service payments for HMO-style plans, however, we are unable to incorporate such plans into our primary analysis. Notably, however, our analysis does incorporate alternative "managed care" arrangements including preferred provider organizations (PPOs) and point-of-service (POS) plans when payments are made and recorded on a fee-for-service basis.

B. Appendix Figure D.1 shows analogous graphs using cross-state variation.

Figure 3 plots event study estimates of the effect of Medicare payment changes using equations (5) and (6). First stage results, marked on the graph with "×" symbols, show that the predicted service-level price changes translate almost one-for-one into observed Medicare payment rates. There also appears to be a slight upward drift associated with gradual increases in Medicare's payments for primary care relative to other services. While this drift makes it important to look closely to the dynamics of private responses, and to check robustness after controlling for a surgery-specific trend, these results give us confidence in our specification of the shock.

The figure also plots reduced form estimates of the shocks' impact on private sector prices. Changes in private prices were uncorrelated with the payment shocks during the years preceding the shock, providing evidence against potentially confounding pre-existing trends driven by changes in technology, demand, or other market conditions. From this point forward, a $1.00 increase in Medicare's predicted payment led, on average, to a $1.30 increase in private payments. The private sector response emerges in full during the year of Medicare's payment change. Changes in technology, demand, and other market conditions, which tend to evolve more gradually, appear quite unlikely to explain the observed change in private prices.

In Table 2, we summarize these results in single coefficients using the framework described by equations (2) through (4). Column 1 reports the first-stage estimates of equation (3). We find $\hat{\pi}$ to be 1.1, which is quite close to 1. The cluster-robust F statistic for testing the null hypothesis that our instrument is weak is 288, which easily satisfies the robust weak instruments pre-test threshold of Olea and Pflueger (2013).[25]

Column 2 shows the reduced form results we obtain when we replace $P_{j,s,t}^{\text{Medicare}}$ with $P_{j,s,t}^{\text{Private}}$

[25]Their Table 1 reports a critical value of 23.11 for the effective F statistic (which, with one instrument, is equal to the cluster-robust F statistic) to reject the null hypothesis of a two-stage least squares bias above 10% of the OLS bias with one instrument in the presence of heteroskedasticity.

as the outcome variable in (3). The coefficient of 1.29 suggests that a one dollar predicted change in Medicare prices translates into a \$1.29 change in private sector prices. Column 3 reports the IV estimate of equation (4), which rescales the private sector change by the actual Medicare response (from column 1). The result is our baseline estimate, which implies that a one dollar change in Medicare payments led to a \$1.16 change in private payments.

Columns 4 through 6 run comparable specifications in which public and private prices are expressed in logs.[26] The instrument in these specifications is a binary indicator for surgical services performed during or after 1998. Column 4 shows that Medicare's elimination of the surgery-specific conversion factor is associated with a decline of 0.22 in the log relative payments for surgeries, which is moderately larger than that called for by the payment reform. The reduced form estimate of the policy change's impact on private prices is -0.11. Column 6 reports the IV estimate, showing that, on average, a 10 percent change in a Medicare payment resulted in a 4.8 percent change in the relevant private payment. This is reconciled with the \$1.16 from column 3 by the fact that average private payments, and in particular payments for surgical services, are higher than the average Medicare payment.[27]

Appendix Table D.1 provides evidence of the robustness of our finding that Medicare prices pass through into the private sector. Column 1 repeats the baseline IV estimate from column 3 of Table 2. Columns 2 and 3 show that the results are not sensitive to dropping our controls for the generosity of the insurance plans represented in the sample.[28] Column

[26]Appendix D.3 examines the effect of Medicare price changes on private sector price dispersion.

[27]Table 1 shows that Medicare pays \$239 on average for surgical services and \$114 for non-surgical. Its surgical payments fell by 10 percent, or \$24, while medical reimbursement rates increased by 5%, or \$6. So the difference fell by $\$24 - \$6 = \$18$. As private non-surgical reimbursements average \$125, and surgical fees average \$374, identical percentage changes to the private sector would have required a \$41 decline in surgical fees and a \$9 increase in medical payments, or a \$30 ($= \$39 - \$9$) relative change. But the private sector cost-following coefficient of 1.16 that we have estimated means that Medicare's \$18 relative change only led to a \$21 ($= 1.16 \times \$18$) relative change in the private sector. Since $\ln\left(1 + \frac{\$21}{\$30}\right) \approx 0.5$, this is the coefficient we estimate in logs in column 6.

[28]These controls are more strongly predictive of private payments in specifications that do not include full sets of state-by-year effects, but even then have little impact on our baseline estimate. State-by-year effects account for most of the variation in plan design contained in the MedStat data.

4 shows that our baseline estimate is modestly sensitive to the inclusion of controls for mid-1990s payment changes targeted at cataract surgery; omitting these controls reduces the magnitude of the cost-following coefficient from 1.2 to 1.0. Column 5 removes the service weights, which reduces the estimate to around 0.7.[29] Column 6 includes a control for the number of Relative Value Units (the quantity metric that appears in equation [1], Medicare's payment formula) assigned to each service. Minor updates to RVU assignments strongly predict Medicare's allowable charges, which they impact formulaically (coefficient not shown). Controlling for these updates has little impact on our baseline result. Finally, column 7 shows that the baseline is robust to controlling directly for a linear trend in private payments for surgical procedures relative to other services. As shown in Figure 3, there is no such trend in private payments.[30]

3 Across-The-Board Payment Shocks

In this section, we analyze payment changes that altered reimbursements across the board in an area. The last term of equation (1), which presents Medicare's payment formula, describes adjustments in payments across areas. While the Conversion Factor is set nationally, the Geographic Adjustment Factor (GAF) varies across payment regions. It is intended to capture differences in input costs, which are estimated using Census and other data on area-level rents, wages, and malpractice insurance premiums.

We analyze the effects of GAF changes driven by an administrative re-shuffling of the areas across which these adjustment are made. Through 1996, payments were differentiated

[29] Accounting for the reductions to payments for cataract surgery improves our ability to correctly track the reduction in payments for surgical procedures relative to other services. Cataract surgery exerts a significant impact on our regressions because it is a very high volume service. Changes in service-specific Part B payments are, in general, implemented in a budgetarily neutral fashion. Appropriately estimating the first stage thus requires weighting each service by its baseline frequency. The unweighted first stage underlying the specification reported in column 5 does a poor job of tracking the Medicare payment change.

[30] When we run comparable analyses on national-level data, the results are very similar. Service-by-year analogues of Tables 2 and D.1 are shown in Appendix Tables D.2 and D.3.

across 210 payment areas, as shown in the top panel of Appendix Figure D.2. In 1997, the federal government consolidated these 210 payment areas into the 89 regions shown in the figure's middle panel. In states where consolidations occurred, the merger of urban and rural areas resulted in budgetarily neutral declines in urban payments and increases in rural payments. We analyzed these payment shocks previously in Clemens and Gottlieb (2014), which provides additional institutional background.

These geography-based changes differ somewhat from the relative price changes we examined in the previous section. If physician specialties provided exclusively either surgical or non-surgical services, these payment changes would be nearly equivalent conceptually. The relative price changes would have implied across-the-board declines in payments for surgical specialties and increases in payments for non-surgical specialties. But in practice, most specialties provide significant quantities of both surgical and non-surgical services. General practitioners, for example, experienced an average payment increase of 2.4 percent. If they provided no surgical services, the increase would have been 5 percent.[31] General surgeons saw their average payment decline by 3.9 percent, while cardiac surgeons experienced a decline of 7.6 percent. An exclusive provider of surgical services would have experienced a 10.4 percent decline. The price changes analyzed in the previous section thus confronted physicians with changes in both their average and relative payments. In contrast, the geography-based shocks that we now analyze affect the average payment level while leaving relative prices for different services nearly unchanged.[32]

As in the previous section, our parameter of interest is a scalar mark-up of private relative to Medicare payments. We express the Medicare payment changes in dollar terms by

[31]Within the Medicare data, roughly 16 percent of the services provided by general practitioners were surgical procedures, which experienced an average payment decrease of 10.4 percent, while roughly 84 percent were non-surgical services, which received an average payment increase of 5 percent.

[32]Equation (1) is a slight simplification of Medicare's full payment formula, and the full version shows that the administrative change would induce slightly different proportional price changes for different services in the same area. Clemens and Gottlieb (2014) provide further details. But these differences across services are small relative to the differences these payment shocks generated across geographic areas.

multiplying the changes in the geographic indices by the average pre-consolidation payment associated with each service in the sample. We denote the resulting area-specific shocks by Payment $\text{Shock}_{a,j}$ and estimate the following equations:

$$\begin{aligned} P^{\text{Medicare}}_{j,a,t} &= \sum_{t \neq 1996} \beta_t \cdot \text{Payment Shock}_{a,j} \cdot \mathbb{1}_t + \gamma_j \mathbb{1}_j + \gamma_a \mathbb{1}_a + \gamma_t \mathbb{1}_t \\ &\quad + \gamma_{j,a} \mathbb{1}_j \cdot \mathbb{1}_a + \gamma_{s(a),t} \mathbb{1}_{s(a)} \cdot \mathbb{1}_t + \zeta' X_{a,s(a),t} + \varepsilon_{j,a,t} \qquad (7) \\ P^{\text{Private}}_{j,a,t} &= \sum_{t \neq 1996} \beta_t \cdot \text{Payment Shock}_{a,j} \cdot \mathbb{1}_t + \delta_j \mathbb{1}_j + \delta_a \mathbb{1}_a + \delta_t \mathbb{1}_t \\ &\quad + \delta_{j,a} \mathbb{1}_j \cdot \mathbb{1}_a + \delta_{s(a),t} \mathbb{1}_{s(a)} \cdot \mathbb{1}_t + \zeta' X_{a,s(a),t} + \varepsilon_{j,a,t}. \qquad (8) \end{aligned}$$

For this analysis, we construct the data set at the service-by-year-by-payment locality level, using the 210 payment localities that existed prior to their consolidation. The specifications include full sets of service, area, and year fixed effects as well as full sets of service-by-area ($\mathbb{1}_j \cdot \mathbb{1}_a$) and state-by-year ($\mathbb{1}_{s(a)} \cdot \mathbb{1}_t$) effects.[33] As all payment area consolidations took place within a state, states are the lowest level of geography at which we can flexibly control for variation over time. The state-by-year effects capture the effects of any potentially relevant changes in state policy, for example in Medicaid reimbursements or eligibility rules. Our robustness analysis explores the relevance of trends correlated with payment area characteristics, such as the extent to which they are rural or urban.

If pre-existing trends in public or private payments are correlated with $\text{PredChg}^{\text{Medicare}}_j$, they will be apparent in the estimates of β_t for years prior to 1997. Estimates for 1997 and beyond will trace out the dynamic effects of Medicare's payment shocks. For Medicare itself, the post-1997 coefficients in equation (5) should hew to 1.

[33] The analysis sample is balanced at the service type-by-payment area level, making the $\gamma_{j,a}$ a standard set of fixed effects at the level of the panel variable.

3.1 Effects of Across-the-Board Payment Changes

The results of estimating equations (7) and (8) appear in Figure 4 and Table 3. The figure shows both the first stage and reduced form estimates. The first stage estimates show that our coding of the payment shocks effectively tracks the policy change. A one unit increase in the payment shock is associated with a one dollar increase in Medicare's allowed charge.

The reduced form estimates plot the private sector response to these public payment changes. The effect of these across-the-board payment changes appears to unfold over several years. As with shocks to relative prices across services, an increase in public payments results in an increase in private payments. Averaging the point estimates across 1997 and subsequent years, a one dollar increase in public payments is associated with a one dollar increase in private payments.

Table 3 summarizes this result in a single coefficient by shifting from the event study framework to a parametric difference-in-differences estimator. The table shows that the baseline result is robust to several potentially relevant specification changes. These include replacing the full set of locality-by-service fixed effects with separate sets of service fixed effects and locality fixed effects, adding interactions between year indicators and proxies for the extent to which the localities are rural or urban, and replacing the full set of state-by-year effects with separate year and locality effects. While precision falls substantially in the last of these specifications, the results are similar throughout.

3.2 Implications for the Allocation of Health Expenditures and Long-Run Entry Incentives

The price changes analyzed in this and the preceding sections have histories linked to perennial payment policy debates. Vigorous, decades-long debates consider both the value of surgical versus medical treatment and the allocation of resources between urban and rural areas. In this section, we conduct an extrapolation exercise similar in spirit to the one that

Hurst, Keys, Seru and Vavra (2014) conduct to study the consequences of geographically heterogeneous mortgage subsidies. Our findings imply that Medicare's decisions on physician pricing are far more consequential than previously recognized.

Consider first the Congressional decision to merge the surgical and non-surgical conversion factors. This change affected the payments for $69.2 billion of medical care in 1998 alone (the first year after the change). By reducing surgical payments by an average of 10.4 percent and increasing medical payments by an average of 5 percent, Congress mechanically shifted $2.6 billion in payments from medical treatment to surgical treatment. Our estimates from section 2 show that the private sector responded by doing more of the same. Extrapolating our estimates to the relevant private physician payments yields an estimate that the private sector magnified this reallocation by a factor of 3.9. In total, we estimate that the merger of surgical and non-surgical Conversion Factors reallocated roughly $10 billion in annual spending from surgical to medical care. We detail this calculation and discuss caveats about external validity in Appendix E.

The consequences of Medicare's across-the-board, geographic payment changes are similarly magnified. The payment reductions experienced by adversely affected urban areas averaged 1.7 percent of their Medicare Part B reimbursements.[34] The change thus mechanically reallocated roughly $282 million between urban and rural areas in 1997 alone. In this context, our estimates (applying the coefficient of 1.0 from column 2 of Table 3) imply a private sector spillover 2.5 times Medicare's direct effect. In total, we estimate that the consolidation of payment localities reallocated $1 billion between urban and rural physicians per year.[35]

One way to benchmark the magnitudes of these reallocations is by comparison with

[34]This is the weighted average of the change in the Geographic Adjustment Factor (GAF) for these 333 counties, when weighted by spending in each county.

[35]As in the previous paragraph, all amounts are inflation-adjusted to 2013 dollars. Appendix E also discusses the caveats associated with this calculation.

similarly motivated but more directly financed programs. For example, the Critical Access Hospital (CAH) program is a prominent means through which Medicare subsidizes rural health care. By obtaining CAH designation, rural hospitals may claim higher payments than they would otherwise be entitled to receive. Unsurprisingly, rural hospitals lobby fiercely to maintain their access to these designations.[36] In 2010, CAH subsidies amounted to approximately $300 million (MedPAC 2012). Our findings imply that the CAH's magnitude is on par with the locality consolidation's direct effects and less than half the size of its private payment spillovers.[37]

Our findings thus illuminate a previously unexplored channel through which Medicare steers health-system resources. In the context of payments for surgical and non-surgical care, this has significant implications for Medicare's influence on long-run resource allocation. The co-movement of public and private prices means that Medicare's payment reforms significantly alter the returns to entering favored and unfavored specialties. A growing body of research finds that the investments of current practitioners (Acemoglu and Finkelstein 2008, Clemens and Gottlieb 2014), the entry decisions of medical residents (Dezee et al. 2011), and the development of new technologies respond significantly to such incentives.[38] The gains from reforming Medicare's payments to reward high value care may thus be immense.

[36]News (Gold 2011, McKee 2013), government reports (Levinson 2013), and advocacy organizations (American Hospital Association 2014a, 2014b, 2014c, 2014d) provide ample evidence of the CAH program's political importance.

[37]Due to the much faster growth of health care spending than overall inflation, the $300 million subsidy in 2010 probably overstates the contemporaneous value of the CAH program in 1997 that would be the appropriate comparison to the $1 billion spillover that we compute from that year.

[38]Acemoglu and Linn (2004), Finkelstein (2004), Blume-Kohout and Sood (2013), Budish et al. (2013), and Clemens (2013) show that innovation responds to potential market sizes.

4 A Model of Private Payments in Medicare's Shadow

We have seen that two, somewhat distinctive, sets of Medicare price changes resulted in significant, same-signed movements in private insurers' payments. To understand why this result could arise, we consider a model in which physicians treat both Medicare beneficiaries and the privately insured. In this framework, Medicare's reimbursements determine the physician's opportunity cost of treating private patients. We focus on how Medicare's reimbursements influence the prices physicians charge to private insurers.

The model abstracts from the rich institutional detail surrounding Medicare's fee schedule. Industry participants note that Medicare's prices serve as benchmarks in many private sector contracts. Under such contracts, changes in Medicare's payments could mechanically influence private prices. The model speaks most directly to actively negotiated prices, but may also help us to understand the rationale for benchmarking private prices to Medicare rates. We return to the institutional details of Medicare-linked contracts in section 7.

We consider a physician with a fixed aggregate treatment capacity K. This capacity constraint could arise if the physician incurs only fixed costs F for providing treatments up to point K, and extremely high variable costs beyond this point. The doctor must allocate this capacity between publicly and privately insured patients. The relative intensity of treating private patients is governed by a parameter α. With m and q denoting the quantities of treatment delivered to Medicare and private patients, respectively, the physician's constraint is $m + \alpha q \leq K$.[39]

Since we focus on health insurers' payments to physicians, we model demand for services as coming from these insurers. Their demand curve reflects patient demand plus any distortions that the insurance market induces. Let $D(p)$ denote demand as a function of the

[39]While a real-world physician likely has some capacity to expand overall treatment, this may be limited in the short run because of staffing constraints such as the physician's limited own time. Even when overall capacity is adjustable, this model may be a reasonable approximation as long as substituting between the two types of patients is easier than expanding overall capacity.

physician's price, p. The doctor takes Medicare's reimbursement rate r_M as given, and sets p to maximize profits.

The physician's general profit function is $\Pi(m, q) = r_M m + pq - F$ whenever $m + \alpha q \leq K$. Assuming she operates at capacity, we can express this as a function of the private price:

$$\pi(p) = pD(p) + r_M[K - \alpha D(p)] - F. \tag{9}$$

Medicare exogenously sets its reimbursement rate r_M, while the physician sets p accounting for private sector demand $q = D(p)$. Because the physician has incurred her fixed costs and operates at capacity, the potential revenue from Medicare patients acts as the opportunity cost of treating private patients. With a fixed Medicare reimbursement of r_M and intensity parameter α, the doctor faces a constant marginal opportunity cost of αr_M. Her pricing decision thus follows a familiar markup rule of $p^* = \frac{\epsilon^D(p^*)}{\epsilon^D(p^*) - 1}\alpha r_M$, where $\epsilon^D(p^*) = -D'(p^*)p^*/D(p^*)$ is the elasticity of demand and the physician will price such that $\epsilon^D(p^*) > 1$.

Our primary interest is in the relationship between physicians' private sector payments and Medicare's reimbursement rate, or $\frac{dp^*}{dr_M} = \frac{\epsilon^D(p^*)}{\epsilon^D(p^*) - 1}\alpha$. The sign of this relationship has been a matter of much debate in a literature on Medicare's influence on hospital prices. Standard considerations, such as those modeled above, imply a positive relationship, which we describe as *cost-following*. The hospital pricing literature finds mixed evidence, often reporting a negative relationship described as *cost-shifting*. Cost-shifting has long featured prominently in the public debate over Medicare payments, to the extent that the *New York Times* takes it for granted. In 2006, the paper reported that "Employers and consumers are paying billions of dollars more a year for medical care to compensate for imbalances in the nation's health care system resulting from tight Medicare and Medicaid budgets, according to Blue Cross officials and independent actuaries" (Freudenheim 2006).[40]

Our framework emphasizes that the private price of our profit-maximizing physician relates positively to her opportunity cost, generating the result that $\frac{dp^*}{dr_M} > 0$. Consistent with the empirical results from prior sections, the framework thus predicts that private payments will move in the same direction as Medicare's fees. But the literature on hospital pricing highlights that the possibility of cost-shifting should not be dismissed. Appendix B.1 shows how an extension of our model that relaxes the capacity constraint and allows for general cost functions can capture this phenomenon. Cost-shifting could also arise through altruism (Cutler 1998, Dranove et al. 2013), income effects (McGuire and Pauly 1991), a change in efficiency, or changes in fixed costs as a physician's scale increases (Kessler 2007).

The relevance of $\frac{dr_p^*}{dr_M}$ for Medicare's influence on aggregate health expenditures, a topic of considerable policy interest, is readily apparent. In a cost-shifting world, Medicare's payment changes are offset by changes in private expenditures—at least in part. In a world of cost-following, Medicare exerts significant influence over total spending.[41] As we emphasize in section 3.2, links between public and private prices have further implications for the health sector's long run development. Cost-following empowers Medicare to shift the distribution of resources across physician specialties, potentially influencing margins including new medical graduates' specialty choices.

We next consider the relationships between market characteristics, average price levels, and cost-following. Consistent with standard intuition about imperfect competition, our model points to a central role for the elasticity of insurers' derived demand for the physician's services. Below we briefly consider cases of constant and non-constant demand elasticities, which diverge in their implications for the link between cost-following and competition.

[40] As Dranove et al. (2013) document, it also shows up in numerous government reports (ProPAC, 1992) and even Supreme Court decisions (Ginsburg 2012).

[41] A growing body of research documents the strain these expenditures cause for federal (Baicker, Shepard and Skinner 2013), state (Baicker, Clemens and Singhal 2012), corporate (Cutler and Madrian 1998), and household (Gross and Notowidigdo 2011) budgets.

Consider first the case in which the demand curve has a constant elasticity of $\bar{\epsilon}$, which implies a constant markup. In this case, the physician's opportunity cost αr_M passes through into prices proportionally, scaled by this markup. Situations that generate a higher $\bar{\epsilon}$, such as more competition among physicians, will induce lower prices and less cost-following.

Now consider a demand function that generates variable markups. Specifically, suppose that $D(p) = (A - p)^\eta$ where $\eta > 1$ scales the demand elasticity.[42] This demand function has a choke point at $p = A$; insurers have an absolute cap of A on what they will pay for this physician's services. We assume that $A \geq \alpha r_M$, as the physician would otherwise only treat Medicare patients. With this demand curve, the physician faces a higher elasticity as prices approach the choke point; specifically, $\epsilon^D(p) = \dfrac{\eta p}{A - p}$. This induces lower markups as prices increase. The physician thus chooses the price to be a weighted average of the choke price and the outside option, where the weights depend on the elasticity parameter η:

$$p^* = \frac{1}{\eta + 1} A + \frac{\eta}{\eta + 1} \alpha r_M. \tag{10}$$

The more elastic demand is, the closer the pricing will be to the physician's outside option of αr_M. Since $A \geq \alpha r_M$, more elastic demand (higher η) reduces prices. But since these lower prices are increasingly pinned down by the outside option, we observe a higher pass-through of Medicare's reimbursement rate as η increases: $\dfrac{dp^*}{dr_M} = \dfrac{\eta \alpha}{\eta + 1}$. So this variable-markups case induces *more* cost-following in the same competitive situations that generate *lower* physician markups.[43]

While the constant and non-constant elasticity demand functions both imply positive co-movement of Medicare reimbursements and private prices, they have different predictions for the relationship between this co-movement and market conditions. In the constant elasticity

[42] Appendix B.2 shows a utility function that gives rise to this demand.

[43] Equation (10) looks like the outcome of a bargaining model in which the insurer has a threat point of A and the physician of αr_M. We discuss this interpretation below in footnote 56.

case, more competition leads to lower prices and less cost-following. In the variable markup case, more competition leads to lower prices but more cost-following. In section 6 we explore heterogeneity in the empirical magnitude of the Medicare-private pricing relationship. We first examine the model's most basic predictions involving the cross-sectional relationship between Medicare and private prices.

5 Basic Relationships Between Public and Private Prices

Before we dig into the mechanics of cost-following, it is useful to see whether our model captures core features of private sector pricing. Section 5.1 compares reimbursement rates between Medicare and the private sector. Section 5.2 introduces measures of concentration on both the physician and insurer sides of the private market. Section 5.3 then examines how these concentration measures relate to pricing. After establishing these core facts, we will be prepared to measure and understand heterogeneity in cost-following in section 6.

5.1 Private and Public Relative Prices

In addition to its predictions about cost-following, the model in section 4 gives guidance about other features of how private and public prices relate. First, the model requires that physicians will never treat privately insured patients unless they pay weakly more than the outside option of treating Medicare patients, scaled by α. While we are not confident about a universal relative cost of treating private versus Medicare patients, evidence from Glied and Graff Zivin (2002) suggests that $\alpha \approx 1$.[44] When this is true, our model predicts that private prices will, in general, exceed Medicare's.

[44]One aspect of the relative cost of treating Medicare versus private patients is the physician's time. Consistent with an assumption of $\alpha \approx 1$, Glied and Graff Zivin (2002) find that doctors spend very similar lengths of time with these two groups of patients. But this is not dispositive; Medicare patients may still be cheaper because they can visit during times of lower private demand, or costlier because of more complex medical situations (In the presence of multiple conditions, providing any particular service for any one of them is likely to be more involved for the physician.)

Table 4 shows how often this prediction is upheld and how often rejected in our data. Row 1 of Panel A compares average prices by state, year and service between the Medicare and MarketScan data. Column 1 shows that 79 percent of these cells have private payments exceeding Medicare's price per service. Subsequent columns reveal that this fraction rises, to between 82 and 89 percent, when cells are weighted to account for service volumes (columns 2 and 3) or spending (columns 4 and 5). Rows 2 and 3 show that private payments exceed public payments more often for surgical than for non-surgical services.[45]

Panel B, and Figure 5, show the magnitudes of the differences between public and private prices. Figure 5 shows the distribution of log public-private price differences at the state-year-service level. For the vast majority of these cells, this difference is strongly positive. Instances in which public payments exceed private payments are relatively few, and the magnitudes swamped by the typical private-over-public mark-up. Panel B of Table 4 quantifies this fact.

To further summarize the cross-sectional relationship between public and private sector prices, Panel C shows results from regressions of private against Medicare prices. The coefficient on Medicare's payments is consistently near 1.4, reflecting the average mark-ups apparent in Panel B and Figure 5. Controlling for year or for state-by-year effects has no effect on this relationship. The simple bivariate regression reported in column 1 has an R^2 of 0.81, revealing the tightness of the relationship between public and private prices.[46] These facts comport well with our model and suggest that it captures some of the key facts about private insurance payments. Thus it might also help us to understand the nature of cost-following, and the heterogeneity that we explore in the next section.

[45]Row 4 accounts for payment variability within each state-year-service cell and finds results very similar to row 1.

[46]Note that in Figure 2 we found an even better fit when estimating this relationship in logs.

5.2 Measuring Physician and Insurer Concentration

We proxy for the degree of physician competition by computing a Herfindahl-Hirschman Index (HHI) for each market. These HHIs are intended to proxy for the demand elasticity ϵ^D in the model. Physicians in relatively competitive markets face relatively elastic demand from insurers for their services. In contrast, when insurers are more concentrated, they should exhibit more elastic demand and have more ability to extract lower payments out of physician groups.

We compute the physician HHIs as follows. We first identify physician groups in the Medicare claims data using the tax identifier associated with each claim. These tax IDs indicate the physician, group, or legal entity that Medicare reimburses for the care. These IDs generally also identify the units that negotiate with insurers.[47] In claims data from a 20 percent sample of all Medicare beneficiaries, we should come close to capturing all Medicare-serving physicians in the country. Treating each Hospital Referral Region (HRR) as the relevant market, we first measure the HHI across physician groups within an HRR.[48] We then average this measure across the HRRs within each state to measure the average degree of competition across the markets within that state. HRRs are an imperfect approximation of the relevant market, but Dranove and Ody (2014) show—for hospitals—that HRR-based HHIs are highly correlated with finer measures of market power. This process gives us our first proxy for physician competition, which varies at the state level.

We next compute a more targeted measure of concentration that varies across specialties as well as states. For this metric we construct HRR-level HHIs separately for each of the

[47] The billing groups may not agree exactly with the negotiating units because of independent practice associations (IPAs), which negotiate as a bloc but bill separately. But the tax IDs should nevertheless be a close approximation. Pope, Trisolini, Kautter and Adamache (2002), Pham, Schrag, O'Malley, Wu and Bach (2007), Centers for Medicare & Medicaid Services (2011), Welch, Cuellar, Stearns and Bindman (2013), Baker, Bundorf, Royalty and Levin (2014), and other authors have previously made the same approximation.

[48] Physician HHI is $\sum_{k=1}^{N} s_{k,i}^2$, where k indexes each of the N physician groups (identified in the claims data via their tax identifiers) operating in Hospital Referral Region i, and where $s_{k,i}$ expresses the number of physicians in group k as a share of all physicians in region i. The measure is constructed such that an index of 1 corresponds to a monopolist and a market approaches perfect competition as the index goes to 0.

32 largest physician specialties. We again average these specialty-specific HHIs across the HRRs within each state. Table 1 reports summary statistics describing both measures of provider consolidation. On average, the specialty-specific HHIs exhibit greater concentration since they consider smaller markets. They also exhibit more variation than the all-physician HHIs.

We measure insurance competition using data from the National Association of Insurance Commissioners (NAIC)'s health insurance reports.[49] Using NAIC data on each insurance carrier's size in each state, we are able to compute state-level HHIs for all states but California.[50] We compute HHIs based on the following four insurer size measures contained in the NAIC reports: enrollment in comprehensive group insurance plans in 2001, enrollment in all plans in 2001, the value of health care provided in 2001, and group comprehensive enrollment in 2002.[51]

5.3 Concentration and Private Prices

Figure 6 provides suggestive evidence that our HHI measures do indeed capture economically relevant aspects of competition. The figure shows a smoothed measure of the average price per service in our sample, averaged across all services, based on the HHI in the area where the service was provided (along the horizontal axis). The two average price lines in Figure 6 are split based on the first NAIC-based measure of insurer concentration. Consistent with Dunn and Shapiro (2012), both curves show that average physician payments are higher in areas with more concentrated physicians. At the same time, more concen-

[49] The earliest comprehensive NAIC reports available are from 2001, and California data are mostly missing and are therefore excluded. For more details on the ultimate sources and issues that arise when computing health insurance market shares, see Dafny, Dranove, Limbrock and Scott Morton (2011). We thank Dafny et al. for useful information on NAIC and other data sources in the paper and via personal communication.

[50] Insurer HHI is $\sum_{k=1}^{N} s_{k,i}^2$, where k indexes each of the N insurers operating in payment area i and where $s_{k,i}$ is insurer k's market share.

[51] Data Source: National Association of Insurance Commissioners, by permission. The NAIC does not endorse any analysis or conclusions based upon the use of its data.

trated insurance markets tend to pay physicians lower reimbursement rates, at any given level of physician concentration. We interpret this as evidence that our measure captures economically meaningful variation in competition among insurance carriers.

The last two regressions of Table 4 summarize these facts. We first standardize the HHI variables as z scores. Column 4 of Panel C includes the physician HHI alongside the services' average Medicare payment. Column 5 adds the insurance HHI and interactions between the HHI measures and the Medicare payment. As in Figure 6, more concentrated insurance markets are associated with lower reimbursements while more concentrated physician markets are associated with higher reimbursements. Since we have not isolated exogenous variation in these measures of market structure, we do not ascribe a causal interpretation to these results. But they do demonstrate the model's general consistency with the data. We therefore believe that section 4 presents a useful framework for understanding cost-following and the heterogeneity that we document next.

6 The Mediating Effects of Concentration

In section 4 we showed theoretically how the effect of Medicare's payment changes could vary across markets based on the level of competition among physician groups and insurers. In this section, we examine the extent to which these relationships emerge in our data. We adapt our empirical specifications to allow for heterogeneity in the size of Medicare's effect and explore its correlation with measures of competition across physician groups and insurance carriers.

To estimate the relationship between provider concentration and the cost-following coefficient, we interact the price shocks with either the all-physician or specialty-specific HHI. Recalling that $\text{PredChg}_j^{\text{Medicare}}$ is the predicted Medicare price change, and letting $HHI_{j,s}$ denote the applicable HHI z-score, we run reduced-form specifications of the following form:

$$
\begin{aligned}
P^{\text{Private}}_{j,s,t} = {} & \beta_1 \cdot \text{PredChg}^{\text{Medicare}}_j \cdot \text{Post1998}_t \\
& + \beta_2 \cdot \text{PredChg}^{\text{Medicare}}_j \cdot \text{Post1998}_t \cdot HHI_{j,s} \\
& + X_{j,s,t}\gamma_1 + X_{j,s,t} \cdot HHI_{j,s}\gamma_2 + \mu^1_j \mathbb{1}_j + \mu^1_s \mathbb{1}_s + \mu^1_t \mathbb{1}_t \\
& + \mu^2_j \mathbb{1}_j \cdot HHI_{j,s} + \mu^2_s \mathbb{1}_s \cdot HHI_{j,s} + \mu^2_t \mathbb{1}_s \cdot HHI_{j,s} \\
& + \mu^1_{j,s} \mathbb{1}_j \cdot \mathbb{1}_s + \mu^1_{t,s} \mathbb{1}_t \cdot \mathbb{1}_s + \mu^2_{j,s} \mathbb{1}_j \cdot \mathbb{1}_s \cdot HHI_{j,s} + e_{j,s,t} \qquad (11)
\end{aligned}
$$

We allow the coefficients on all time-varying controls to vary with the relevant HHI variable.[52]

6.1 Heterogeneity by Physician Market Power

Table 5 presents the estimates. Column 1 shows that the average cost-following coefficient is roughly 1.4, but that it varies significantly with the all-physician and specialty-specific HHIs. The coefficient of -0.5 on the physician HHI interaction implies that as HHI increases by 1 standard deviation, the cost-following coefficient falls by two-fifths of its value at the mean; the point estimate is statistically distinguishable from zero at the $p < 0.01$ level. Cost following in relatively uncompetitive markets is thus much weaker than in the most competitive markets.

In column 2, we add an interaction between the predicted payment shocks and the number of physicians in a market (also measured as a z-score). This variable enters significantly, but with little impact on the coefficient associated with the HHI interaction. The HHI coefficient is thus not merely capturing differences in the absolute sizes of the relevant markets. As we discuss below, the total number of physicians correlates strongly with the extent to which

[52] We also graphically report results from specifications in which we divide the sample into terciles of provider consolidation. Estimation on sub-samples implicitly interacts all controls with the HHI variables at no additional computational cost. In equation (11) we have omitted interactions between the HHI variables and the state-by-service code fixed effects ($\mathbb{1}_j \cdot \mathbb{1}_s \cdot HHI_{j,s}$), of which there are in excess of 50,000. Note that because the first stage coefficient in Table 2's levels regression was nearly 1, the reduced-form and IV results are nearly identical. Here and for the remainder of the paper we run reduced form specifications to further ease the computational burden.

markets are urban or rural.

Since they vary both across states and across specialties within each state, the specialty-specific HHIs will give us our most compelling look, in terms of econometric identification, at the role of market power in mediating the effect of Medicare's price changes on private markets. But this comes at a cost in terms of the number of services included in the sample. Specifically, incorporating specialty-specific HHIs requires restricting attention to services that tend to be provided primarily by members of a particular specialty.[53]

Columns 3 through 5 conduct a similar analysis using HHIs measured at the specialty-by-market level. The results are statistically strong and consistent with the all-physician HHI results. The point estimate of interest is robust to controlling for interactions with the number of physicians, either within a specialty or throughout the market.[54] Column 6 includes both the all-physician and specialty-specific interactions. When included jointly, both concentration measures remain strong predictors of the strength of Medicare's price transmission. The results uniformly support the view that Medicare is more relevant in competitive markets than in markets with concentrated providers.

Panel A of Figure 7 reports the first stage and cost-following coefficients separately for each tercile of the specialty HHI distribution. The cost-following coefficient falls from just over 2 in the most competitive tercile to around zero in the most concentrated.

Because our variation in physician concentration is not quasi-experimental, the heterogeneity of our effects is not identified as persuasively as our estimates of Medicare's average effect on private payments. To address concerns that this variation may reflect unobserved

[53]This is true because the private sector claims data say little about the physicians associated with each service. The construction of specialty-specific HHIs and the linking of service codes with particular specialties could only be done consistently in the Medicare claims data. Consequently, the number of distinct service codes in our analysis sample falls from 2,149 to 1,303 for our analysis of specialty-specific consolidation.

[54]Because of the geographic element in our HHI measures, we also show in Appendix Table D.4 that the significance is robust to clustering by state instead of service code. This table also explores robustness to clustering at more aggregated levels of services. The results maintain their precision when we allow for correlated price shocks across distinct, but similar services (as defined using three different levels of Betos codes).

area characteristics, we include state-specific interactions with the payment shock in column 7. In this column, we only identify the coefficients on the specialty-specific market conditions based on cross-specialty variation within a state. The coefficients are virtually unchanged from column 6.

In order to investigate the robustness of the more general result in column 6, Appendix Table D.5 adds a variety of additional controls. Specifically, we control for interactions between the payment shocks and area characteristics including Census region indicators, income per capita, and measures of population density. Note that these controls enter the regressions in the exact same manner, and with the same degree of flexibility, as the HHIs. Even so, these controls have little effect on the coefficients associated with our measures of provider consolidation. This contrasts with the coefficients on the total numbers of physicians, which are less stable. This supports our initial interpretation that the physician count is a proxy for an area's urban status; it is strongly correlated with county characteristics such as total population and population density. Even after adding these controls, our finding that Medicare exerts particularly strong influence over the payments made to competitive physician groups appears to be quite robust.

6.2 Insurance Competition and Price Transmission

We now explore the relevance of competition on the insurers' side of the market. Competition in insurance markets likely has the opposite impact from competition in physician markets. While competition among physicians should reduce each physician's market power, or increase her perceived elasticity of demand, competition among insurers should do the opposite. Higher insurer concentration implies that the physician faces more elastic demand, and hence prices more aggressively, than with competitive insurers. As section 4 demonstrated, the impact of these differences on cost-following depends on the shape of the demand curve.

As in section 6.1, we convert HHIs into z-scores and run regressions paralleling equation (11). Table 6 presents the results. The reduced sample size relative to Table 2 reflects the omission of California from the insurance market data. Columns 1 through 3 reveal a positive relationship between insurance concentration and the magnitude of Medicare's effect on private payments. The precision varies depending on which measure of market share is used to compute HHIs. Depending on the measure, a one standard deviation increase in concentration is associated with a \$0.15 to \$0.36 increase in the cost-following coefficient.[55]

The distribution of HHIs is asymmetric. The mean HHI in our sample is 0.25 and the standard deviation is 0.17. The fifth percentile of insurer HHI is 0.08, which is associated with a cost-following coefficient of 0.85. The ninety-fifth percentile HHI is 0.78, with an implied cost-following coefficient of 2.2. We graphically display results by tercile of concentration in Panel B of Figure 7.

Columns 4 through 6 of Table 6 show that coefficient on the insurer-HHI interaction is robust to controlling for interactions with the number of insurers in each state as well as our measure of physician concentration. Appendix Table D.6 reports further evidence on the robustness of this result. The estimate associated with the insurer HHI is stable when controlling for interactions between the payment shocks and a variety of state economic and demographic characteristics.

Thus far, this section's results have used insurance concentration data from 2001, as it is the earliest year for which we have comprehensive insurer enrollment data. Appendix Table D.7 explores the results' robustness to substituting measures constructed using HMO enrollment data from earlier and later years. When we measure concentration in 1997, 1998, or 2002, the estimated effect of insurance competition is unchanged. The HHI from 1996 has a lower and noisier coefficient, and is indistinguishable from zero. If we measure

[55]We obtain these results by intermingling data computed from the NAIC insurer reports with other data sources. These results are not NAIC information and NAIC is not responsible for any analysis or conclusions drawn as a result of this intermingling.

insurance concentration in 2006, using either NAIC data or a separate measure available from the American Medical Association (2007), the estimated coefficient is close to 0. Since the 2006 market shares are estimated nearly a decade following the surgical/non-surgical payment shock, they are unlikely to accurately reflect conditions at the time of our natural experiment.

The results in this and the previous section tell a consistent story that informs our interpretation of the theoretical model. As we saw in Figure 6, increases in doctors' market power—whether due to physician group concentration or insurer competition—are associated with charging higher prices to insurers. Now we have also seen that these same conditions are associated with weaker cost-following. These results are consistent with the variable-markups version of our model in section 4. In this case, the physician's pricing is given by equation (10), which predicts exactly these patterns.[56] In order to better understand how these prices are set in practice, we next discuss institutional details of the price-setting process and examine their implications.

7 The Mechanics of Medicare's Influence On Private Prices

Thus far we have interpreted our results through a standard price-theoretic lens. Here we explore further considerations associated with institutional details emphasized by practitioners (*e.g.* Nandedkar 2011, Gesme and Wiseman 2010, Mertz 2004). The market for physician care faces numerous frictions and a morass of complex transactions (Cutler and Ly 2011). We

[56]This equation also looks like the outcome of a bargaining model in which the physician has bargaining power of $1/(\eta + 1)$ and the insurer $\eta/(\eta + 1)$. This would generate similar predictions. Without richer data on specific physician-insurer relationships, we cannot distinguish between a bargaining story and our model from section 4 in which the physician sets prices. (The private claims data do not provide any information about the specific insurer that processes each claim, and we cannot link physicians across firms, so we can't observe whether the physician and insurer reach agreement or break apart negotiations and exercise their outside options.) But given the similarity in the pricing equation, the bargaining model would also yield similar predictions.

thus consider how transaction costs might shape the structure of physician-insurer contracting (Coase 1937, 1960). The resulting insights contribute to our understanding of the overall magnitude of Medicare's effects on private prices, and especially its influence on services' relative valuations.

Practitioners describe two modes of negotiation between providers and private insurers. Insurers typically offer small physician groups contracts based on a fixed fee schedule. This may be Medicare's schedule itself, or a schedule that the insurer has modified. The parties then negotiate a constant scaling of this schedule. In contrast, insurers are said to negotiate in more detail with hospitals and large groups over service-specific pricing.[57]

Adopting Medicare's fee schedule may be optimal due to the substantial negotiation and coordination costs in this setting. But if Medicare's menu has inefficiencies, the value of an insurer's product can potentially be improved through more detailed negotiations over service- or bundle-specific prices.[58] The essential insight for rationalizing practitioners' descriptions of these negotiations is that the value of such improvements rises with the scale of the physician group. When insurers contract with small physician groups, the transaction costs associated with detailed negotiations may regularly outweigh the cost of inefficiencies in Medicare's menu. By contrast, a great deal of value may be at stake when insurers contract with large physician groups.

The implications of this institutional story differ subtly from the price theory considerations we modeled in section 4. Specifically, it emphasizes the absolute size of physician groups, rather than their market power, as a factor that would mediate cost-following's

[57]Their own descriptions can be found in Appendix A. Documentary evidence for this adoption can be found in many insurers' monthly newsletters for participating providers (*e.g.* Blue Cross and Blue Shield of Texas 2010, Anthem Blue Cross and Blue Shield 2012b, BlueCross BlueShield of Illinois 2013, Blue Cross Blue Shield of Michigan 2013)

[58]Since Medicare's payments are cost-based, they likely deviate from the efficient price for service j. In this context, cost-based means the average cost of care at observed quantities. Since Medicare beneficiaries, in particular the 90 percent with supplemental insurance (MedPAC 2011), are comprehensively insured, there may be a substantial wedge between marginal cost and marginal benefit at these quantities.

magnitude.[59]

To take this view to the data, we proxy for the presence of small physician groups using two variables. The first is the share of physicians working as sole practitioners, measured at the level of specialties and hospital referral regions. The second, similarly constructed, is the share of physicians working in groups with five or fewer members. We use these variables to estimate specifications in the form of equation (11), in which we allow the effect of Medicare's price shocks to vary with the fraction of physicians practicing in small groups.

Table 7 shows the results of this exercise. The prevalence of small physician groups is strongly associated with the extent to which Medicare influences private prices. A one standard deviation increase in either proxy is associated with an increase of 0.8 in the cost-following coefficient. In columns 2 and 4 we augment the specification with full sets of interactions with our measure of physician market power. The measures of absolute size and market power separately predict variation in the cost-following coefficient in the expected directions.

Properly interpreting the results in columns 2 and 4 is difficult for two reasons. First, HHI and the small group market share are quite correlated with one another, and are imperfect proxies for the relevant economic forces. Second, the coefficient on each variable is estimated conditional on the other. Increasing the small group market share conditional on HHI, for example, requires increasing the size of the market's other, larger groups. It is thus not obvious that the magnitudes of both forces can be estimated simultaneously. Consequently, these specifications can only suggestively speak to the distinctive relevance of market power and absolute group size.

[59]In Appendix D.2, we show that Medicare also has a larger influence when it commands a larger share of the market for a particular service.

8 Conclusion

This paper aims to advance our understanding of price determination in the important, complex, and opaque markets for physician services. We show that Medicare exerts widespread and quantitatively substantial influence over the rates that private insurers pay. We estimate that a $1 change in Medicare's payments for one service relative to another results in a $1.16 change in private payments. Section 3.2 showed that these payment changes reoriented billions of both public and private sector health dollars. We find that Medicare similarly moves the level of private payments when it alters fees across the board. In aggregate, our estimates imply that Medicare's pricing decisions can appreciably move both health-sector and overall inflation (Clemens et al. 2014).

This analysis lays the groundwork for a new strand of research into the economics of physician pricing. Our evidence on how public programs influence private reimbursement rates creates at least three distinct opportunities for future work. First, public-private linkages may also be of interest with respect to state Medicaid programs. Anecdotal evidence, coupled with the economics of outside options, suggests that Medicaid's rates may be particularly relevant for plans sold on the Affordable Care Act's exchanges.[60] Second, our analysis sample involves payments made on an exclusively fee-for-services basis. While we know less about physician payments within HMOs, documentary evidence suggests that such payments also draw heavily on Medicare's menu.[61] Further work that directly analyzes major HMOs' contracts with physicians will be instrumental in determining the pervasiveness of Medicare's

[60]Numerous sources discuss the limited number of providers accepting insurance plans purchased on the exchanges (*e.g.* Pear 2013, Bauman, Coe, Ogden and Parikh 2014, Blumenthal and Collins 2014). Gruber and McKnight (2014) show directly that these "narrow network" plans offer lower payment rates for physician visits, potentially making Medicaid more relevant as an outside option for participating physicians. Anecdotally, Harvey (2014) discusses California exchange plans paying 80 percent of the state's Medicaid rates and Pittman (2013) makes similar claims.

[61]Provider newsletters for non-HMO and HMO plans explicitly describe their use of Medicare's relative values (*e.g.* Blue Cross and Blue Shield of Texas 2010, Anthem Blue Cross and Blue Shield 2012a). These are sizable examples, as the HMO Blue Texas plan advertised having 38,000 physicians in its network as of 2009 (Blue Cross and Blue Shield of Texas 2014)

influence in this important setting. Third, it remains to be seen whether Medicare influences private insurers' generation and adoption of novel payment models. For example, how do private plans react to public insurers' adoption of "bundled payment" mechanisms, or to transitions from cost- to value-based payments?

Separately, our analysis opens a line of questions regarding the mechanisms underlying the magnitude and pervasiveness of Medicare's influence. We emphasize that Medicare likely exerts sway through multiple channels. As a large market participant, it competes with private insurers for physicians' resources. Practitioners further highlight that Medicare's menu serves a benchmarking function in many private insurers' contracts. The latter influence suggests that Medicare may be an essential participant in meaningful payment-system experimentation and reform. Further analyses of these mechanisms may generate insights that enrich our understanding of Medicare's capacity to shape the U.S. health system.

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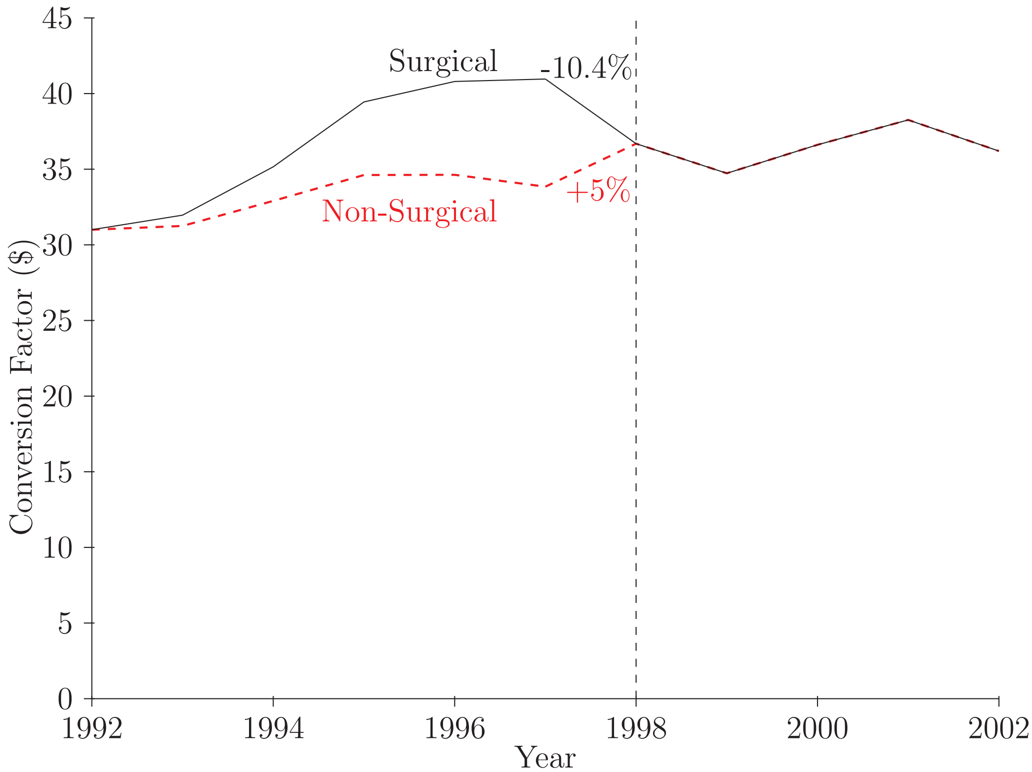
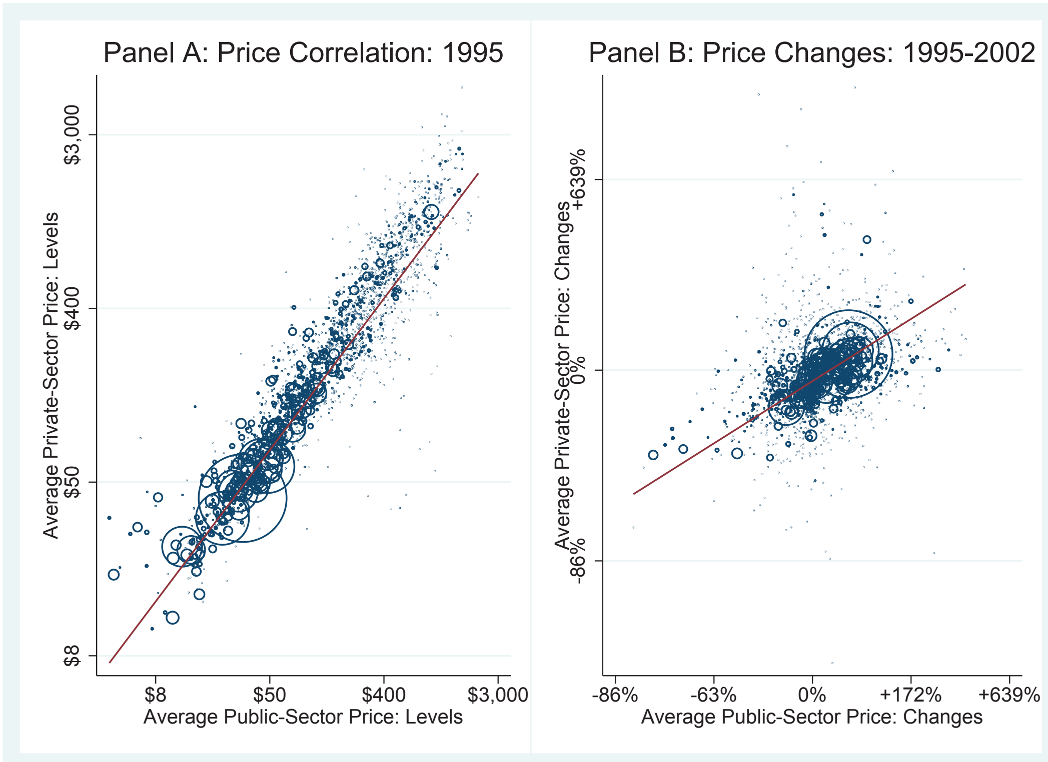
Figure 1: Evolution of Medicare Surgical and Non-Surgical Conversion Factors



This figure shows the nominal Conversion Factors that Medicare applied to surgical and general non-surgical services for each year from 1992 through 2002. Source: American Academy of Pediatrics (2012).

Figure 2: Cross-Service Relationship Between Private and Medicare Prices



Note: This figure shows the raw cross-service relationships between average private reimbursements and average Medicare reimbursements. The values shown are the average payments we observe in our public (Medicare) and private (Medstat) sector claims data, plotted on a log scale. Panel A presents these average payments for 1995 while Panel B shows the changes in these average payments from 1995 to 2002. Circle sizes are proportional to the number of times a code is observed in the Medicare data. The best-fit line shown in Panel A results from estimating

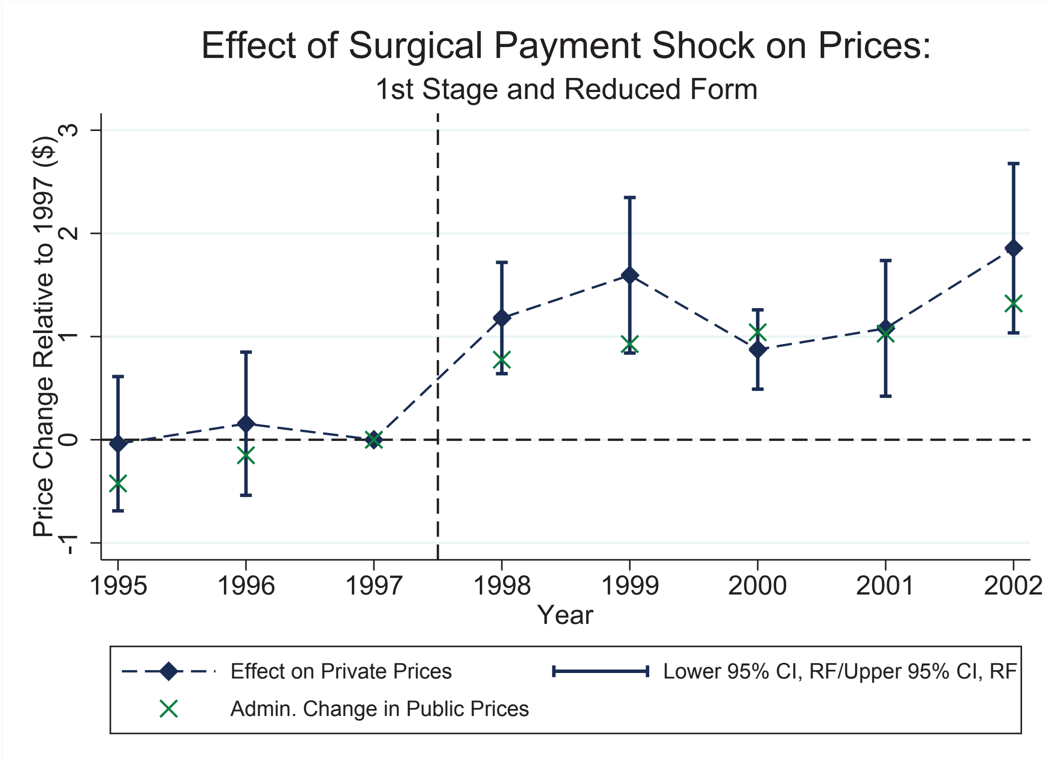
$$\ln(P_j^{\text{Private}}) = \beta_0 + \beta_1 \ln(P_j^{\text{Medicare}}) + u_j$$

across services j, weighted by the code's frequency. The regression yields a coefficient of $\beta_1 = 0.87$ and $R^2 = 0.89$ with $N = 2,194$. The best-fit line shown in Panel B results from estimating

$$\triangle \ln(P_j^{\text{Private}}) = \gamma_0 + \gamma_1 \triangle \ln(P_j^{\text{Medicare}}) + v_j,$$

again weighted by the code's frequency. The regression yields a coefficient of $\gamma_1 = 0.65$ and $R^2 = 0.60$ with $N = 2,194$. Note that the regressions are run in logs and the values shown along the axes are computed by exponentiating the log values.

Figure 3: Effects of Medicare's Elimination of the Surgical Conversion Factor



Note: This figure presents the δ_t coefficients, with associated 95% confidence intervals, from estimates of the equation below:

$$P^{\text{Private}}_{j,s,t} = \sum_{t \neq 1997} \delta_t \cdot \mathbb{1}_t \cdot \text{PredChg}^{\text{Medicare}}_j + X_{j,s,t}\alpha + \upsilon_j \mathbb{1}_j + \upsilon_s \mathbb{1}_s + \upsilon_t \mathbb{1}_t$$
$$+ \upsilon_{j,s} \mathbb{1}_j \cdot \mathbb{1}_s + \upsilon_{t,s} \mathbb{1}_t \cdot \mathbb{1}_s + \upsilon_{j,s,t}.$$

where $\text{PredChg}^{\text{Medicare}}_j$ are the predicted changes in Medicare payments associated with the elimination of the separate surgical conversion factor. The figure also plots the point estimates from the associated first stage, showing that our coding of $\text{PredChg}^{\text{Medicare}}_j$ correctly tracks the Conversion Factors' merger. The dependent variable is the level of the average private payment, calculated at the service-by-state-by-year level, that we observe in our data on private sector claims. Controls include full sets of service-by-state ($\mathbb{1}_j \cdot \mathbb{1}_s$) and state-by-year ($\mathbb{1}_t \cdot \mathbb{1}_s$) fixed effects, corresponding direct effects, as well as indicator variables that account for sharp reductions in Medicare's payments for cataract surgery that occurred during the mid-1990s. Also included are two variables accounting for the insurance plan types associated with our data on private sector claims. Standard errors are clustered at the service code level. Sources: Authors' calculations using Medicare and Thompson Reuters MarketScan data.

Figure 4: Effect of Geographic Payment Shocks on Private Prices

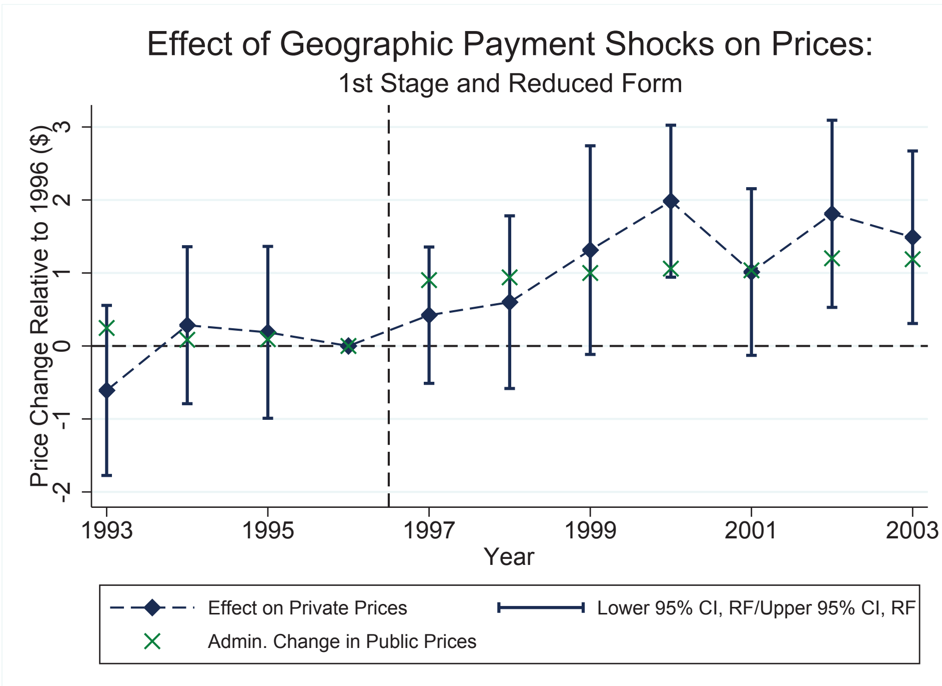
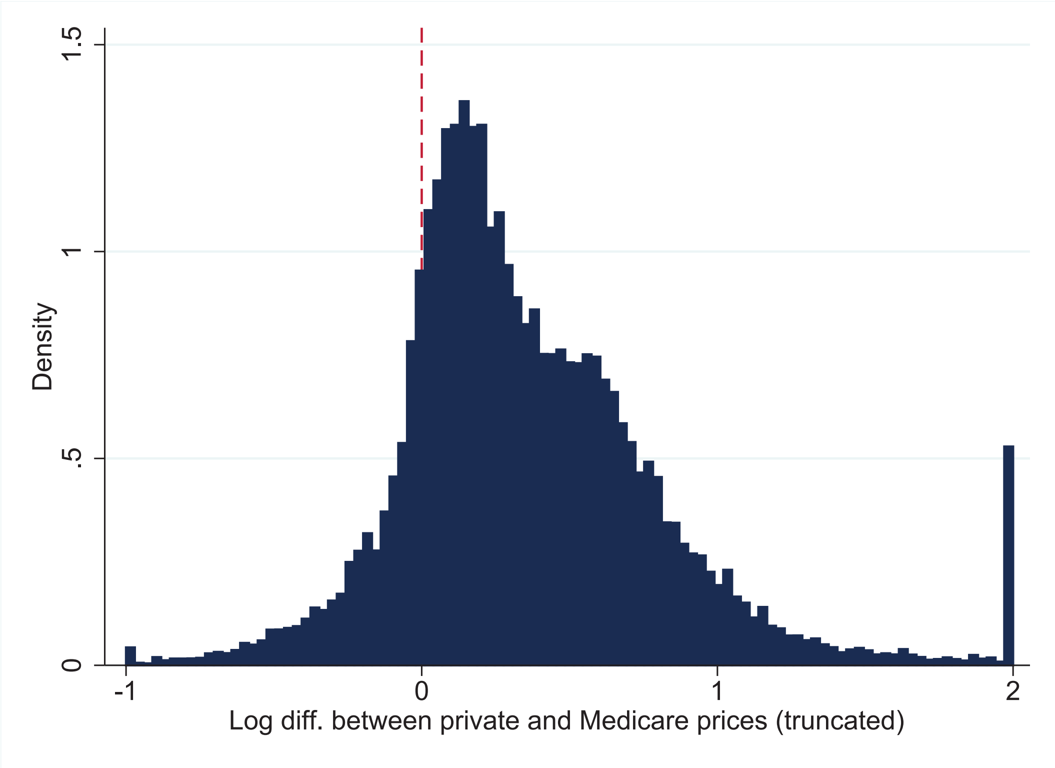


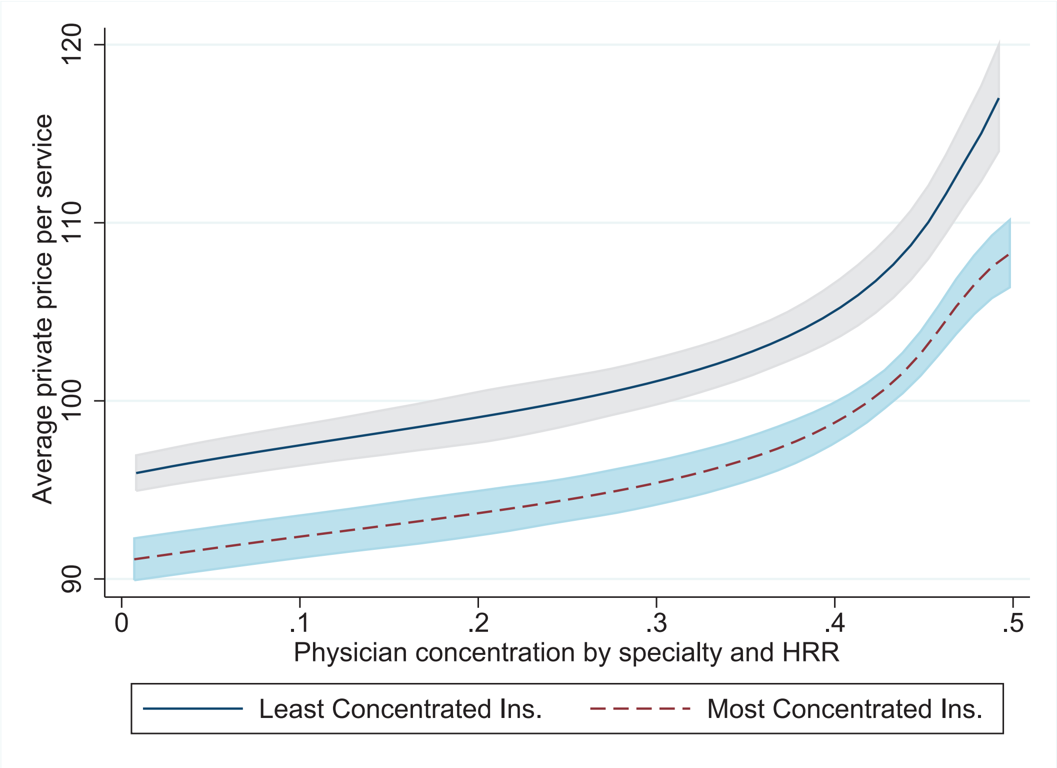
Figure shows the the results from estimating equations (7) and (8) as described in section 3.1. The payment shocks are constructed such that a one unit change in the payment shock should correspond to a one dollar increase in Medicare's payments. This is confirmed by the point estimates labeled "Admin. Change in Public Prices." Estimates labeled "Effect on Private Prices" are the corresponding estimates associated with the relationship between Medicare's payment shocks and private sector prices. Sources: *Federal Register*, various issues; Authors' calculations using Medicare claims, Thompson Reuters MarketScan data, and Ruggles et al. (2010).

Figure 5: Distribution of Private-Medicare Price Difference



Note: This figure shows the distribution of the difference between the log average prices in the private and Medicare databases, across all services, states and years. Values are winsorized at -1 and +2. Sources: Authors' calculations using Medicare claims and Thompson Reuters MarketScan data.

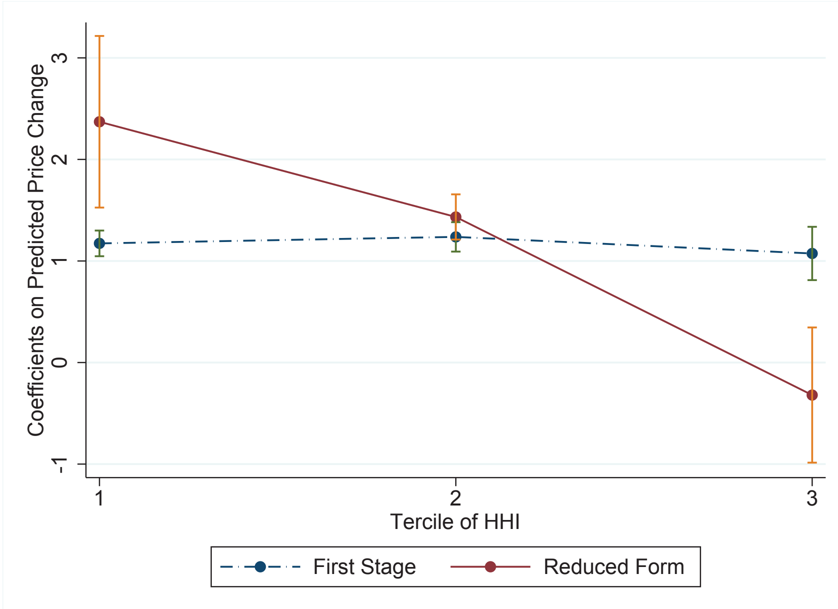
Figure 6: Variation in Private Prices with Provider and Insurer Market Power



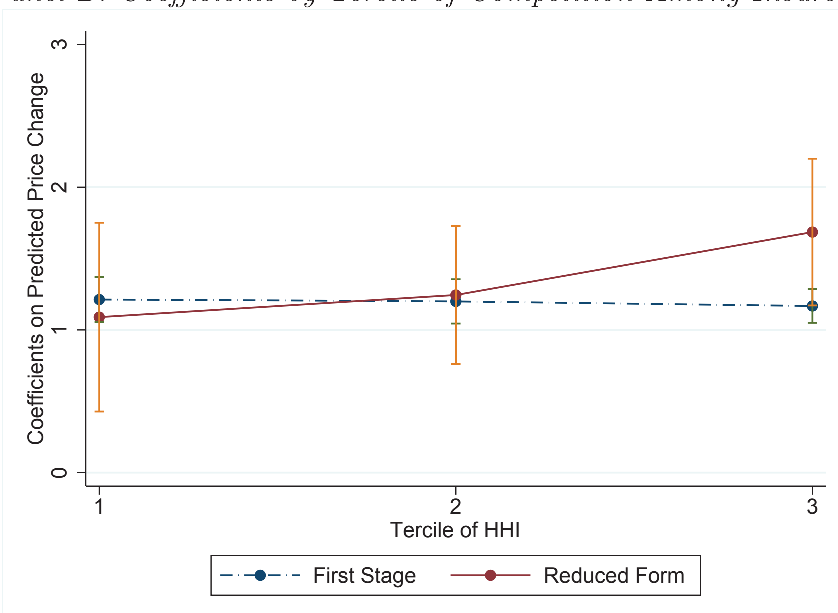
Note: This figure shows the average private sector payments separately in low-concentration (blue solid line) and high-concentration (red dashed line) insurance markets, based on the degree of provider concentration (along the x-axis). The private payments are averaged across all years, states, and services.

Figure 7: First Stage and Reduced Form Coefficients by Tercile of Competition

Panel A: Coefficients by Tercile of Competition Among Physicians



Panel B: Coefficients by Tercile of Competition Among Insurers



This figure shows coefficients of Medicare price and private prices on the predicted price change interacted with years following its implementation, from specifications based on equation (3), with associated 95% confidence intervals. Coefficients are estimated separately when cutting the sample by the HHI of (A) physician groups, computed at the specialty-by-state level, and (B) insurance carriers, computed by state. In each panel, the dashed line shows first-stage coefficients indicating the impact on Medicare payments. The solid line shows reduced form coefficients indicating the impact on private insurer reimbursement rates.

Data Source for the insurance concentration measures: National Association of Insurance Commissioners, by permission. The NAIC does not endorse any analysis or conclusions based upon the use of its data.

Table 1: Summary Statistics

		Non-Surgical Care $N = 140,716$	Surgical Services $N = 163,012$
Private	Mean/SD:	$125.17 ($133.35)	$374.48 ($441.33)
Payments per Service	Range:	[$1.10, $6,967]	[$0.30, $21,891]
Medicare	Mean/SD:	$114.44 ($148.39)	$239.23 ($248.05)
Payments per Service	Range:	[$3.28, $1,150]	[$3.51, $2,112]
Standard Deviation	Mean/SD:	$84.07 ($127.61)	$257.76 ($403.80)
Private Payments per Service	Range:	[$0.00, $9,503]	[$0.00, $34,634]
Standard Deviation	Mean/SD:	$17.60 ($27.03)	$78.05 ($106.09)
Medicare Payments per Service	Range:	[$0.41, $260.65]	[$0.20, $1,320]
Plan Type	Mean/SD:	83.06 (43.41)	237.43 (142.34)
Payment Generosity	Range:	[15.52, 586.00]	[19.06, 670.95]
Service Specific	Mean/SD:	0.226 (0.188)	0.180 (0.180)
Cost Sharing	Range:	[0.00, 1.00]	[0.00, 1.00]
State level	Mean/SD:	0.023 (0.024)	0.022 (0.020)
Physician HHI	Range:	[0.003, 0.14]	[0.003, 0.14]
Physician Specialty	Mean/SD:	0.157 (0.100)	0.147 (0.109)
HHI	Range:	[0.007, 0.79]	[0.007, 1.00]
Private Market	Mean/SD:	61.60 (309.61)	11.43 (46.57)
Volume (1000s)	Range:	[0.01, 4898.08]	[0.01, 721.61]
Medicare	Mean/SD:	7.28 (20.94)	2.66 (4.62)
Relative Size	Range:	[0.003, 459.83]	[0.015, 136.87]
State Level	Mean/SD:	0.258 (0.187)	0.227 (0.167)
Insurer HHI (NAIC)	Range:	[0.00, 0.95]	[0.00, 0.95]
State Level	Mean/SD:	0.335 (0.135)	0.324 (0.122)
Insurer HHI (AMA)	Range:	[0.15, 0.69]	[0.15, 0.69]

Note: This table shows summary statistics for our data on public and private payments, characteristics of the private plans we observe, and the characteristics of the geographic and service-specific markets that we use to explore heterogeneity in the effect of Medicare price changes on public prices. Observations are constructed at the service-by-state-by-year level and the panel is balanced in the sense that each service-by-state panel is only included if public and private prices are available for each year from 1995 through 2002. Private and Medicare Payments Per Service are expressed in dollars and are the average payment within each service-by-state-by-year cell. The standard deviations are correspondingly standard deviations of claims-level payments within these cells. The construction of "Plan Type Payment Generosity" and "Out of Pocket Share" is described in section 1.4, and that of the HHI variables in section 5.2. The first insurance market HHI variable comes from authors' calculations on data obtained from the National Association of Insurance Commissioners (NAIC), and NAIC is not responsible for these calculations. The second insurance market HHI variable is provided directly by the American Medical Association (2007), which does not provide HHIs for the following states: KS, ND, MS, PA, SD, WV, and DC. "Private Market Volume" expresses (in tens of thousands of dollars) the total payments associated with each service in private sector claims data. "Medicare Relative Size" is the ratio of the number of times a service appears in the Medicare claims data and in the private-sector claims data. Sources: Medicare claims and Thompson Reuters MarketScan data (lines 1–10). Line 11: National Association of Insurance Commissioners, by permission. The NAIC does not endorse any analysis or conclusions based upon the use of its data. Line 12: American Medical Association (2007).

Table 2: **Baseline Estimates of the Effect of Medicare Price Changes on Private Sector Prices**

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable:	Payment Levels			Log Payments		
	Public 1st Stage	Private Red. Form	Private IV	Public 1st Stage	Private Red. Form	Private IV
Payment Shock × Post 1997	1.192**	1.385**				
	(0.072)	(0.261)				
Public Payment			1.162**			
			(0.225)			
Surgical Procedure × Post 1997				-0.225**	-0.108**	
				(0.032)	(0.027)	
Ln(Public Payment)						0.480**
						(0.062)
Plan Type Payment Generosity	6.861	-3.543	-11.514	0.074	0.009	-0.026
	(8.988)	(22.334)	(28.184)	(0.041)	(0.021)	(0.028)
Service Specific Cost Sharing	-0.972+	-3.439	-2.310	-0.008	-0.005	-0.001
	(0.578)	(6.008)	(6.144)	(0.006)	(0.020)	(0.020)
N	303,728	303,728	303,728	303,120	303,120	303,120
Number of Clusters	2,194	2,194	2,194	2,192	2,192	2,192

Note: **, *, and + indicate statistical significance at the 0.01, 0.05, and 0.10 levels respectively. The table shows the results of OLS and IV specifications of the forms described in section 1.2. Columns 1 and 2 report estimates of equations (3) and its associated reduced form respectively, where the payment shock and outcome variables are expressed in dollar terms. Column 3 reports an estimate of equation (4). Columns 4 through 6 report otherwise equivalent specifications in which the dependent variables are expressed in logs and the instrument is an indicator for surgical procedures performed in years following 1997. Observations are constructed at the service-by-state-by-year level. In columns 1 through 3, observations are weighted according to the number of times the service is observed in Medicare claims in 1997. In columns 4 through 6, the weights reflect each service's average share of payments made through Medicare Part B in 1997. The panel is balanced in the sense that each service-by-state panel is only included if public and private prices are available for each year from 1995 through 2002. Standard errors are calculated allowing for arbitrary correlation among the errors associated with each service. Additional features of each specification are described within the table. The construction of all variables is further described in the note to Table 1 and in the main text. Sources: Authors' calculations using Medicare claims and Thompson Reuters MarketScan data.

Table 3: **Estimates of the Effect of Across-the-Board, Area-Specific, Medicare Payment Shocks on Private Sector Prices**

	(1)	(2)	(3)	(4)
	Private Payment Level			
Payment Shock × Post-1996	1.268*	0.990*	1.267**	0.846
	(0.500)	(0.466)	(0.477)	(0.902)
N	176,960	176,960	176,960	176,960
Number of Clusters	199	199	199	199
State By Year FE	Yes	Yes	Yes	No
Year FE	No	No	No	Yes
HCPCS By Old MPL FE	Yes	Yes	No	Yes
Old MPL FE	No	No	Yes	No
HCPCS FE	No	No	Yes	No
Pop. By Year Controls	No	Yes	No	No
Panel Balanced	Yes	Yes	Yes	Yes

Note: **, *, and + indicate statistical significance at the 0.01, 0.05, and 0.10 levels respectively. The table shows the results of OLS, reduced form specifications taking the form of equation (8). Observations are constructed at the service-by-year-by-payment locality level. The panel is balanced in the sense that each service-by-payment locality panel is only included if public and private prices are available for each year from 1993 through 2003. Standard errors are calculated allowing for arbitrary correlation among the errors associated with each payment locality, which is the level at which the relevant payment shocks occur. Additional features of each specification are described within the table. Sources: Authors' calculations using Medicare claims, Thompson Reuters MarketScan data, and Ruggles et al. (2010).

Table 4: Summarizing the Differences Between Medicare and Private Prices

Panel A: Share of Observations with Private Rates Exceeding Medicare

		(1)	(2)	(3)	(4)	(5)
	Weighting:	None	Service Count in		Total Spending in	
			Medicare	Private	Medicare	Private
(1)	Based on mean price	0.793	0.885	0.825	0.844	0.821
(2)	Surgical only	0.854	0.911	0.883	0.908	0.915
(3)	Non-surgical only	0.722	0.880	0.810	0.814	0.756
(4)	Based on t-test probabilities	0.793	0.885	0.825	0.844	0.819

Panel B: Mean Difference Between Log Private and Medicare Prices

		(1)	(2)	(3)	(4)	(5)
	Weighting:	None	Service Count in		Total Spending in	
			Medicare	Private	Medicare	Private
	Among services with:					
(1)	$P^{\text{Medicare}} > P^{\text{Private}}$	-0.309	-0.122	-0.159	-0.162	-0.176
(2)	$P^{\text{Medicare}} < P^{\text{Private}}$	0.548	0.394	0.413	0.408	0.493

Panel C: Regression of Private on Medicare Prices

	(1)	(2)	(3)	(4)	(5)
Public payment	1.447***	1.447***	1.449***	1.504***	1.426***
	(0.054)	(0.054)	(0.053)	(0.061)	(0.090)
Physician HHI				1.806***	2.930*
				(0.406)	(1.443)
Insurance HHI					-3.188*
					(1.266)
Public payment					-0.023
× Specialty HHI					(0.023)
Public payment					0.067**
× Insurance HHI					(0.024)
N	253,632	253,632	253,632	253,632	253,632
Number of Clusters	1,364	1,364	1,364	1,364	1,364
R^2	0.813	0.814	0.818	0.814	0.827
Fixed Effects	None	Year	State-Year	Year	Year & Specialty
Physician HHI Measure				All MDs	Specialty

Note: **, *, and + indicate statistical significance at the 0.01, 0.05, and 0.10 levels respectively. Panel A shows the share of state-year-service observations in which the average private sector payment exceeds the average Medicare payment. Rows 1–3 simply count the number of state-year-service cells, while applying various different weights, according to the respective column heading, to each cell. Row 4 splits each cell based on the probability ascribed to having an underlying private mean payment above the Medicare mean payment based on a one-tailed t-test of the null hypothesis $P^{\text{Medicare}} \geq P^{\text{Private}}$. Panel B shows the log price difference between the average private payment and average Medicare payment within each cell, split based on which is higher. Panel C regresses the observed average private price on the average Medicare price and, in columns 4 and 5, measures of concentration. Regressions are weighted by service count as in Table 2. Sources: Authors' calculations using Medicare claims and Thompson Reuters MarketScan.

Table 5: **Heterogeneity in Surgical CF Shock's Effect by Provider Concentration**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent Variable:				Private Payment Level			
Payment Shock × Post-1997	1.429**	1.476**	1.272**	1.327**	1.345**	1.410**	
	(0.265)	(0.290)	(0.190)	(0.218)	(0.211)	(0.235)	
Payment Shock × Post-1997	-0.538**	-0.511**				-0.511**	
× Physician HHI	(0.092)	(0.092)				(0.063)	
Payment Shock × Post-1997			-0.891**	-0.846**	-0.827**	-0.516*	-0.599**
× Specialty HHI			(0.286)	(0.275)	(0.278)	(0.224)	(0.222)
Payment Shock × Post-1997		0.353**		0.369**		0.064	
× Physician Count		(0.065)		(0.071)		(0.159)	
Payment Shock × Post-1997					0.408**	0.640**	0.480+
× Specialty Count					(0.066)	(0.202)	(0.248)
Payment Shock × Post-1997							
× State Fixed Effects							Yes
N	240,264	240,264	240,264	240,264	240,264	240,264	240,264
Number of Clusters	1,303	1,303	1,303	1,303	1,303	1,303	1,303
Weighted	Yes	Yes	Yes	Yes	Yes	Yes	Yes
State By Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
HCPCS By State FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Fully Interacted	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Eye Procedure Reductions	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Plan Type Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Panel Balanced	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Merged Sample Restrictions	Spec	Spec	Spec	Spec	Spec	Spec	Spec

Note: **, *, and + indicate statistical significance at the 0.01, 0.05, and 0.10 levels respectively. The table shows the results of reduced form specifications of the form described by equation (11) in section 5.2. Observations are constructed at the service-by-state-by-year level. The panel is balanced in the sense that each service-by-state panel is only included if public and private prices are available for each year from 1995 through 2002. Observations are weighted according to the number of times the service is observed in Medicare claims in 1997. The dependent variable in all columns is the level of the average private payment. Standard errors are calculated allowing for arbitrary correlation among the errors associated with each service. The "HHI" and "Count" variables have been converted to z-scores, and further details of the construction of all variables are described in the note to Table 1 and in the main text. Sources: Authors' calculations using Medicare claims and Thompson Reuters MarketScan data.

Table 6: **Heterogeneity in Surgical CF Shock's Effect by Insurer Concentration**

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable:	Private Payment Level					
Insurance HHI Measure:	Group Plan Mbrs., 2001	Tot. Mbrs., 2001	Total Med. Spend, 2001	Group Plan Mbrs., 2001	Group Plan Mbrs., 2001	Group Plan Mbrs., 2001
Payment Shock × Post-1997	1.296**	1.391**	1.375**	1.295**	1.317**	1.189**
	(0.214)	(0.260)	(0.254)	(0.215)	(0.218)	(0.144)
Payment Shock × Post-1997 × Insurer HHI	0.392*	0.390	0.169	0.440	0.395*	0.564**
	(0.169)	(0.287)	(0.107)	(0.293)	(0.164)	(0.186)
Payment Shock × Post-1997 × Insurer Count, 2001				0.070		
				(0.254)		
Payment Shock × Post-1997 × Physician HHI					-0.565**	-0.084
					(0.087)	(0.183)
Payment Shock × Post-1997 × Specialty HHI						-0.885*
						(0.404)
N	293,688	293,688	293,688	293,688	291,952	245,456
Number of Clusters	2,194	2,194	2,194	2,194	2,194	1,364
Weighted	Yes	Yes	Yes	Yes	Yes	Yes
State By Year FE	Yes	Yes	Yes	Yes	Yes	Yes
HCPCS By State FE	Yes	Yes	Yes	Yes	Yes	Yes
Fully Interacted	Yes	Yes	Yes	Yes	Yes	Yes
Eye Procedure Reductions	Yes	Yes	Yes	Yes	Yes	Yes
Plan Type Controls	Yes	Yes	Yes	Yes	Yes	Yes
Panel Balanced	Yes	Yes	Yes	Yes	Yes	Yes
Other Sample Restrictions	No CA	No CA	No CA	No CA	No CA	Spec Merge, No CA

Note: **, *, and + indicate statistical significance at the 0.01, 0.05, and 0.10 levels respectively. The table shows the results of reduced form specifications of the form described by equation (11) as modified in section 6.2. Observations are constructed at the service-by-state-by-year level. The panel is balanced in the sense that each service-by-state panel is only included if public and private prices are available for each year from 1995 through 2002. Observations are weighted according to the number of times the service is observed in Medicare claims in 1997. The dependent variable in all columns is the level of the average private payment. Standard errors are calculated allowing for arbitrary correlation among the errors associated with each service. The "HHI" and "Count" variables have been converted to z-scores, and further details of the construction of all variables are described in the note to Table 1 and in the main text. Sources: Authors' calculations using Medicare claims, Thompson Reuters MarketScan data, American Medical Association (2007), and data obtained from the National Association of Insurance Commissioners, by permission. The NAIC does not endorse any analysis or conclusions based upon the use of its data.

Table 7: **Heterogeneity in Surgical CF Shock's Effect by Prevalence of Small Physician Groups**

	(1)	(2)	(3)	(4)
Dependent Variable:	Private Payment Level			
Payment Shock	1.242**	1.216**	1.187**	1.217**
	(0.201)	(0.170)	(0.175)	(0.170)
Payment Shock	0.863**	0.479**		
× Share in Solo Practice	(0.220)	(0.129)		
Payment Shock			0.759**	0.337**
× Share in Small Groups			(0.230)	(0.116)
Payment Shock		-0.713**		-0.758**
× Specialty HHI		(0.272)		(0.278)
N	253,632	253,632	253,632	253,632
Number of Clusters	1,364	1,364	1,364	1,364
Weighted	Yes	Yes	Yes	Yes
State By Year FE	Yes	Yes	Yes	Yes
HCPCS By State FE	Yes	Yes	Yes	Yes
Fully Interacted	Yes	Yes	Yes	Yes
Eye Procedure Reductions	Yes	Yes	Yes	Yes
Plan Type Controls	Yes	Yes	Yes	Yes
Panel Balanced	Yes	Yes	Yes	Yes
Other Sample Restrictions	Spec Merge	Spec Merge	Spec Merge	Spec Merge

Note: **, *, and + indicate statistical significance at the 0.01, 0.05, and 0.10 levels respectively. The table shows the results of reduced form specifications of the form described by equation (11) in section 5.2. Observations are constructed at the service-by-state-by-year level. The panel is balanced in the sense that each service-by-state panel is only included if public and private prices are available for each year from 1995 through 2002. Observations are weighted according to the number of times the service is observed in Medicare claims in 1997. The dependent variable in all columns is the average private payment. Standard errors are calculated allowing for arbitrary correlation among the errors associated with each service. "Public-Private Ratio" and "Medstat Volume" are expressed as percentile ranks (across all services observed within a given market) minus 0.5; the variables thus have a mean of 0 and range from -0.5 to 0.5. Further details regarding the construction of all variables are described in the note to Table 1 and in the main text. Sources: Authors' calculations using Medicare claims and Thompson Reuters MarketScan data.

Appendix For Online Publication Only

A Background on Physician-Insurer Negotiations

In this appendix, we present practitioner descriptions of negotiations between physicians and insurers. The depictions come largely from physicians and consultants who represent physicians in these negotiations. Perhaps unsurprisingly, the latter sometimes seek to dispel small physician groups' concerns regarding their prospects for success in such negotiations. Two themes are emphasized regularly: the importance of the Medicare fee schedule and the importance of market power. Below we present samples of consultants' discussions of each.

A.1 The Role of Medicare's Fee Schedule

Practitioners' views of physician-insurer negotiations frequently emphasize the role of Medicare's fee schedule as a starting point from which negotiations take place. Some emphasize the relevance of "the fee schedule" in general, while placing varying degrees of relevance on Medicare itself. Examples follow:

- "All insurance companies will offer a fixed fee-for-service schedule. For some carriers, you may only be allowed to request a certain percentage above Medicare rates. Others may accept number values" (Nandedkar 2011).
- "The fee schedule will be the platform for negotiation" (Nandedkar 2011).
- "Anthem's fee schedule is based on the CMS Resource Based Relative Value Scale ('RBRVS'). The RBRVS is based on the resources a physician typically uses for each procedure and service, from physical, intellectual and emotional effort to overhead and training... Throughout this Manual, Anthem's method of reimbursement will be referred to as the current Anthem fee schedule, which is a combination of the modified RBRVS values, the services not evaluated by RBRVS and the Anthem conversion factor." (Anthem Blue Cross and Blue Shield 2012a)
- "Today, most health plans operate with fixed fee schedules. Often these schedules have little in common with the RBRVS, and while some are roughly based on a percentage of what Medicare pays, they may be tied to payment levels that are three or more years old. Most physicians who question this methodology for paying for professional services are told to take it or leave it." (Mertz 2004).
- "The fee schedule in many contracts is stated as a percentage of the Medicare rate. All individuals interviewed for this article recommended specifying a year to be used for the Medicare rate to protect against potential Medicare cuts" (Gesme and Wiseman 2010).

- "The ParPlan, BlueChoice® and HMO Blue® Texas (Independent Provider Network only) maximum allowable fees for practitioners will be updated to reflect 2010 CMS values effective July 1, 2010." (Blue Cross and Blue Shield of Texas 2010)

Negotiating consultants recommend that physicians be wary of negotiating over payments for specific codes rather than negotiating over average payments. This line of advice is directly linked to the fee schedule's complexity. Consultants express the concern that insurers' negotiating sophistication, in particular relative to that of small physician groups, will give insurers an advantage when trading off increases and decreases in payments for individual service codes:

- "Why do we focus on Revenue per Visit and not, say, the fee schedule of your most important codes? For one very simple reason: Focusing on the fees for specific procedure codes plays right into the shell game the insurance companies love to play" (Reckenen 2013).
- "One difficulty in negotiating a fee schedule is the sheer number and variety of codes that may be covered within a negotiation. Companies may make this more difficult by offering irregular payment schedules that don't correspond to standard fee schedules like Medicare or an RVU based system" (Fontes 2013).
- "A physician should beware of companies that state average reimbursements either in terms of RVU or a Medicare fee schedule. One may find that the fee for a frequently used CPT code is well below average and CPT codes rarely billed are several multiples higher to skew the average. An effective method to counter this tactic is for the practice to submit its top 30 CPT codes by volume and have the insurance company specifically define the fee schedule for these high-volume codes" (Fontes 2013).
- "Bob Phelan, chief executive officer of Integrated Community Oncology Network (Jacksonville, FL), a multispecialty cancer services network spanning four northeast Florida counties, explains why his network initially assesses the aggregated fees: 'The payers try to slide the money from one bucket to another. They'll increase E&M [evaluation and management] codes by 20%, but that's really only approximately 12% to 13% of business. At the same time, they decrease drug reimbursement by 2%, which offsets the E&M increase" (Gesme and Wiseman 2010).

Physicians who opt to negotiate over code-specific payments are encouraged to ensure that the codes over which they negotiate account for the bulk of their practice's revenues:

- "Be sure the codes on your list account for at least 75 percent of total practice charges.... Whatever method you choose, be sure to update your fee schedule annually based on changes to the Medicare fee schedule" (Mertz 2004).

While commenting on the evolution of provider networks, one consultant concludes with emphasis on one of the industry's few certainties:

- "It is not clear how or when these evolving provider structures and systems will be rewarded or remunerated. What is clear is that there will be complex negotiation occurring in the near future as result" (Fontes 2013).

A.2 The Importance of Market Power

Market power emerged as a common theme, both as a determinant of whether it makes sense to negotiate it all, and as a source of leverage over a negotiation's course:

- "Unless you dominate your market, payers are unlikely to grant sweeping fee increases. However, you may be able to negotiate increases for individual services if you can demonstrate inequities using your data analysis" (Mertz 2004).
- "Before negotiating a contract with any insurance company, first look at the state of your own company. Why should any carrier negotiate with you? What makes your practice unique relative to your competitors? What do you have that the carrier wants?" (Nandedkar 2011)
- "Negotiating strength comes from robust patient relationships..." (Nandedkar 2011)
- "If a health plans payment levels are extremely low, you may be tempted to bypass negotiations and simply no longer accept patients from that plan. Whether this is a sound strategy depends on your local market. For example, if you practice in a highly competitive market, those patients will easily find another physician and you will simply lose market share. However, in less competitive markets, patients may complain to their employers that the loss of your practice has created a hardship and they may pressure the insurance company to return to the bargaining table" (Mertz 2004).

Only the most optimistic of consultants actively encourage sole practitioners to pursue active negotiations:

- "Can a solo physician or small group practice really negotiate their payer contract language and increase reimbursement rates? The answer is YES!" (Glassman 2012).
- "I am told everyday that the large healthcare insurance companies (such as Blue Cross, Blue Shield, Aetna, United Healthcare, Health Net, Cigna and Independent Physicians Organizations (IPAs), do not negotiate with solo physicians and small group practices. Although the health plans would love for you to believe that, it simply is not true" (Glassman 2012).

B Further Details on the Model of Section 4

B.1 Generalizing the Physician Cost Function

According to the model presented in section 4, physicians choose their supply to the two sectors, private and Medicare, to maximize profits. Here, we relax the assumption of a fixed capacity and replace it with a general cost function $C(m, q)$ that depends on the number of patients treated in each sector. So profits are given by

$$\Pi(m, p) = r_M m + pq - C(m, q) \tag{B.1}$$

where q and m are quantities in the two markets respectively, and $C(m, q)$ is the total cost function. Note that we have defined the profit function in terms of the Medicare quantity but the private sector price. The cost function is defined in terms of both sectors' quantities.

The physician faces a demand curve $q = D(p)$ and sets the private price p taking this into account. We assume r_M is a parameter set exogenously by Medicare. To explicitly incorporate the demand curve that the physician faces, we rewrite profits as

$$\Pi(m, p) = r_M m + pD(p) - C(m, D(p)). \tag{B.2}$$

The physician's first-order conditions are:

$$r_M = C_m(m^*, D(p^*)) \tag{B.3}$$
$$0 = p^* D'(p^*) + D(p^*) - C_q(m^*, D(p^*))D'(p^*) \tag{B.4}$$

where C_m and C_q denotes the marginal costs of supplying Medicare and private patients, respectively, and asterisks denote equilibrium choices or outcomes. The physician chooses m^* and p^* (or, equivalently through the demand curve, q^*). Since the physician takes Medicare's payment r_M as given, equation (B.3) says that the marginal cost of treating a Medicare patient must equal the reimbursement rate. In contrast, the marginal cost of treating a privately insured patient must equal the marginal revenue from such a patient, or $p^* D'(p^*) + D(p^*)$, as equation (B.4) shows.

We now take comparative statics with respect to r_M, suppressing the arguments of the second derivatives:

$$1 = C_{mm} \frac{dm^*}{dr_M} + C_{mq} D'(p^*) \frac{dp^*}{dr_M} \tag{B.5}$$
$$\begin{aligned} 0 = {} & p^* D''(p^*) \frac{dp^*}{dr_M} + 2D'(p^*) \frac{dp^*}{dr_M} - C_q(m^*, D(p^*))D''(p^*) \frac{dp^*}{dr_M} \\ & - D'(p^*) C_{mq} \frac{dm^*}{dr_M} - D'(p^*)^2 C_{qq} \frac{dp^*}{dr_M} \end{aligned} \tag{B.6}$$

We solve these equations jointly for $\frac{dr_P^*}{dr_M}$:

$$\frac{dp^*}{dr_M} = \frac{C_{mq}}{-D'(p^*)[C_{qq}C_{mm} - C_{mq}^2] + \left\{[p^* - C_q(m^*, D(p^*))]\frac{D''(p^*)}{D'(p^*)} + 2\right\} C_{mm}} \tag{B.7}$$

Equation (B.7) gives the general expression for cost-following or cost-shifting. Since the denominator is generally positive, the overall sign depends on C_{mq} in the numerator. First consider the case where C_{mq} is positive; treating additional Medicare patients raises the marginal cost of treating private patients. This case is illustrated in Figure B.1, which shows why competitive market intuitions prevail. An increase in Medicare's reimbursement rate leads the physician to treat more Medicare patients,[62] which shifts the private sector marginal cost curve to the left (from C_q^1 to C_q^2). This moves the equilibrium higher up the marginal cost curve, so private prices increase.

Now consider the case of declining marginal costs. This could arise in the case of a natural monopoly with large fixed costs and constant or declining marginal costs. Or perhaps a practice that expands would be able to invest in an efficiency-enhancing technology. For either reason, the cross partial C_{mq} is negative. Figure B.2 depicts this case. Suppose that Medicare increases its reimbursement rate, again increasing the supply of care to Medicare beneficiaries. This movement along the Medicare supply curve shifts the private sector C_q curve to the left. This shift reduces the equilibrium private sector price; the increase in supply to Medicare beneficiaries lowers the marginal cost of care because, by assumption, $C_{mq} < 0$. Such cost curves may be relatively plausible in the hospital setting, where fixed costs are large and many markets are highly concentrated.

B.2 Utility Function Generating Variable Markups

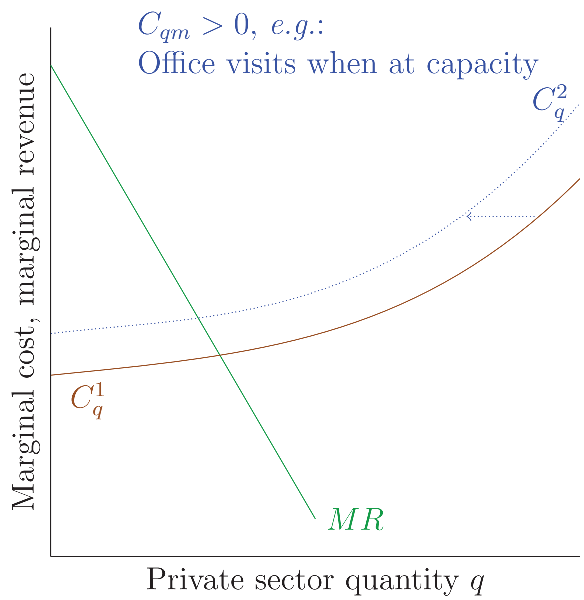
Section 4 shows an example demand function of $D(p) = (A - p)^\eta$, where $\eta > 1$, that can generate variable physician markups. This demand reflects the following utility function:

$$U(x, q) = x + Aq - \frac{\eta}{\eta + 1} q^{\frac{\eta+1}{\eta}} \tag{B.8}$$

where q represents the consumption of health care and x that of the numeraire composite commodity. Note, however, that the demand function in the text is that of the insurers who represent the patients in the local market. Depending on the structure of competition, underlying patient demand need not be directly reflected in the insurers' ultimate demand.

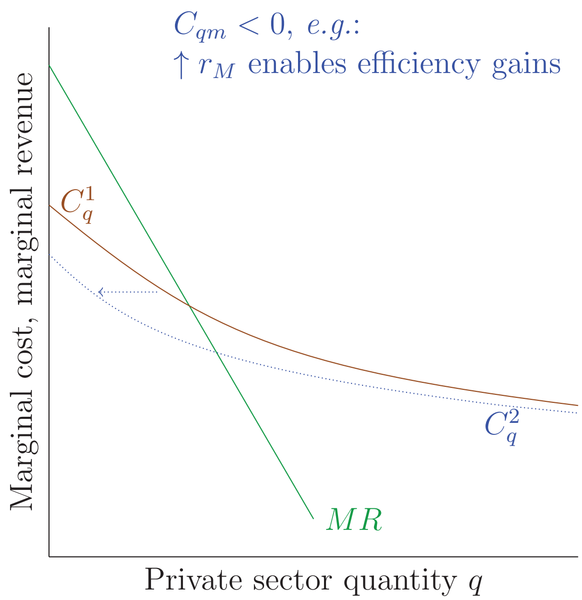
[62] Although own-price supply responses are positive in most settings, this is a subject of longstanding debate in markets for health care services. Papers finding evidence of positive own-price supply responses include Hadley and Reschovsky (2006) and Clemens and Gottlieb (2014). Evidence consistent with backward bending supply curves can be found in Rice (1983) and, more recently, Jacobson, Chang, Newhouse and Earle (2013). An assumption of backward bending labor supply is embedded in the federal budgeting process (Codespote, London and Shatto 1998), where it is described as a source of "volume offsets." The literature on physician behavior has extensively studied the target income hypothesis, which is shown by McGuire and Pauly (1991) to require large and unusually patterned income effects.

Appendix Figure B.1: Cost-Following in Response to a Medicare Price Increase



This figure shows the cost-following that we would expect to see in response to a Medicare price increase when the physician's cost curve slopes up. In this case, an increase in Medicare reimbursement rates, combined with an upward-sloping *Medicare* supply curve (not pictured), shifts the *private sector* marginal cost curve left (from C_q^1 to C_q^2, as depicted here). This leads to higher private sector reimbursement rates. Hence we find *cost-following* in this scenario.

Appendix Figure B.2: Cost-Shifting in Response to a Medicare Price Increase



This figure shows the cost-shifting that we would expect to see in response to a Medicare price increase when the physician's cost curve slopes up. In this case, an increase in Medicare reimbursement rates, combined with an upward-sloping *Medicare* supply curve (not pictured), shifts the *private sector* marginal cost curve left (from C_q^1 to C_q^2, as depicted here). This leads to lower private sector reimbursement rates. Hence we find *cost-shifting* in this scenario.

C Data Appendix

This Data Appendix serves two purposes. In section C.1 we describe our core datasets in further detail. In section C.2 we document the data loss that arises from our Medicare-private merge procedure in order to determine how comprehensive and representative our final dataset is.

C.1 Data Sources

Our Medicare claims data are provided by the Research Data Assistance Center (ResDAC) in Minneapolis, Minn., on behalf of the Centers for Medicare and Medicaid Services. These data come directly from the claims that physicians file with the Medicare carriers who process reimbursements on behalf of CMS in each state. Because Medicare is a centralized national program, these data are directly comparable across locations.

We measure payments using a variable in the carrier claims file, called the "Line Allowed Charge Amount." This is defined as "The amount of allowed charges for the line item service on the noninstitutional claim. This charge is used to compute pay to providers or reimbursement to beneficiaries." Note that "The amount includes beneficiary-paid amounts (i.e., deductible and coinsurance)." So this reflects the full amount that physicians are allowed to bill for the service, and hence what Medicare pays them.

The MarketScan data are somewhat more complex. This database compiles health care claims processed by insurers for "a selection of large employers, health plans, and government and public organizations." It incorporates around 100 payers and 500 million individual claims annually. According to the documentation, "These data represent the medical experience of insured employees and their dependents for active employees, early retirees, COBRA continues and Medicare-eligible retirees with employer-provided Medicare Supplemental plans." We use the outpatient services portion of the Commercial Claims and Encounters Database in MarketScan.

The benefit of MarketScan is its ability to pool data from numerous separate firms and other insurance providers. While these data represent care provided by numerous insurers and ultimate payers, Thompson Reuters standardizes the files and variables so as to be comparable across firms. We pool together claims from all firms in the MarketScan files. It is important to keep in mind that the number and composition of firms changes over time, and these firms use a range of different insurers. So when we analyze heterogeneity in private price responses we can only do so at the aggregate level, and cannot distinguish between individual insurers.

In both datasets we eliminate claims of less than $1 or with quantities of 100 or more. In MarketScan, we also eliminate claims associated with capitated payment arrangements, as recorded in the data. In general, the payments associated with such arrangements cannot be meaningfully linked to identifiable units of care.

C.2 Comprehensiveness of the Medicare-MarketScan Merge

Analyzing the relationships between private and public prices requires merging the Medicare and MarketScan databases. This merge is made possible by the fact that Medicare and private insurers both make payments using the Healthcare Common Procedure Coding System (HCPCS). The HCPCS is, in turn, linked to the American Medical Association's Current Procedure Terminology (CPT). Importantly, CPT is designed to characterize the universe of services provided by physicians; it is not catered specifically to care for the elderly, young, or working-age adult population.

We merge the Medicare and MarketScan databases on the HCPCS codes. Inclusion in our analysis sample requires satisfying two criteria. First, the codes must be observed in both the Medicare and MarketScan databases. Second, we require that our panel be balanced in the following sense: a state-by-service pair is only included if it appears in each year from 1995 through 2002.

Our estimation sample includes 2,194 of the 12,729 unique HCPCS codes observed in MarketScan during our sample period. While our sample accounts for a minority of codes, these codes account for a majority of the total care provided. This reflects the fact that the more commonly used codes are more likely to satisfy our criteria. Lost codes include codes that are never or rarely provided to the elderly, the non-elderly, or both. They also include codes that were introduced or eliminated over the course of our sample. Appendix Table C.1 presents details on the data loss associated with each step of the merge process.

The table focuses on data from 1995, namely the first year of our sample, and progressively eliminates codes in four steps. Row C presents the total number of distinct codes appearing in each data set in 1995. 6,037 distinct HCPCS codes were submitted within our Medicare claims data, while 8,781 were submitted within MarketScan. Since row C presents a full accounting of the codes in each database in 1995, the codes are associated with a full 100 percent of each dataset's care in terms of both dollars spent and unique services counted. Row D eliminates codes that do not exist in the official Medicare RVU files. This eliminates zero Medicare claims. In column 6, we see that this first exclusion eliminates 30.1 percent of the unique MarketScan codes. Columns 4 and 5 reveal that these codes represent a relatively small portion of overall private spending—column 5 shows that they account for only 3.1 percent of services and column 4 shows that they represent only 8.6 percent of spending. It is reassuring that matchable codes, where public and private payment rates could conceivably be compared, represent an overwhelming majority of the care provided.

The next three rows show why additional data are eliminated as we progress to the final estimation sample. Row E imposes the criterion that a code be used at some point in time in the complementary dataset. The percentages for Medicare thus show the share of codes that appear in the MarketScan data, and vice versa. We again see virtually no exclusions from the Medicare data. We also see minimal further data loss from MarketScan. Among codes actively in use in either database, more than 98 percent make at least one appearance in the complementary database.

Row F imposes the requirement that the Medicare claims data be balanced. That is, it drops all remaining codes that do not appear in the Medicare data in each year from

1995 to 2002. This panel-balance requirement results in significant loss of codes in both the Medicare and MarketScan data. Just under 50 percent of the remaining service codes are lost in both data sets. These codes account for 6.5 percent of Medicare spending and 21 percent of remaining private sector spending (that is, 19 of the remaining 91 percent). These include irregularly use codes as well as codes that were either eliminated or introduced over the course of the sample.

Finally, row G imposes that the panel be balanced in both the Medicare and MarketScan databases. This reduces the MarketScan files to 63.5 percent of their initial spending and the Medicare files to 76.8 percent of their initial spending. The data loss associated with the last two steps reflect our effort to construct a balanced panel. The codes that we lose through these steps are those that are not continually relevant to medical providers during our sample period. Importantly, they do not reflect services for which either Medicare or private insurers selectively opted out of determining prices altogether.

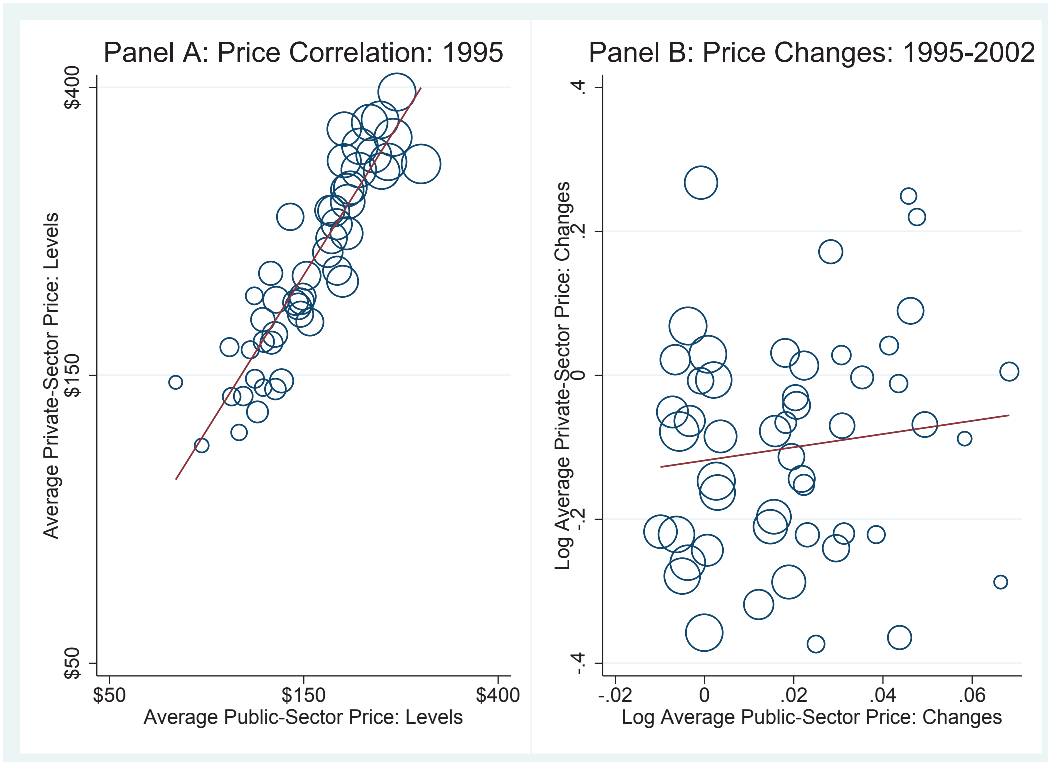
Appendix Table C.1: **Measures of Merge Comprehensiveness**

	(1)	(2)	(3)	(4)	(5)	(6)
(A) Initial sample:	Starting merge with Medicare data			Starting merge with MarketScan data		
(B) Size measure:	Spending	Quantity	Codes	Spending	Quantity	Codes
(C) Initial dataset	100%	100%	6,037	100%	100%	8,781
(D) Code exists	100%	100%	100%	91.4%	96.9%	69.9%
(E) Used in complementary dataset	100%	100%	99.9%	90.8%	96.0%	68.3%
(F) Medicare balanced panel	93.5%	87.3%	54.6%	71.7%	77.4%	36.7%
(G) Panel balanced	76.8%	63.4%	35.3%	63.5%	59.9%	25.0%

Note: This table presents the comprehensiveness of our merge procedure between the Medicare claims data and private sector data from MarketScan introduced in section 1.4 using data from 1995. Line (A) indicates which file we start with; columns 1 and 2 show the merge overlap when starting with Medicare and then merging in the MarketScan private data while columns 3 and 4 conduct the merge in the opposite direction. Line (B) distinguishes between measurements of the overlap in dollar terms (labeled "Spending") and in service count (labeled "Quantity"). The subsequent lines show the share of the initial dataset that survives each step of the merge procedure. Line (C) starts with the full dataset. Line (D) shows the share remaining after matching codes with the opposite dataset from the one listed in line (A). Line (E) shows the final share remaining after balancing the panel. Sources: Authors' calculations using Medicare claims and Thompson Reuters MarketScan data.

D Further Results and Robustness Tests

Appendix Figure D.1: Cross-State Relationship Between Private and Medicare Prices



Note: This figure shows the raw cross-state relationships between average private reimbursements and average Medicare reimbursements. The payments are the natural logs of the average payment we observe in our public (Medicare) and private (Medstat) sector claims data. Panel A presents these average payments for 1995 while Panel B shows the changes in these average payments from 1995 to 2002. Circle sizes are proportional to Medicare spending in each state. The best-fit line shown in Panel A results from estimating

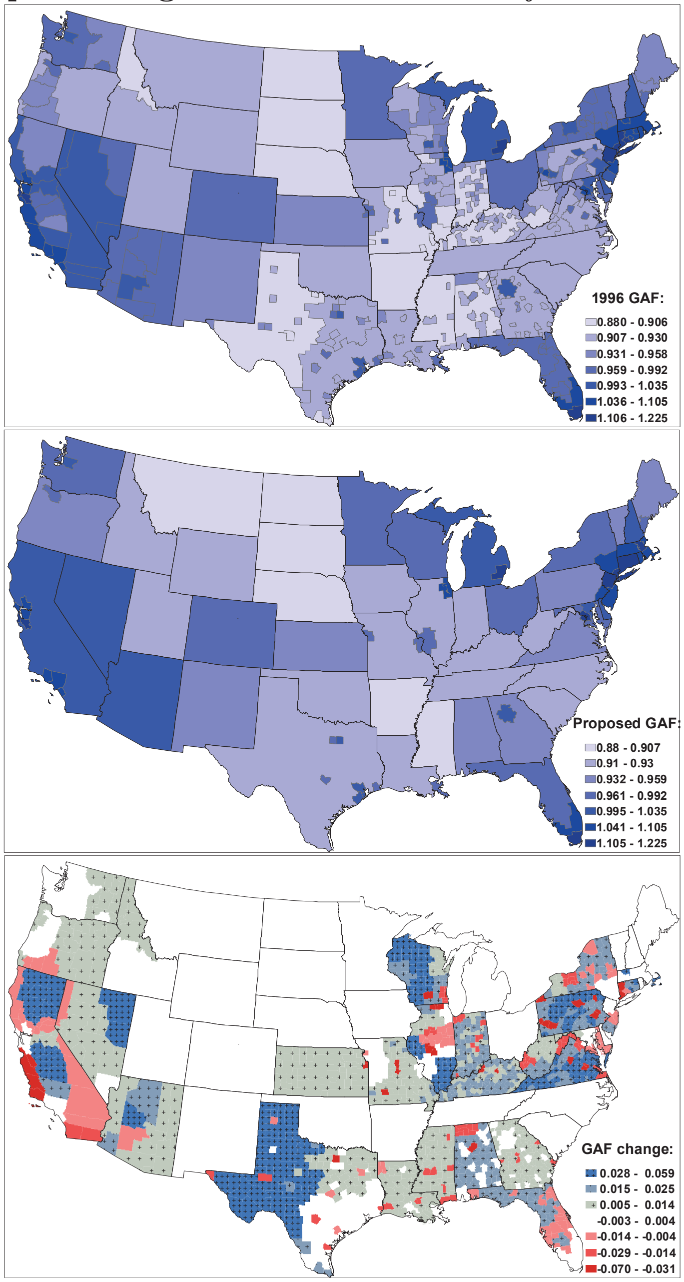
$$\ln(P_s^{\text{Private}}) = \beta_0 + \beta_1 \ln(P_s^{\text{Medicare}}) + u_s$$

across states s, weighted by each state's Medicare spending. The regression yields a coefficient of $\beta_1 = 1.07$ and $R^2 = 0.81$, with $N = 50$. The best-fit line shown in Panel B results from estimating

$$\triangle \ln(P_s^{\text{Private}}) = \gamma_0 + \gamma_1 \triangle \ln(P_s^{\text{Medicare}}) + v_s,$$

again weighted state spending. The regression yields a coefficient of $\gamma_1 = 1.00$ (statistically indistinguishable from zero) and $R^2 = 0.02$ with $N = 50$. Note that the regressions are run in logs and the values shown along the axes are computed by exponentiating the log values.

Appendix Figure D.2: Medicare Payment Areas



The first panel shows the 206 Medicare fee schedule areas in the continental United States as of 1996 and the second shows the 85 such localities after the consolidation in 1997. (These totals exclude Alaska, Hawaii, Puerto Rico, and the U.S. Virgin Islands, each of which was its own unique locality throughout this period.) The colors indicate the Geographic Adjustment Factors (GAF) associated with each Payment Locality, with darker colors indicating higher reimbursement rates. The third panel shows the change in GAF for each county due to the payment region consolidation that took place in 1997. Source: Clemens and Gottlieb (2014), based on data from the *Federal Register*, various issues.

D.1 Robustness Checks

Appendix Tables D.1 through D.7 report various robustness checks as discussed in the main text, including cost-following measurements using nationally aggregated data.

D.2 Service Market Size and Price Transmission

We now consider a dimension of heterogeneity in Medicare's effect that most models predict. This is Medicare's market share, or its size relative to the private market. In the model of section 4, a setting where Medicare is larger likely makes it easier for physicians to find private patients to treat, relative to Medicare patients. The model would account for this through a higher relative cost parameter α, and hence an increase in cost-following ($e.g$ in equation [10]).

In a more general cost structure, which we discuss in Appendix B.1, we also see that growth in Medicare's share of the patient pool increases the magnitude of cost-following. In this case, the force driving the relationship is Medicare's larger role in total costs. As the public sector grows, a pricing change that shifts physicians along their Medicare supply curve will have a larger effect on the residual supply for privately insured patients, and hence on the price in that sector.

Finally, the institutional story that we discussed in section 7 also predicts that Medicare's market share should affect its pricing influence. As the private market's size grows, the gains available by negotiating away from Medicare's default price increase.

We test this rather universal prediction by using two proxies for the relative sizes of the private and public markets. First, we look at the relative sizes for each service and, second, at the number of Medicare beneficiaries as a fraction of each state's population. As in previous sections, we examine the relationship between these variables and the extent of Medicare's influence on private prices.

Our measure of the relative sizes of the public and private markets for each service is based on the service counts we observe in our databases of public and private sector claims. Specifically, we construct the variable "Medicare Relative Size" as the ratio of the number of times a service appears in a single year of the Medicare claims data and the number of times it appears in a single year of the MedStat data. Because MedStat is a non-random sample of the private market, with time-varying size, the variable would poorly characterize the actual relative sizes of public and private markets. Nonetheless, it should form a reasonable basis for dividing services into those with relatively large and small Medicare market shares. This variable is strongly right skewed; the lower bound of the relevant z-scores is roughly -0.2 for Private Market Volume and -0.4 for Medicare Relative Size. Consequently, we normalize it using percentile ranks rather than z-scores. We subtract 0.5 from the percentile ranks so that the resulting variables are symmetric about 0. Our measure of Medicare's share of the patient pool is constructed directly from state-level beneficiary and population counts. Since this variable is relatively symmetric, we normalize it using z-scores. We then interact these variables with the price shocks and controls as in equation (11).

Appendix Table D.8 presents these results. Column 1 shows that the public-private ratio enters significantly, with a coefficient of 1.3. Moving from the first to the 99th percentile of the Medicare Relative Size distribution is associated with moving from a price transmission coefficient of 0.3 to a cost-following coefficient of 1.5. The larger the relative size of the Medicare market, the larger the cost-following coefficient. Column 2 shows the relationship between cost-following and Medicare's share of the potential patient pool. The estimate in this case is statistically indistinguishable from zero.

D.3 Does Payment Reform Affect Private-Sector Price Dispersion?

In this section we briefly explore additional pricing consequences of Medicare payment policy. One outcome potentially of interest is price dispersion. Dispersion in private payments for ostensibly similar services is substantial, and its determinant are not fully understood. We estimate the extent to which price dispersion responded to our natural experiment, which involved a substantial reduction in payments for surgical procedures relative to other services. It may also have resolved a degree of uncertainty surrounding the future of Medicare's payments, at least temporarily.

Appendix Table D.9 reports the results, which involve specifications taking the same form as those reported in columns 2 and 5 of Table 2. The dependent variables measure price dispersion at the service-by-state-by-year level. In columns 1 and 2, the dependent variable is the standard deviation of prices within these markets, while in columns 3 and 4 it is the coefficient of variation. The results imply that increases in payments are associated with increases in dispersion. In column 3, the coefficient of variation is uncorrelated with the magnitude of the payment shocks, while column 4 shows that the overall level of price dispersion did increase for surgical procedures.

Appendix Table D.1: Robustness Checks on the Effect of Medicare Price Changes on Private Sector Prices

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Private Payment Level						
Public Payment	1.162**	1.157**	1.161**	0.972**	0.673**	1.324**	1.116**
	(0.225)	(0.212)	(0.226)	(0.107)	(0.112)	(0.261)	(0.250)
Plan Type Controls	-11.514	-0.657	-7.540	5.415	-13.455	-8.536	
	(28.184)	(14.557)	(28.435)	(5.849)	(24.134)	(16.669)	
Cost Sharing Fraction	-2.310	-2.252		-2.403	1.767**	-3.093	-2.389
	(6.144)	(6.214)		(6.024)	(0.190)	(5.963)	(6.126)
N	303,728	303,728	303,728	303,728	303,728	303,728	303,728
Number of Clusters	2,194	2,194	2,194	2,194	2,194	2,194	2,194
Weighted	Yes	Yes	Yes	Yes	No	Yes	Yes
State By Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
HCPCS By State FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Eye Procedure Reductions	Yes	Yes	Yes	No	Yes	Yes	Yes
RVUs Per Service Control	No	No	No	No	No	Yes	No
Trend by Procedure	No	No	No	No	No	No	Yes
Panel Balanced	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Other Sample Restrictions	None	None	None	None	None	None	None

Note: **, *, and + indicate statistical significance at the 0.01, 0.05, and 0.10 levels respectively. The table shows the results of IV specifications based on those in column 3 of Table 2. Observations are constructed at the service-by-state-year level. Unless noted, observations are weighted according to the number of times the service is observed in Medicare Part B in 1997. The panel is balanced in the sense that each service-by-state pairing is only included if public and private prices could be estimated for each year from 1995 through 2002. Standard errors are calculated allowing for arbitrary correlation among the errors associated with each service. Additional features of each specification are described within the table. The construction of all variables is further described in the note to Table 1 and in the main text. Sources: Authors' calculations using Medicare claims and Thompson Reuters MarketScan data.

Appendix Table D.2: **The Effect of Medicare Price Changes on Private Sector Prices, National Regressions**

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable:	Payment Levels			Log Payments		
	Public	Private	Private	Public	Private	Private
	1st Stage	Red. Form	IV	1st Stage	Red. Form	IV
Payment Shock × Post 1997	1.122**	1.305**				
	(0.067)	(0.225)				
Public Payment			1.163**			
			(0.210)			
Surgical Procedure × Post 1997				-0.228**	-0.116**	
				(0.031)	(0.027)	
Ln(Public Payment)						0.508**
						(0.061)
Plan Type Control	9.707	19.600	8.312	0.113+	-0.000	-0.057
	(13.707)	(17.664)	(19.059)	(0.066)	(0.036)	(0.045)
Cost Sharing Fraction	-17.324+	-102.019**	-81.874*	-0.140	-0.276*	-0.205*
	(10.323)	(36.129)	(39.342)	(0.089)	(0.133)	(0.102)
N	17,552	17,552	17,552	17,536	17,536	17,536
Number of Clusters	2,194	2,194	2,194	2,192	2,192	2,192
Weighted	Yes	Yes	Yes	Yes	Yes	Yes
HCPCS FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Eye Procedure Reductions	Yes	Yes	Yes	Yes	Yes	Yes
RVUs Per Service Control	No	No	No	No	No	No

Note: **, *, and + indicate statistical significance at the 0.01, 0.05, and 0.10 levels respectively. The table shows the results of specifications analogous to those in Table 2, except that here observations are constructed at the service-by-year level. Unless noted, observations are weighted according to the number of times the service is observed in Medicare Part B in 1997. The panel is balanced in the sense that each service is only included if public and private prices could be estimated for each year from 1994 through 2002. All specifications include service code and year fixed effects. Standard errors are calculated allowing for arbitrary correlation among the errors associated with each service. Additional features of each specification are described within the table. The construction of all variables is further described in the note to Table 1 and in the main text. Sources: Authors' calculations using Medicare claims and Thompson Reuters MarketScan data.

Appendix Table D.3: **Robustness Checks on the Effect of Medicare Price Changes on Private Sector Prices Nationally**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent Variable:	Private Payment Level						
Public Payment	1.163**	1.167**	1.148**	0.965**	0.181	1.370**	1.111**
	(0.210)	(0.207)	(0.216)	(0.096)	(0.412)	(0.261)	(0.243)
Plan Type Control	8.312		12.109	-25.287	31.678*	24.241	7.645
	(19.059)		(17.931)	(38.700)	(13.197)	(14.779)	(18.693)
Cost Sharing Fraction	-81.874*	-82.765*		-86.306*	18.937	-99.272*	-83.596*
	(39.342)	(39.400)		(38.231)	(12.135)	(41.228)	(39.028)
N	17,552	17,552	17,552	17,552	17,552	17,552	17,552
Number of Clusters	2,194	2,194	2,194	2,194	2,194	2,194	2,194
Weighted	Yes	Yes	Yes	Yes	No	Yes	Yes
HCPCS FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Eye Procedure Reductions	Yes	Yes	Yes	No	Yes	Yes	Yes
RVUs Per Service Control	No	No	No	No	No	Yes	No
Trend by Procedure	No	No	No	No	No	No	Yes

Note: **, *, and + indicate statistical significance at the 0.01, 0.05, and 0.10 levels respectively. The table shows the results of IV specifications based on those in column 3 of Table D.2. Observations are constructed at the service-by-year level. Unless noted, observations are weighted according to the number of times the service is observed in Medicare Part B in 1997. The panel is balanced in the sense that each service is only included if public and private prices could be estimated for each year from 1994 through 2002. All specifications include service code and year fixed effects. Standard errors are calculated allowing for arbitrary correlation among the errors associated with each service. Additional features of each specification are described within the table. The construction of all variables is further described in the note to Table 1 and in the main text. Sources: Authors' calculations using Medicare claims and Thompson Reuters MarketScan data.

Appendix Table D.4: Heterogeneity in Surgical CF Shock's Effect by Provider Concentration, Alternative Clustering

	(1)	(2)	(3)	(4)
Dependent Variable:	Private Payment Level			
Payment Shock × Post-1997	1.345**	1.345**	1.345**	1.345**
	(0.312)	(0.234)	(0.317)	(0.082)
Payment Shock × Post-1997	-0.827*	-0.827**	-0.827*	-0.827**
× Specialty HHI	(0.336)	(0.305)	(0.347)	(0.113)
Payment Shock × Post-1997	0.408+	0.408**	0.408**	0.408**
× Specialty Count	(0.219)	(0.068)	(0.079)	(0.050)
N	240,264	240,264	240,264	240,264
Number of Clusters	49	75	20	4
Weighted	Yes	Yes	Yes	Yes
State By Year FE	Yes	Yes	Yes	Yes
HCPCS By State FE	Yes	Yes	Yes	Yes
Fully Interacted	Yes	Yes	Yes	Yes
Eye Procedure Reductions	Yes	Yes	Yes	Yes
Plan Type Controls	Yes	Yes	Yes	Yes
Panel Balanced	Yes	Yes	Yes	Yes
Merged Sample Restrictions	Spec	Spec	Spec	Spec
Cluster Level	State	Betos	Betos	Betos
		(3-digit)	(2-digit)	(1-digit)

Note: **, *, and + indicate statistical significance at the 0.01, 0.05, and 0.10 levels respectively. The table shows the results of reduced form specifications based on column 4 of Table 5. Observations are constructed at the service-by-state-year level. The panel is balanced in the sense that each service-by-state pairing is only included if public and private prices could be estimated for each year from 1995 through 2002. Observations are weighted according to the number of times the service is observed in Medicare Part B in 1997. The dependent variable in all columns is the level of the average private payment. Standard errors are calculated allowing for arbitrary correlation within the clusters indicated in each column. The construction of all variables is further described in the note to Table 1 and in the main text. Sources: Authors' calculations using Medicare claims and and Thompson Reuters MarketScan data.

Appendix Table D.5: Robustness Checks on Heterogeneity in Surgical CF Shock's Effect by Provider Concentration

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable:			Payment Levels			
Payment Shock × Post-1997	1.418**	1.323**	1.505**	1.440**	1.443**	1.406**
	(0.226)	(0.201)	(0.314)	(0.247)	(0.254)	(0.243)
Payment Shock × Post-1997	-0.485**	-0.441**	-0.546**	-0.518**	-0.553**	-0.486**
× Physician HHI	(0.056)	(0.046)	(0.073)	(0.066)	(0.074)	(0.063)
Payment Shock × Post-1997	-0.589*	-0.675*	-0.418*	-0.451*	-0.434*	-0.540*
× Specialty HHI	(0.276)	(0.310)	(0.191)	(0.220)	(0.192)	(0.229)
Payment Shock × Post-1997	-0.316	-0.216	-0.942*	-0.248	-0.235	-0.210
× Physician Count	(0.253)	(0.193)	(0.411)	(0.159)	(0.160)	(0.159)
Payment Shock × Post-1997	0.904*	0.321	0.814**	0.638**	0.632**	0.659**
× Specialty Count	(0.425)	(0.268)	(0.268)	(0.202)	(0.203)	(0.205)
N	240,264	240,264	240,264	240,264	240,264	240,264
Number of Clusters	1,303	1,303	1,303	1,303	1,303	1,303
Census Region	Yes	Yes	No	Yes	No	No
Census Division	No	Yes	No	No	No	No
Log Population	No	No	Yes	No	No	No
Log Density	No	No	No	Yes	No	No
Log Income Per Capita	No	No	No	No	Yes	No
Education (HS and BA Compl.)	No	No	No	No	No	Yes
Sample Restrictions	Phys Merge	Phys Merge	Phys Merge	Phys Merge	Phys Merge	Phys Merge

Note: **, *, and + indicate statistical significance at the 0.01, 0.05, and 0.10 levels respectively. The table shows the results of reduced form specifications based on column 6 of Table 5. Observations are constructed at the service-by-state-year level. The panel is balanced in the sense that each service-by-state pairing is only included if public and private prices could be estimated for each year from 1995 through 2002. Observations are weighted according to the number of times the service is observed in Medicare Part B in 1997. The dependent variable in all columns is the level of the average private payment. Standard errors are calculated allowing for arbitrary correlation among the errors associated with each service. In columns 1 and 2, "Payment Shock × Post-1997" is interacted with a full set of region or division fixed effects, and the coefficient shown is the average of those interactions. The construction of all variables is further described in the note to Table 1 and in the main text. Sources: Authors' calculations using Medicare claims and Thompson Reuters MarketScan data, and Ruggles et al. (2010).

Appendix Table D.6: **Robustness Checks on Heterogeneity in Surgical CF Shock's Effect by Insurer Concentration**

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable:			Payment Levels			
Payment Shock × Post-1997	1.338**	1.280**	1.424**	1.326**	1.305**	1.338**
	(0.205)	(0.211)	(0.282)	(0.227)	(0.216)	(0.235)
Payment Shock × Post-1997	0.309+	0.487*	0.711**	0.474**	0.501**	0.368*
× Insurer HHI	(0.163)	(0.204)	(0.228)	(0.169)	(0.191)	(0.176)
N	293,688	293,688	293,688	293,688	293,688	293,688
Number of Clusters	2,194	2,194	2,194	2,194	2,194	2,194
Census Region	Yes	Yes	No	Yes	No	No
Census Division	No	Yes	No	No	No	No
Log Population	No	No	Yes	No	No	No
Log Density	No	No	No	Yes	No	No
Log Income Per Capita	No	No	No	No	Yes	No
Education (HS and BA Compl.)	No	No	No	No	No	Yes
Other Sample Restrictions	No CA	No CA	No CA	No CA	No CA	No CA

Note: **, *, and + indicate statistical significance at the 0.01, 0.05, and 0.10 levels respectively. The table shows the results of reduced form specifications based on column 6 of Table 5. Observations are constructed at the service-by-state-year level. The panel is balanced in the sense that each service-by-state pairing is only included if public and private prices could be estimated for each year from 1995 through 2002. Observations are weighted according to the number of times the service is observed in Medicare Part B in 1997. The dependent variable in all columns is the level of the average private payment. Standard errors are calculated allowing for arbitrary correlation among the errors associated with each service. In columns 1 and 2, "Payment Shock × Post-1997" is interacted with a full set of region or division fixed effects, and the coefficient shown is the average of those interactions. The construction of all variables is further described in the note to Table 1 and in the main text. Sources: Authors' calculations using Medicare claims, Thompson Reuters MarketScan data, Ruggles et al. (2010), and data obtained from the National Association of Insurance Commissioners, by permission. The NAIC does not endorse any analysis or conclusions based upon the use of its data.

Appendix Table D.7: Insurer Concentration Measured in Various Years

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable:	Private Payment Level					
Insurance HHI Measure:	NAIC 1996	NAIC 1997	NAIC 1998	NAIC 2002	NAIC 2006	AMA 2006
Payment Shock × Post-1997	1.246**	1.289**	1.309**	1.257**	1.378**	1.429**
	(0.227)	(0.206)	(0.232)	(0.185)	(0.267)	(0.252)
Payment Shock × Post-1997	0.102	0.411+	0.302**	0.521*	0.039	-0.051
× Insurer HHI	(0.205)	(0.213)	(0.082)	(0.240)	(0.062)	(0.166)
N	286,568	293,688	288,072	293,688	293,688	275,728
Number of Clusters	2,194	2,194	2,194	2,194	2,194	2,193
Other Sample Restrictions	No CA	No CA	No CA	No CA	No CA	AMA Data

Note: **, *, and + indicate statistical significance at the 0.01, 0.05, and 0.10 levels respectively. The table shows the results of reduced form specifications based on column 6 of Table 5. Observations are constructed at the service-by-state-year level. The panel is balanced in the sense that each service-by-state pairing is only included if public and private prices could be estimated for each year from 1995 through 2002. Observations are weighted according to the number of times the service is observed in Medicare Part B in 1997. The dependent variable in all columns is the level of the average private payment. Standard errors are calculated allowing for arbitrary correlation among the errors associated with each service. In columns 1 and 2, "Payment Shock × Post-1997" is interacted with a full set of region or division fixed effects, and the coefficient shown is the average of those interactions. The construction of all variables is further described in the note to Table 1 and in the main text. Sources: Authors' calculations using Medicare claims, Thompson Reuters MarketScan data, and data obtained from the National Association of Insurance Commissioners, by permission. The NAIC does not endorse any analysis or conclusions based upon the use of its data.

Appendix Table D.8: **Heterogeneity in Surgical CF Shock's Effect by Medicare Market Share**

	(1)	(2)
Dependent Variable:	Private Payment Level	
Payment Shock	0.991**	1.414**
	(0.135)	(0.276)
Payment Shock	1.331**	
× Public-Private Ratio	(0.483)	
Payment Shock		-0.178
× Medicare Population Share		(0.180)
N	303728	303728
Number of Clusters	2,194	2,194
Weighted	Yes	Yes
State By Year FE	Yes	Yes
HCPCS By State FE	Yes	Yes
Fully Interacted	Yes	Yes
Eye Procedure Reductions	Yes	Yes
Plan Type Controls	Yes	Yes
Panel Balanced	Yes	Yes
Other Sample Restrictions	None	None

Note: **, *, and + indicate statistical significance at the 0.01, 0.05, and 0.10 levels respectively. The table shows the results of reduced form specifications of the form described by equation (11) in section 5.2. Observations are constructed at the service-by-state-year level. The panel is balanced in the sense that each service-by-state pairing is only included if public and private prices could be estimated for each year from 1995 through 2002. Observations are weighted according to the number of times the service is observed in Medicare Part B in 1997. The dependent variable in all columns is the average private payment. Standard errors are calculated allowing for arbitrary correlation among the errors associated with each service. "Public-Private Ratio" is expressed in percentile ranks (across all services observed within a given market) minus 0.5; it thus has a mean of 0 and range from -0.5 to 0.5. "Medicare Population Share" is expressed in z-scores. Further details regarding the construction of all variables are described in the note to Table 1 and in the main text. Sources: Authors' calculations using Medicare claims and Thompson Reuters MarketScan data.

Appendix Table D.9: **Baseline Estimates of the Effect of Medicare Price Changes on Price Variation**

	(1)	(2)	(3)	(4)
Dependent Variable:	Private Payment SD		Private Payment CV	
Payment Shock × Post-1997	0.681*		-0.0002	
	(0.292)		(0.0004)	
Surgical Procedure × Post-1997		2.763		0.123*
		(8.569)		(0.049)
Plan Type Control	29.550	37.817	0.046	0.056
	(24.710)	(30.637)	(0.074)	(0.077)
Cost Sharing Fraction	24.051	24.380	0.184+	0.188+
	(30.134)	(30.110)	(0.103)	(0.103)
N	293,714	293,714	293,714	293,714
Number of Clusters	2,194	2,194	2,194	2,194
Weighted	Yes	Yes	Yes	Yes
State By Year FE	Yes	Yes	Yes	Yes
HCPCS By State FE	Yes	Yes	Yes	Yes
Eye Procedure Reductions	Yes	Yes	Yes	Yes
Play Type Controls	Yes	Yes	Yes	Yes
RVUs Per Service Control	No	No	No	No
Panel Balanced	Yes	Yes	Yes	Yes
Other Sample Restrictions	Num. Obs.	Num. Obs.	Num. Obs.	Num. Obs.

Note: **, *, and + indicate statistical significance at the 0.01, 0.05, and 0.10 levels respectively. The table shows the results of reduced form specifications of the forms described in section 1.2, but with measures of price dispersion, rather than average prices, as the dependent variables. Columns 1 and 3 report estimates that take the same form as that reported in column 2 of Table 2, while columns 2 and 4 report estimates that take the same form as that reported in column 5 of Table 2. In columns 1 and 2 the dependent variables is the standard deviation of payments, as calculated at the service-by-state-year level. In columns 3 and 4 the dependent variables is the coefficient of variation of payments, again calculated at the service-by-state-year level. Observations are weighted according to the number of times the service is observed in Medicare Part B in 1997. The panel is balanced in the sense that each service-by-state pairing is only included if public and private prices could be estimated for each year from 1995 through 2002. Standard errors are calculated allowing for arbitrary correlation among the errors associated with each service. Additional features of each specification are described within the table. The construction of all variables is further described in the note to Table 1 and in the main text. Sources: Authors' calculations using Medicare claims and Thompson Reuters MarketScan data.

E Estimates of the Overall Reallocation Resulting from Our Payment Shocks

Medicare spending on surgical care amounted to $21.9 billion (inflation-adjusted to 2013 dollars) after the surgical payment cut, and that on medical care $47.3 billion (also in 2013 dollars) after the payment increase. [63] Private insurers spent $175 billion on physician and clinical services in 1998 (Centers for Medicare and Medicaid Services (CMS) 2013). Assuming the same medical-surgical split as in Medicare, and using our baseline cost-following estimate of 1.16, yields a reallocation of $7.6 billion in private insurance payments due to the medical-surgical pricing change.

This estimate of the private sector spillover is subject to two caveats. The first is a standard concern of our estimates' external validity. As we observed in section 1.4, the MarketScan data represent the universe of claims associated with a selected set of plans provided by large employers. Small employers' plans or individual market insurance may be more or less likely to pay physicians according to fee schedules influenced by Medicare's relative payments.

Second, Medicare may exert more influence over fee-for-service payments than over capitated payments. On this point it is important to keep in mind that, although managed care was pervasive during the period we study, it typically did not translate into capitated payment of physician groups. The Community Tracking Study (CTS) reveals that throughout the period we study, roughly 40 percent of the revenue of physicians' practices was linked to managed care contracts. A much smaller share of revenue was prepaid or capitated. In 1996 this share was 16 percent (CSHSC 1999, 56), while in 2004 it was 13 percent (CSHSC 2006, 4-29). Throughout our sample, capitated payments thus represented less than one-third of the revenues associated with managed care. Additionally, the share of revenue reported in the CTS as being capitated includes payments under both Medicaid and Medicare managed care arrangements, which are sizable. The capitated share of physician revenues from private sector payers is thus not particularly large. In the context of Medicare's payments for surgical procedures relative to other services, conservative adjustments for these factors would leave our estimate of the ratio of the private sector spillover to Medicare's direct effect on the order of 3, rather than the 3.9 that we present in section 3.2.

We must also consider external validity in the context of our estimates of the effects of changes in Medicare's system of geographic adjustments. Here too, Medicare's influence on payments in small-group and individual market plans may be either greater or smaller than its influence on the payments of plans provided by large employers. But the concern about capitation that we raised in the previous paragraph is less relevant here. Recall that these geographic payment changes altered payments across the board rather than differentially across services. In fee-for-service settings, benchmarking to Medicare's relative payment

[63] The data come from Medicare's Physician/Supplier Procedure Summary Master File, which is available from the Centers for Medicare and Medicaid Services at `http://www.cms.gov/Research-Statistics-Data-and-Systems/Files-for-Order/NonIdentifiableDataFiles/PhysicianSupplierProcedureSummaryMasterFile.html` (accessed August 11, 2014).

menu creates a strong, mechanical link between public and private payments for one service relative to another, which would not be relevant in capitated contracts. But since this mechanism is less applicable when considering across-the-board payment changes, our results in this context likely reflect Medicare's influence on the physician's opportunity cost, which would be relevant for both capitated and fee-for-service payments. So there is less need to exclude capitated payments when calculating the spillovers from broad-based geographic payment changes.