In the Shadow of a Giant: Medicare's Influence on Private Physician Payments

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Abstract

We analyze Medicare's influence on private insurers' payments for physicians' services. Using a large administrative change in reimbursements for surgical versus medical care, we find that private prices follow Medicare's lead. A $1 increase in Medicare's fees increases corresponding private prices by $1.16. A second set of Medicare fee changes, which generates area-specific payment shocks, has a similar effect on private reimbursements. Medicare's influence is strongest in areas with concentrated insurers and competitive physician markets, consistent with insurer-doctor bargaining. By echoing Medicare's pricing changes, these payment spillovers amplify Medicare's impact on specialty choice and other welfare-relevant aspects of physician practices.

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1 Introduction

The United States spends 3.5 percent of GDP, or nearly $600 billion annually, on physician care and similar medical services.[1] The markets and public programs that allocate this care thus have substantial welfare implications (Chandra, Jena, and Skinner, 2011). When prices signal relationships between production costs and consumers' willingness to pay, they steer markets towards efficient outcomes. But most medical services are purchased through insurance, which can sever consumers from the price mechanism (Gaynor, Haas-Wilson, and Vogt, 2000; Baicker and Goldman, 2011).

We ask how physicians and private insurers determine the prices that insurers pay on their beneficiaries' behalf. In particular, we show that the payment rates set by Medicare, the federal insurer of the elderly and disabled, influence private insurers' payments. We also assess the economic forces likely driving this relationship.

We use two overhauls of Medicare's administrative payment mechanisms to overcome the concern that private and Medicare prices covary because of underlying productivity or demand shocks. Our central analysis exploits a sharp reduction to Medicare's payments for surgical procedures relative to nonsurgical services. We also examine across-the-board payment changes that differ by geographic location.

To study these changes, we construct a novel link between databases of Medicare and private sector claims. We construct a rich panel of public and private prices that vary across years, geography, and individual medical services. The services we consider are defined quite precisely. For example, a 20-minute office visit is distinct from a 30-minute office visit, and coronary artery bypass grafts (CABG) are counted differently depending on the number of grafts and whether arterial grafts are used in addition to venous grafts.

Exploiting the surgical payment reduction, we estimate that a $1 decrease in Medicare's

[1]This figure comes from the "Physician and Clinical Services" line of the National Health Expenditure Data for 2013 (Centers for Medicare and Medicaid Services (CMS), 2014). The non-physician part of this category includes freestanding outpatient clinics and some laboratories.

payment for a surgical service causes a \$1.16 decline in private payments for that service. This response emerges within one year of Medicare's administrative change. The private prices show no pre-trends prior to the Medicare payment changes. Exploiting a broad overhaul of geographic adjustments, we estimate that, over the medium to long run, a \$1 decrease in Medicare's fees induces a \$1.12 decrease in private payments. Private sector responses to these broad-based rate changes appear to unfold over several years.

To better understand the economic forces behind these results, we explore the characteristics of markets in which Medicare's rates have the largest effects. We find that Medicare's influence is particularly strong in areas with relatively concentrated insurance markets and with relatively competitive provider groups. Caution is warranted in interpreting this heterogeneity, as markets' baseline characteristics are not randomly assigned. Nonetheless, these results are robust to controlling flexibly for the relationships between Medicare's price changes and economic, demographic, and medical care market characteristics.

We interpret our findings through the lens of a bargaining framework in which Medicare's payment rates can affect physicians' outside options. Our baseline results and the heterogeneity across markets are consistent with such a model. Improvements in physicians' outside options enable them to bargain for higher private payments. Medicare's relevance is greatest when physicians can be kept close to their outside option, namely in markets with low physician concentration or high insurer concentration.

The results have direct implications for Medicare's effects on the returns to practicing in particular specialties or geographic areas. The surgical-medical payment change reduced surgeons' average effective wage by around 10 percent, while increasing it for general practitioners. Using nationally representative physician surveys, we find that these reimbursement changes are followed by declines in surgeons' willingness to take new patients. We also find that new medical school graduates become less likely to enter surgical specialties, in magnitudes consistent with prior work on specialty choice (Nicholson and Souleles, 2001). These estimates speak to Medicare's long-run influence on real resource allocation and, by extension, economic welfare.

Our findings counter the conventional wisdom in many health policy discussions. This conventional wisdom, often labeled "cost shifting," holds that reductions in Medicare's payment rates will be partially offset by private payment increases. The academic literature has considered this question almost exclusively in the context of hospitals, where cost-shifting may arise because of non-profits' behavior in the presence of high fixed costs (Dranove, 1988).[2] While cost shifting is theoretically less plausible in the context of physicians' practices, it has nonetheless been assumed in recent policy discussions.[3]

The paper proceeds as follows. In section 2 we present institutional background on the structure of Medicare and private insurance payments. We present a conceptual framework for thinking about Medicare's influence on private payments in section 3. We describe our data in section 4, empirical strategy in section 5, and results in section 6. Section 7 discusses welfare-relevant implications of our findings, and section 8 concludes.

2 Background

Public and private payments for health care services are set through very different mechanisms. In the physician setting we study, Medicare determines its payments using a centralized administrative apparatus. The Centers for Medicare and Medicaid Services (CMS) attempts to measure the resources required to provide each service, and compensates physicians accordingly.

U.S. private sector health care prices are largely unregulated.[4] Private reimbursements

[2]For overviews of the extensive cost-shifting literature (which includes Cutler, 1998; Kessler, 2007; Wu, 2010), see Frakt (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 from price shocks (White, 2013; White and Wu, 2014).

[3]For example, Congressional testimony during debate over the Affordable Care Act asserted that private payment increases offset 40 percent of decreases in public payments to both hospitals and physicians (Shiels, 2009).

[4]Some exceptions apply to this statement. For instance, all hospital payment rates in Maryland are set by a state government board.

are negotiated between insurance carriers and the providers with whom they contract, in markets with varying degrees of competition (Dafny, Duggan, and Ramanarayanan, 2012). Negotiated prices are often unknown to final consumers and can vary substantially, for ostensibly similar services, across both providers and insurers (Dunn and Shapiro, 2014).

Existing research illuminates some determinants of private health care prices. Some price variation reflects differences between the rates negotiated by HMOs and by traditional health insurance plans (Cutler, McClellan, and Newhouse, 2000). Price variation also stems from producer heterogeneity, with more attractive hospitals commanding higher prices (Ho, 2009; Moriya, Vogt, and Gaynor, 2010; Gowrisankaran, Nevo, and Town, 2013; Lewis and Pflum, 2015). Insurance market competition increases payments to physicians and hospitals (Town and Vistnes, 2001; Dafny, 2005; Dafny, Duggan, and Ramanarayanan, 2012), while competition among provider networks reduces them (Dunn and Shapiro, 2014). Showalter (1997) finds a positive cross-sectional relationship between state Medicaid fees and private insurers' physician reimbursements.

Aside from standard price-theoretic considerations, institutional details can have significant influence on physician pricing. Specifically, practitioners describe two modes of negotiation between providers and private insurers. Large providers can engage in detailed bargaining with insurers over service-specific pricing. In contrast, insurers typically offer small physician groups contracts based on a fixed fee schedule. This may be Medicare's schedule of relative rates or a customized fee schedule. The parties then negotiate a dollars-per-unit scaling, known as a conversion factor, which can itself be negotiated relative to Medicare's Conversion Factor. In an article advising physicians on the negotiation process, Gesme and Wiseman (2010) explain that, "The fee schedule in many contracts is stated as a percentage of the Medicare rate."

Fee schedules can incorporate a variety of modifications. Blue Cross Blue Shield of Michigan (2013) explains that, "Most maximum payment levels are based on the Resource Based Relative Value Scale (RBRVS) developed by the Centers for Medicare and Medicaid Services

(CMS)." In cases when this is not true, "Other factors that may be used in setting maximum payment levels include, but are not limited to, comparison to similar services, corporate medical policy decisions, analysis of historical charge data and geographic anomalies."

The benchmarking of private payments to Medicare's menu can generate a mechanical relationship between changes in Medicare's *relative* payments and corresponding private payments. Over the long run, renegotiations of insurer-physician contracts should tend to reverse mechanical price changes that the parties find deleterious. The extent of these subsequent revisions is an empirical question to which our analysis can speak.

Benchmarking to Medicare's payments could have three related economic rationales. First, Medicare's size in the markets for physicians' services makes it a relevant source of patients in lieu of those from a given private insurer. That is, Medicare will often be relevant as a doctor's outside option. Second, Medicare's relative value scale contains a comprehensive, if controversial (Ginsburg and Berenson, 2007), accounting of treatments' relative input costs. Contracts benchmarked to Medicare's relative rates seamlessly incorporate updated estimates of physicians' costs. Third, benchmarking reflects physicians' and insurers' desire to negotiate over a small number of contract parameters rather than payments for thousands of distinct service codes (Reckenen, 2013; Fontes, 2013). Regardless of the rationale, the long-run relationship between Medicare and private prices determines the extent to which Medicare influences the overall returns to different types of medical practice.

3 Conceptual Framework

We consider payment rate negotiations between one insurer and one physician group. This group joins the insurer's network if the parties agree on a reimbursement rate r^* for the care that the group provides to the insurer's patients. Let r_M be the corresponding payment that the group would receive if it provided the same care to Medicare patients instead.

We assume that the insurer and physician group choose r^* through Nash bargaining, and the insurer has a constant, exogenous bargaining weight $\theta \in [0,1]$. In order to examine the partial equilibrium problem of bargaining with this one group, we assume that the insurer's outside option is some fixed, exogenous value of care to its patients, v_I. Let u_{MD} be the physician's outside option if negotiations fail, which we will describe below. When $v_I > u_{MD}$, it is efficient for the group to join the insurer's network. The agreed payment rate is:

$$r^* = (1-\theta)v_I + \theta u_{MD}. \tag{1}$$

Empirical Implications

We consider three possibilities for the physician's outside option. First, suppose the physician has a constant marginal cost c per unit of care and faces no capacity constraint. Thus her outside option is the savings from not providing treatment, namely $u_{MD} = c$. In this case, Medicare does not affect the outside option so $\frac{dr^*}{dr_M} = 0$.

The second case is when the doctor faces increasing marginal costs of providing care. Changes in Medicare's reimbursements may influence her supply of care to Medicare patients, thereby altering the marginal cost of treating private patients. We capture this case by writing $u_{MD} = f(r_M)$, with $f'(r_M) > 0$. Thus Medicare's influence on private prices is $\frac{dr^*}{dr_M} = \theta f'(r_M) > 0$. We term this positive relationship between Medicare and private payments *price-following.*

In the third case, the physician group operates at capacity. Treating one additional private patient means treating α fewer Medicare patients, where α could be above or below 1.[5] In this case, the physician's outside option is the revenue from treating Medicare patients, $u_{MD} = \alpha r_M$. It follows that $\frac{dr^*}{dr_M} = \alpha\theta > 0$.

This framework enables us to interpret empirical estimates of the price-following coefficient that we will estimate, $\frac{dr^*}{dr_M}$. If the data show significant price-following, we can infer

[5]Glied and Graff Zivin (2002) find that doctors spend very similar lengths of time with private and Medicare patients, suggesting that $\alpha \approx 1$.

that physicians face costs of scaling up their practice. In these cases, we would expect to see price-following increasing in the insurer's bargaining weight (θ).

The contracting institutions discussed in the previous section also have a natural interpretation within this bargaining framework. When contracts are benchmarked to Medicare's rates, prices are $r^* = \phi r_M$. Contracts of this form can update automatically to reflect CMS's changing assessment of physicians' opportunity costs. With contracts of this form, $\frac{dr^*}{dr_M} = \phi$.

Our framework emphasizes that the private payment to our profit-maximizing physician relates positively to her opportunity cost, generating price-following. But the literature on hospital pricing highlights that the opposite possibility, which it calls *cost-shifting*, should not be dismissed. Cost-shifting could arise through altruism (Cutler, 1998; Dranove, Garthwaite, and Ody, 2013), income effects (McGuire and Pauly, 1991), a change in efficiency, or changes in fixed costs as a physician's scale increases (Kessler, 2007). If physician groups are operating below minimum efficient scale, then an increase in Medicare payments could counterintuitively reduce their marginal cost of treating private patients, *i.e.* $f'(r_M) < 0$.

Welfare Implications

The welfare consequences of price-following depend on its short- and long-run implications for real resource allocations. Short run welfare implications depend in part on whether the marginal benefits of care are higher for the elderly relative to the privately insured. These relative marginal benefits are not known. Further, while the empirical literatures on specialty choice (Nicholson and Souleles, 2001) and other extensive margins consistently find positive supply responses,[6] the literature on short-run, intensive-margin supply is more mixed.[7]

Our primary empirical setting involves changes to Medicare's payments for surgical relative to nonsurgical care. Long-run welfare effects of these changes depend primarily on the

[6] Acemoglu and Finkelstein (2008) and Clemens and Gottlieb (2014), for example, study investment decisions, while Acemoglu and Linn (2004), Finkelstein (2004), Blume-Kohout and Sood (2013), Budish, Roin, and Williams (2015), and Clemens (2013) show that innovation responds to potential market sizes, which affect the return to practice more generally.

[7] See, for example, Jacobson, Earle, Price, and Newhouse (2010) and Clemens and Gottlieb (2014).

marginal benefits of surgical relative to nonsurgical care and on how the policy change affects the availability of this care. Policy makers and researchers have long expressed concern that Medicare's relative rates do not reflect the relative marginal benefits of these two treatment categories, and that the U.S. health system provides inefficiently little primary care relative to intensive procedures (Newhouse, 2002; Cutler, 2011).

We thus examine empirical outcomes including new physicians' specialty choices and practicing physicians' willingness to take new patients. The relationship between Medicare and private payments has unambiguous relevance for Medicare's ability to shift these margins. With price-following, private payments augment Medicare's influence on the returns to practicing in surgery relative to primary care. By contrast, under cost-shifting, private payment changes blunt Medicare's influence.

4 Data and Measures

4.1 Health Care Price Data

We study the public sector's influence on private sector health care prices by linking insurance claims data across the two environments. In both settings, providers request reimbursement by submitting claims to the relevant third-party payer. We use Medicare claims from a 5 percent random sample of the 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 that Medicare purchases for these beneficiaries.[8] For pricing purposes, the data include the Health Care Procedure Coding System (HCPCS) code for each service along with Medicare's payment (the "allowed charge"). We compute

[8] The medical payments system in the U.S. compensates hospitals through a completely separate system. For example, a surgery performed in a hospital operating room generates a payment to the surgeon and a separate payment to the hospital for use of the facility. We do not examine hospital pricing in this paper, and our data would correspondingly include the former payment but not the latter.

Medicare's observed payment rates by averaging the allowed charges for each service in each year.

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

The MarketScan data are a selected sample and do not represent the full population of private insurance claims. They comprise claims data from around 100 large payers, such as large employers, who cover employees', dependents', and some other constituencies' health care. Our results are thus most directly applicable to the self-insured employer segment of the insurance market. This segment is substantial, comprising 55 percent of all privately insured individuals as of 2008 (Fernandez, 2010). Even if the empirical magnitudes vary in other settings, the underlying economic forces likely remain relevant. Nearly all plans compete with Medicare and other insurers to procure physicians' resources, and must thus offer competitive rates. Appendix B.1 further discusses the breadth of our results' applicability.

Our baseline sample includes 2,194 unique 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: an area-by-service pair is only included in the sample if it appears in each year from 1995 through 2002. The Data Appendix provides further detail on these primary data sources and our merge procedure. Appendix B.2 discusses the insurance plan types underlying our private payments data. Our estimates are robust to relaxing the panel balance criterion or altering the level of aggregation.[9]

[9] Appendix B.4 further discusses the construction of control variables associated with the insurance plan

Table 1 presents summary statistics for Medicare and private sector prices across services and states, separately for surgical and nonsurgical services. The average surgery payment is \$172 in Medicare and \$279, or over 60 percent higher, in the private market. The average nonsurgical service is reimbursed \$54 in Medicare and \$70 in the private sector. Figure 1 displays raw correlations between public and private payments across services. The correlations are quite strong in both levels (panel A) and changes (panel B).[10] A cross-sectional regression of average private prices against Medicare rates yields a coefficient of 1.45.

4.2 Measuring Physician and Insurer Concentration

To explore how Medicare's influence varies across markets, we construct measures of insurer and physician concentration using standard Herfindahl-Hirschman Indices (HHIs). To compute physician HHIs, we follow Baker et al. (2014) in using the group tax identifiers available in the Medicare claims.[11] We also compute a more targeted measure of concentration that varies across both specialties and areas. We measure insurance competition using data from the National Association of Insurance Commissioners (NAIC)'s health insurance reports,[12] which allow us to compute state-level HHIs for all states except California.[13]

Figure 2 provides suggestive evidence that our HHI measures do indeed capture economically relevant aspects of competition. The figure shows a smoothed measure of the average price per service in our sample, pooling across all services, based on the HHI in the area where the service was provided (along the horizontal axis). The two curves in Figure 2 di-

types underlying the payments we observe. Appendix C confirms the reduced form results at levels of aggregation ranging from individual claims (N = 144 million) to Hospital Service Area, Hospital Referral Region, state, or national aggregation.

[10]Appendix Figure B.2 shows analogous graphs using cross-state variation.

[11]While tax identifiers may not correspond directly with negotiating units, Baker et al. (2014) and other authors find the resulting HHIs to have significant economic content. Table 1 reports summary statistics on our concentration variables, and Appendix B.5 further details their construction.

[12]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.

[13]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).

vide the price data based on insurer concentration in that state.[14] Consistent with Dunn and Shapiro (2014), both curves show that average physician payments are higher in areas with more concentrated physicians. At the same time, more concentrated insurance markets tend to pay physicians lower reimbursement rates, at any given level of physician concentration.

4.3 Physician Data from the Community Tracking Study

Using a third data source, the Community Tracking Study (CTS), we examine a variety of welfare-relevant dimensions along which physician practices may respond to the price changes we study. The CTS is a biennial survey of around 12,000 physicians per wave in 60 geographic areas across the United States (CSHSC 1999). The survey was conducted in 1996-97, in 1998-99, in 2000-01, and in 2004-05, and covers topics including physicians' specialties, graduation years, willingness to accept new patients, maintenance of board certification, and career satisfaction. When examining changes in practice characteristics, we exclude the 916 survey respondents from the 2004-2005 CTS whose graduation years were between 1996 and 2004, as these individuals' specialty choices may have been affected by the payment changes we analyze. Appendix B.6 further describes the questions we use in our analysis.

5 Empirical Model

5.1 Two Shocks to the Relative Prices of Outpatient Services

Since 1992, Medicare has paid physicians and other outpatient providers through a system of centrally administered prices, based on a national fee schedule. This fee schedule, known as the Resource-Based Relative Value Scale (RBRVS), assigns relative values to more than 10,000 distinct billing codes according to the resources CMS believes the services to require. Medicare scales these relative valuations by multipliers called Conversion Factors (CFs). Our

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

first natural experiment involves a large, administrative change in the CFs for surgical and nonsurgical services. Because input costs vary across areas, the fee schedule further adjusts payments to partially offset such differences. Our second natural experiment exploits an administrative change in this system of geographic adjustments. For service j, supplied in year t by a provider in payment area a, the provider's fee is approximately:

$$\begin{aligned}\text{Reimbursement}_{a,j,t} = \text{Conversion Factor (CF)}_{t,c(j)} \times \text{Relative Value Units (RVU)}_{j,t} \\ \times \text{Geographic Adjustment Factor (GAF)}_{a,t}. \end{aligned} \tag{2}$$

Shock to Surgical *versus* Medical Payments

The Conversion Factor is a national adjustment factor, updated annually and identical across broad categories of services, $c(j)$. In the early 1990s, wrangling over payments across specialties led to the introduction of separate CFs for surgical procedures and other services. Surgeons argued that slower growth in the use of procedures relative to other medical services should be rewarded. Congress implemented this plan, and CMS first distinguished between the CFs for surgery, primary care, and other services in 1993.[15]

From 1993 to 1995, care volumes evolved such that payments for surgical procedures grew relative to payments for other services. CFs were then relatively stable from 1995 to 1997, with an average bonus of 15.5 percent for surgical RVUs relative to primary care and other nonsurgical RVUs. These unequal payments for equal RVUs spawned political discontent among non-surgeons. In 1998, this 15.5 percent bonus was eliminated through a budgetarily neutral merger of the CFs, which was plausibly exogenous with respect to changes in beneficiary demographics, other determinants of demand, and changes in the resource costs of providing care.[16] The evolution of the surgical and nonsurgical Conversion Factors during this era is shown in Figure 3.

[15] We owe our knowledge of this political history to Newhouse (2002).
[16] 62 *Federal Register* 59048, 59102 (1997).

To take two illustrative examples, consider Medicare's payments for a coronary artery bypass graft (CABG) and a cardiac stress test with nuclear imaging (SPECT). In 1997, Medicare's fee for CABG (CPT code 33533) averages \$1,428. In 1998, the average fee falls to \$1,283, or by just over 10 percent. In 1997, SPECT (CPT code 78465) generates an average Medicare fee of \$475. Because SPECT is an imaging service rather than a surgical procedure, its average fee rises to \$513 (an increase of 8 percent).

We use Medicare payment data to construct an instrument based on these price shocks. We compute the average price $\overline{P^{\text{Medicare}}}_{j,\text{pre}}$ for each service j prior to the policy change. Specifically, we use data from 1995 to 1997, a baseline period when the surgical-medical Conversion Factor difference was stable.[17] We then construct a variable that captures the price change implied by the 1998 CF merger:

$$\text{PredChg}_j^{\text{CF}} = \overline{P^{\text{Medicare}}}_{j,\text{pre}} \cdot \begin{pmatrix} -0.104 \cdot \text{Surgical}_j \\ +0.05 \cdot \text{Nonsurgical}_j \end{pmatrix}, \tag{3}$$

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

Across-the-Board Payment Shocks

We next analyze payment changes that varied across geographic areas, and altered reimbursements across the board within those areas. While the Conversion Factor is set nationally, the last term of equation (2) incorporates a Geographic Adjustment Factor that varies across payment regions. It is intended to capture differences in input costs, which are estimated using Census and other data on area-level rents, wages, and malpractice insurance premiums.

We analyze the effects of GAF changes driven by an administrative re-shuffling of the

[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 of heterogeneity in Medicare's effects across services and states. Appendix C shows that our results remain similar when using national level observations.

areas across which these adjustment are made. Until 1996, payments were differentiated across 210 payment areas, as shown in the top panel of Appendix Figure B.3. In 1997, the federal government consolidated these 210 regions into the 89 larger ones shown in the figure's middle panel. These mergers were budget-neutral within each state, and the consolidations generally reduced urban payments and increased rural payments. Clemens and Gottlieb (2014) provides additional institutional background.

To construct our second instrument, let GAF_a denote the GAF in area a prior to the payment area merger. When payment localities are merged, a joins a new larger region $A \supset a$ with a common payment factor, computed by averaging over all of the constituent sub-regions: $GAF_A = \overline{GAF_a}$. Using this information, we define the payment shock in pre-merger region a (shown in Panel C of Appendix Figure B.3) as:

$$\text{PredChg}_a^{\text{Geo}} = \overline{P^{\text{Medicare}}}_{a,\text{pre}} \cdot (GAF_A - GAF_a). \tag{4}$$

Interpreting the Payment Shocks

These payment shocks speak to somewhat distinct aspects of Medicare's influence. The first instrument involves relative price changes across services. Its effects are relevant for analyzing the allocation of spending across types of health care. The second instrument exploits broad-based pricing changes, and thus provides estimates useful for predicting the consequences of across-the-board changes in payment levels.

Because of the way insurers react to Medicare pricing, comparing responses to the two shocks can help illuminate the mechanisms behind price-following. Even when payment contracts rely on explicit Medicare benchmarking, as discussed in section 2, they often disregard some parts of the Medicare fee schedule. Blue Cross and Blue Shield of Texas (2010) explicitly disregards the geographic adjustments, while other provider manuals (Anthem Blue Cross and Blue Shield, 2012; Blue Cross Blue Shield of Michigan, 2013) do not mention these adjustments. Miller, Zuckerman, and Gates (1993) argue that most private insurers disregard

Medicare's GAFs.[18]

5.2 Estimation Framework for Price Responses

We exploit both payment shocks in a standard instrumental variables framework, in which $\text{PredChg}_{j,a}^{\text{Medicare}}$ is a variable that can represent either one of these instruments, or a vector including both simultaneously. Let $\text{PostImplementation}_t$ be an indicator for years after the respective change was implemented (or a vector of such indicators). With observations at the service (j), by area (a), by year (t) level, we use two-stage least squares to estimate:

$$\begin{aligned} P_{j,a,t}^{\text{Medicare}} &= \pi \cdot \text{PredChg}_{j,a}^{\text{Medicare}} \times \text{PostImplementation}_t + X_{j,a,t}\phi_1 + \mu_j \mathbb{1}_j + \mu_a \mathbb{1}_a + \mu_t \mathbb{1}_t \\ &\quad + \mu_{j,a} \mathbb{1}_j \cdot \mathbb{1}_a + \mu_{t,s} \mathbb{1}_t \cdot \mathbb{1}_s + e_{j,a,t} \qquad (5) \\ P_{j,a,t}^{\text{Private}} &= \beta \cdot \widehat{P_{j,s,t}^{\text{Medicare}}} + X_{j,s,t}\phi_2 + \nu_j \mathbb{1}_j + \nu_a \mathbb{1}_a + \nu_t \mathbb{1}_t \\ &\quad + \nu_{j,a} \mathbb{1}_j \cdot \mathbb{1}_a + \nu_{t,s} \mathbb{1}_t \cdot \mathbb{1}_s + \varepsilon_{j,a,t}. \qquad (6) \end{aligned}$$

When examining the first payment shock, the geographic unit (a) corresponds to state (s). When using the second instrument, areas a are the smaller pre-consolidation payment localities. To account for each service's total contribution to Medicare spending, we weight by the number of times a service is performed in 1997. When using the first instrument we cluster standard errors by service codes. With the second instrument we cluster by payment locality. In both cases, this corresponds to the dimension along which the payment shocks vary.

The coefficient π in the first stage equation (5) describes how a \$1 predicted Medicare price change flows into the actual Medicare payment for a service. We would estimate $\hat{\pi} = 1$ in the absence of measurement error and correlated reimbursement changes. The coefficient β in the second-stage equation (6) measures how private payments for a service respond, in

[18] In a different context, Chambers, Chenoweth, Thorat, and Neumann (2015) find that insurers adopt a wide range of policies for determining whether to cover medical devices approved by Medicare—approximately half adopt Medicare's policies while half develop their own.

dollars, to a \$1 change in Medicare reimbursements.

We control for service, area, and year fixed effects as well as service-by-region ($\mathbb{1}_j \cdot \mathbb{1}_a$) and state-by-year ($\mathbb{1}_s \cdot \mathbb{1}_t$) effects. The vector of additional controls ($X_{j,a,t}$) includes indicators for major individual-service payment changes. Specifically, our first stage most cleanly tracks the policy changes of interest when we control separately for major mid-1990s payment changes for cataract surgery.[19,20]

The predicted Medicare prices are valid instruments under the following assumptions. First, the predicted change $\text{PredChg}_{j,a}^{\text{Medicare}}$ must be reflected in the actual Medicare prices in the first stage equation (5). Second, the shocks used to generate these predicted prices must be conditionally independent of other sources of change in private sector payment rates, or $\text{Cov}\left(\varepsilon_{j,a,t}, \text{PredChg}_{j,a}^{\text{Medicare}} \times \text{PostImplementation}_t \middle| X_{j,a,t}, \mathbb{1}_j, \mathbb{1}_a, \mathbb{1}_t, \mathbb{1}_j \cdot \mathbb{1}_a, \mathbb{1}_t \cdot \mathbb{1}_s\right) = 0$. The conditioning variables are intended to capture potential confounds such as technology shocks, demand shocks, and other changes in market conditions.

Parametric Event Study

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:

$$
\begin{aligned}
P_{j,a,t}^{\text{Medicare}} = &\sum_{t \neq t_0} \gamma_t \cdot \mathbb{1}_t \times \text{PredChg}_{j,a}^{\text{Medicare}} + X_{j,a,t}\psi_1 + \mu_j \mathbb{1}_j + \mu_a \mathbb{1}_a + \mu_t \mathbb{1}_t \\
&+ \mu_{j,a} \mathbb{1}_j \cdot \mathbb{1}_a + \mu_{t,s} \mathbb{1}_t \cdot \mathbb{1}_s + u_{j,a,t} \qquad (7) \\
P_{j,a,t}^{\text{Private}} = &\sum_{t \neq t_0} \delta_t \cdot \mathbb{1}_t \times \text{PredChg}_{j,a}^{\text{Medicare}} + X_{j,a,t}\psi_2 + \upsilon_j \mathbb{1}_j + \upsilon_a \mathbb{1}_a + \upsilon_t \mathbb{1}_t \\
&+ \upsilon_{j,a} \mathbb{1}_j \cdot \mathbb{1}_a + \upsilon_{t,s} \mathbb{1}_t \cdot \mathbb{1}_s + v_{j,a,t}. \qquad (8)
\end{aligned}
$$

[19] Cataract surgery was subjected to significant payment reductions in the years preceding the 1998 payment shock on which we focus.

[20] Appendix B.4 describes further controls for the types of insurance plans in our private data.

If pre-existing trends in either public or private payments are correlated with $\text{PredChg}^{\text{Medicare}}_{j,a}$, they will be apparent in estimates of γ_t and δ_t for years prior to the payment shock, $t < t_0$. When using the first instrument, we omit the interaction for $t_0 = 1997$. For the second instrument, $t_0 = 1996$. Estimates of γ_t and δ_t for $t > t_0$ will trace out the dynamic relationship between Medicare's payment shocks and public and private payments, respectively.

Null Hypotheses

Our setting suggests three distinct benchmarks worth considering when estimating price-following. Specifically, we test the null hypotheses of $\beta = 0$, $\beta = 1$, and $\beta = 1.45$. The test of $\beta = 0$ corresponds with the model case in which physicians face constant marginal costs. The test of $\beta = 1$ can be interpreted as testing full pass-through of changes in Medicare's payments, as when Medicare's payment is the physicians' outside option and insurers have full bargaining power. The test of $\beta = 1.45$ corresponds with the mechanical pass through we would expect if contracts are benchmarked directly to Medicare's relative payments, since 1.45 is the average scaling of private to Medicare payments in our data.

5.3 Estimation Framework for Physician Outcomes

We use the CTS to investigate how work hours, propensity to take new patients, career satisfaction, and maintenance of board certification evolved for surgeons relative to other physicians following the reductions in surgeons' relative payments. Using y_{it} to denote any one of these outcomes, we estimate the following difference-in-differences specification on physician-level data from all four CTS waves:

$$y_{it} = \kappa\,\text{Surgeon}_i \times \text{PostImplementation}_t + \lambda\,\text{Surgeon}_i + \chi_t\,\text{SurveyWave}_t + \varepsilon_{it}. \qquad (9)$$

In this regression, Surgeon_i is an indicator for whether respondent i is in a surgical specialty, SurveyWave_t is an indicator for a survey wave t, and $\text{PostImplementation}_t$ indicates years

1998 and beyond. The omitted physician category is non-surgeons and the omitted survey wave is 1996–97.

6 Effects of Medicare Prices on Private Prices

6.1 Surgical *versus* Medical Price Changes

Figure 4 plots estimates of the effect of Medicare's changes to payments for surgical procedures relative to medical services. First stage estimates of $\hat{\gamma}_t$ from equation (7), the parametric event study specification, are marked on the graph with "$\times$" symbols. They show that the dollar value of the predicted price change for a service translates almost one-for-one into realized Medicare payment rates. There also appears to be a slight upward drift associated with gradual increases in Medicare's payments for primary care relative to other services. While this drift makes it important to look closely at the dynamics of private responses, and to check robustness after controlling for a surgery-specific trend, these results give us confidence in our specification of the shock.

Figure 4 also plots the $\hat{\delta}_t$ estimates from equation (8), which are reduced form measures of the payment change's effect on private prices. Changes in private prices are uncorrelated with the payment shocks during the years preceding the shock, providing evidence against potentially confounding pre-existing trends driven by changes in technology, demand, or other market conditions. From this point forward, a \$1.00 increase in Medicare's predicted payment for a service leads, on average, to a \$1.40 increase in private payments per service. The private sector response emerges in full during the year of Medicare's payment change. The initial changes are not reversed in the long-run, as we might have seen if adjustments driven by Medicare-benchmarked contracts were subsequently undone. Figure 5 plots the raw means of Medicare and private payments for surgical and nonsurgical care, showing that these patterns are also visible in the raw data without any regression controls.

In Table 2, we summarize these results using the instrumental variables framework of equations (5) and (6). Column 1 reports the first-stage estimate of equation (5). We estimate $\hat{\pi} \approx 1.2$, which is close to the value we would expect absent measurement error ($\pi = 1$). The cluster-robust F statistic for testing the null hypothesis that our instrument is weak is 294, which easily satisfies the robust weak instruments pre-test threshold of Olea and Pflueger (2013).[21]

Column 2 shows the reduced form result we obtain when we replace $P_{j,s,t}^{\text{Medicare}}$ with $P_{j,s,t}^{\text{Private}}$ as the outcome variable in (5). The coefficient of 1.39 means that a one dollar predicted change in Medicare prices translates into a $1.39 change in private sector prices. Column 3 reports the IV estimate of equation (6), which simply rescales the private sector change by the actual Medicare change from column 1. This result is our baseline estimate, which implies that a one dollar change in actual Medicare payments for a service leads to a $1.16 change in private payments for that service. This estimate is strongly statistically distinguishable from 0, allowing us to reject the null hypothesis that Medicare's payments have no effect on private payments. It is statistically indistinguishable from 1, which corresponds with the hypothesis that Medicare is the outside option of capacity-constrained physicians. It is also statistically indistinguishable from 1.45, which represents full pass-through of Medicare's relative price changes scaled by the average markup of private payments to public payments.

Appendix Table C.1 evaluates the robustness of our finding that Medicare prices pass through strongly into the private sector. Columns 1 through 3 show that our results are robust to augmenting the baseline IV specification with controls for insurance plan characteristics. Column 4 shows that our baseline estimate is modestly sensitive to omitting controls for mid-1990s cataract surgery payment changes, which reduces the price-following coefficient from 1.16 to 0.97. Column 5 removes the service weights, which reduces the estimate to around 0.7.[22] Column 6 controls for the number of Relative Value Units (the

[21] 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 weak instruments in the presence of heteroskedasticity or clustering.

[22] Accounting for the reductions to payments for cataract surgery improves our ability to correctly track

quantity metric that appears in Medicare's payment formula) assigned to each service. Minor updates to RVU assignments strongly predict Medicare's allowable charges (coefficient not shown), but controlling for these updates has little impact on our baseline result. Finally, column 7 shows that the baseline result is robust to controlling directly for a linear trend in private payments for surgical procedures relative to other services. As shown in Figure 4, there is no such trend in private payments. Appendix Table C.2 shows that the baseline results are also robust to estimation on data that we have aggregated to the national level.

6.2 Across-the-Board Payment Changes

We next consider across-the-board payment changes that result from the overhaul of Medicare's payment localities. Figure 6 mirrors Figure 4 and presents parametric event study estimates of $\hat{\gamma}_t$ and $\hat{\delta}_t$ from equations (7) and (8). The first stage estimates of $\hat{\gamma}_t$ show that our coding of the payment shocks effectively tracks the policy change. A one dollar increase in the predicted payment shock is associated with a one dollar increase in Medicare's allowed charge for a service. The reduced form estimates plot the private sector response to these public payment shocks. As with shocks to relative prices across services, an increase in public payments generates an increase in private payments. The effect of these across-the-board payment changes appears to unfold over several years.

We summarize these results in columns 4 through 6 of Table 2. Column 4 reports a first stage coefficient of 0.89 and column 5 reports a reduced form coefficient of 1.0. Column 6 reports the resulting IV estimate of 1.12, which is nearly identical to our price-following estimate from the surgical-nonsurgical payment change in column 3.

The sample on which we are able to analyze these geographic payment shocks is notably smaller than the one we used with the first payment change (N = 128,694 in the former

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

case vs. N = 303,728 in the latter). This reflects the fact that the pre-consolidation payment localities are substantially smaller than states, and thus have fewer distinct services provided in all eight years of our sample. The precision of these estimates is correspondingly lower. In Appendix Table C.4, we re-estimate the effect of the surgical-nonsurgical payment change, but on the smaller, less geographically aggregated, sample on which we analyze the geographic payment shocks. The results are similar. The IV estimate of 0.9 is statistically indistinguishable from the baseline estimate. In columns 4 through 6 we simultaneously include both sets of Medicare shocks as instruments. This too has no meaningful impact on the results. The effects of the two payment shocks are statistically indistinguishable from one another.

Comparing the short- and long-run effects of Medicare's geographic adjustments to the dynamics after the surgical-nonsurgical payment change casts light on the mechanisms underlying our estimates. The shocks' short-run effects are economically quite different; the relative price change appears to feed immediately into private payments, likely because of contractual benchmarking, while the across-the-board payment change has more gradual influence. The long-run estimates are similar and are, in both instances, economically substantial. Medicare's payments thus appear to exert large and persistent influence on the outcomes of active physician-insurer contract negotiations.

6.3 Heterogeneity by Market Characteristics

The conceptual framework of section 3 suggests that the strength of price-following depends on physicians' and insurers' relative bargaining power. We explore this prediction by estimating the strength of Medicare's influence in subsamples split based on our measures of physician and insurer concentration.

Panel A of Figure 7 reports price-following estimates separately for each tercile of the physician concentration distribution. The graph shows a very clear pattern of heterogeneity across markets, with the least concentrated markets exhibiting the strongest price-following relationship in every year following the Conversion Factor merger. This group's prices reflect the Medicare changes beginning immediately in 1998, and are relatively stable thereafter. The intermediate concentration group exhibits a slightly delayed response, and of smaller magnitude. Finally, the most concentrated group shows no response.

We next examine the relevance of competition on the insurer side of the market. Just as competition among physicians should reduce physicians' relative bargaining power, competition among insurers should increase it. We display graphical results by tercile of insurer concentration in Panel B of Figure 7. This figure reveals a positive relationship between insurance concentration and the magnitude of Medicare's effect on private payments. Both physician and insurer concentration thus appear to mediate price-following as predicted by the simple bargaining framework of section 3.

Appendix D further analyzes these dimensions of heterogeneity. We move from subsample analyses to interacted specifications, presented in Appendix D.1, which allow us to formally test for differences in the strength of Medicare's influence between markets with high and low levels of physician and insurer concentration. Columns 1 and 2 of Appendix Table D.1 show that the heterogeneity associated with both physician and insurer concentration is statistically significant. Subsequent columns report that the strength of Medicare's impact is relatively large in markets with more HMO penetration, with relatively small physician groups, when Medicare accounts for a relatively large share of spending on a service, and when there is relatively more *ex ante* price dispersion. These results hold up when allowing for heterogeneity along all of these dimensions simultaneously.

Market characteristics are not randomly assigned, and could thus be correlated with other factors. Appendix Tables D.2 and D.3 thus explore the robustness of our heterogeneity results to simultaneously including interactions between the Medicare payment shocks and a variety of economic and demographic characteristics. The results are robust to including a range of such controls.

6.4 Evidence on Medicare-Benchmarked Contracts

Section 2 presented anecdotal evidence that negotiated payment schedules are sometimes benchmarked directly to Medicare's relative rates. We cannot directly estimate the prevalence of such contracts in our main sample because MarketScan data do not track payments between identifiable insurer-physician group pairs over time. Appendix E analyzes claims data from Blue Cross Blue Shield of Texas (BCBS-TX) in which we can overcome this hurdle.

In the BCBS-TX data, we estimate that benchmarking to Medicare's relative rates contributes 0.4 to 0.5 to the short-run relationship of 1.2 between Medicare and private payments (based on the first treatment year in Figure 4). We thus estimate that the initial effect on private prices of a one dollar change Medicare payments would be around \$0.70 to \$0.80 rather than \$1.20 in the absence of this benchmarking. Because the \$1.20 estimate persists over subsequent years, we infer that insurers and physician groups largely agree to maintain these mechanical changes in subsequent years' negotiations.

6.5 Consequences for Physician Behavior and Specialty Choice

We conclude our empirical analysis by investigating whether the surgical-nonsurgical payment change alters physicians' practices. We first analyze service quantities in the Medicare and MarketScan data. Consistent with the mixed empirical literature on short-run responses to physicians' financial incentives, the data do not support strong conclusions.[23]

For evidence on extensive margin adjustments with consequences for long-run welfare, we turn to the Community Tracking Study. In Table 3 we show results describing the differential evolution of surgeons' practices relative to those of other physicians following the surgical-nonsurgical payment change. Each column estimates equation (9) with a different independent variable. Column 1 reports a small and statistically insignificant change in

[23] Appendix F presents and further discusses the Medicare and MarketScan quantity data. While the provision of surgical relative to nonsurgical care is lower at the sample's end than at its beginning, the timing of the observed changes makes it difficult to attribute them with confidence to the surgical-nonsurgical payment change.

hours of work. Columns 2 and 3 show statistically strong evidence that, relative to other physicians, surgeons reduced their propensity to accept new patients from both Medicare and private insurers. The changes are 6 to 8 percent of the baseline levels. Column 4 shows that surgeons become less likely to report being very satisfied with their careers, though the result is only marginally significant. Finally, column 5 shows that surgeons became less likely to pursue the continuing education needed to maintain board certification. Magnitudes for both of these final results are around 5 percent of the baseline levels.

The results in Table 3 suggest that the reduction in surgeons' relative payments induced a retreat of surgeons' practices along various extensive margins. Figure 8 provides evidence of further impacts on the health-sector's long-run composition. Using the 2004–05 wave of the CTS, Figure 8 tracks specialty choice across cohorts of medical school graduates. Following several periods of stability, the figure reveals that physicians graduating between 1996 and 2004 were less likely to pursue surgical specialties than earlier cohorts, by nearly 7 percentage points. In the next section we consider what this development means for welfare and we relate it to prior research on physician specialty choice.

7 Implications for Resource Allocation

7.1 Physician Income and Specialty Choice

The allocation of medical resources between surgical and nonsurgical care is an area of active policy debate. Perhaps the most visible element of this policy debate surrounds physician specialty choice.[24] Here we show how our estimates of price-following help to rationalize the changes in specialty choice depicted in Figure 8.

Table 4 shows several specialties' exposure to the conversion factor overhaul under three

[24]For example, while Petterson et al. (2012) predict an impending shortage of 52,000 primary care physicians by 2025, HHS (2013) predicts a shortfall as low as one-eight that magnitude. The American Association of Medical Colleges (Dall et al., 2015) predicts a shortage of *non*-primary care physicians twice as large as that for primary care, while Fisher et al. (2003) find that additional specialists are associated with worse clinical outcomes. For a range of estimates of the medical returns to more intense care, compare Fuchs (2004) and Skinner, Staiger, and Fisher (2006) with Chandra and Staiger (2007) and Doyle (2011).

assumptions about private payment responses: (1) no private payment response, (2) a 40 percent private offset (as assumed by Shiels, 2009), and (3) price-following with a coefficient of 1.16. Appendix G.1 further describes the methodology underlying these calculations.

Column 1 of Table 4 presents the direct effects of Medicare price changes on take-home pay for physicians in family practice, general practice, general surgery, and dermatology. These are equal to the total effects if there is no private market response. Accounting for the composition of each specialty's patients, treatments, and practice expenses, the specialties face net income shocks ranging from −5.5% for dermatologists to +2.9% for family practitioners. Column 2 adopts the cost-shifting assumption used by Shiels (2009) and others in policy debates. Under this assumption, the Medicare CF changes would have little effect on take-home pay differentials across specialties. Even surgeons and dermatologists would see their net incomes fall by less than 2 percent under the 40% offset assumption, as private insurers offset most of Medicare's direct effects on income. Finally, column 3 shows that the price-following we document dramatically amplifies Medicare's direct effects. The net income shocks now range from −16.6% for dermatologists to +11.2% for those in family practice.

Panel E relates these income changes to specialty choice. Nicholson and Souleles (2001) find that a $10,000 increase in the peak lifetime income difference between non-primary care and primary care results in a 1.4 percentage point increase in new medical school graduates' probability of entering primary care. The estimates in column 3 of Table 4 show that, with the price-following we estimate, the surgical-nonsurgical payment change increases the income of family practice physicians by $15,300 and decreases the income of general surgeons by $23,800. Nicholson and Souleles's estimate suggests that this $40,000 relative income change would drive a 5.5 percentage point shift towards family practice. Figure 8 shows that this prediction lines up well with the observed specialty choice data. The "no spillover" and "40% offset" scenarios would support shifts of just 1.9 and 0.6 percentage points. As Figure 8 also shows, these predictions differ meaningfully from the observed shift.

To place these estimates in context, the Department of Health and Human Services (2013) has expressed concern that maintaining socially desired levels of primary care access over the next decade will require an additional 20,000 physicians to choose primary care specialties, beyond those currently projected to do so. Recent years have seen around 18,000 medical graduates annually. A 5.5 percentage point increase in primary care' share of graduates thus represents around 1,000 additional primary care physicians per cohort. Over two decades, increased flows of this magnitude would accommodate the projected demand.

7.2 Magnitudes of Resource Allocation

Beyond specialty choice, a growing body of research finds that the investments of current practitioners (Acemoglu and Finkelstein, 2008; Clemens and Gottlieb, 2014) and the development of new technologies respond significantly to their expected returns.[25] These margins further illustrate the long-run relevance of Medicare's capacity to steer resources.

Consider first the income transfers from merging the surgical and nonsurgical Conversion Factors. In aggregate, this payment change directly shifts \$1.8 billion in Medicare payments from medical treatment to surgical treatment. Assuming our estimates apply broadly, and ignoring any quantity response, we estimate that private payment responses would amplify this reallocation by an additional factor of 3.3. In total, we estimate that this payment change reallocates \$5.9 billion in annual private sector spending from surgical to medical care.[26]

Next consider the geographic payment changes we analyze. The payment reductions in adversely affected (generally urban) areas average 1.7 percent of their Medicare Part B reimbursements.[27] These changes thus mechanically reallocate roughly \$282 million between urban and rural areas in 1997. Again assuming that our results are broadly applicable, we

[25] Acemoglu and Linn (2004), Finkelstein (2004), Blume-Kohout and Sood (2013), Budish, Roin, and Williams (2015), and Clemens (2013) show that innovation responds to potential market sizes.

[26] We detail this calculation and discuss caveats about external validity in Appendix G.2.

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

estimate that private payment responses magnify this change by a factor of 3.3. In total, Medicare's payment change thus reallocated more than $1 billion from urban to rural areas. This is over three times as large as the Critical Access Hospital (CAH) program, Medicare's explicit subsidy for rural health care.[28]

Finally, if Medicare were to change its payment rates nationwide, this would likely have impacts similar to the across-the-board geographic shocks we study. Under this assumption, a 1 percent increase in Medicare payments for all physician and clinical spending would directly cost $1.3 billion annually. Our estimates imply that private insurers would spend an additional $3.5 billion annually due to price-following.

8 Conclusion

We show that Medicare exerts widespread and quantitatively substantial influence over the rates that private insurers pay. A $1 change in Medicare's payments for one service relative to another, or in one geographic region relative to another, drives a change of just over $1 in private payments. These payment changes reorient billions of public and private sector health dollars across locations and types of care. In aggregate, our estimates imply that Medicare's pricing decisions can appreciably move both health-sector and overall inflation (Clemens, Gottlieb, and Shapiro, 2014).

Our analysis sets the stage for several additional strands of research on the economics of physician pricing. First, state Medicaid programs may have similar private market spillovers, especially for plans sold on the Affordable Care Act's exchanges.[29] Second, our analysis only speaks directly to fee-for-service payments, but documentary evidence suggests that

[28]The CAH program itself is prominent and politically important in rural areas, as we document in Appendix G.3. This appendix also discusses the caveats associated with this calculation.

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

some HMOs' physician payments also draw on Medicare's menu.[30] Third, does Medicare influence private insurers to develop or adopt novel payment models? For example, how do private plans react to a public insurer's adoption of "bundled payment" mechanisms, or to transitions from resource- to value-based payments? Fourth, recent work on price-following in hospital markets (White, 2013; White and Wu, 2014) could be expanded to explore Medicare's influence on relative pricing across hospital services—which may plausibly be different than in the physician context.

Separately, our results raise questions about the mechanisms underlying Medicare's pervasive influence. Medicare could exert sway through multiple channels. As a large market participant, it competes with private insurers for physician resources. Practitioners further emphasize that many private insurers' contracts are benchmarked directly to Medicare's menu. Both channels of influence make Medicare an important participant in payment-system experimentation and reform. Further analyses of these mechanisms may enrich our understanding of Medicare's capacity to shape the U.S. health system.

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

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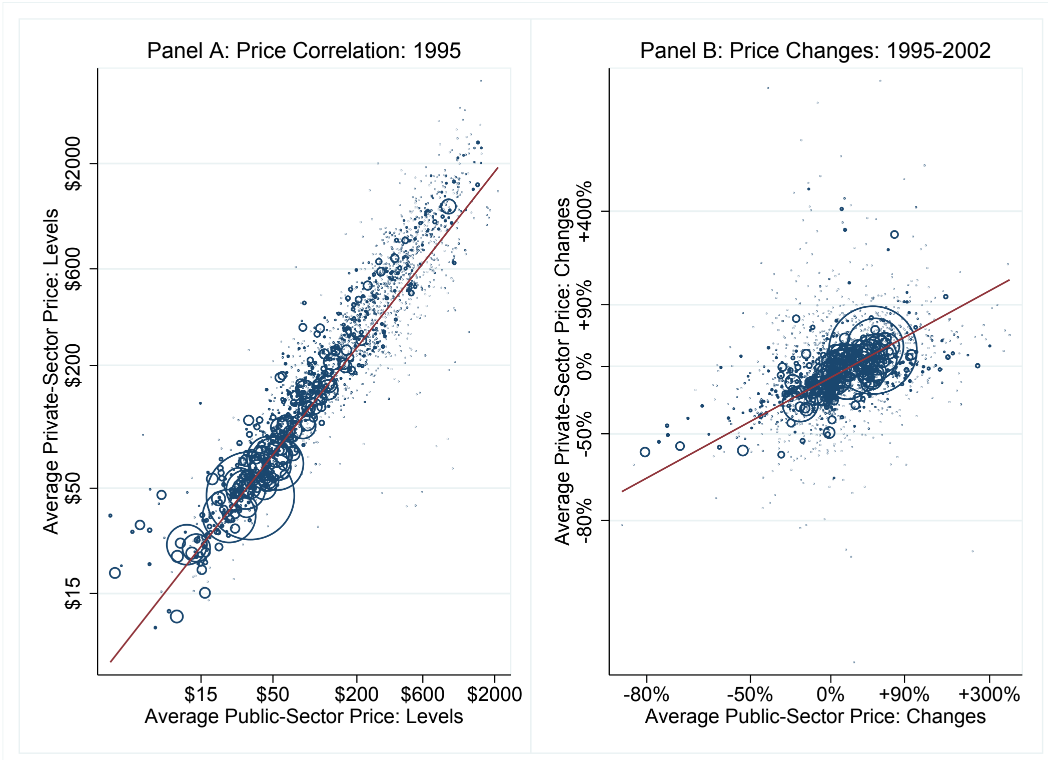
Figure 1: Cross-Service Relationship Between Private and Medicare Prices



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

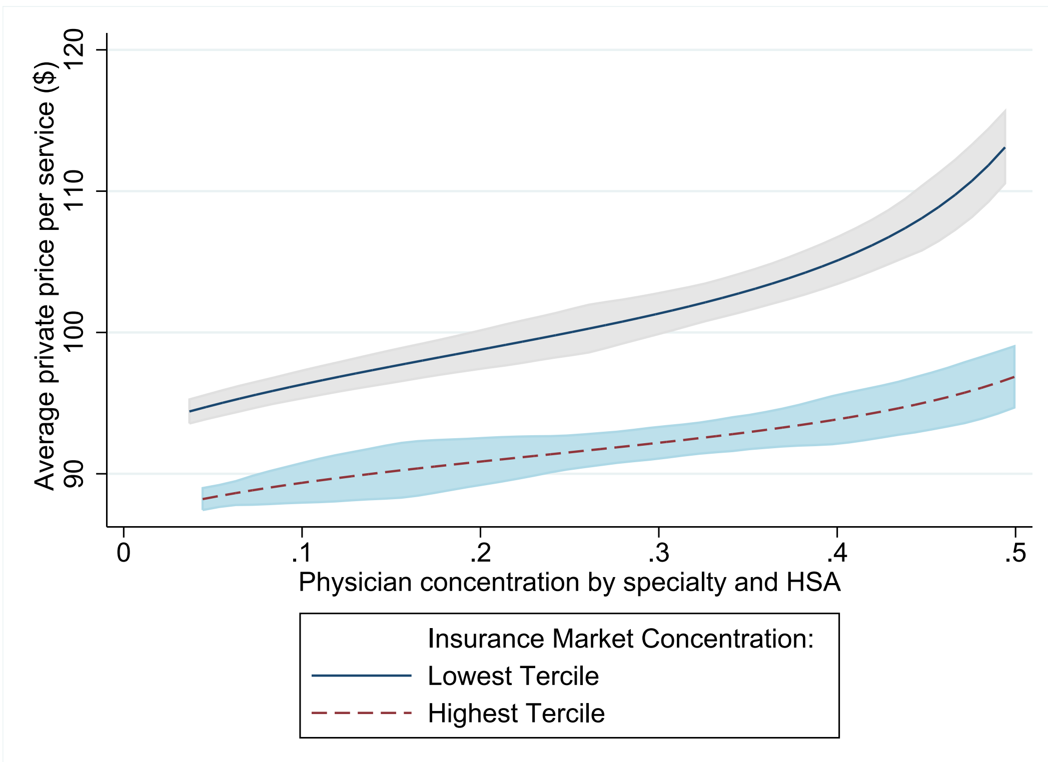
$$\ln(P_j^{\text{Private}}) = \beta_0 + \beta_1 \ln(P_j^{\text{Medicare}}) + u_j$$

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

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

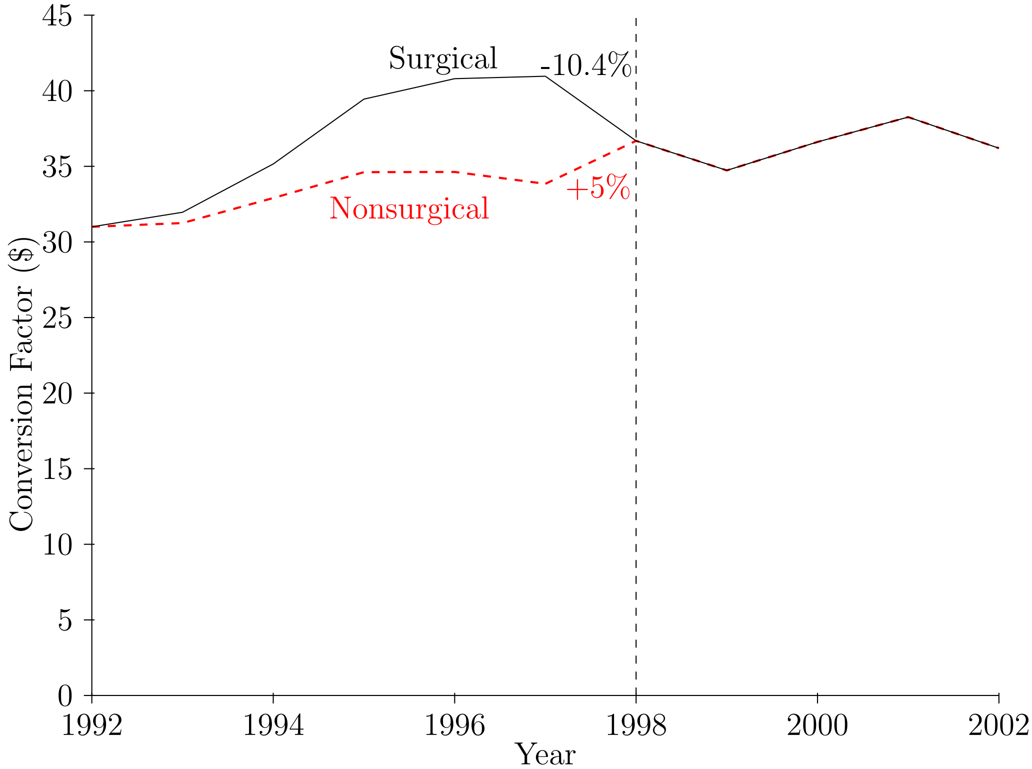
again weighted by the code's frequency. The regression yields a coefficient of $\gamma_1 = 0.65$ and $R^2 = 0.60$ with $N = 2,194$. Note that the regressions are run in logs and the values shown along the axes are computed by exponentiating the predicted values.

Figure 2: Variation in Private Prices with Provider and Insurer Market Power



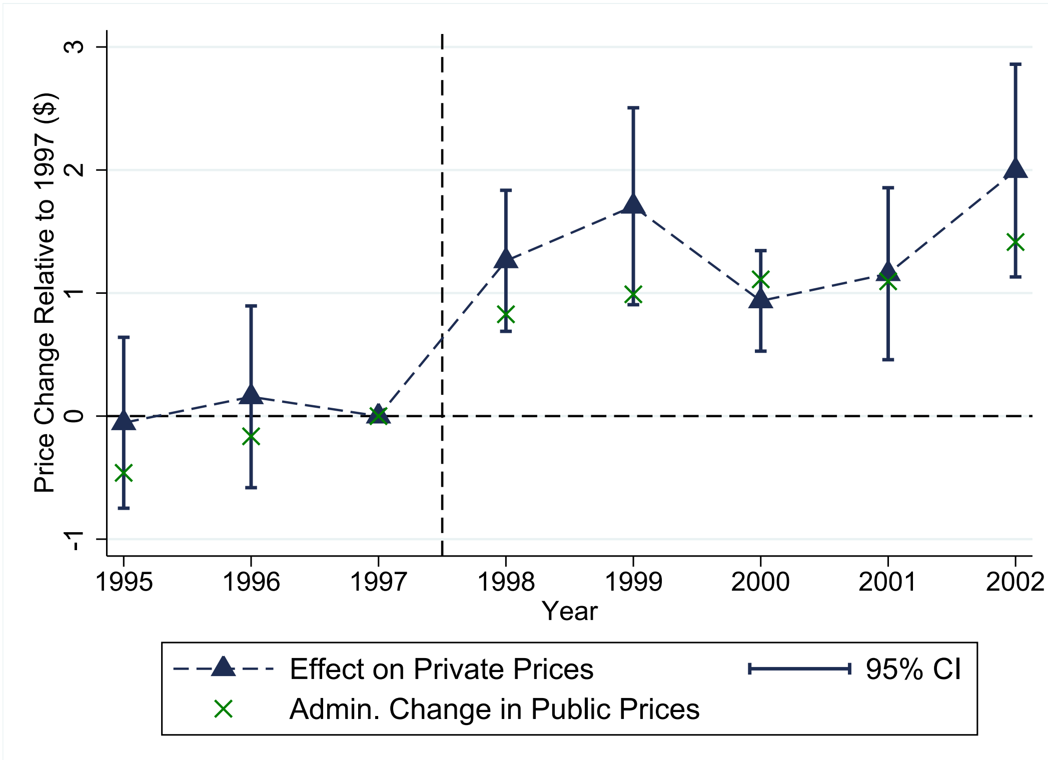
Note: This figure shows the average private sector payments separately in low-concentration (blue solid line) and high-concentration (red dashed line) insurance markets, based on the degree of provider concentration (along the x-axis). Details on construction of the concentration measures are in Appendix B.5. The private payments are averaged across all years, states, and services and smoothed using an Epanechnikov kernel with bandwidth 0.2. Source: Authors' calculations using Medicare claims and ThompsonReuters MarketScan data.

Figure 3: Evolution of Medicare Surgical and Nonsurgical Conversion Factors



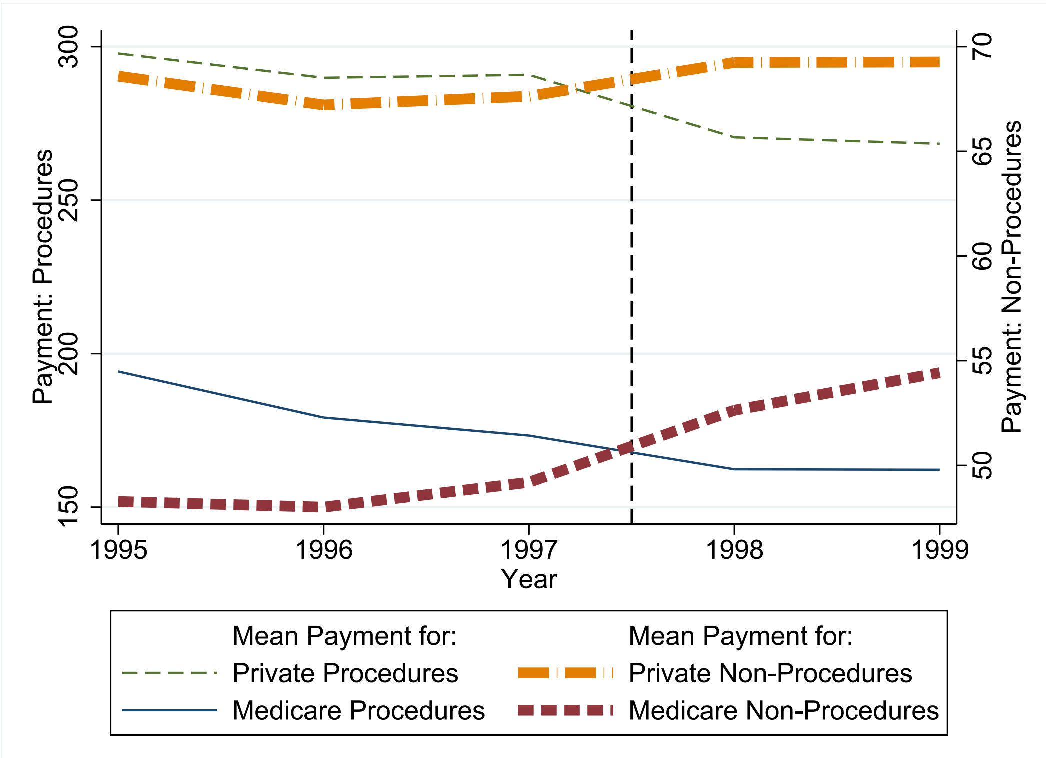
Note: This figure shows the nominal Conversion Factors that Medicare applied to surgical and general nonsurgical services for each year from 1992 through 2002. Source: *Federal Register*, various issues.

Figure 4: Effects of Medicare's Elimination of the Surgical Conversion Factor



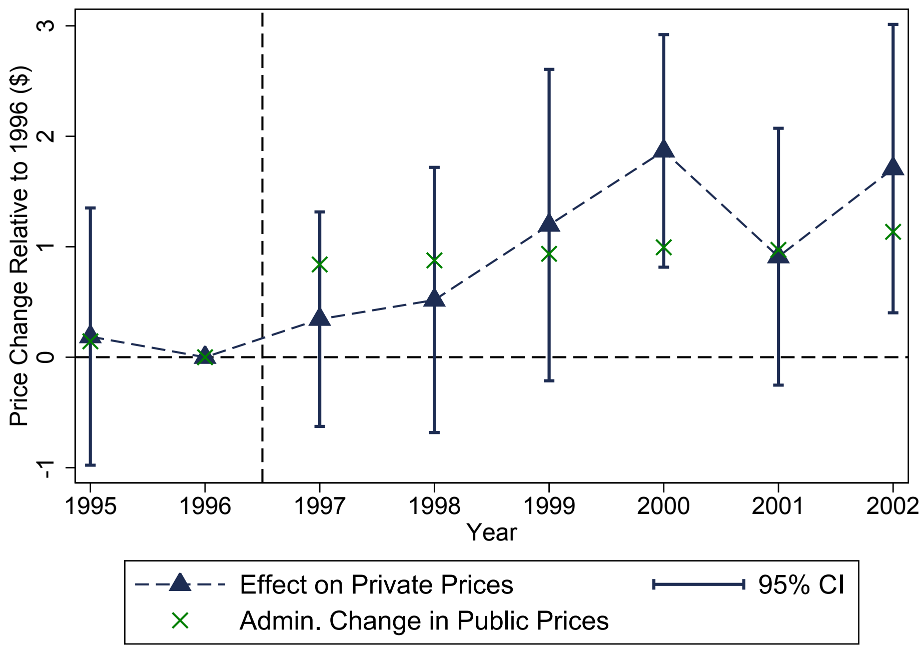
Note: This figure shows the estimates of $\hat{\gamma}_t$ and $\hat{\delta}_t$ from estimating equations (7) and (8) in section 5.2. The payment shock is based on the elimination of the separate surgical and nonsurgical Conversion Factors, and is defined in equation (3). The payment shocks are constructed such that a one unit change in the independent variable (the predicted 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, which are presented along with their 95 percent confidence intervals. Sources: *Federal Register*, various issues; Authors' calculations using Medicare claims, Thompson Reuters MarketScan data, and Ruggles et al. (2010).

Figure 5: Raw Mean Payments Around Medicare's Elimination of the Surgical Conversion Factor



Note: This figure shows the raw mean payments for surgical and nonsurgical services, and for Medicare and private insurance payments, from 1995 through 1999. Sources: Authors' calculations using Medicare and Thompson Reuters MarketScan data.

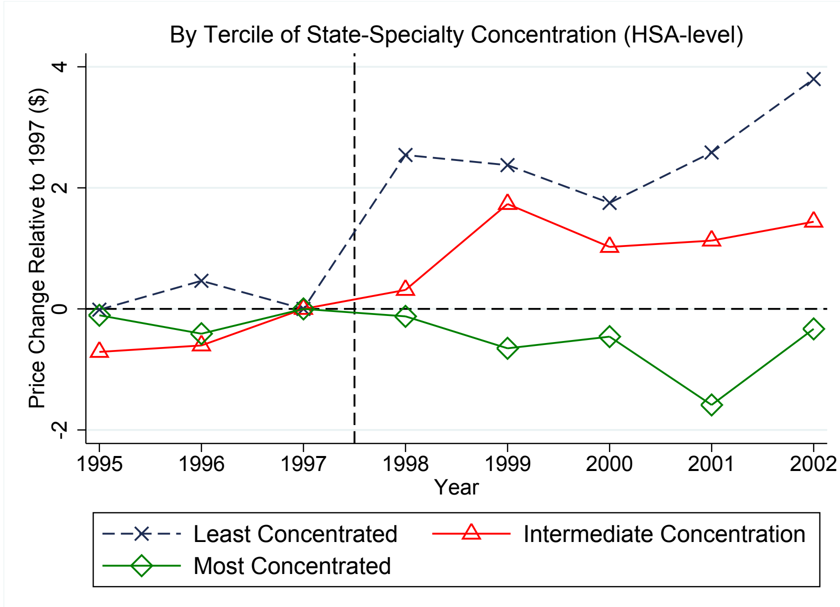
Figure 6: Effect of Geographic Payment Shocks on Private Prices



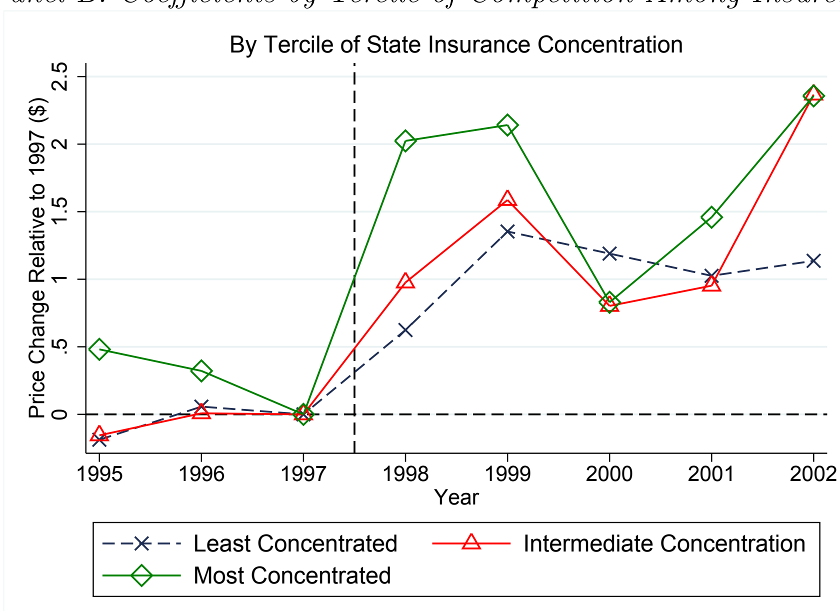
Note: This figure shows the estimates of $\hat{\gamma}_t$ and $\hat{\delta}_t$ from estimating equations (7) and (8) in section 5.2. The payment shock is based on the overhaul of Geographic Adjustment Factors, and is defined in equation (4). The payment shocks are constructed such that a one unit change in the independent variable (the predicted 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, which are presented along with their 95 percent confidence intervals. Sources: *Federal Register*, various issues; Authors' calculations using Medicare claims, Thompson Reuters MarketScan data, and Ruggles et al. (2010).

Figure 7: Reduced Form Coefficients by Tercile of Competition

Panel A: Coefficients by Tercile of Competition Among Physicians



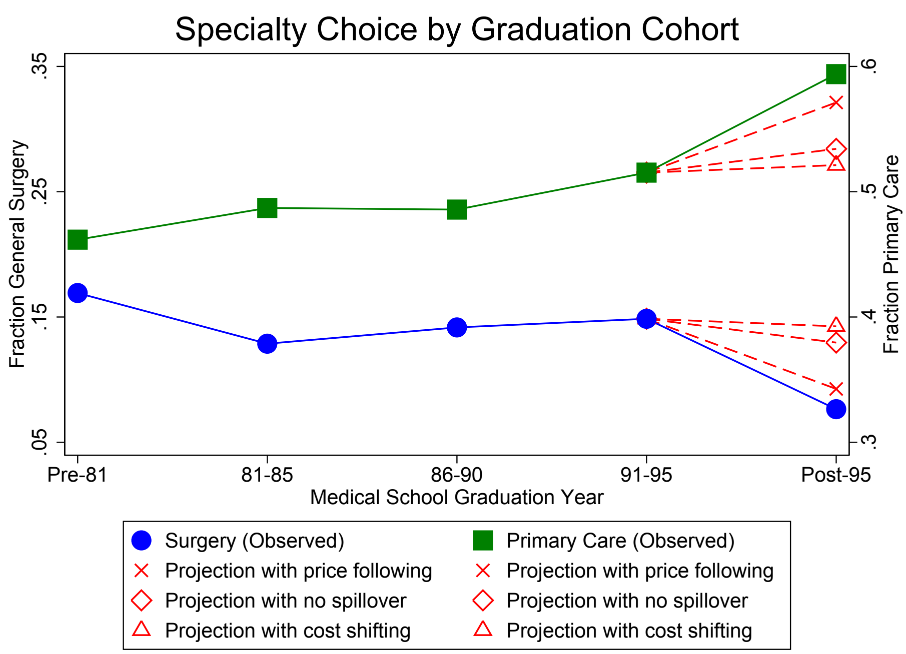
Panel B: Coefficients by Tercile of Competition Among Insurers



Note: 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). Coefficients are estimated separately when cutting the sample by (A) the HHI of physician groups, computed at the specialty-by-HSA level, and (B) the HHI of insurance carriers, computed by state.

Sources: Authors' calculations using Medicare and Thompson Reuters MarketScan data. Data Source for the insurance concentration measures: National Association of Insurance Commissioners, by permission. The NAIC does not endorse any analysis or conclusions based upon the use of its data.

Figure 8: Effect of Surgical Conversion Factor Elimination on Specialty Choice



Note: The solid lines show the shares of physicians choosing different categories of specialty, depending on their year of medical school graduation. The dashed lines show projections for the post-1995 graduation cohort, based on the three different assumptions for private market pricing used in Table 4. To generate these projections we add the numbers from Panel E of Table 4 to the primary care share from 1991–95, and subtract them from the surgery share from 1991–95. Source: Authors' calculations using Community Tracking Survey microdata from the 2004–05 wave (CSHSC 2006).

Table 1: Summary Statistics

	Mean	Std. Dev.	Min	Max
Full Sample ($N = 303,728$)				
Private Payment Per Service	\$95.34	\$169.34	\$0.30	\$21,891
Medicare Payment Per Service	\$68.38	\$104.12	\$3.28	\$2,112
Specialty-Specific Physician HHI	0.072	0.074	0.005	1.00
State Level Insurer HHI	0.300	0.204	0.077	0.95
Nonsurgical Care ($N = 140,376$)				
Private Payment Per Service	\$69.93	\$53.92	\$1.10	\$6,967
Medicare Payment Per Service	\$54.08	\$45.57	\$3.28	\$1,150
Surgical Services ($N = 163,352$)				
Private Payment Per Service	\$278.71	\$419.94	\$0.30	\$21,891
Medicare Payment Per Service	\$171.59	\$248.91	\$3.51	\$2,112

Note: This table shows summary statistics for our data on public and private payments and the characteristics of the geographic and service-specific markets that we use to explore heterogeneity in the effect of Medicare price changes on public prices. Observations are constructed at the service-by-state-by-year level and the panel is balanced in the sense that each service-by-state panel is only included if public and private prices are available for each year from 1995 through 2002. Private and Medicare Payments Per Service are expressed in dollars and are the average payment within each service-by-state-by-year cell. The construction of HHI variables is described in section 4.2. The 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. Sources: All except for line 4—Medicare claims and Thompson Reuters MarketScan data. Line 4—Authors' calculations based on data obtained from National Association of Insurance Commissioners, by permission. The NAIC does not endorse any analysis or conclusions based upon the use of its data.

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

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable:	Medicare Payment	Private Payment		Medicare Payment	Private Payment	
	1st Stage	Red. Form	IV	1st Stage	Red. Form	IV
Payment Shock	1.201**	1.386**		0.887**	0.997*	
× Post-Implementation	(0.070)	(0.258)		(0.087)	(0.470)	
Instrumented Medicare Payment			1.155**			1.124*
			(0.212)			(0.493)
N	303,728	303,728	303,728	128,694	128,694	128,694
Number of Clusters	2,194	2,194	2,194	199	199	199
Number of Services	2,194	2,194	2,194	156	156	156
Geographic Unit	State	State	State	Pre-Consolidation Payment Area		
Additional Significance Tests—p-value Against the Following Nulls:						
H_0 : coefficient = 1	0.004	0.13	0.47	0.18	0.99	0.80
H_0 : coefficient = 1.45		0.81	0.16		0.34	0.51

Note: **, *, and + indicate coefficients statistically different from zero 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 5.2. Columns 1 and 2 report estimates of equations (5) 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 (6). Columns 1 through 3 use the payment change due to the overhaul of Medicare's Conversion Factors, defined in equation (3) as the instrument. Columns 4 through 6 repeat similar specifications, but using the payment change due to the overhaul of Medicare's Geographic Adjustment Factors, defined in equation (4), as the instrument. Observations are constructed at the service-by-state-by-year level in columns 1–3, and service-by-payment locality-by-year level in 4–6. In columns 1 through 3, observations are weighted according to the number of times the service is observed in Medicare claims in 1997. The panel is balanced in the sense that each service-by-geographic unit panel is only included if public and private prices are available for each year from 1995 through 2002. Standard errors are calculated allowing for arbitrary correlation among the errors associated with each service, in columns 1–3, and arbitrary correlation among the errors associated with each payment locality, in columns 4–6. These clustered standard errors are used in the hypothesis tests shown in 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: Physician Practice Patterns

Dependent Variable:	(1) Log Hours Worked	(2) Accepting New Medicare Pts.?	(3) Accepting New Private Pts.?	(4) Career Satisfaction	(5) Board Certification
Surgeon x Post-1997	0.008	-0.039**	-0.054***	-0.026+	-0.040***
	(0.009)	(0.012)	(0.013)	(0.015)	(0.009)
N	42,950	42,950	42,950	42,950	42,776
R-Squared	0.015	0.020	0.007	0.001	0.012
Dependent Variable Mean	3.95	0.65	0.69	0.42	0.83

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 as shown in equation (9). The unit of observation is an individual physician's response in one survey wave. The sample excludes the 916 survey respondents from the 2004-2005 CTS whose graduation years were between 1996 and 2004, as these individuals' specialty choices may have been affected by the payment changes we analyze. The sample in column 5 is smaller than the sample in other columns because board certification status is sometimes reported as "inapplicable" or "not ascertained." The dependent variable is listed at the top of each column. Variable details are provided in Appendix B.6. Heteroskedasticity-robust standard errors are shown in parenthesis. Source: Authors' calculations using Community Tracking Survey microdata.

Table 4: **Effect of Conversion Factor Shock on Physician Incomes, Depending on Private Price Response**

		Assumption on private payments: (1) No spillover	(2) 40% offset	(3) Price-following
	Panel A: Predicted income change in Family Practice			
(1)	Income change in \$1,000:	\$4.0	\$0.1	\$15.3
(2)	Income change in percent:	2.9%	0.08%	11.2%
	Panel B: Predicted income change in General Practice			
(1)	Income change in \$1,000:	\$2.4	\$0.1	\$9.1
(2)	Income change in percent:	1.8%	0.05%	6.7%
	Panel C: Predicted income change in General Surgery			
(1)	Income change in \$1,000:	−\$9.4	−\$4.5	−\$23.8
(2)	Income change in percent:	−3.9%	−1.9%	−9.9%
	Panel D: Predicted income change in Dermatology			
(1)	Income change in \$1,000:	−\$13.5	−\$4.1	−\$40.9
(2)	Income change in percent:	−5.5%	−1.7%	−16.6%
	Panel E: Predicted entry into Primary Care			
	Change in share (percentage points):	1.9	0.6	5.5

This table calculates the change in incomes of different specialties that the Medicare Conversion Factor (CF) shocks in 1998 would generate based on different assumptions about the private sector's response, shown in the three columns Let SurgicalShare_s be the share of specialty s's Medicare revenue that comes from surgical services. Then $\text{SpecialtyShock}_s = -0.104 \times \text{SurgicalShare}_s + 0.05 \times (1 - \text{SurgicalShare}_s)$ is the proportional change in the specialty's Medicare reimbursements after the CF merger.

Column 1 calculates the income shock as $\text{IncomeShock}_s^A = \text{SpecialtyShock}_s \times \text{MedicareRevenue}_s$ where MedicareRevenue_s is the specialty's average revenue from Medicare patients in 1997, prior to the policy change (Gonzalez and Zhang, 1998, Table 96). In line 1 of Panels A–D, values are shown in thousands of 1997 dollars. The percent income shock in line 2 is $\text{IncomeShock}_s^A/\text{NetIncome}_s$, where NetIncome_s (from Gonzalez and Zhang, 1998, Table 83) is the specialty's average net income after expenses (which are substantial, at around half of gross revenue) in 1997.

Column 2 follows Shiels (2009) and many others in assuming that Medicare "cost-shifts" onto private payers, with the latter adjusting physician payments by 40 percent in the opposite direction from Medicare. Thus $\text{IncomeShock}_s^B = \text{SpecialtyShock}_s \times \text{MedicareRevenue}_s - 0.4 \times \text{SpecialtyShock}_s \times \text{PrivateRevenue}_s$, where PrivateRevenue_s is from Gonzalez and Zhang (1998, Tables 98 and 99). The percent income shock is calculated just as in column 1.

Column 3 uses our price-following estimate of 1.16 to calculate $\text{IncomeShock}_s^3 = \text{SpecialtyShock}_s \times \text{MedicareRevenue}_s + 1.16 \times \text{SpecialtyShock}_s \times \text{PrivateRevenue}_s$. The percent income shock is again calculated similarly.

Data Appendix

In section A.1 we describe our core datasets in further detail. In section A.2 we assess the quality of our Medicare-private merge procedure in order to determine how comprehensive and representative our final dataset is.

A.1 Data Sources

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

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

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

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

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

A.2 Comprehensiveness of the Medicare-MarketScan Merge

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

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

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

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

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

Row F imposes the requirement that the Medicare claims data be balanced. That is, it drops all remaining codes that do not appear in the Medicare data in each year from 1995 to 2002. This panel-balance requirement results in significant loss of codes in both the Medicare and MarketScan data. Just under 50 percent of the remaining service codes are lost in both data sets. These codes account for 6.5 percent of Medicare spending and 21 percent of remaining private sector spending (that is, 19 of the remaining 91 percent). These include irregularly used codes as well as codes that were eliminated over the course of the sample.

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

Appendices B through G are available in the supplemental materials online.

Appendix Table A.1: Measures of Merge Comprehensiveness

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

Note: This table presents the comprehensiveness of our merge procedure between the Medicare claims data and private sector data from MarketScan introduced in section 4.1 using data from 1995. Line (A) indicates which file we start with; columns 1 and 2 show the merge overlap when starting with Medicare and then merging in the MarketScan private data while columns 3 and 4 conduct the merge in the opposite direction. Line (B) distinguishes between measurements of the overlap in dollar terms (labeled "Spending") and in service count (labeled "Quantity"). The subsequent lines show the share of the initial dataset that survives each step of the merge procedure. Line (C) starts with the full dataset. Line (D) shows the share remaining after matching codes with the opposite dataset from the one listed in line (A). Line (E) shows the final share remaining after balancing the panel. Sources: Authors' calculations using Medicare claims and Thompson Reuters MarketScan data.