Bargaining in the Shadow of a Giant: Medicare's Influence on Private Payment Systems

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Abstract

We analyze Medicare's influence on private payments for physicians' services. Using a large administrative change in surgical relative to medical reimbursements, we find that private prices follow Medicare's lead. A $1 change in Medicare's relative payments moves private payments by $1.20. Results are similar when Medicare alters overall reimbursement levels. Medicare thus strongly influences both relative valuations and aggregate expenditures. Medicare's price transmission is strongest when physician groups are numerous and competitive. Transaction and bargaining costs may lead the development of payment systems to suffer from a classic coordination problem. Improvements in Medicare's payment models may therefore be public goods.

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Textbook treatments of Coase (1960) emphasize that, absent transaction costs, resources flow to their efficient uses irrespective of initial conditions. As Coase noted, however, costless transactions are "very unrealistic."[1] When transactions are costly, default arrangements may significantly influence final outcomes. From this perspective, we analyze the outcomes of bargaining between physicians and private health insurers.

Advance negotiations determine how insurers pay physicians for treating insured patients. These negotiations take place on ostensibly open, moderately competitive markets (Dafny, Duggan and Ramanarayanan 2012). The employers and beneficiaries who purchase private plans have much at stake, both medically and financially (Enthoven and Fuchs 2006). It is thus puzzling that insurers' payments to providers are rarely rooted in either the medical benefits or cost-effectiveness of care (Baicker and Chandra 2011, Cutler 2011).

One explanation for the scarcity of value-oriented payment systems involves the influence of Medicare, the federal insurer of the elderly and disabled. Medicare may influence private markets through multiple channels. First, as the largest buyer of physicians' services, Medicare competes with private insurers for medical resources (Foster 1985). High Medicare rates may bid up private fees, resulting in co-movement between public and private payments.

Second, the environment is replete with complex transactions. Millions of physician-insurer pairings could, in principle, negotiate payments for thousands of recognized treatments. The expense of bargaining and subsequent claims billing (Cutler and Ly 2011) may leave private players in search of payment models around which they can coordinate. As a large, publicly administered payer, Medicare may naturally provide such a focal point. Medicare itself is legislatively bound to base payments on input costs (Newhouse 2002), potentially driving the prevalence of cost-based payment.

Empirically, we examine how private payments respond to administrative changes in Medicare's fee schedule. Our main analysis studies a 17 percent shock to Medicare's valu-

[1] Coase (1960, 15). The size of legal and financial institutions makes the point (Wallis and North 1986).

ations of surgical procedures relative to other medical services.[2] This one-time change was implemented in 1998 and varied substantially in dollar terms across individual services. We also examine a set of across-the-board payment shocks that varied across geographic areas.

We find that private payments move tightly with Medicare's payments. On average, a $1 decrease in Medicare's payment for a surgical service led to a $1.20 decline in private payments for that service. In response to across-the-board payment changes, we find that a $1 decrease in Medicare's payments led to a $1 decrease in private payments. These findings support the view that Medicare's pricing decisions exert substantial influence over private payments. Medicare strongly influences both relative valuations of and aggregate expenditures on physicians' services.

Private prices reflect the outcomes of negotiations between physicians and private insurers. Medicare's influence on these prices must thus be mediated by the bargaining process. Since knowledge of this process may shed light on the welfare effects of Medicare's payments, we investigate the mechanisms underlying our baseline results.

Industry participants describe two modes of negotiation between providers and private insurers.[3] Insurers often make take-it-or-leave-it offers, based on fixed fee schedules, to small providers. These schedules tend to be based in large part on Medicare's payment menu, perhaps with a scalar markup.[4] In contrast, they say that insurers negotiate with hospitals and large provider groups over payments for bundles of services. We show in section 4 that both arrangements are readily rationalized. The value of an insurer's product can be enhanced by improving on Medicare's cost-based prices, but complex negotiations are costly.[5]

[2]Administratively, the payment change was associated with the "conversion factors" used to determine the generosity of payments across broad classes of services within Medicare Part B. We provide a detailed characterization of the relevant institutions in section 2.

[3]In Appendix A we allow industry participants to characterize these negotiations in their own words.

[4]This point is made specifically by Nandedkar (2011), Gesme and Wiseman (2010) and Mertz (2004).

[5]The absence of "health care entrepreneurs" (Cutler 2011) is particularly puzzling in light of the health sector's ubiquitous transaction costs and coordination problems, both emphasized by Coase (1937) as rationales for folding activities under the umbrella of a firm.

Expected gains from actively negotiated payments, which rise with the provider's scale, must exceed these substantial transaction costs for active negotiations to take place.

We find empirical evidence consistent with a role for both modes of negotiation. If payments to small group practices work directly from Medicare's menu, then the transmission of Medicare's rate changes should be particularly powerful in markets, and among specialties, with low levels of provider concentration. Conversely, highly concentrated specialties and markets should exhibit greater independence. We find empirically that this is indeed the case. Medicare's prices are transmitted most strongly in low-concentration markets, as measured both across specialties within a geographic area and across geographic areas. We also find relatively strong price transmission in markets with large numbers of providers. This too is consistent with a role for transaction costs.

In a world of active bargaining, price transmission depends on the relative sizes of public and private markets. When Medicare comprises a larger share of the market for a service, public payment changes lead to significant shifts in the resources available for private sector care. Large Medicare markets should thus predict relatively strong transmission of Medicare's payments into private prices. We find this to be the case, providing evidence for traditional bargaining considerations.

The welfare implications of public payment changes differ under the two modes of private sector bargaining. For physicians that receive take-it-or-leave-it offers linked to Medicare's fees, those fees have significant and immediate importance. When Medicare pays generously for low value services, incentives for this portion of the private sector echo that mistake. The value of improvements in public payment schedules will be similarly magnified.

When insurers bargain with large providers, the welfare consequences of Medicare payment reforms are more subtle. When Medicare influences private markets by shifting resources across sectors, a reduction in payments for a low value service may optimally increase its short-run supply to the private sector. Over the long run, however, the co-movement of

public and private prices unambiguously worsens the returns to the affected specialties. Future entry (Dezee et al. 2011), investment decisions (Acemoglu and Finkelstein 2008, Clemens and Gottlieb 2014), and innovation[6] reflect changes in the returns to practice.

Recent decades have seen a trend towards consolidation on the part of health care providers (Gaynor and Haas-Wilson 1999, Dunn and Shapiro 2012). Our analysis highlights the relevance of this trend for Medicare's role in health care markets. We find that large-scale providers are less likely to follow Medicare's lead. The development of value-oriented payment systems may suffer from a classic public goods problem; diffuse private players appear to have insufficiently strong incentives to innovate beyond Medicare's menu. Overcoming these negotiation costs may be an under-emphasized benefit of provider consolidation.[7]

1 Payment Reform with Private-Market Spillovers

As the largest U.S. purchaser of health care, Medicare may exert significant influence over private markets for physicians' services. We begin with a general characterization of the channels through which Medicare's payments may influence welfare. We maintain sufficient generality to nest a variety of possibilities. These include models in which private markets mimic Medicare's payments, models of cost-shifting, and models in which Medicare influences physician/insurer bargaining by shifting resources across sectors.

1.1 The Costs and Benefits of Public- and Private-Sector Care

Medicare pays for health care at reimbursement rate r_M, leading providers to offer q_M units of care for each of its N_M beneficiaries. This care has a per-beneficiary benefit $B_M(q_M)$,

[6] 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.

[7] The following passage from Coase (1960, 16–17) reads as though it were intended as a description of health insurance contracts. "But where contracts are peculiarly difficult to draw up and an attempt to describe what the parties have agreed to do or not to do... would necessitate a lengthy and highly involved document, and, where, as is probable, a long-term contract would be desirable, it would be hardly surprising if the emergence of a firm or the extension of the activities of an existing firm was not the solution adopted..."

with marginal benefits denoted $b_M(q_M)$. Private sector quantities and benefits are similarly defined, with a private reimbursement rate of r_p, quantity of q_p, and patient population of size N_p. Both the public and private quantity may exhibit own- and cross-price responses, leading us to write $q_M = q_M(r_M, r_p)$ and $q_p = q_p(r_M, r_p)$.

Since Medicare's reimbursement rate r_M is the relevant policy parameter, we write the social welfare function as:

$$U(r_M) = N_M B_M(q_M(r_M, r_p)) + N_p B_p(q_p(r_M, r_p)) - C(q_p N_p, q_M N_M), \tag{1}$$

where $C(q_p N_p, q_M N_M)$ describes production costs. We allow production costs to depend generally on both the private and public sector quantities. We write the marginal costs for private and public care as c_p and c_M and take no stand on the form of the cost function.

1.2 The Welfare Effects of Payment Reform

A change in Medicare's reimbursements may affect welfare through both public and private channels. Differentiating equation (1) yields a welfare impact of:

$$U'(r_M) = \underbrace{(b_M - c_M) N_M \left[\frac{\partial q_M}{\partial r_M} + \frac{\partial q_M}{\partial r_p} \frac{dr_p}{dr_M} \right]}_{\text{Public-Sector Impact}} + \underbrace{(b_p - c_p) N_p \left[\frac{\partial q_p}{\partial r_p} \frac{dr_p}{dr_M} + \frac{\partial q_p}{\partial r_M} \right]}_{\text{Private-Sector Impact}}. \tag{2}$$

We characterize the public and private components of equation (2) in turn. The first term on the right-hand side describes the public-sector consequences of Medicare's payment changes. Had public payments been optimized, this term would vanish as $b_M - c_M = 0$. This is unlikely, as Medicare is legislatively bound to reimburse on the basis of average cost rather than value. If payments have induced inefficiently large quantities of care, then $b_M < c_M$ and an increase in care provision will reduce social welfare.[8]

[8]We take care in our choice of words to allow for the possibility that own-price supply responses are negative in markets for health care services. This view, though non-standard in most settings, is embedded

The magnitude of the public-sector response depends significantly on how public prices move private prices $\left(\frac{dr_p}{dr_M}\right)$. The most natural expectation is that the cross-price elasticity $\frac{\partial q_M}{\partial r_P}$ is negative. Under this assumption, if private prices move *with* public prices $\left(\frac{dr_p}{dr_M} > 0\right)$, the impact of reimbursement changes on public sector quantities would be muted. If private prices move against public prices $\left(\frac{dr_p}{dr_M} < 0\right)$ the impact on public sector quantities may be magnified. The latter scenario is conventionally known as cost-shifting.[9] In addition to raising questions of access, as it implies an expanded wedge between public and private payments, cost shifting has pessimistic implications for Medicare's capacity to constrain aggregate costs.

The second term of equation (2)'s right-hand side describes the private-sector consequences of Medicare's payment changes. As in the public sector, the welfare change vanishes if private payments are optimized, as $b_p - c_p = 0$. This is what one might anticipate in a world of efficient, undistorted insurance markets.[10] Two broad classes of reasons make this scenario unlikely.

A first, traditional set of distortions may lead the wedge between private costs and benefits to be either positive or negative. Insurance carriers' market power could lead to

in the federal budgeting process through CBO's "volume offset" assumption (Codespote, London and Shatto 1998). The volume offset assumption is closely related to the target income hypothesis, which is shown by McGuire and Pauly (1991) to require physicians to exhibit large income effects. Papers finding evidence of volume offsets include Rice (1983) and, more recently, Jacobson, Chang, Newhouse and Earle (2013). Papers finding evidence of positive own-price supply responses include Hadley and Reschovsky (2006) and Clemens and Gottlieb (2014).

[9]For overviews of the cost-shifting literature, see Frakt (2011, 2013). Evidence in favor of cost-shifting comes primarily from the hospital context (Cutler 1998, Kessler 2007, Wu 2010, Robinson 2011). 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 (White 2013, White and Wu 2013).

[10]The optimal insurance problem is typically characterized in terms of coinsurance and demand. In this characterization, there is no implementable first-best efficient contract. Our interest in this paper centers on the other side of the problem, namely provider payments and supply. In principle, optimal insurance considerations impose no constraint on selecting the efficient reimbursement rate that leads providers to supply care until the point at which marginal benefit equals marginal cost. Determining this rate is made difficult by the need to know the relevant portions of the marginal benefit and marginal cost curves; in most settings, aggregation of such information is left to market mechanisms.

monopoly pricing and inefficiently low levels of care consumption (Dafny et al. 2012). Adverse selection may similarly result in inefficiently little insurance and thus inefficiently little care consumption from an *ex ante* perspective (Cutler and Reber 1998). Alternatively, the tax exclusion for employer-provided health insurance may drive system-wide excesses in both insurance generosity and care provision (Feldstein 1973).

A second type of distortion more mechanically involves Medicare itself. Given its status as a large, public payer, Medicare's prices may serve as benchmarks, or even defaults, for private sector negotiations. At present we wish only to raise this possibility, saving additional detail for section 4. With Medicare bound legislatively to reimburse on the basis of cost, distortions of this form likely imply reimbursement rates that deviate from marginal benefits.

We summarize all relevant distortions from the second-best as a wedge between benefits and costs of care. This wedge is a function $\phi(q_p)$ that drives consumers away from their marginal benefit curve. Instead of experiencing the marginal benefit of q_p units of care as $b_p(q_p)$, consumers subject to the distortion perceive the benefit as $b_p(q_p) + \phi(q_p)$. In equilibrium, marginal costs equal the distorted marginal benefits of care, so that

$$b_p(q_p^*) + \phi(q_p^*) = c_p(q_p^* N_p, q_M^* N_M) \tag{3}$$

at equilibrium care levels q_p^* and q_M^*. Given such a wedge, the welfare effects of Medicare payment reform include a private-sector spillover given by

$$\text{Welfare Spillover} = -\phi(q_p^*) N_p \left[\frac{\partial q_p^*}{\partial r_p^*} \frac{dr_p^*}{dr_M} + \frac{\partial q_p^*}{\partial r_M} \right]. \tag{4}$$

The size of the welfare spillover depends on the size of the relevant distortion, the size of the private market, the manner in which Medicare's payments are transmitted to the private sector, and the responsiveness of supply. Given that private markets for physicians' services are 2.5 times that of Medicare, the welfare implications of private-sector spillovers may exceed the consequences within Medicare itself.

Medicare payments could also influence private medical expenditures, according to:[11]

$$\text{Expenditure Spillover} = \frac{dr_p^*}{dr_M} q_p^* N_p + r_p^* N_p \left[\frac{\partial q_p^*}{\partial r_p^*} \frac{dr_p^*}{dr_M} + \frac{\partial q_p^*}{\partial r_M} \right]. \tag{5}$$

The relevance of the price-transmission process for Medicare's influence on aggregate health expenditures is readily apparent. A cost-shifting world, where $\frac{dr_p^*}{dr_M} < 0$, is a world in which Medicare's payment changes are, at least in part, mechanically offset by changes in private expenditures. A world of positive price transmission, $\frac{dr_p^*}{dr_M} > 0$, is a world in which Medicare exerts significant influence over total spending. This will be particularly true over the long run, when $\frac{\partial q_p^*}{\partial r_p^*}$ and $\frac{\partial q_M^*}{\partial r_M}$ reflect entry decisions. This paper seeks to characterize the price-transmission relationship and understand its underlying mechanisms.

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

To estimate Medicare's influence on private sector pricing, we exploit a large, administrative change in Medicare's reimbursement rates. While Medicare's fees are set according to administrative rules, these rules can be changed by acts of Congress. Such changes deliver variation in payment rates that may be independent of patient demand, technological change, and supply-side market pressures. Before describing the relevant shocks, we first characterize price determination in markets for health care services.

2.1 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

[11] Though not a welfare measure *per se*, the magnitude of aggregate health expenditures make them of considerable independent interest. 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.

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 according to measured resource utilization, as in Medicare, they are agreed upon through negotiations between insurance carriers and the provider networks 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). Providers themselves may have little information about payments received by others and hence of the "competitive" rate. The details of these negotiations are also not transparent, and our limited knowledge about private sector prices comes from claims data that reveal the reimbursements actually paid for specific services.

Previous work sheds light on the economic determinants of health care pricing. Cutler, McClellan and Newhouse (2000) find significant differences between the prices negotiated by HMOs and traditional health insurance plans, with HMOs paying 30 to 40 percent less for comparable services. 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). 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).

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

2.2 A Large Shock to the Relative Prices of Outpatient Services

Compared to the private sector, Medicare's pricing is transparent. 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 health care services according to the resources they are believed to require. It also recognizes that goods and services have different production costs in different parts of the country; Congress mandates price adjustments to offset these differences in input costs. For service j, supplied by a provider in payment area i, the provider's fee is approximately:

$$\begin{aligned}\text{Reimbursement}_{i,j,t} = \text{Conversion Factor (CF)}_{t,c(j)} \times \text{Relative Value Units (RVU)}_{j,t} \\ \times \text{Geographic Adjustment Factor (GAF)}_{i,t}. \qquad (6)\end{aligned}$$

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 successfully that lower prior growth in procedure use than in the use of other medical services should be rewarded. Congress implemented this plan, and distinguished between the CFs for surgery, primary care, and other non-surgical services from 1993 through 1997. 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 17 percent for surgical RVUs relative to primary care and other non-surgical RVUs. This spawned political discontent among non-surgeons.[14] In 1998, this 17 percent bonus was eliminated through a budgetarily neutral merger of the CFs.[15]

[14]The American Medical Association presents data on historical Conversion Factor rates at `http://www.ama-assn.org/ama1/pub/upload/mm/380/cfhistory.pdf` (accessed March 26, 2011).

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

The CF merger resulted in a large change to relative payments across broad categories of services, with substantial service-level variation in the dollar value of the payment shocks. In Appendix B we present additional analysis of across-the-board payment changes associated with a separate overhaul of the system of geographic adjustments. While complementary, these natural experiments are best suited for answering somewhat different questions. Shocks to relative prices are best suited for assessing the link between Medicare and the private sector's relative valuation of services; resulting estimates will thus speak to Medicare's role as a driver of cost-based reimbursement. Across-the-board payment shocks more directly affect physicians' bottom lines; resulting estimates are thus relevant to questions related to cost shifting and to Medicare's effects on aggregate health expenditures.

2.3 Estimation Strategy

We use these administrative changes to see how actual Medicare payments affect private sector reimbursement rates. To do this, we use the conversion factor shock as an instrument for the average Medicare payments observed in the claims data. Our private sector claims data, described in section 2.4 below, do not permit identification of individual insurers or physician groups. As we therefore cannot estimate a structural bargaining model, we summarize the data using average prices across firms for identifiable services.

Practitioner characterizations of private payment negotiations inform our estimation framework (e.g. Gesme and Wiseman 2010, Mertz 2004). These characterizations suggest that private prices respond linearly to Medicare, for instance according to

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

where b is a positive scalar. Under this model, the parameter of interest, b, must be estimated using the levels of Medicare payments as opposed to logs.[16] This is especially true if there

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

is economically interesting heterogeneity in b. Estimation in levels is also consistent with traditional models of bargaining over a fixed surplus. These considerations drive the specific framework laid out below. We also estimate the relationship between log public and log private prices. The latter estimates are qualitatively similar, but with inferior model fit.

Stage 0: Compute the Instrument: Predicted Price Change

Using Medicare payment data from 1997 and before, we compute each service's average price $\overline{P^{\text{Medicare}}}_{j,\text{pre}}$ before the policy change. We then use the merger of Conversion Factors in 1998 to predict the Medicare price change in following years. Specifically, we define

$$\text{PredChg}_j^{\text{Medicare}} = \overline{P^{\text{Medicare}}}_{j,\text{pre}} \times \begin{pmatrix} -0.11 \times \text{Surgical}_j \\ +0.06 \times \text{Non-Surgical}_j \end{pmatrix} \tag{7}$$

where the factors -0.11 and 0.06 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 $\text{PredChg}_j^{\text{Medicare}}$ as an instrument for the actual Medicare reimbursement rate. Specifically, we run the following first stage regression:

$$\begin{aligned} P_{j,s,t}^{\text{Medicare}} = \pi \cdot \text{PredChg}_j^{\text{Medicare}} \times \text{Post1998}_t + X_{j,s,t}\psi + \mu_j I_j + \mu_s I_s + \mu_t I_t \\ + \mu_{j,s} I_j \times I_s + \mu_{t,s} I_t \times I_s + e_{j,s,t} \end{aligned} \tag{8}$$

at the service (j), by state (s), by year (t) level.[17] We weight observations by the number of times the service was performed in 1997.[18] Because the payment changes vary significantly

[17] 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 that uses state-level heterogeneity. Appendix C.1 shows that our results remain similar when using national level observations.

[18] 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

across services, we cluster standard errors at that the service-code level.

Equation (8) 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-by-state ($I_j \times I_s$) and state-by-year ($I_s \times I_t$) fixed effects, and direct service, state, and year 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.[19] We further include controls, defined in section 2.4, for the types of insurance plans associated with the data.

Stage 2: Second Stage

The Medicare price predicted in equation (8) then serves as an instrument for actual Medicare prices in the following second stage equation:

$$P_{j,s,t}^{\text{Private}} = \beta \cdot \widehat{P_{j,s,t}^{\text{Medicare}}} + X_{j,s,t}\phi + \nu_j I_j + \nu_s I_s + \nu_t I_t + \nu_{j,s} I_j \times I_s + \nu_{t,s} I_t \times I_s + \varepsilon_{j,s,t} \quad (9)$$

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 (8). Second, the shock used to generate predicted prices must be conditionally independent of other sources of change in private sector payment rates

service by its service count times its assigned number of relative value units.

[19] 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 17 percent reduction in the surgical CF.

$\varepsilon_{j,s,t}$. These include technology shocks, demand shocks, and 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:

$$P_{j,t}^{\text{Medicare}} = \sum_{t \neq 1997} \gamma_t \cdot I_t \times \text{PredChg}_j^{\text{Medicare}} + X_{j,s,t}\psi + \mu_j I_j + \mu_s I_s + \mu_t I_t + \mu_{j,s} I_j \times I_s + \mu_{t,s} I_t \times I_s + u_{j,s,t} \quad (10)$$

$$P_{j,s,t}^{\text{Private}} = \sum_{t \neq 1997} \delta_t \cdot I_t \times \text{PredChg}_j^{\text{Medicare}} + X_{j,s,t}\alpha + \upsilon_j I_j + \upsilon_s I_s + \upsilon_t I_t + \upsilon_{j,s} I_j \times I_s + \upsilon_{t,s} I_t \times I_s + \upsilon_{j,s,t} \quad (11)$$

Estimates of δ_t and γ_t for 1997 and before would find any pre-existing trends, while estimates for 1998 and beyond trace out the dynamic effects of Medicare's payment shocks. For Medicare itself, the post-1997 coefficients in equation (10) should hew to 1.

2.4 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 the claims for service j in year t and computing the average allowed charge, $P_{j,t}^{\text{Medicare}}$.[20] We measure private sector prices similarly, separately for each state s, using private insurance claims data from the ThompsonReuters MarketScan ("MedStat") database. Private insurers use procedure codes that overlap substantially with the HCPCS system. Participating insurers submit those codes, along with service-level payment rates and additional information, to MarketScan. 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.[21] We aggregate claims to the code-by-state-by-year level and compute $P_{j,s,t}^{\text{Private}}$.

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.[22] 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 service claims came from Major Medical or Comprehensive Insurance (CI) plans, 52 per-

[20] We first eliminate claims with payments of \$1 or less, or service quantities of 100 or above.

[21] We again eliminate claims of less than \$1 or with quantities of 100 or more. We also eliminate claims associated with capitated payment arrangements, which do not reflect the per unit prices of interest.

[22] Although our estimation sample includes only 2,194 of the 12,729 unique HCPCS codes observed in MarketScan during our sample period, the codes included represent the majority of care provided. This is because the commonly used codes are more likely to (a) be officially recognized by public and private payors, (b) appear in both our Medicare and MarketScan samples, and (c) have a balanced panel. Of the 12,729 codes observed in MarketScan, 4,306 match Medicare claims with complete pricing information, and they represent 66 percent of services provided. When we balance the panel across years we are left with our final estimation sample, representing 58 percent of the MarketScan services.

cent 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 driven by differential shifts towards different plan types over time, we construct a control variable "Plan Type Payment Generosity" to capture the evolution of generosity in plan types. The variable is constructed by regressing payments on plan types and aggregating the resulting predicted payments at the state-by-year-by-service level. We also construct a control for plan generosity based on 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.

3 Empirical Effect of Medicare Prices on Private Prices

Panel A of Figure 1 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 B. Appendix Figure C.1 shows analogous graphs using cross-state variation.

Figure 2 plots event study estimates of the effect of Medicare payment changes using equations (10) and (11). First stage results, marked on the graph with "$\times$" symbols, show that the predicted service-level price changes translate approximately one-for-one into observed

Medicare payment rates. This gives 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 pre-existing trends driven by changes in technology, demand, or other market conditions. From this point forward, a $1.00 increase in Medicare's payment led, on average, to a $1.20 increase in private payments.

In Table 2, we summarize these results in single coefficients using the framework described by equations (7) through (9). Column 1 reports the first-stage estimates of equation (8). 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).[23]

Column 2 shows the reduced form results we obtain when $P_{j,s,t}^{\text{Private}}$ replaces $P_{j,s,t}^{\text{Medicare}}$ as the outcome variable in (8). 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. The IV estimate of equation (9) in column 3 rescales this private sector response by the actual Medicare response (from column 1), and represents our best estimate of how the private sector reacts to Medicare pricing. Each dollar change in Medicare payments leads to a $1.15 change in private reimbursements, in the same direction as the Medicare change, confirming Figure 2.

Columns 4 through 6 run comparable specifications in which public and private prices are expressed in logs.[24] The instrument in these specifications is a binary indicator for surgical services performed in or after 1998. Column 4 shows that Medicare's elimination of the surgery-specific conversion factor resulted in a 22 log point decline in relative payments for surgeries. The point estimate is statistically indistinguishable from the 17 percent decline

[23]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 absence of homoskedasticity.

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

called for by the payment reform. We report a reduced form estimate of this policy change's impact on private prices of −0.11 in column 5. Column 6 reports the IV estimate, showing that, on average, a 10 percent change in Medicare's payment for a service resulted in a 4.8 percent change in private payments for that service. This is reconciled with the $1.15 from column 3 by the fact that the average private payment for a service is significantly higher than the average Medicare payment.[25]

Appendix Table C.1 demonstrates the robustness of our main 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 insurance plans represented in the sample.[26] Column 4 shows that our baseline estimate is not qualitatively sensitive to our controls for mid-1990s payment changes targeted at cataract surgery, although omitting them changes the magnitude of the price transmission coefficient from 1.2 to 1.0. Column 5 removes the service weights, which reduces the estimate to around 0.7 but maintains precision.[27] Column 6 includes a control for the number of Relative Value Units (the quantity metric that appears in equation [6], Medicare's payment formula) assigned to each service. Minor updates to RVU assignments strongly predict Medicare's allowable charges, which they impact formulaically (coefficient

[25]Table 1 shows that Medicare pays $239 on average for surgical services and $114 for non-surgical. Its surgical payments fell by 11 percent, or $26, while medical reimbursement rates increased by 7%, or $8. So the difference fell by $\$26 - \$8 = \$18$. 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 1.15 that we have estimated means that Medicare's $18 relative change only led to a $21 ($= 1.15 \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.

[26]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.

[27]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 must, as a general rule, be implemented in a budgetarily neutral fashion, making it essential to weight each service by its baseline frequency.

not shown). These updates are modest predictors of changes in private prices, however, and controlling for them has little impact on our baseline result. Finally, column 7 shows that the baseline is robust to controlling directly for a linear trend in payments for surgical procedures relative to other services. As Figure 2 shows, there is no such trend.[28]

We also estimate the effects of a set of across-the-board payment changes that varied across geographic areas. We present this analysis in Appendix B. The estimated price-transmission coefficient is on the order of 1, and is thus qualitatively quite similar to the transmission of relative price shocks. As discussed in section 2.2, these responses to across-the-board payment shocks are relatively direct evidence against cost shifting. Taken together, the estimates suggest that Medicare's payment changes exert significant influence over both relative valuations of and aggregate spending on physicians' services. We next explore the mechanisms underlying the price transmission process.

4 What Underlies Medicare's Impact On Private Prices?

Private prices reflect the outcomes of negotiations between physicians and private insurers. Medicare's influence on these prices must thus be mediated by the bargaining process. Here we develop a bargaining framework that matches practitioners' characterizations of their negotiations. Their own descriptions can be found in Appendix A.

Practitioners describe two modes of negotiation between providers and private insurers. Insurance carriers typically offer small providers, such as sole practitioners and small physician groups, contracts based on a fixed fee schedule. Whether it is copied directly from Medicare or modified by the insurer, the parties negotiate a constant scaling of this schedule (Nandedkar 2011, Gesme and Wiseman 2010, Mertz 2004). In contrast, insurers are said to negotiate in more detail with hospitals and large groups over specific payment rates.

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

Adoption of Medicare's fee schedule may be optimal due to the substantial negotiation and coordination costs in our setting (Coase 1937, 1960).[29] The value of the insurance product can potentially be improved, however, through more detailed negotiations over service- or bundle-specific prices. Such negotiations may reduce inefficiencies in the incentives that Medicare's pricing schedule offers.[30] To incorporate these deviations from the Medicare menu, we model negotiations as taking place in two distinct steps. The insurer and physician group first determine an average payment rate. They then choose whether to base relative payments on Medicare's menu or adopt alternative relative prices.

4.1 Determining the Average Payment Level

Insurer i and physician group g first negotiate the overall price level. This is expressed as a constant markup φ_i^g of Medicare's fee schedule, such that when Medicare pays $r_{M,j}$ for service j the private payment is $r_{i,j}^g = \varphi_i^g r_{M,j}$. The markup φ_i^g is determined through Nash bargaining. The physician's outside option is his expected wage from treating other patients. We assume that the physician would treat patients randomly in proportion to the other insurers' market shares, including Medicare (M). His reservation value is thus

$$\overline{u}^g = \lambda_M^g \overline{r}_M + \sum_{\iota \neq i} \lambda_\iota^g \overline{r}_\iota^g \tag{12}$$

where ι indexes all insurers other than the one involved in the current negotiation, λ_ι^g is insurer ι's market share, and $\overline{r}$ represents the average payment rate across services.

[29] Medicare's position as the single-largest payer for health care services further reinforces its relevance as a setter of the default menu. The Medicare menu may be particularly relevant for relative prices across services. Practitioners describe the offers made by insurers to sole practitioners, for example, as being take-it-or-leave it, scalar mark-ups (or occasionally slight mark-downs) of Part B prices. Providers themselves may find deviating from Medicare's menu costly due to increasing in the non-trivial administrative expenses associated with billing (Cutler and Ly 2011). Regulations requiring insurers to pay sufficiently to ensure access to "medically necessary" services may also contribute to such a role for public players in these markets.

[30] 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 those with supplemental insurance, are comprehensively insured, there may be a substantial wedge between marginal cost and marginal benefit at these quantities.

We define the insurer's reservation payment rate as an exogenous parameter $\underline{u}_i$. This parameter represents the amount the insurer would pay to its next-best option if it fails to reach agreement with group g.[31] The insurer's Nash bargaining parameter is θ, and the physician group's is $1-\theta$. The parties will thus settle on an average reimbursement rate of

$$\overline{r}_i^g = \varphi_i^g \overline{r}_M = \theta \left(\lambda_M^g \overline{r}_M + \sum_{\iota \neq i} \lambda_\iota^g \overline{r}_\iota^g \right) + (1-\theta)\underline{u}_i. \tag{13}$$

Since each insurer's average rate depends on all of the others, we solve for the reimbursements jointly (see Appendix D). In the case with I symmetric insurers, each with market share $\overline{\lambda} = \frac{1-\lambda_M}{I}$ and paying a reimbursement rate $\overline{r}$, the average private payment rate is

$$\overline{r} = \frac{\theta \lambda_M \overline{r}_M + (1-\theta)\underline{u}}{1-\theta+\theta\lambda_M + \theta\frac{1-\lambda_M}{I}}. \tag{14}$$

Its response to a change in the overall Medicare payment level is

$$\frac{\partial \overline{r}}{\partial \overline{r}_M} = \frac{\theta \lambda_M}{1-\theta+\theta\lambda_M + \theta\frac{1-\lambda_M}{I}}. \tag{15}$$

Medicare's average payment is positively transmitted to the private sector, consistent with our empirical evidence. The magnitude of the transmission coefficient is decreasing in physicians' bargaining power $(1-\theta)$ and increasing in Medicare's market share (λ_M increases). The expression yields an ambiguous prediction regarding insurer market power; a decline in the number of competitors I directly increases the price transmission coefficient, but would also tend to decrease θ, which pushes in the opposite direction. To the extent that physician groups specialize in surgical vs. non-surgical services, equation (15) also has implications for relative prices. Because the price transmission coefficient is bounded between 0 and 1,

[31] We think of $\underline{u}_i$ as also being adjusted to account for differences in quality, and profit losses if patients leave insurer i should its network exclude provider g.

however, it cannot fully explain the results from section 3.[32]

4.2 When To Reference Price and When To Bargain

We now consider the physician and insurer's choice to adopt or deviate from Medicare's menu of relative payments. If the parties adopt Medicare's menu, the price for service g is given by $r^g_{i,j} = \varphi^g_i r_{M,j}$, with the constant markup φ^g_i, given by:

$$\varphi^g_i = \frac{\overline{r}}{\overline{r}_M} = \frac{\theta\lambda_M + (1-\theta)\frac{\underline{u}}{\overline{r}_M}}{1-\theta+\theta\lambda_M+\theta\frac{1-\lambda_M}{I}}. \tag{16}$$

In such negotiations, a \$1 change in Medicare's relative payments for two services will be scaled by φ^g_i. In our data the average ratio of private payments to Medicare payments is 1.4. This is suggestively close to section 3's baseline estimate of relative price transmission.

Because Medicare's payment model has inefficiencies, we model deviation from Medicare's menu as a way to improve the insurance contract quality. This could result from better aligning service-specific payments with their associated health benefits. Insurer i's quality is Ξ_i. Consumers have preferences over Ξ_i and the cost of insurance, captured by a management fee f_i. Specifically, we assume that the demand curve is:

$$D_i(f_i, \boldsymbol{N}_i) = K\zeta_i \left[\frac{\Xi_i(\boldsymbol{N}_i)}{f_i}\right]^\epsilon \tag{17}$$

$$\text{where} \quad \Xi_i(\boldsymbol{N}_i) = \prod_{g=1}^{G} \left(1 + N_{ig}\Lambda_g\xi_g\right). \tag{18}$$

In equation (17), K is a constant, ζ_i represents consumers' average unconditional preference for insurer i (as in Starc, forthcoming), f_i is the management fee charged by the insurer,[33]

[32] Because physician groups can provide a diverse mix of services, some classified as surgical and others as non-surgical, their average Medicare payments would have moved moderately with the price shocks analyzed in section 3. The transmission coefficient for Medicare's relative price changes, as implied by equation (15), would thus be further below 1.

[33] We focus on the insurer's management fee, which tends to be a fixed price per enrollee, and abstract away from the incurred medical losses. In our empirical setting, this approximation may be appropriate, as the

and the demand elasticity is $\epsilon \geq 0$.[34] The quality Ξ_i depends on the negotiation decision with each group g. Groups vary in the share of the insurer's patients they serve, Λ_g. There may also be variation in the value ξ_g to consumers of improving the reimbursement schedule. If Medicare's payments for angioplasty are particularly unhinged from value, for example, both ξ_g and Λ_g will be large when insurers negotiate with a large cardiology group.

The binary variable $N_{ig} \in \{0, 1\}$ indicates i's choice to deviate from Medicare when negotiating with g, at cost c_i, and $\boldsymbol{N}_i$ the vector of choices. Insurers' management cost is m_i per patient. The insurer's profit, expressed as a function of negotiation choices, is thus:

$$\pi_i(f_i, \boldsymbol{N}_i) = (f_i - m_i)D_i(f_i, \boldsymbol{N}_i) - c_i \sum_g N_{ig}. \tag{19}$$

It makes sense to deviate from Medicare's menu with group g whenever

$$\Lambda_g \xi_g \geq \chi^* \equiv \left[1 + \frac{c_i}{K\zeta_i \tilde{\epsilon}_i \Xi_i(\boldsymbol{N}_{i,-g})^\epsilon}\right]^{\frac{1}{\epsilon}} - 1 \tag{20}$$

where $\tilde{\epsilon}_i$ depends on the management cost m_i and the demand elasticity ϵ, and $\Xi_i(\boldsymbol{N}_{i,-g})$ denotes the part of quality calculated from (18) ignoring group g. Appendix D derives this inequality and develops additional characteristics of Medicare's influence on private-sector bargaining.

The model's empirical implications for the transmission of Medicare's relative payments to the private sector hinge largely on inequality (20). The central role of this decision is intuitive, as private prices mechanically track Medicare's prices when physicians and insurers work directly from Medicare's menu. If inequality (20) holds then a physician-insurer pair no longer follows Medicare's relative prices.

Equation (20) shows the conditions under which deviations from Medicare's menu are

plans we study are self-insured plans provided by large employers. Thus the management fee is the relevant price from the insurer's perspective. Whether this simplification is innocuous depends on whether the insurer is simply passing costs through onto its patients, as in a perfectly competitive market, or potentially marking them up, as when it has market power or is a two-sided platform (Bardey and Rochet 2010).

[34] Appendix D presents one way to motivate this demand system from underlying consumer preferences.

likely. First, innovating beyond Medicare's menu generates significant value when a negotiation impacts a large number of consumers (Λ_g is large), and thus when insurers encounter large physician groups representing many patients. Second, insurers benefit more from improving the value of their product when consumers' demand is more elastic (ϵ is large). This will be true in competitive markets, where consumers have alternative insurers they can choose. Finally, deviations from Medicare's menu are relatively likely when insurers have low per-service negotiation costs (c_i), more patients (higher K or ζ_i), and when Medicare's deviation from the potential negotiated fee schedule is large (ξ_g is large).

4.3 Welfare Implications

Medicare's payment policies have substantial welfare implications when physicians and private insurers adopt Medicare's menu. When Medicare pays generously for low value services, incentives for this portion of the private sector echo Medicare's mistake. The value of improvements in Medicare's payment policies will be similarly magnified.

When physicians and insurers negotiate actively over service-specific payments, the response of private payments to Medicare's payments may have neutral efficiency implications. That is, when prices are actively negotiated, Medicare moves private prices and quantities because it shifts private-sector supply. Suppose, for example, that Medicare improves policy by reducing payments for services that have low medical value. The payment reductions lead the relevant specialty to shift resources away from Medicare and thus, as illustrated in Figure 3, towards the private sector. Importantly, this shift both reduces the optimal private payment *and* increases the optimal provision of the service to the privately insured because of lower costs. Consequently, it would be a mistake to view these increases in private supply as evidence that Medicare's change has had an adverse effect.

Cross-sector effects and income effects may blunt efforts to reduce the utilization of low-value services in the short run. But theory speaks unambiguously to the long-run effect of

reducing Medicare's payments for low value services. The co-movement of Medicare and private payments unambiguously reduces the returns to practicing in the relevant specialty or making associated investments. In the long run, this reduction in the returns to practice will shift the supply of physicians away from the specialty providing the targeted services.

5 Provider Concentration and Price Transmission

We next explore the empirical conditions under which Medicare's influence on private prices is weaker or stronger. Estimating heterogeneity in the strength of Medicare's price transmission serves two primary purposes. First, it allows us to explore the relevance of the previous section's framework. Second, we take advantage of an opportunity to describe outcomes associated with multilateral bargaining. Specifically, we provide evidence on how bargaining equilibria are altered by changes in parties' outside options. The facts generated by this analysis thus inform our understanding of bargaining in markets in which neither demand nor supply is perfectly competitive.

5.1 Measures of Physician Market Power

We begin by examining the importance of provider consolidation. We measure the degree of competition among physician groups using a Herfindahl-Hirschman Index (HHI) constructed with Medicare claims data. The claims data report both a unique physician identifier and the tax identifier of the group with which each physician is associated. In claims data from a 20 percent sample of all Medicare beneficiaries, we will come quite close to having this information for all Medicare-serving physicians in the country. Our first measure is constructed by estimating the HHI for all physician groups within a Hospital Referral Region (HRR).[35] We then average this measure across the HRRs within each state

[35]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

to measure the average degree of competition across the markets within that state.

We next compute a more targeted measure of concentration that is allowed to vary across specialties as well as states. For this metric we construct HRR-level HHIs separately for each of the 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 (alternatively, less competition) for largely mechanical reasons. More importantly, they exhibit a great deal more variation than the all-physician HHIs.

Incorporating specialty-specific HHIs into our analysis requires restricting attention to services that tend to be provided primarily by members of a particular specialty.[36] Since they vary both across states and across specialties within each state, the specialty-specific HHIs 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.

To estimate the influence of provider consolidation on Medicare's price transmission, we interact the price shocks with either the all-physician or specialty-specific HHI. We first standardize the HHI variables as z scores. Because the first stage coefficient in Table 2's levels regression was nearly 1, making the IV and reduced-form results nearly identical, we now focus on reduced form estimates. Recalling that $\text{PredChg}_j^{\text{Medicare}}$ is the predicted Medicare price change, and using $HHI_{j,s}$ to denote the applicable HHI z-score, we run:

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.

[36] This is 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 provider consolidation.

$$
\begin{aligned}
P_{j,s,t}^{\text{Private}} = {} & \beta_1 \cdot \text{PredChg}_j^{\text{Medicare}} \times \text{Post1998}_t \\
& + \beta_2 \cdot \text{PredChg}_j^{\text{Medicare}} \times \text{Post1998}_t \times HHI_{j,s} \\
& + X_{j,s,t}\gamma_1 + X_{j,s,t} \times HHI_{j,s}\gamma_2 + \mu_j^1 I_j + \mu_s^1 I_s + \mu_t^1 I_t \\
& + \mu_j^2 I_j \times HHI_{j,s} + \mu_s^2 I_s \times HHI_{j,s} + \mu_t^2 I_s \times HHI_{j,s} \\
& + \mu_{j,s}^1 I_j \times I_s + \mu_{t,s}^1 I_t \times I_s + \mu_{j,s}^2 I_j \times I_s \times HHI_{j,s} + e_{j,s,t} \qquad (21)
\end{aligned}
$$

We allow the coefficients on all time-varying controls to vary with the relevant HHI variable.[37]

5.2 Heterogeneity by Physician Market Power

Table 3 presents the estimates. Column 1 shows that the average price-transmission coefficient is roughly 1.3, but that it varies dramatically with physician HHI. The coefficient of -0.5 on the physician HHI interaction implies that as HHI increases by 1 standard deviation, the price-transmission 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. Price transmission 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. Thus the HHI coefficient is not merely capturing differences in the absolute sizes of the relevant markets. It is of independent interest that, conditional on HHI, the number of physicians is strongly associated with the strength of Medicare's price transmission. Together, the results suggest that fragmented markets are relatively likely to follow Medicare's cues, perhaps because they are markets in which the gains from active bargaining are unlikely to outweigh its costs.

[37] 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 (21) we have omitted interactions between the HHI variables and the state-by-service code fixed effects ($I_j \times I_s \times HHI_{j,s}$), of which there are in excess of 50,000.

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 statewide 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. 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, and the impact of physician numbers loads on the specialty-specific variable. The results uniformly support the view that Medicare is more relevant in competitive markets than in markets with concentrated providers.

Panel A of Figure 4 reports the first stage and price-transmission coefficients separately for each tercile of the specialty-HHI distribution. The price transmission falls from above 2 in the most competitive tercile to around zero in the most concentrated. Appendix Table C.4 demonstrates the robustness of this variation across HHIs to the inclusion of additional controls. We interact various area characteristics—ranging from Census region indicators to income per capita—with the predicted payment shock. These controls have little effect on the coefficients associated with our measures of provider consolidation.

These results show that Medicare is especially relevant to payments for small and competitive physician groups. This is exactly as our model predicts.

6 Insurance Competition and Price Transmission

We next explore the relevance of insurance market conditions. Competition in insurance markets has two potential consequences for transmission of Medicare's prices into the private sector. First, small insurers—much like small provider groups—could have relatively high costs of developing novel fee schedules. This would predict that areas with more insurance market competition should have stronger price transmission.

Alternatively, competition among insurers could drive the opposite result. Relatively mo-

nopolistic insurers may have diluted incentives to create value, and thus to actively improve upon Medicare's payment rates. Areas with more competition among insurers would thus have weaker price transmission than areas with concentrated insurance markets.

6.1 Measuring Insurance Competition

Our primary measures of insurance competition are computed from health insurance reports obtained from the National Association of Insurance Commissioners (NAIC).[38] Based on the NAIC reports on each insurance carrier's size in each state, we compute state-level Herfindahl-Hirschman Indices (HHIs) for health insurance markets in each state except California.[39] 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.[40] We also compute the number of active insurers in each state in 2001.[41]

We also use a second measure of insurer competition taken directly from American Medical Association (2007). The AMA reports HHIs for all states but Kansas, North Dakota, Mississippi, Pennsylvania, South Dakota, West Virginia, and Washington, DC, from 2006.[42]

[38]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.

[39]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. 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.

[40]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.

[41]An insurer is "active" if it has positive enrollment, premiums collected, and health care expenditures.

[42]The AMA data on insurance carrier HHI have a number of problems, many of which have been documented by Dafny et al. (2011). Most significantly, they measure competition among carriers for fully-insured health plans, while the private sector data from Thompson Reuters are for self-insured companies. Second, the state-level HHI will naturally decline with the geographic size of the state, even if any one sub-state geographic market has limited competition. Third, these data are implausibly volatile over time, suggesting that observations from any one year are subject to significant measurement error. These issues suggest that regressions based on the AMA concentration data are likely to be subject to measurement error and may well underestimate the importance of concentration.

As in Section 5, we convert HHIs into z-scores and run regressions paralleling equation (21).

6.2 Impact of Insurance Competition on Price-Setting

Table 4 shows these regressions. The reduced sample size compared with Table 2 reflects the omission of California from the insurance market data. Columns 1 through 3 all show a strong positive relationship between insurance concentration and the magnitude of price transmission. The magnitude 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 price transmission coefficient.[43]

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, and is thus associated with a price transmission coefficient of around 0.85. The ninety-fifth percentile HHI is 0.78 and has cost-following of 2.2. This result is shown graphically, by tercile of concentration, in Panel B of Figure 4.

Columns 4 through 6 of Table 4 show that the interaction is robust to also controlling for interactions with the number of insurers in the state and with physician HHI. In the appendix, we further investigate the robustness of this result. Appendix Table C.5 includes as controls interactions between the predicted price change and various state-level demographic measures. The effect of insurer HHI is robust to all of these controls.

Thus far, this section's results have used insurance concentration data from 2001, as it is the earliest year with comprehensive insurer enrollment data. Appendix Table C.7 shows the robustness of column 1 when using earlier data, which are computed only from HMO enrollment, as well as later years' concentration. When we measure concentration in 1997, 1998, or 2002, the effect of insurance competition is unchanged. The HHI from

[43]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.

1996 has a lower and noisier coefficient, and is indistinguishable from zero. If we measure insurance concentration in 2006, with either NAIC or AMA data, the result disappears. 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.

6.3 Interpreting the Insurance Competition Results

Section 4.2 suggests that insurer competition drives deviations from Medicare's menu through insurers' efforts to improve quality. If so, we would expect payment-system quality to be correlated with the insurer HHI. There are many barriers to interpreting payment differences as differences in quality. We thus consider this issue briefly, adopting the conventional wisdom that surgical payments are excessive relative to payments for other services.

We regress payment rates per service on the Medicare rates, physician concentration, insurance concentration, and the interaction between insurance concentration and a surgical procedure indicator:

$$\ln P_{j,s,t}^{\text{Private}} = \underset{(0.02)}{0.94} \ln P_j^{\text{Medicare}} - \underset{(0.002)}{0.007} \cdot \text{Insurance HHI}_s + \underset{(0.008)}{0.024} \cdot \text{Insurance HHI}_s \times \text{Surgical}_j$$
$$+ \underset{(0.002)}{0.024} \cdot \text{Physician HHI}_s + \underset{(0.05)}{0.47} \cdot \text{Surgical}_j + \underset{(0.10)}{0.48}, \tag{22}$$

where the HHI variables are z-scores, and the standard errors (in parenthesis) are clustered at the service code level. As in Figure 5, which is consistent with Dunn and Shapiro (2012), average physician payments are higher in areas with more competitive insurers. This is not true, however, for surgical services. This result hints at more concentrated effort by competitive insurers to reduce payments for care that is widely perceived as overpaid.

7 Service Market Size and Price Transmission

We next explore the relevance of the size of markets for individual medical services. Equation (15) points to the relative size of the relevant public and private markets. As Medicare's relative size grows, one would expect cross-price responses to increase. Large public markets may have strong price transmission. Equation (20) points out that the absolute size of private markets may influence whether or not prices deviate from Medicare's menu. Common services may thus have private pricing that is relatively independent of Medicare. While negotiations take place at the group level, as emphasized in section 4, these factors will influence relative prices across services when physicians specialize in the provision of particular services.

Our measure of private market size simply adds all instances in which each service appears during a baseline year of the MedStat database ("Private Market Volume"). We also construct a metric that proxies for the relative sizes of the Medicare market and private markets ("Medicare Relative Size"). This metric is 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, this variable may poorly measure the absolute level of the relative public and private market sizes. Nonetheless, it should form a reasonable basis for dividing services into those with relatively large and small Medicare market shares.

We present summary statistics describing Private Market Volume and Medicare Relative Size in Table 1. Both of these variables are 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 them using percentile ranks rather than z-scores. We subtract 0.5 from the percentile ranks so that the resulting variables are symmetric about 0. We then interact these variables with the price shocks and controls as in equation (21).

Table 5 presents these results. Column 1 shows that the public-private ratio enters significantly, with a coefficient of 1.3. Moving from the 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 price transmission coefficient of 1.5. The larger the relative size of the Medicare market, the stronger the transmission. Column 2 tests the impact of Medstat private market volume alone. The estimate in this case is statistically indistinguishable from zero.

8 Specialty Exposure to Payment Reductions

Finally, we use the Medicare claims data to calculate the share of each specialty's revenue that comes from surgical procedures as opposed to other medical services. We then link the resulting measure of specialty "exposure" to the payment shocks to the specialty-dominated services described in Section 5.1. We convert this variable into z-scores and interact it with the predicted payment shocks and controls as in equation (21).

Column 3 of Table 5 presents estimates of the relevance of specialties' exposure to the downside of the reduction in payments for surgical services. The cost-shifting hypothesis suggests that income effects will lead Medicare payment reductions to result in private payment increases. The bargaining framework from Section 4.1 has the opposite implication, while income effects are irrelevant when the Medicare menu is adopted by default. The negative estimate here indicates some reduction in cost-following for more affected specialties, but it is not large enough to drive cost-shifting. This is consistent with the results presented in Appendix B, which we take as more direct evidence on the relevance of cost-shifting in markets for physicians' services. The analysis presented in Appendix B involves across-the-board payment shocks that varied across geographic areas. The price transmission coefficients are quite similar to those associated with the relative price shocks analyzed above. Income effects appear to mediate the price-transmission process only weakly.

9 Conclusion

We assess Medicare's influence on private fees for physicians' services, and find its influence to be substantial. A $1 change in Medicare's relative payments across services leads to a $1.20 change in private payments. When Medicare mistakenly pays generously for low-value services, much of the private sector follows its lead. Medicare similarly moves the level of private payments when it alters fees across the board. Medicare thus influences both the relative valuation of, and aggregate expenditures on, physicians' services.

Medicare's influence varies significantly across markets. We find it to be greatest where physician groups are small and insurance markets concentrated. The importance of physician group sizes has implications for ongoing trends towards health care provider consolidation (Kletke, Emmons and Gillis 1996). A world of diffuse providers and concentrated insurers provides only weak incentives to innovate relative to Medicare's menu. Gains associated with reductions in provider fragmentation may thus extend beyond improvements in care coordination (Cutler 2010).

Medicare's influence derives from multiple sources. First, as a large market participant, Medicare competes with private insurers for physicians' resources. Second, Medicare's payment menu provides the benchmarks from which bargaining begins. Bargaining costs and the expense of complex billing operations contribute to this role in establishing benchmarks and setting defaults. Finally, health-care payment systems have the essential properties of public goods; public payers may thus be essential participants in payment-system experimentation and reform. Improvements in our understanding of these sources of influence should prove valuable as policy makers reckon with the high cost of health care and aim to improve delivery.

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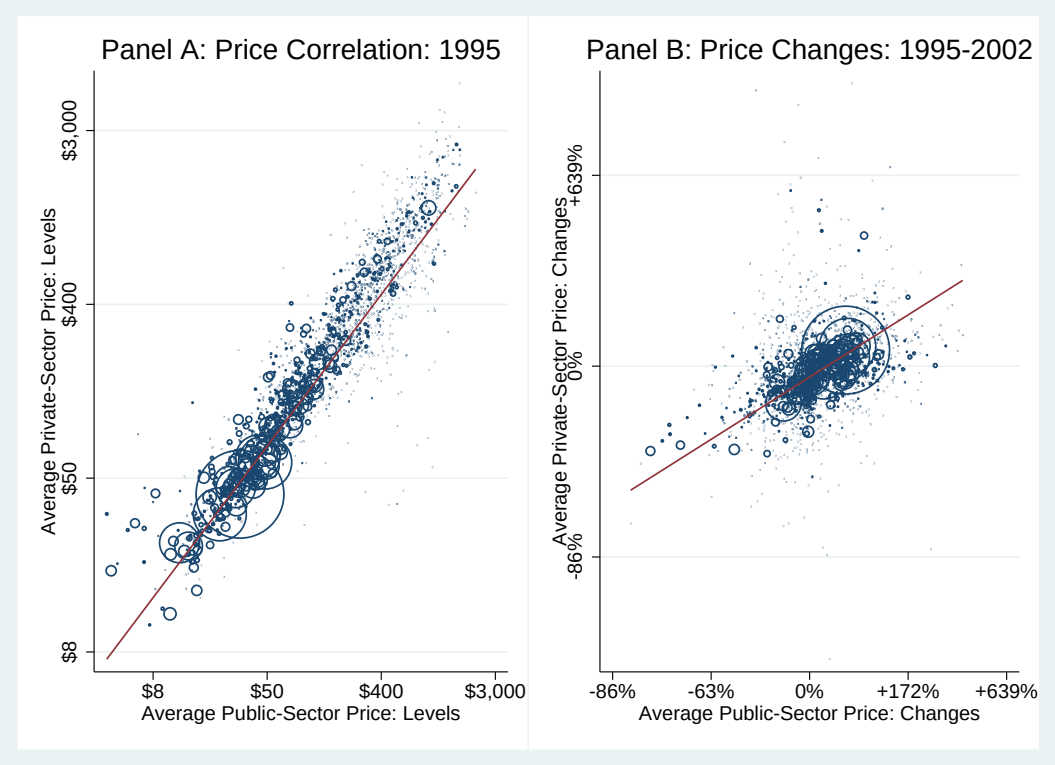
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Figure 1: Cross-Service Relationship Between Private and Medicare Prices



Note: This figure shows the raw cross-service relationships between average private reimbursements and average Medicare reimbursements. The 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 the number of times a code is observed in the Medicare data. The best-fit line shown in Panel A results from estimating

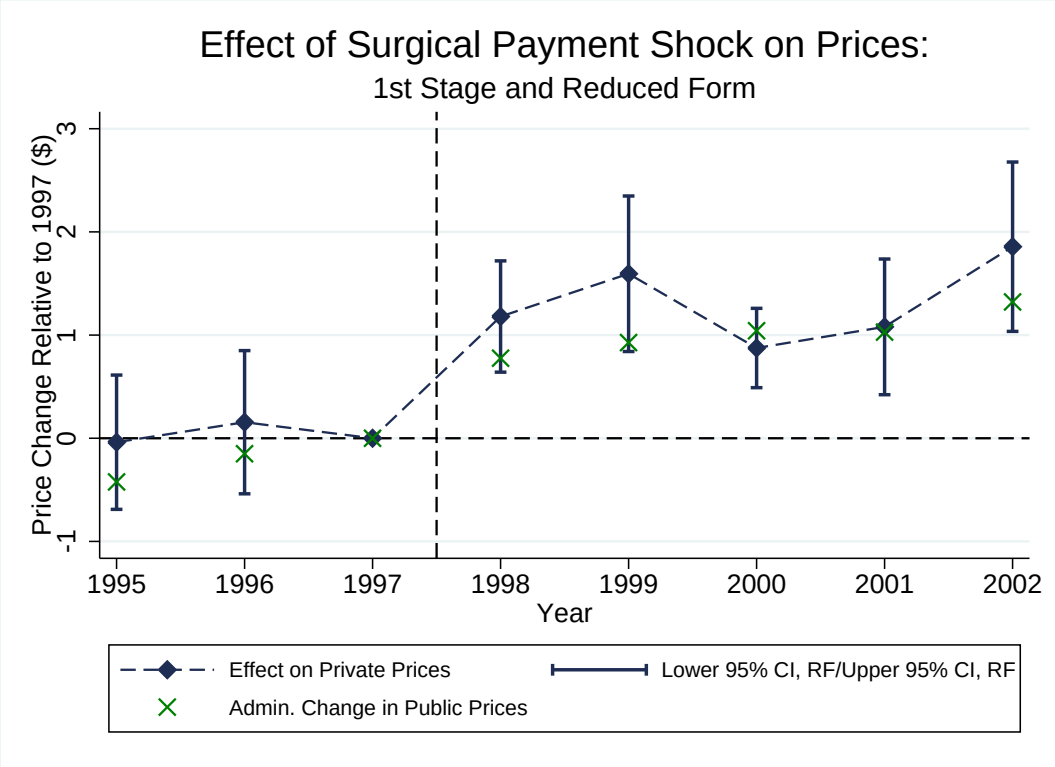
$$\ln(P_j^{\text{Private}}) = \beta_0 + \beta_1 \ln(P_j^{\text{Medicare}}) + u_j$$

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 2: Effects of Medicare's Elimination of the Surgical Conversion Factor

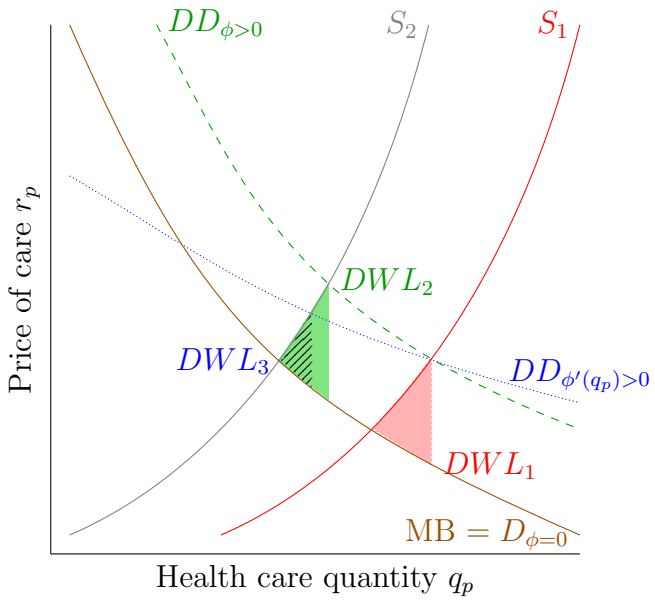


Note: This figure presents the δ_t coefficients, with associated 95% confidence intervals, from estimates of the equation below:

$$P_{j,s,t}^{\text{Private}} = \sum_{t \neq 1997} \delta_t \cdot I_t \times \text{PredChg}_j^{\text{Medicare}} \times \text{Post1998}_t + X_{j,s,t}\alpha + \mu_j I_j + \mu_s I_s + \mu_t I_t + \mu_{j,s} I_j \times I_s + \mu_{t,s} I_t \times I_s + u_{j,s,t}$$

where $\text{PredChg}_j^{\text{Medicare}}$ are the predicted changes in Medicare payments associated with the elimination of the surgical conversion factor. The figure also plots the point estimates from the associated first stage, showing that our coding of $\text{PredChg}_j^{\text{Medicare}}$ 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 ($I_j \times I_s$) and state-by-year ($I_s \times I_t$) 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 3: Private Health Care Market With a Distortion

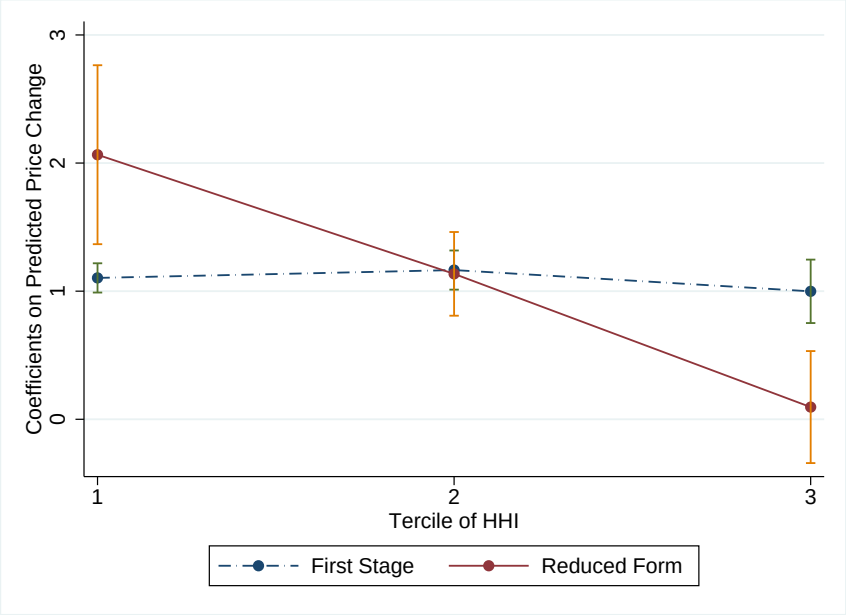


This figure illustrates the equilibrium and deadweight losses in the private market with and without a distortion ϕ. The true marginal benefit (MB) curve for health care is denoted $D_{\phi=0}$. One possible distortion is a demand curve exhibiting a constant shift relative to the MB curve. This is illustrated with the dashed curve, denoted $DD_{\phi>0}$. As this curve is above the MB curve, it could be illustrating a situation where the tax exclusion for employer-provided insurance is the dominant distortion. When the supply curve is S_1, the distortion leads to a deadweight loss denoted DWL_1. If the private sector supply curve moves to S_2, perhaps because of increased Medicare reimbursement rates, the new deadweight loss is given by the full DWL_2 area (both the dashed and solid portions) and is as large as DWL_1.

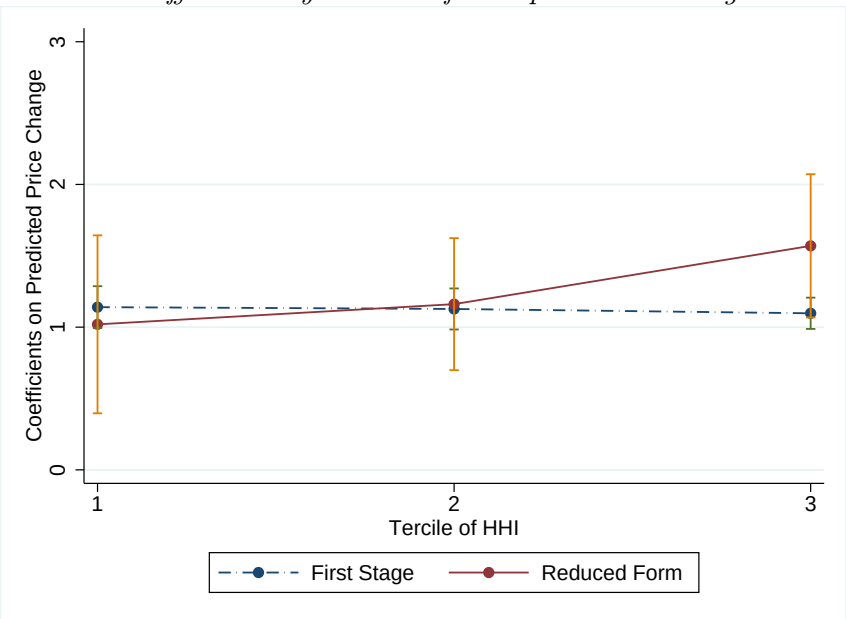
Now consider a distorted demand curve when $\phi'(q_p) > 0$. This variable distortion could arise if adverse selection dominates when quantities are low and excess moral hazard when they are high. In this case, illustrated in the dotted curve denoted $DD_{\phi'(q_p)>0}$, the demand curve is more elastic and deadweight loss is lower under supply curve S_2 (the dashed area marked as DWL_3) than S_1 (where the deadweight loss remains DWL_1). If the supply curve shifted from S_1 to S_2 because of a Medicare price increase, cost-following is lower under this distorted demand curve than under either the $D_{\phi=0}$ or $DD_{\phi>0}$ demand curves.

Figure 4: First Stage and Reduced Form Coefficients by Tercile of Competition

Panel A: Coefficients by Tercile of Competition Among Physicians

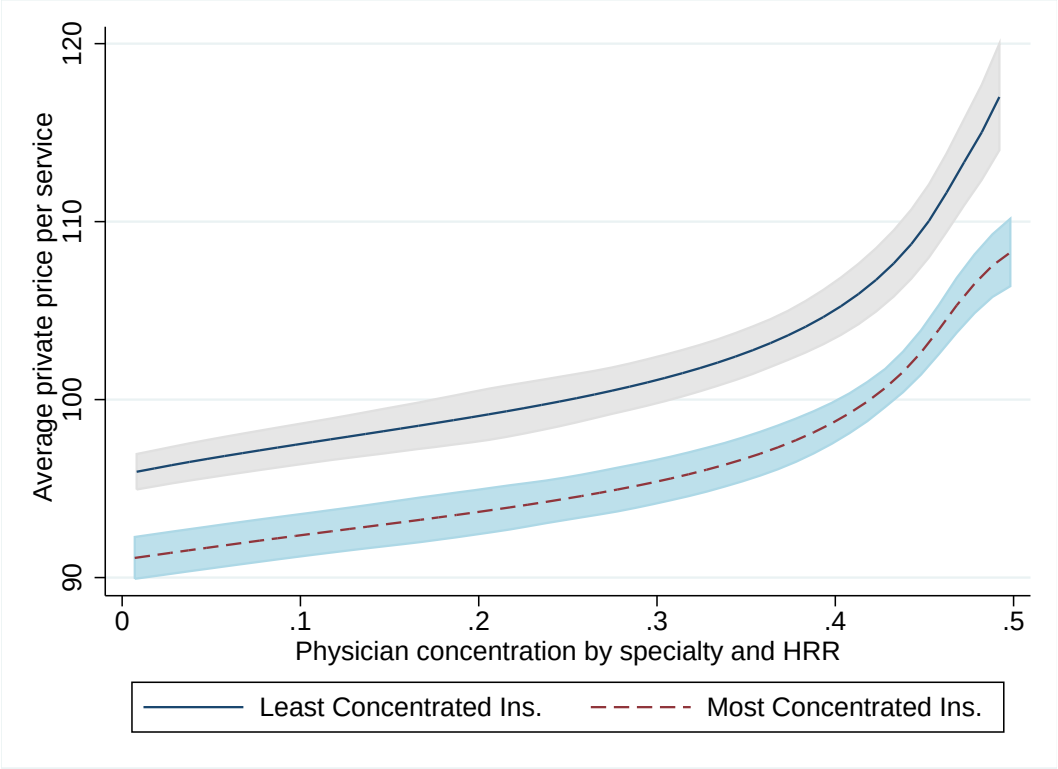


Panel B: Coefficients by Tercile of Competition Among Insurers



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 (8), 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, taken from American Medical Association (2007). 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.

Figure 5: Variation in Private Prices with Provider and Insurer Market Power



Note: This figure shows the relationship between average private sector payments between low-concentration (blue solid line) and high-concentration (red dashed line) insurance markets, with variation shown by the degree of provider concentration (x-axis). The private payments are averaged across all years, states, and services.

Table 1: **Summary Statistics**

		Non-Surgical Care ($N = 140,716$)			Surgical Services ($N = 163,012$)		
		Mean	Std. Dev.	Range	Mean	Std. Dev.	Range
1)	Private Payment Per Service	\$125.17	\$133.35	(\$1.10, \$6,967)	\$374.48	\$441.33	(\$0.30, \$21,891)
2)	Medicare Payment Per Service	\$114.44	\$148.39	(\$3.28, \$1,150)	\$239.23	\$248.05	(\$3.51, \$2,112)
3)	Std. Dev. of Private Pmt./Svc.	\$84.07	\$127.61	(\$0, \$9,503)	\$257.76	\$403.80	(\$0, \$34,634)
4)	Std. Dev. of Medicare Pmt./Svc.	\$17.60	\$27.03	(\$0.41, \$260.65)	\$78.05	\$106.09	(\$0.20, \$1,320)
5)	Insurance Plan Type Control	83.06	43.41	(15.52, 586.00)	237.43	142.34	(19.06, 670.95)
6)	Out of Pocket Share	0.226	0.188	(0.00, 1.00)	0.193	0.180	(0.00, 1.00)
7)	State Level Physician HHI	0.023	0.024	(0.003, 0.14)	0.022	0.020	(0.003, 0.14)
8)	Physician Specialty HHI	0.157	0.100	(0.007, 0.79)	0.147	0.109	(0.007, 1.00)
9)	Private Market Volume (1000s)	61.6	309.61	(0.01, 4898.08)	11.43	46.57	(0.01, 721.61)
10)	Medicare Relative Size	7.28	20.94	(0.003, 459.83)	2.66	4.62	(0.015, 136.87)
11)	State Level Insurer HHI (NAIC)	0.258	0.187	(0, 0.95)	0.227	0.167	(0, 0.95)
12)	State Level Insurer HHI (AMA)	0.335	0.135	(0.15, 0.69)	0.324	0.122	(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 pairing is only included if public and private prices could be estimated 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 service-by-state-by-year cells. The construction of "Plan Type Payment Generosity" and "Out of Pocket Share" is described in section 2.4. "State Level Physician HHI" is a physician market Herfindahl-Hirschman Index (HHI) constructed using information, from a 20 percent sample of Medicare claims for 1999, by aggregating physicians under groups associated with a common tax identification number; the measure was first constructed at the level of Hospital Referral Regions (HRRs), then averaged across the HRRs within each state. "Specialty-Specific Physician HHI" is similar to "State Level Physician HHI," but varies within each state at the level of 32 distinct physician specialties. This variable is only constructed for services that are provided predominantly by a single specialty, resulting in fewer observations than are associated with other variables described in the table. "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. 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. 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**

Dependent Variable:	(1)	(2)	(3)	(4)	(5)	(6)
	Payment Levels			Log Payments		
	Public 1st Stage	Private Red. Form	Private IV	Public 1st Stage	Private Red. Form	Private IV
Payment Shock × Post 1997	1.120**	1.292**				
	(0.066)	(0.251)				
Public Payment			1.154**			
			(0.227)			
Surgical Procedure × Post 1997				-0.225**	-0.108**	
				(0.032)	(0.027)	
Ln(Public Payment)						0.480**
						(0.062)
Plan Type Control	0.074+	0.009	-0.026	6.907	-3.453	-11.421
	(0.041)	(0.021)	(0.028)	(9.000)	(22.307)	(28.123)
Cost Sharing Fraction	-0.008	-0.005	-0.001	-0.977+	-3.441	-2.314
	(0.006)	(0.020)	(0.020)	(0.573)	(6.012)	(6.142)
N	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

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 2.2. Columns 1 and 2 report estimates of equations (8) 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 (9). 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-year level. In columns 1 through 3, observations are weighted according to the number of times the service is observed in Medicare Part B 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 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.

Table 3: Heterogeneity in Surgical CF Shock's Effect by Provider Concentration

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable:	Private Payment Level					
Payment Shock × Post-1997	1.327**	1.402**	1.103**	1.193**	1.230**	1.309**
	(0.250)	(0.291)	(0.153)	(0.191)	(0.193)	(0.235)
Payment Shock × Post-1997	-0.507**	-0.471**				-0.488**
× Physician HHI	(0.077)	(0.076)				(0.074)
Payment Shock × Post-1997			-0.883**	-0.840**	-0.793**	-0.494**
× Specialty HHI			(0.238)	(0.237)	(0.230)	(0.168)
Payment Shock × Post-1997		0.353**		0.369**		0.064
× Physician Count		(0.074)		(0.080)		(0.122)
Payment Shock × Post-1997					0.563**	0.479*
× Specialty Count					(0.133)	(0.201)
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
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
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 of the form described by equation (21) in section 5.1. 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. 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 4: 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	1.209**	1.297**	1.282**	1.208**	1.223**	1.069**
	(0.206)	(0.250)	(0.244)	(0.207)	(0.206)	(0.132)
Payment Shock × Post-1997	0.364*	0.363*	0.157	0.411	0.345*	0.412**
× Insurer HHI	(0.161)	(0.270)	(0.101)	(0.275)	(0.159)	(0.139)
Payment Shock × Post-1997				0.067		
× Insurer Count, 2001				(0.235)		
Payment Shock × Post-1997					-0.505**	-0.118
× Physician HHI					(0.074)	(0.137)
Payment Shock × Post-1997						-0.773**
× Specialty HHI						(0.290)
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	2,194
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 (21) as modified in section 6.1. 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. 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 5: Heterogeneity in Surgical CF Shock's Effect by Service Market Characteristics

	(1)	(2)	(3)
Dependent Variable:	Private Payment Level		
Payment Shock × Post-1997	0.916**	1.192**	1.436**
	(0.126)	(0.332)	(0.236)
Payment Shock × Post-1997	1.283**		
× Public-Private Ratio	(0.462)		
Payment Shock × Post-1997		0.354	
× Medstat Volume		(0.525)	
Payment Shock × Post-1997			-0.517**
× Specialty Income Exposure			(0.181)
N	303,728	303,728	234,800
Number of Clusters	2,194	2,194	1,261
Weighted	Yes	Yes	Yes
State By Year FE	Yes	Yes	Yes
HCPCS By State FE	Yes	Yes	Yes
Fully Interacted	Yes	Yes	Yes
Eye Procedure Reductions	Yes	Yes	Yes
Plan Type Controls	Yes	Yes	Yes
Panel Balanced	Yes	Yes	Yes
Other Sample Restrictions	None	None	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 (21) in section 5.1. 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 log of 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 section, we present practitioner characterizations of negotiations between physicians and insurers. The characterizations come largely from physicians and consultants who work as physicians' representatives 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 were regularly emphasized, namely the importance of the Medicare fee schedule and the importance of market power. Below we present consultants' characterizations of each.

A.1 The Role of Medicare's Fee Schedule

Practitioner characterizations 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).
- "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).

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 Analysis of Additional Payment Shocks

In this section, we present complementary analysis of an additional source of payment shocks. Recall the formula characterizing Medicare's payments for physician services, which we reproduce below. For service j, supplied by a provider in payment area i, the provider's fee is approximately:

$$\begin{aligned}\text{Reimbursement}_{i,j,t} = \text{Conversion Factor}_{t,c(j)} \times \text{Relative Value Units}_{j,t} \\ \times \text{Geographic Adjustment Factor}_{i,t}.\end{aligned} \tag{B.1}$$

The Conversion Factor (CF) is a national adjustment factor, updated annually and generally identical across broad categories of services, $c(j)$. The Relative Value Units (RVUs) associated with service j are intended to measure the resources required to provide that service; the normalization of units is such that a brief office visit amounts to roughly a single RVU. RVUs are constant across areas while varying across services. The RVUs associated with each service are updated on a rolling basis to account for technological and regulatory changes that alter their resource intensity. Finally, the Geographic Adjustment Factor (GAF) is the federal government's adjustment for differences in input costs across payment regions. The adjustments are derived from Census and other data on area-level rents, wages, and malpractice insurance premiums. In summary, the payment for a service depends on its resource-intensity (RVUs), a local price index (the GAF), and program-wide budgetary limits (expressed through the CF).

In the main text we analyzed price changes driven by the elimination of separate conversion factors for surgical procedures and other forms of care. This resulted in a 17 percent reduction in the surgical payments relative to other payments, with substantial cross-service variation in changes in the dollar value of Medicare's payments. Here we analyze shocks associated with the system of geographic adjustments.

The main text's focus on the shock associated with the conversion factor is driven by its suitability for assessing this paper's central questions. Its virtues include its size, with 17 percent constituting a massive change in relative payments across broad classes of services, as well as its make-up and motivation, which are tightly related. The elimination of the surgical conversion factor was motivated by the concern, echoed in recent policy discussions, that payments for surgical procedures were excessive, and that the returns to primary care and other medical services needed to be improved. The changes in geographic adjustments are conceptually distinct in that, within each geographic area, they were implemented on an across-the-board basis. Importantly, these payment changes are thus well-suited for addressing the cost-shifting hypothesis. They are not well suited, however, for estimating the likely effect of Medicare moving from cost-based payments towards value-based payments.

B.1 Price Variation from Payment Region Consolidation

We begin our analysis of supplemental payment shocks using variation driven by an administrative shift in the system of geographic adjustments. These are the same payment

shocks used in Clemens and Gottlieb (2013), from which we quote liberally and from which we have reproduced the maps in Appendix Figure B.1.

> In 1997, the Health Care Financing Administration consolidated the payment regions in many states, leading to reimbursement rate shocks that vary across the pre-consolidation regions. The 210 payment areas that existed as of 1996 were consolidated to 89 distinct regions, as shown in Appendix Figure B.1. The top panel of Appendix Figure B.1 presents the regions as of 1996, with darker colors indicating higher GAFs; the middle panel shows the post-consolidation payment regions. As the maps indicate, the consolidation of payment regions dramatically changed the county groupings in many states, leading to differential price shocks.

As in the analysis in the main text, the parameter of interest is a scalar mark-up relative to Medicare's payments. We thus express the payment changes in dollar terms by multiplying the changes in the geographic indices by the average pre-consolidation payment associated with the services in the sample. We denote the resulting, area-specific shocks Payment Shock$_a$. We proceed to estimate

$$P_{j,a,t}^{\text{Medicare}} = \sum_{p(t)\neq 1996} \beta_t \cdot \text{Payment Shock}_a \times I_t + \gamma_{j,a} + \gamma_{s(a),t} + \zeta' X_{a,s(a),t} + \varepsilon_{j,a,t}, \tag{B.2}$$

where I_t is an indicator for observations from year t $\gamma_{j,a}$ are a set of service type-by-payment area effects and $\gamma_{s(a),t}$ are a set of state-by-year effects. 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. The state-by-year effects subsume standard year effects. 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 other changes to payment policies and the structure of medical care that took place during this time period. We can further control for characteristics of the payment areas, such as the extent to which they are rural or urban, with little impact on the results presented below.[44]

The coefficients of interest are the β_t. Estimates of β_t for years prior to 1996 provide a sense for the importance of pre-existing trends. Estimates of β_t for years following 1996 trace out the effects of the payment shocks. In equation (B.2) above, which is essentially a first stage, coefficients of 0 prior to 1996 and of 1 following 1996 would indicate that we have effectively picked up the policy of interest. When we turn to private privates by estimating equation (B.3), written out below, the β_t become estimates of the effect of Medicare's payment changes on private prices.

[44] As illustrated in the results below, the consolidation-induced payment shocks were not correlated with pre-existing trends in private prices. This is not the case in the context of care utilization, as emphasized in Clemens and Gottlieb (2013).

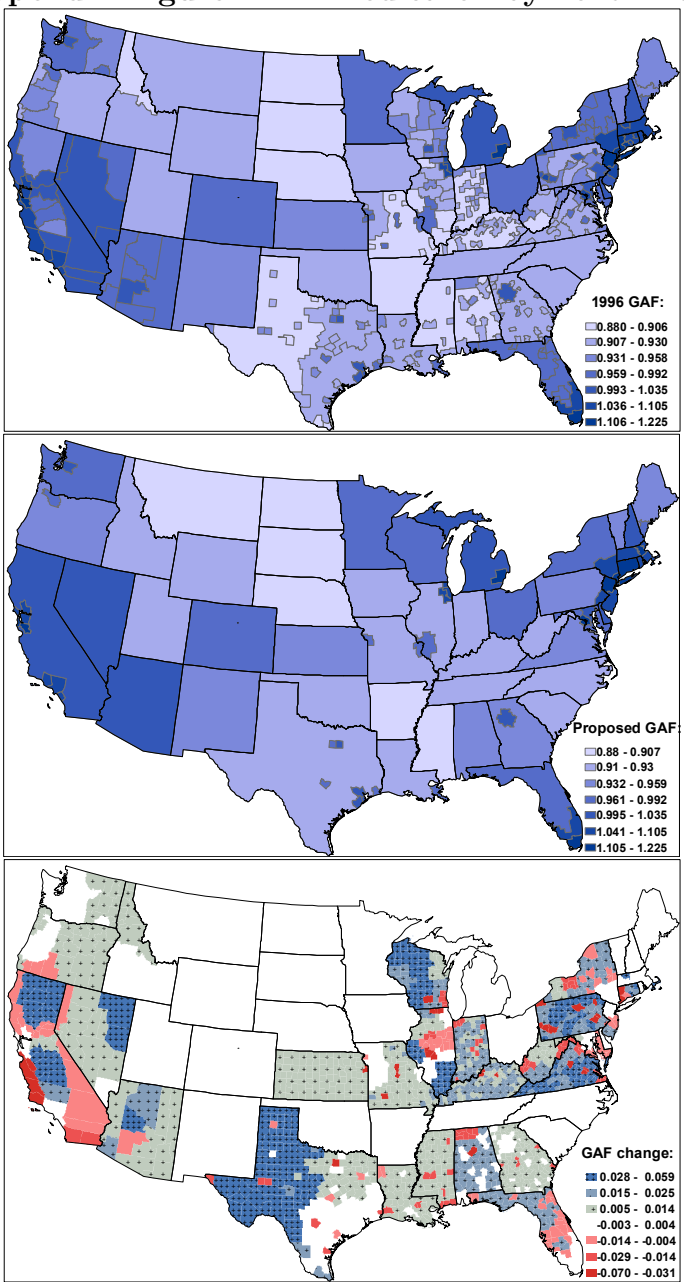
$$P_{j,a,t}^{\text{Private}} = \sum_{p(t)\neq 1996} \beta_t \cdot \text{Payment Shock}_a \times I_t + \delta_{j,a} + \delta_{s(a),t} + \zeta' X_{a,s(a),t} + \varepsilon_{j,a,t}. \quad \text{(B.3)}$$

B.2 Effects of Across-the-Board Payment Changes

The results of estimating equations (B.2) and (B.3) appear in Appendix Figure B.2 and Appendix Table B.1. The figure shows both the first stage and reduced form estimates. The first stage estimates show that our coding of the payment shocks has effectively tracked the policy change as it was meant to be implemented. A one unit increase in the payment shock is associated with a one dollar increase in Medicare's allowed charge for each service. The reduced form estimates plot out the private sector response to these public payment changes. In contrast with the results associated with the change in relative prices for surgical and non-surgical services, the effect of these across-the-board payment changes appears to unfold over a couple of years. The end result, however, is indistinguishable. An increase in public payments is associated with decreases in private payments, and vice versa. Averaging across the point estimates associated with years after 1996, the estimates suggest that a one dollar increase in the public payment is associated with an increase in private payments of just over one dollar.

Appendix Table B.1 condenses the result of interest into a single coefficient by shifting from parametric event study specifications to a more standard 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, controlling additionally interactions between year dummy variables and proxies for the extent to which the localities are rural or urban, and replacing the full set of state-by-year effects with year effects alone (the state fixed effects are subsumed by payment locality effects). While precision falls substantially in the last of these specifications, the results are similar throughout.

Appendix Figure B.1: 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: *Federal Register*, various issues.

Appendix Figure B.2: Effect of Geographic Payment Shocks on Private Prices

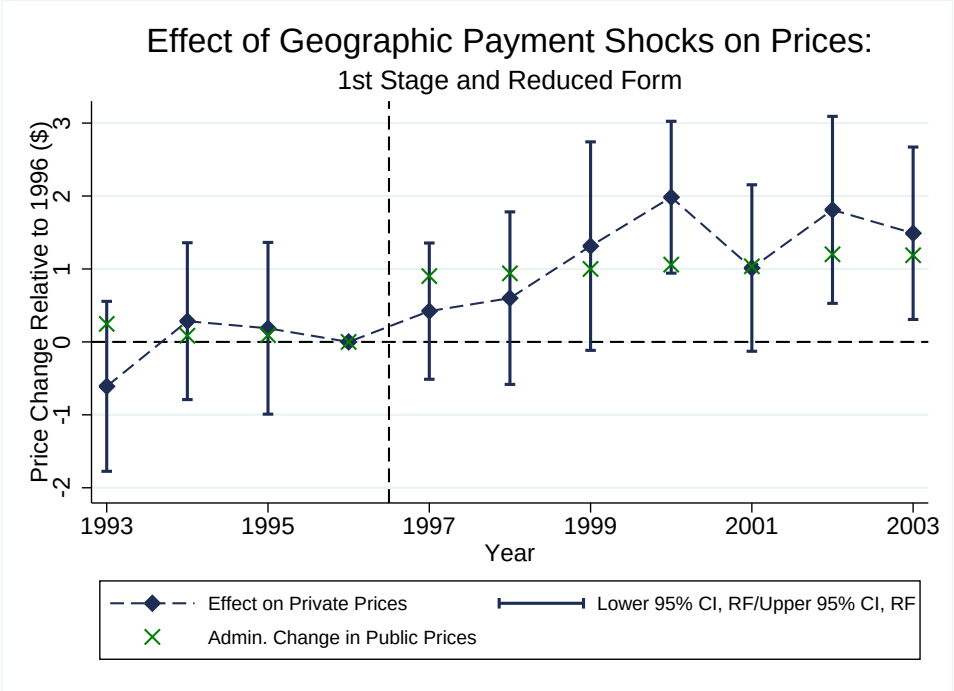


Figure shows the the results from estimating equations (B.2) and (B.3) as described in Appendix B.1. The payment shocks are constructed such that a one unit change in the payment shock should correspond to a one dollar increase in Medicare's payments. This is confirmed by the point estimates labeled "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).

Appendix Table B.1: 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 (B.3). Observations are constructed at the service-by-payment locality level. The panel is balanced in the sense that each service-by-payment locality pairing is only included if public and private prices could be estimated 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).

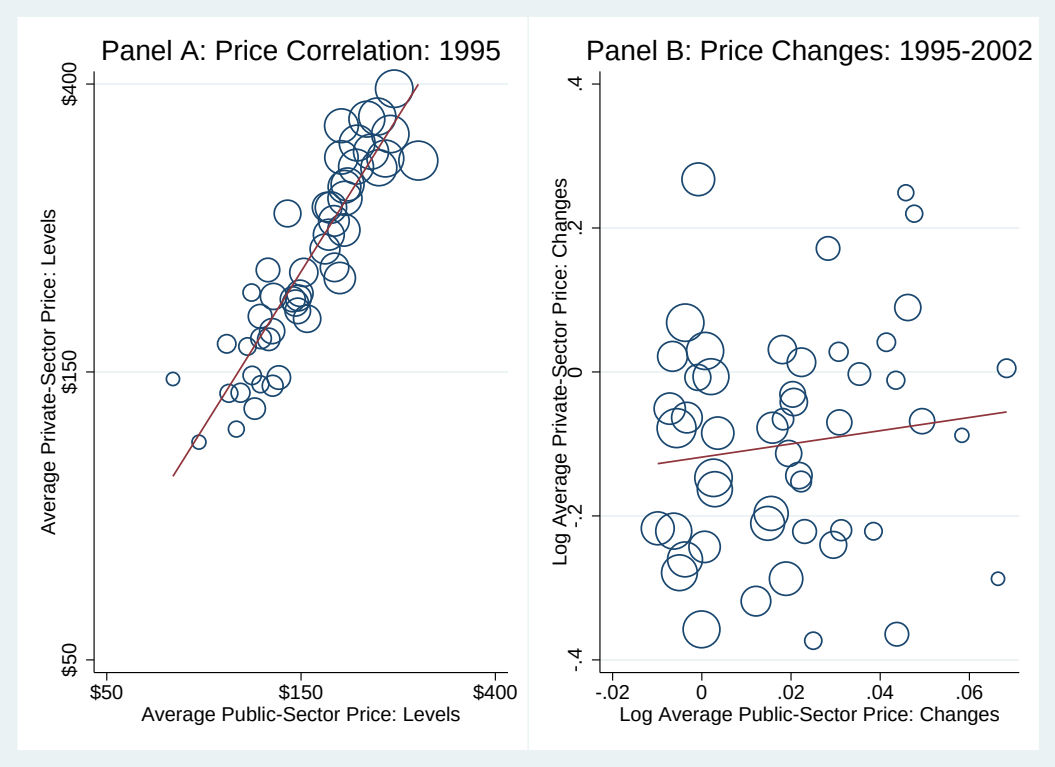
C Further Results and Robustness Tests

C.1 Robustness Checks

Appendix Tables C.1 through C.5 report various robustness checks as discussed in the main text, including cost-following measurements using nationally aggregated data.

C.2 Cross-Sectional Private and Public Pricing

Appendix Figure C.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.

C.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.

Table C.6 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 C.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	1.154**	1.149**	1.153**	0.968**	0.670**	1.311**	1.106**
	(0.227)	(0.213)	(0.227)	(0.108)	(0.111)	(0.263)	(0.251)
Plan Type Controls	-80.618		-11.357	-32.341	22.711**	-3.010	-11.779
	(78.781)		(28.178)	(40.440)	(4.758)	(17.838)	(27.982)
Cost Sharing Fraction	-2.314	-2.257		-2.405	1.766**	-3.081	-2.397
	(6.142)	(6.211)		(6.024)	(0.190)	(5.957)	(6.123)
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 C.2: **The Effect of Medicare Price Changes on Private Sector Prices, National Regressions**

Dependent Variable:	(1)	(2)	(3)	(4)	(5)	(6)
	Payment Levels			Log Payments		
	Public 1st Stage	Private Red. Form	Private IV	Public 1st Stage	Private Red. Form	Private IV
Payment Shock × Post 1997	1.122**	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	73.469**	63.485+	-10.297	0.259*	0.057	-0.081
	(26.409)	(35.768)	(55.927)	(0.127)	(0.066)	(0.096)
Cost Sharing Fraction	-8.986	-121.264+	-112.240+	-0.203	-0.389*	-0.280*
	(18.769)	(67.348)	(62.818)	(0.133)	(0.165)	(0.130)
N	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
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 C.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 C.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 C.4: 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	1.314**	1.201**	1.350**	1.332**	1.339**	1.305**
	(0.217)	(0.170)	(0.271)	(0.247)	(0.256)	(0.239)
Payment Shock × Post-1997	-0.438**	-0.401**	-0.532**	-0.495**	-0.538**	-0.470**
× Physician HHI	(0.066)	(0.051)	(0.088)	(0.079)	(0.098)	(0.067)
Payment Shock × Post-1997	-0.585**	-0.702**	-0.506**	-0.446**	-0.416**	-0.515**
× Specialty HHI	(0.221)	(0.271)	(0.186)	(0.164)	(0.144)	(0.174)
Payment Shock × Post-1997	0.112	0.039	-0.689+	-0.008	0.001	0.058
× Physician Count	(0.129)	(0.139)	(0.390)	(0.119)	(0.121)	(0.120)
Payment Shock × Post-1997	0.530+	0.218	0.525*	0.486*	0.464*	0.495*
× Specialty Count	(0.270)	(0.204)	(0.219)	(0.199)	(0.195)	(0.207)
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 3. 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 C.5: 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	1.247**	1.194**	1.331**	1.237**	1.218**	1.248**
	(0.199)	(0.203)	(0.270)	(0.218)	(0.208)	(0.226)
Payment Shock × Post-1997	0.287+	0.453*	0.664**	0.440**	0.466*	0.341*
× Insurer HHI	(0.154)	(0.194)	(0.217)	(0.162)	(0.183)	(0.168)
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 3. 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 C.6: 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 2.2, 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.

Appendix Table C.7: Insurer Concentration Meausred 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	1.108**	1.151**	1.174**	1.119**	1.273**	1.317**
	(0.215)	(0.177)	(0.205)	(0.149)	(0.239)	(0.221)
Payment Shock × Post-1997	0.172	0.475**	0.399**	0.637**	0.042	-0.047
× Insurer HHI	(0.295)	(0.151)	(0.120)	(0.191)	(0.089)	(0.155)
N	286,568	293,688	293,688	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 3. 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.

D Further Details on the Model of Section 4

D.1 Step 1: Determining the Average Reimbursement

This appendix elaborates on the model from section 4. Recall that the model features two modes of negotiation, taking place in two distinct steps. First, the insurer and physician group settle on an average markup over Medicare rates. Then they negotiate service-specific relative prices to improve the quality of care provided. In each step, we make a number of simplifications away from a full model of multilateral insurer-physician bargaining. We do this for both theoretical and empirical reasons. In the actual setting we consider, each market has many insurers and many physician groups. The literature on multilateral bargaining has shown that a plethora of solutions can arise, depending on specifics such as the timing and information structure of the game (Cai, 2000; Guo and Iyer, 2012; Krishna and Serrano, 1996; Stole and Zwiebel, 1996). But in our context, these complications may not be terribly problematic. Dunn and Shapiro (2012) show that prices increase with physician market power, exactly as intuition would suggest. In a closely related setting, the literature on hospital price negotiations has also found evidence for intuitive consequences of both physician and provider market power (Town and Vistnes 2001, Dafny 2005, Ho 2009, Moriya et al. 2010, Gowrisankaran et al. 2013).

The second reason we adopt a stylized approach rather than estimate a comprehensive bargaining game is data limitations. Unlike in many hospital-based datasets, our data offer no indication of the specific insurance company responsible for an insured patient or a claim (even a masked identifier). We also cannot match physicians across insurers or observe a physician practice's overall size. So we consider here a stylized model of these negotiations.

We begin with the negotiations over the overall price level, from section 4.1. To recap, the physician's reservation value is

$$\overline{u}^g = \lambda_M^g \overline{r}_M + \sum_{\iota \neq i} \lambda_\iota^g \overline{r}_\iota^g \tag{D.1}$$

where ι indexes all insurers other than the one involved in the current negotiation, λ_ι^g is insurer ι's market share, and $\overline{r}$ represents the average reimbursement rate across services (weighted by their shares).

We define the insurer's reservation value as an exogenous parameter $\underline{u}_i$. We assume that the insurer's Nash bargaining parameter is θ, and the physician group's is $1-\theta$. The parties will thus settle on a reimbursement rate of

$$\overline{r}_i^g = \theta \left(\lambda_M^g \overline{r}_M + \sum_{\iota \neq i} \lambda_\iota^g \overline{r}_\iota^g \right) + (1-\theta)\underline{u}_i. \tag{D.2}$$

When we average this across all insurers, and then differentiate with respect to the Medicare

rate, we find that the responses of all insurers together must satisfy:

$$\sum_i \left[\lambda_i^g(1-\theta+\theta\lambda_M^g)+\theta\left(\lambda_i^g\right)^2\right]\overline{r}_i^g = \theta\lambda_M^g(1-\lambda_M^g)\overline{r}_M + (1-\theta)\sum_i \lambda_i^g \underline{u}_i \tag{D.3}$$

$$\sum_i \left\{\lambda_i^g\left[1-\theta\left(1-\lambda_M^g\right)\right]+\theta\left(\lambda_i^g\right)^2\right\}\frac{\partial \overline{r}_i^g}{\partial r_M^g} = \theta\lambda_M^g(1-\lambda_M^g). \tag{D.4}$$

The cost-following is defined implicitly by equation (D.4).

To make the price behavior more transparent, consider first a symmetric equilibrium of I insurers, each with $\underline{u}_i = \underline{u}$. In this case, equation (D.3) reduces to:

$$N\left[\overline{\lambda}(1-\theta+\theta\lambda_M^g)+\theta\overline{\lambda}^2\right]\overline{r}^g = \theta\lambda_M^g(1-\lambda_M^g)\overline{r}_M + I(1-\theta)\overline{\lambda}\underline{u} \tag{D.5}$$

where $\overline{\lambda} = \frac{1-\lambda_M^g}{I}$ is each symmetric insurer's market share and $\overline{r}$ is the (again symmetric) reimbursement rate. We solve this for $\overline{r}^g$:

$$\overline{r}^g = \frac{\theta\lambda_M^g\overline{r}_M + (1-\theta)\underline{u}}{1-\theta+\theta\lambda_M^g+\theta\frac{1-\lambda_M^g}{I}}. \tag{D.6}$$

In the case where I is large, the final term in the denominator disappears and equation (D.6) is just a weighted average of the outside options, where Medicare's importance is reduced relative to the insurers' outside option based on its market share λ_M. As the number of insurers I falls, the symmetric reimbursement rate falls, consistent with less competition among insurers. Mathematically, this is due to less reflection (in the Manski [1993] sense) as we move away from a continuum of small insurers, because insurer i's price doesn't reflect itself in equation (D.2) and it has mass $\frac{1-\lambda_M^g}{I}$. In the symmetric case, average cost-following is:

$$\frac{\partial\overline{r}^g}{\partial\overline{r}_M} = \frac{\theta\lambda_M^g}{1-\theta+\theta\lambda_M^g+\theta\frac{1-\lambda_M^g}{I}} \tag{D.7}$$

which is increasing in the insurer's bargaining power and in Medicare's size.

To determine how heterogeneity among insurers affects this result, we subtract equation (D.5) from equation (D.3):

$$(1-\theta+\theta\lambda_M^g)\sum_i\left[\lambda_i^g\overline{r}_i^g-\overline{\lambda}\overline{r}^g\right]+\theta\sum_i\left[\left(\lambda_i^g\right)^2\overline{r}_i^g-\overline{\lambda}^2\overline{r}^g\right] = (1-\theta)\sum_i\left[\lambda_i^g\underline{u}_i-\overline{\lambda}\underline{u}\right]. \tag{D.8}$$

Hence

$$\frac{1}{1-\lambda_M^g}\sum_i \lambda_i^g \overline{r}_i^g = \overline{r}^g - \frac{\theta}{(1-\lambda_M^g)(1-\theta+\theta\lambda_M^g)}\sum_i \left[(\lambda_i^g)^2\,\overline{r}_i^g - \overline{\lambda}^2\overline{r}^g\right] + \frac{1-\theta}{(1-\lambda_M^g)(1-\theta+\theta\lambda_M^g)}\sum_i \left[\lambda_i^g \underline{u}_i - \overline{\lambda}\underline{u}\right]. \tag{D.9}$$

The average reimbursement observed in this case deviates from the symmetric one given in (D.6) by terms related to the covariance of insurer size with payment generosity or outside options. The second term on the right shows that a higher insurer Herfindahl reduces payments less than a simple average would suggest if larger insurers pay less than smaller ones.

D.2 Step 2: Negotiating Away From Medicare's Default

In step 2, the insurer and physician group can negotiate away from the default relative prices set by Medicare. They do this based on the benefits from quality improvements in each service. This section motivates the model of these choices presented in section 4.2 and derives the results presented there.

The representative consumer has preferences over insurance companies and the composite commodity C. These preferences take into account the quality as defined in equation 18, namely

$$\Xi_i(\boldsymbol{N}_i) = \prod_{g=1}^{G}\left(1 + N_{ig}\Lambda_g\xi_g\right). \tag{D.10}$$

Insurer i's quality depends on whether it negotiates with each of the G groups indexed by g. N_{ig} is an indicator for its choice to negotiate with this group, and deviate from a constant markup over Medicare. $\boldsymbol{N}_i$ is simply the vector of such choices. Groups vary according to the share of the insurer's patients who use the group, $\Lambda_g \geq 0$, and the value to consumers of achieving the efficient reimbursement schedule, $\xi_g \geq 0$.

Given this quality, consumers consume the composite commodity C and ι_i units of insurance from each company i. These companies lie along a continuum from 0 to 1, and preferences are CES:

$$U(\boldsymbol{\iota}, C) = C^{1-\omega}\left[\int_0^1 \zeta_i^{\frac{1}{\epsilon}}\Xi_i\iota_i^{\frac{\epsilon-1}{\epsilon}}di\right]^{\frac{\omega\epsilon}{\epsilon-1}}$$

Insurer i has a fee of f_i and C is the numeraire. Given the immense concentration in insurance markets shown in Table 1, and studied in depth by the existing literature (*e.g.* Cutler and Reber 1998, Dafny et al. 2012), assuming a continuum of small non-strategic insurers is clearly counterfactual. But we cannot identify individual insurers in our empirical work, so in our setting there would seem to be little advantage to the extra complexity that would

arise when considering strategic behavior. This stylized model is intended to offer some straightforward comparative statics on the decision to accept Medicare's default prices, or negotiate service-specific pricing, so we proceed with this tractable approach.

Assume that consumers have fixed income of Y, so the budget constraint is

$$C + \int_0^1 f_i \iota_i di \leq Y$$

Each insurer's demand is thus

$$\iota_i = \frac{\zeta_i \Xi_i^\epsilon}{f_i^\epsilon} \frac{\omega Y}{\tilde{F}^{1-\epsilon}} \tag{D.11}$$

where the ideal price index is

$$\tilde{F} = \left[\int_0^1 f_i^{1-\epsilon} \zeta_i \Xi_i^\epsilon di \right]^{\frac{1}{1-\epsilon}}.$$

Equation (D.11) is exactly equal to equation (17) in the text, where $K = \frac{\omega Y}{\tilde{F}^{1-\epsilon}}$ collects the constants.

To repeat equation (19) in section 4.2, insurer profits are:

$$\begin{aligned}
\pi_i(f_i, \boldsymbol{N}_i) &= (f_i - m_i) D_i(f_i, \boldsymbol{N}_i) - c_i \sum_g N_{ig} \\
&= (f_i - m_i) K \zeta_i \frac{\Xi_i (\boldsymbol{N}_i)^\epsilon}{f_i^\epsilon} - c_i \sum_g N_{ig} \\
&= K \zeta_i \Xi_i (\boldsymbol{N}_i)^\epsilon (f_i^{1-\epsilon} - m_i f_i^{-\epsilon}) - c_i \sum_g N_{ig}.
\end{aligned}$$

For any given quality level it chooses, the firm sets prices as a fixed markup over m_i, namely $f_i^* = \frac{\epsilon}{\epsilon - 1} m_i$. Letting $\tilde{\epsilon}_i = m_i^{1-\epsilon} \left[\left(\frac{\epsilon}{\epsilon-1} \right)^{1-\epsilon} - \left(\frac{\epsilon}{\epsilon-1} \right)^{-\epsilon} \right]$, we can rewrite the profit as

$$\pi_i(f_i^*, \boldsymbol{N}_i) = K \zeta_i \tilde{\epsilon}_i \Xi_i (\boldsymbol{N}_i)^\epsilon - c_i \sum_g N_{ig}.$$

Given this profit-maximizing price, the gain from negotiating with group g (expressed as $\boldsymbol{N}_i = \boldsymbol{N}_{i,g=1}$) relative to accepting Medicare's default ($\boldsymbol{N}_i = \boldsymbol{N}_{i,g=0}$) is thus:

$$\begin{aligned}
\pi_i(f_i^*, \boldsymbol{N}_{i,g=1}) - \pi_i(f_i, \boldsymbol{N}_{i,g=0}) &= K \zeta_i \tilde{\epsilon}_i \Xi_i (\boldsymbol{N}_{i,-g})^\epsilon \left(1 + \Lambda_g \xi_g\right)^\epsilon - c_i - K \zeta_i \tilde{\epsilon}_i \Xi_i (\boldsymbol{N}_{i,-g})^\epsilon \\
&= K \zeta_i \tilde{\epsilon}_i \Xi_i (\boldsymbol{N}_{i,-g})^\epsilon \left[\left(1 + \Lambda_g \xi_g\right)^\epsilon - 1 \right] - c_i.
\end{aligned}$$

This is positive iff

$$K\zeta_i\tilde{\epsilon}_i\Xi_i(\boldsymbol{N}_{i,-g})^\epsilon\left[(1+\Lambda_g\xi_g)^\epsilon-1\right]>c_i$$
$$(1+\Lambda_g\xi_g)^\epsilon>1+\frac{c_i}{K\zeta_i\tilde{\epsilon}_i\Xi_i(\boldsymbol{N}_{i,-g})^\epsilon}. \tag{D.12}$$

Let $\chi_g=\Lambda_g\xi_g$ be a summary index representing importance to the insurer of negotiations with g. We first show the monotonicity of inequality (D.12) in χ_g. That is, if the inequality is satisfied for a group g, it is also satisfied for every group k with $\chi_k>\chi_g$.

Assume that group g satisfies inequality (D.12):

$$(1+\chi_g)^\epsilon>1+\frac{c_i}{K\zeta_i\tilde{\epsilon}_i\Xi_i(\boldsymbol{N}_{i,-g})^\epsilon}. \tag{D.13}$$

Holding constant the decision regarding all services except for g and k, consider a service k with $\chi_k>\chi_g$. Suppose that, for this service k, inequality (D.12) is violated, so:

$$(1+\chi_k)^\epsilon\leq 1+\frac{c_i}{K\zeta_i\tilde{\epsilon}_i\Xi_i(\boldsymbol{N}_{i,-k})^\epsilon}. \tag{D.14}$$

Because we are holding all other services constant, and we know that the firm will negotiate over service g, we can rewrite $\Xi_i(\boldsymbol{N}_{i,-k})^\epsilon$ as

$$\Xi_i(\boldsymbol{N}_{i,-k})^\epsilon=\Xi_i(\boldsymbol{N}_{i,-g})^\epsilon(1+\chi_g)^\epsilon.$$

Using this equality in inequality (D.14) yields

$$(1+\chi_k)^\epsilon(1+\chi_g)^\epsilon\leq(1+\chi_g)^\epsilon+\frac{c_i}{K\zeta_i\tilde{\epsilon}_i\Xi_i(\boldsymbol{N}_{i,-g})^\epsilon}. \tag{D.15}$$

Since $\chi_k\geq 1$ and $\epsilon>0$, this implies

$$(1+\chi_k)^\epsilon(1+\chi_g)^\epsilon\leq(1+\chi_g)^\epsilon+(1+\chi_k)^\epsilon\frac{c_i}{K\zeta_i\tilde{\epsilon}_i\Xi_i(\boldsymbol{N}_{i,-g})^\epsilon}. \tag{D.16}$$

To compare this with inequality (D.17), multiply the latter by $(1+\chi_k)^\epsilon$ on each side:

$$(1+\chi_k)^\epsilon(1+\chi_g)^\epsilon>(1+\chi_k)^\epsilon+(1+\chi_k)^\epsilon\frac{c_i}{K\zeta_i\tilde{\epsilon}_i\Xi_i(\boldsymbol{N}_{i,-g})^\epsilon}. \tag{D.17}$$

This yields a contradiction, and establishes the monotonicity of inequality (D.12) in χ_g.

Given this monotonicity, let us re-order the physician groups g according to χ_g, from lowest to highest. The monotonicity of the negotiation decision implies the existence of a cutoff χ_g^* such that all services with $\chi_g>\chi_g^*$ will have negotiated prices. Note that nothing prohibits a corner solution, where no services or all services have negotiated prices. From

re-arranging inequality (D.12), we can determine the cutoff to be:

$$\chi_g^* = \left[1 + \frac{c_i}{K\zeta_i\tilde{\epsilon}_i\Xi_i(\boldsymbol{N}_{i,-g})^\epsilon}\right]^{\frac{1}{\epsilon}} - 1$$

where the g in $\boldsymbol{N}_{i,-g}$ refers to the service g with $\chi_g = \chi_g^*$. If there is no such service, then this can be taken to refer to the service with the highest χ_g below χ_g^*. This proves inequality (20) in the text and completes our characterization of the insurer's problem.

D.3 Step 3: Cost-Following Overall and Across Services

Based on the choices made in step 1 and step 2, we can predict the cost-following that will result both from a change in the overall level of Medicare's reimbursements and a service-specific change.

To express the result summarized in equation (20), we first define a service-specific index that governs the negotiation decision. Let $\chi_j = \Lambda_j\xi_j$, let $H(\cdot)$ be the distribution of χ_j, and let χ^* be the cutoff value at which equation (20) holds with equality. For services with $\chi_j \geq \chi^*$, we assume that prices are negotiated to their surplus-maximizing level $\hat{r}_{i,j}^g$. This is independent of Medicare's reimbursement schedule. But the average price $\overline{r}_i^g$ agreed upon in the first step has to be maintained. So a second scaling factor ϕ_i^g is applied uniformly across all services, to bring the average back to $\overline{r}_i^g$.[45] The reimbursement rate for service j with a negotiated price is thus $\phi_i^g\hat{r}_{i,j}^g$.[46] For a service k where Medicare's default is used, the price is $\phi_i^g\varphi_i^g r_{M,k}$.

These price outcomes point to an important distinction in understanding price transmission. When the overall Medicare price level changes, equation (15) governs the response of average private reimbursements. When relative prices change across services, the private response is more complex. Services that follow Medicare's benchmark will naturally reflect the Medicare change:

$$\forall j : \chi_j < \chi^* \qquad \frac{\partial r_{i,j}^g}{\partial r_{M,j}} = \phi_i^g\varphi_i^g + \left(\phi_i^g\frac{\partial\varphi_i^g}{\partial r_{M,j}} + \varphi_i^g\frac{\partial\phi_i^g}{\partial r_{M,j}}\right) r_{M,j} \tag{D.19}$$

This response has two terms. The immediate effect of the Medicare change is passed through

[45] Recalling that φ_i^g is the overall markup agreed by insurer i and group g, $r_{M,j}$ is the Medicare reimbursement for service j, and $\hat{r}_{i,j}^g$ is the optimal price for service j, the new scaling factor

$$\phi_i^g = \frac{\varphi_i^g\overline{r}_M}{\sum_{\{j:\chi_j\geq\chi^*\}}\lambda_j\hat{r}_{i,j}^g + \sum_{\{j:\chi_i<\chi^*\}}\lambda_j\varphi_i^g r_{M,j}} \tag{D.18}$$

ensures that the average reimbursement rate remains at the level agreed in the first step.

[46] We assume that each service is sufficiently small as a share of the physician's overall practice that its effect on the average markup ϕ_i^g is ignored when setting the price for j. In another data set, which contains a university of claims from one insurer (and is not otherwise used in this paper), we find that the average service is provided by a group that provides 523 services.

to the private insurer with the same constant of proportionality as the overall markup, $\phi_i^g \varphi_i^g$. The second term kicks in if the Medicare change is large enough to influence the markups themselves. A revenue-neutral Medicare change will not affect the average markups, unless it switches some services between negotiated and default prices.

In the case where Medicare's change affects the markups for other services, it can also influence those that are actively negotiated. Because the overall influences those services as well, they respond according to:

$$\forall j : \chi_i > \chi^* \qquad \frac{\partial r_{i,j}^g}{\partial r_{M,j}} = \frac{\partial \phi_i^g}{\partial r_{M,j}} \hat{r}_j. \tag{D.20}$$

These services' reimbursements reflect any changes in the markup that take place.

Finally, for services exactly at the point of indifference between negotiation and default pricing, there can be a a discontinuous change of $\pm \phi_i^g (\hat{r}_j - \varphi_i^g r_{M,j})$ as the change in Medicare's price increases or decreases the value of negotiations ξ_j, and hence moves the service across the $\Lambda_j \xi_j = \chi^*$ threshold.