Online Appendix to: "In the Shadow of a Giant: Medicare's Influence on Private Physician Payments"

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Appendix For Online Publication Only

B Additional Data Descriptions

In this Appendix, section B.1 discusses the external validity of our results. Section B.2 describes the types of insurance plans represented in our MarketScan sample, and section B.3 presents details on relative public and private prices in the cross-section. Section B.4 describes the construction of some control variables, and then section B.5 goes into further detail on our construction of HHI measures. Finally, section B.6 introduces an additional data source, the Community Tracking Study (CTS). CTS data allow us to examine directly welfare-relevant outcomes involving physicians' behavior and choice of specialty.

B.1 Applicability Beyond Large Self-Insured Employers

The MarketScan data come from a particular segment of the insurance market, namely self-insured plans operated by large employers. In 2008, coverage under such plans accounted for 55 percent of all privately insured individuals Fernandez (2010). Our estimates' direct applicable to this market segment this gives them non-trivial scope. Note, however, that there is good reason to expect the underlying economic forces to be relevant in other market segments. Whether the coverage is through a large, self-insuring employer or a small-group plan, for example, physician payments must be competitive with the rates paid by other insurers, of which Medicare is the largest. Further, to the extent that Medicare benchmarking reflects a desire to limit contract complexity, it may well be even more pervasive in the contracts associated with small-scale insurance arrangements than in those associated with large firms. While the magnitudes of either force may thus be larger or smaller in other market segments, one should expect them to be relevant in some form or another across a broad range of insurance arrangements.

Second, consider the case of HMOs. HMOs are generally lower-cost insurance plans, and if we would expect any type of insurance plan to use a different payment scheme, HMOs would be first in line. Even so, documentary evidence indicates that at least some notable HMOs use Medicare-linked payments. For example, some insurers' provider newsletters make explicit their use of Medicare's relative values within HMO plans (*e.g.* Blue Cross and Blue Shield of Texas, 2010; Anthem Blue Cross and Blue Shield, 2012). These are non-trivial applications, as the HMO Blue Texas plan advertised having 38,000 physicians in its provider network as of 2009 (Blue Cross and Blue Shield of Texas, 2014).

B.2 Insurance Plan Types in the MarketScan Data

The MarketScan data contain private sector prices that come from a range of insurance plan types. In 1996, 38 percent of service claims came from Major Medical or Comprehensive Insurance (CI) plans, 52 percent from less generous Preferred Provider Organization (PPO) plans, and 10 percent from even more restrictive Point of Service (POS) plans. By 2006, 8 percent of MarketScan 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 two control variables designed to capture the evolution of generosity in plan types across space and over time. They are described in Appendix B.4.

The Community Tracking Study (CTS), discussed in more detail in Appendix B.6, provides important context about the payment methods used by various types of insurance plans. Despite the trend away from comprehensive insurance, and towards more restrictive managed care plans, CTS data reveal that these changes had little impact on insurers' methods for paying physician groups. Between 1995 and 2004, the fraction of physicians' revenues associated with capitated, as opposed to fee-for-service, payments *declined* from 16 percent to 13 percent (CSHSC 1999, 56; 2006, 4-29). This reflects the fact that many managed care arrangements ultimately pay for at least some physician care through relatively traditional fee for service arrangements.[1]

B.3 A Descriptive Look at the Relationship between Medicare and Private Payments

In addition to its predictions about price-following, the model in section 3 gives guidance about other features of how private and public prices relate. As long as physicians have the outside option of treating Medicare patients, they will only reach agreement with the private insurers to treat the latter's patients when these insurers pay at least Medicare rates, scaled by α. While we are not confident about a universal relative cost of treating private versus Medicare patients, footnote 5 in the paper offers evidence suggesting that $\alpha \approx 1$. When this is true, private prices will generally exceed Medicare's.

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

Panel B, and Figure B.1, show the magnitudes of the differences between public and private prices. Figure B.1 shows the distribution of log public-private price differences at the state-year-service level. For the vast majority of these cells, this difference is strongly

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

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

positive. Instances in which public payments exceed private payments are relatively few, and the magnitudes swamped by the typical private-over-public mark-up. Panel B of Table B.1 quantifies this fact.

To further summarize the cross-sectional relationship between public and private sector prices, Panel C of Table B.1 shows results from regressions of private against Medicare prices in levels. The coefficient on Medicare's payments is consistently near 1.45, reflecting the average mark-ups apparent in Panel B and Figure B.1. Controlling for year or for state-by-year effects has no effect on this relationship. These facts comport well with our model and suggest that it captures some of the key facts about private insurance payments. Thus it might also help us to understand the nature of price-following.

B.4 Variable Construction

To control for plan characteristics, we construct a variable called "Insurance Plan Type Control" by first regressing payments on plan type indicators. Using the resulting coefficients and changes in the plan type composition, we generate predicted payments that we aggregate to the state-by-year-by-service level. We also construct a control for plan generosity based on patient cost sharing. This variable, "Cost Sharing Fraction," 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. Summary statistics for these variables are shown in Table B.2. Because changes in plan types unfolded relatively smoothly over time, we would expect any associated concerns to reveal themselves in the event study estimates of equation (8).

Different types of physician groups may have different outside options, so the framework of section 3 predicts different levels of price-following depending on a region's composition of large *versus* small groups.

To take this view to the data, we will proxy for the presence of small physician groups using two variables. The first is the share of physicians working as sole practitioners, measured at the level of specialties and hospital referral regions. The second, similarly constructed, is the share of physicians working in groups with five or fewer members.

We will also examine Medicare's size relative to the private market. We measure this as ratio of the number of times a service appears in a single year of the Medicare claims data to the number of times it appears in a single year of the MarketScan data. Because MarketScan is a non-random sample of the private market, with time-varying size, the variable would poorly characterize the actual relative sizes of public and private markets. Nonetheless, it should form a reasonable basis for dividing services into those with relatively large and small Medicare market shares. This variable is strongly right skewed; the lower bound of the relevant z-scores is roughly -0.2 for Private Market Volume and -0.4 for Medicare Relative Size. Consequently, we normalize it using percentile ranks rather than z-scores. We subtract 0.5 from the percentile ranks so that the resulting variables are symmetric about 0.

B.5 Detailed Description of the Construction of the Physician and Insurer HHI Variables

We proxy for the degree of physician competition by computing a Herfindahl-Hirschman Index (HHI) for each market. These HHIs are intended to proxy for the bargaining weight θ in the framework of section 3. When physicians are relatively concentrated, they are likely in a position to extract most of the surplus from joining an insurer's network. In contrast, when insurers are more concentrated, the carriers will likely obtain most of the surplus.

We compute the physician HHIs as follows. We first identify physician groups in the Medicare claims data using the tax identifier associated with each claim. These tax IDs indicate the physician, group, or legal entity that Medicare reimburses for the care. These IDs generally also identify the units that negotiate with insurers.[3] In claims data from a 20 percent sample of all Medicare beneficiaries, we should come close to capturing all Medicare-serving physicians in the country. Treating each Hospital Service Area (HSA) as the relevant market, we first measure the HHI across physician groups within an HSA.[4] We then average this measure across the HSAs within each state to measure the average degree of competition across the markets within that state. HSAs are an imperfect approximation of the relevant market, so we have confirmed that the results are similar when measured using the larger Hospital Referral Regions (HRRs).[5] Dranove and Ody (2014) show that, for hospitals at least, HRR-based HHIs are highly correlated with finer measures of market power. This process gives us our first proxy for physician competition, which varies at the state level.

We next compute a more targeted measure of concentration that varies across specialties as well as states. For this metric we construct HSA-level HHIs separately for each of the 32 largest physician specialties. We again average these specialty-specific HHIs across the HSAs within each state. Table 1 reports summary statistics describing both measures of provider consolidation. On average, the specialty-specific HHIs exhibit greater concentration since they consider smaller markets. They also exhibit more variation than the all-physician HHIs.

We measure insurance competition using data from the National Association of Insurance Commissioners (NAIC)'s health insurance reports.[6] Using NAIC data on each insurance carrier's size in each state, we are able to compute state-level HHIs for all states but Cal-

[3]The billing groups may not agree exactly with the negotiating units because of independent practice associations (IPAs), which negotiate as a bloc but bill separately. But the tax IDs should nevertheless be a close approximation. Pope et al. (2002), Welch et al. (2013), Baker et al. (2014), and other authors have previously made the same approximation.

[4]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 Service Area i, and where $s_{k,i}$ expresses the number of physicians in group k as a share of all physicians in region i. The measure is constructed such that an index of 1 corresponds to a monopolist and a market approaches perfect competition as the index goes to 0.

[5]Appendix Figure C.1 shows that Figure 7 from the paper is virtually unchanged when switching to HRR-based concentration measures.

[6]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 et al. (2011). We thank Dafny et al. for useful information on NAIC and other data sources in the paper and via personal communication.

ifornia.[7] We compute HHIs based on NAIC data on enrollment in comprehensive group insurance plans in 2001.[8]

Concentration and Private Prices

Figure 2 from the main text provides suggestive evidence that our HHI measures do indeed capture economically relevant aspects of competition. The last two regressions of Table B.1 further flesh out the relationship between private payments and our concentration metrics. these facts. We first standardize the HHI variables as z scores. Column 4 of Panel C includes the physician HHI alongside the services' average Medicare payment. Column 5 adds the insurance HHI and interactions between the HHI measures and the Medicare payment. As in Figure 2, more concentrated insurance markets are associated with lower reimbursements while more concentrated physician markets are associated with higher reimbursements. Since we have not isolated exogenous variation in these measures of market structure, we do not ascribe a causal interpretation to these results. But they do demonstrate the model's general consistency with the data. We therefore believe that section 3 presents a useful framework for exploring possible of price-following.

B.6 Community Tracking Study Questions

Our analysis of welfare-relevant outcomes comes from the physician surveys conducted by the Community Tracking Study during its 1996–97, 1998–99, 2000–01 and 2004–05 waves. We obtain these data from ICPSR, where they are available under study numbers 2597, 3267, 3820, and 4584, respectively. We rely on the following questions:

- PCPFLAG: A flag indicating that the physician spends most of his/her patient care time in a primary care specialty.
- SPECX: Physician's specialty based on responses to questions A8 and A10 and grouped into seven types of specialties. Ob/Gyn and Psychiatrists are separated out because these types of physicians were asked specific questions in the survey.
- GRADYRX: Year physician graduated from medical school. For confidentiality reasons, years before 1966 were bottom coded (GRADYRX=1), and years after 1995 were top coded (GRADYRX=8). For confidentiality reasons, this is a categorical variable which groups years together into approximate 5 year intervals.
- NWMCARE: Is the practice accepting all, most, some, or no new patients who are insured through Medicare, including Medicare managed care patients?
  (We code this variable as: All = 1, any other response = 0)

[7] 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.

[8] 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.

- NWPRIV: Is the practice accepting all, most, some, or no new patients who are insured through private or commercial insurance plans including managed care plans and HMOs with whom the practice has contracts? This includes both fee for service patients and patients enrolled in managed care plans with whom the practice has a contract. It excludes Medicaid or Medicare managed care.
  (We code this variable as: All = 1, any other response = 0)

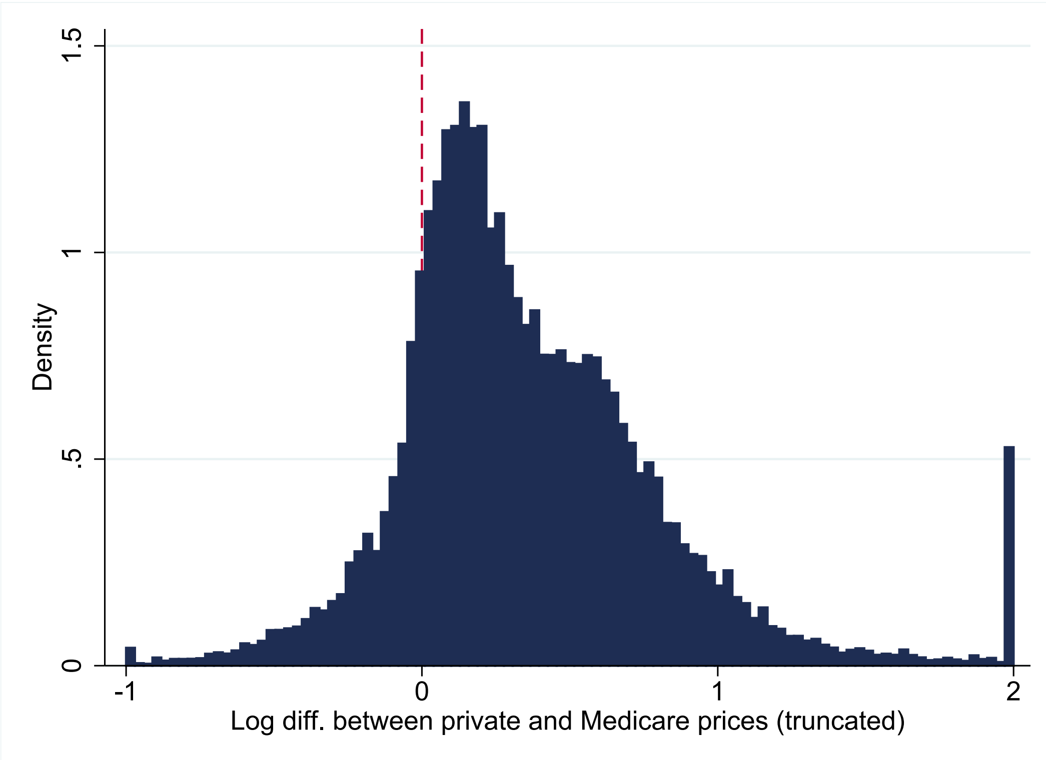
- CARSAT: Many of the remaining questions are about your practice and your relationships with patients. Before we begin those questions, let me ask you: Thinking very generally about your satisfaction with your overall career in medicine, would you say that you are CURRENTLY: Very satisfied, Somewhat satisfied, Somewhat dissatisfied, Very dissatisfied, Neither satisfied/dissatisfied?
  (We code this variable as: Very satisfied = 1, any other response = 0)

- BDCERT: Board certification status of physician. This variable summarizes the certification/eligiblity [sic] status of the physician in any specialty or subspecialty.
  (We code this variable as: Board certified = 1, any other response = 0)

- HRSPATX: Number of hours physician spent in direct patient care activities during last complete week of work.

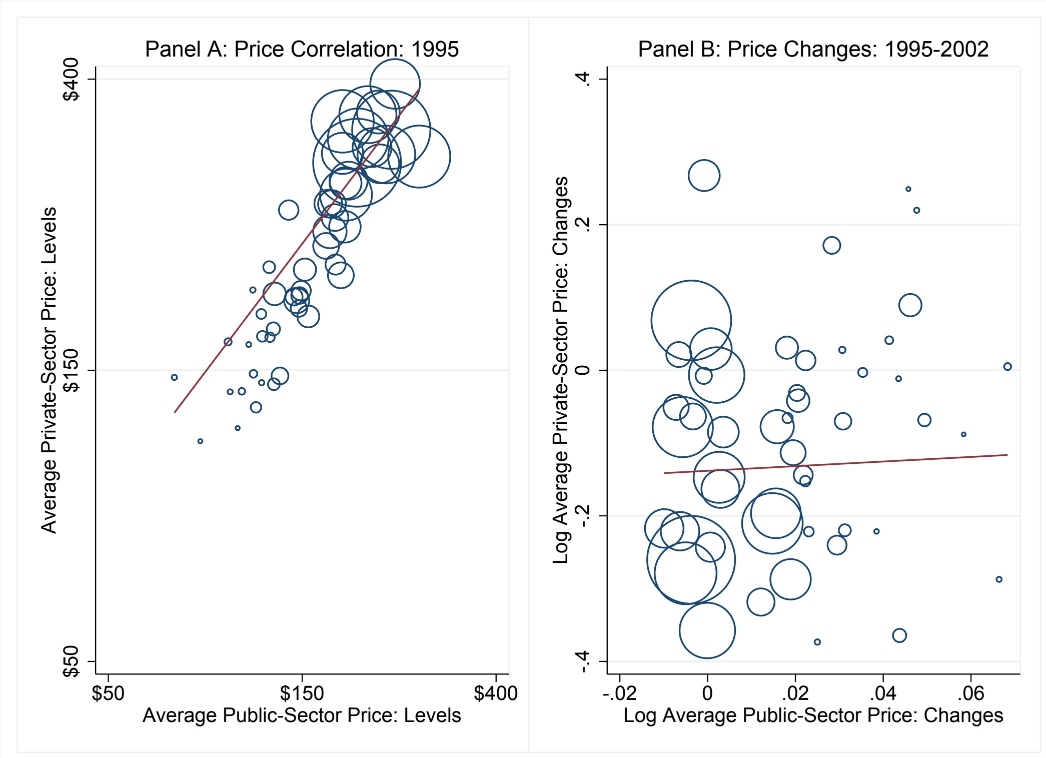
- HRSMEDX: Number of hours physician spent in medically related activities during last complete week of work.

Appendix Figure B.1: Distribution of Private-Medicare Price Difference



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

Appendix Figure B.2: 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 (MarketScan) 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

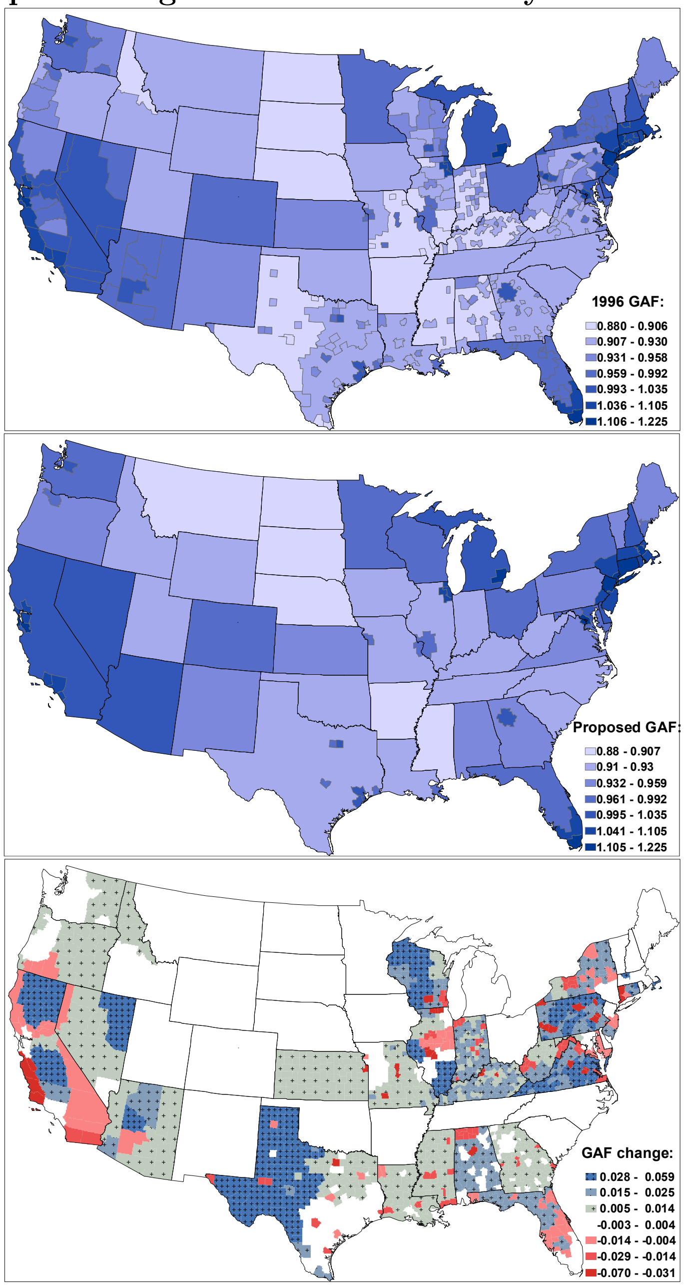
$$\ln(P_s^{\text{Private}}) = \beta_0 + \beta_1 \ln(P_s^{\text{Medicare}}) + u_s$$

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 predicted values.

Appendix Figure B.3: Medicare Payment Areas



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

Appendix Table B.1: **Summarizing the Differences Between Medicare and Private Prices**

Panel A: Share of Observations with Private Rates Exceeding Medicare

		(1)	(2)	(3)	(4)	(5)
	Weighting:	None	Service Count in		Total Spending in	
			Medicare	Private	Medicare	Private
(1)	Based on mean price	0.793	0.885	0.825	0.844	0.821
(2)	Surgical only	0.853	0.911	0.882	0.908	0.915
(3)	Non-surgical only	0.722	0.880	0.810	0.814	0.756
(4)	Based on t-test probabilities	0.793	0.885	0.825	0.844	0.819

Panel B: Mean Difference Between Log Private and Medicare Prices

		(1)	(2)	(3)	(4)	(5)
	Weighting:	None	Service Count in		Total Spending in	
			Medicare	Private	Medicare	Private
	Among services with:					
(1)	$P^{\text{Medicare}} > P^{\text{Private}}$	-0.309	-0.122	-0.155	-0.162	-0.176
(2)	$P^{\text{Medicare}} < P^{\text{Private}}$	0.548	0.394	0.413	0.408	0.493

Panel C: Regression of Private on Medicare Prices

	(1)	(2)	(3)	(4)	(5)
Public payment	1.455**	1.457**	1.457**	1.457**	1.397**
	(0.057)	(0.056)	(0.056)	(0.056)	(0.062)
Physician HHI				1.918**	3.511*
				(0.443)	(1.430)
Insurance HHI					-4.239**
					(1.394)
Public payment					-0.040+
× Specialty HHI					(0.022)
Public payment					0.085**
× Insurance HHI					(0.026)
N	222,040	222,040	222,040	222,040	222,040
Number of Clusters	1,364	1,364	1,364	1,364	1,364
R^2	0.811	0.812	0.816	0.812	0.826
Fixed Effects	None	Year	State-Year	Year	Year & Specialty
Physician HHI Measure				All MDs	Specialty

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

Appendix Table B.2: **Additional Summary Statistics For Robustness Analysis**

	Mean	Std. Dev.	Min	Max
HMO Share 1997	20.48	12.40	0.000	47.20
Specialty Sole Practice Share	0.67	0.13	0.000	1.00
Medicare Relative Size Proxy	23.39	44.14	0.003	459.83
Insurance Plan Type Control	-0.68	0.53	-1.13	3.80
Cost Sharing Fraction	0.17	0.15	0.00	1.00

Note: This table shows additional summary statistics relevant for the analyses in this appendix. Sources: The HMO share variable is from Insterstudy as reported in the Statistical Abstract of the United States. Specialty Sole Practice Share is the share of sole practitioners by specialty, based on authors' calculations using Medicare claims data. Medicare Relative Size Proxy is based on authors' calculations using Medicare and ThompsonReuters MarketScan data. The final two rows are based on authors' calculations using MarketScan data. Further details on variable construction are provided in section B.4.

C Robustness Tests

This appendix presents a variety of robustness checks mentioned in the paper's main text. Appendix Figure C.1 reproduces Panel A of Figure 7, but with concentration measured at the HRR level rather than the HSA level. Appendix Tables C.1 through C.5 report various robustness checks on our baseline results. With the exception of Table C.3, which we describe below, these tables are described in sufficient detail in the section 6 of the paper.

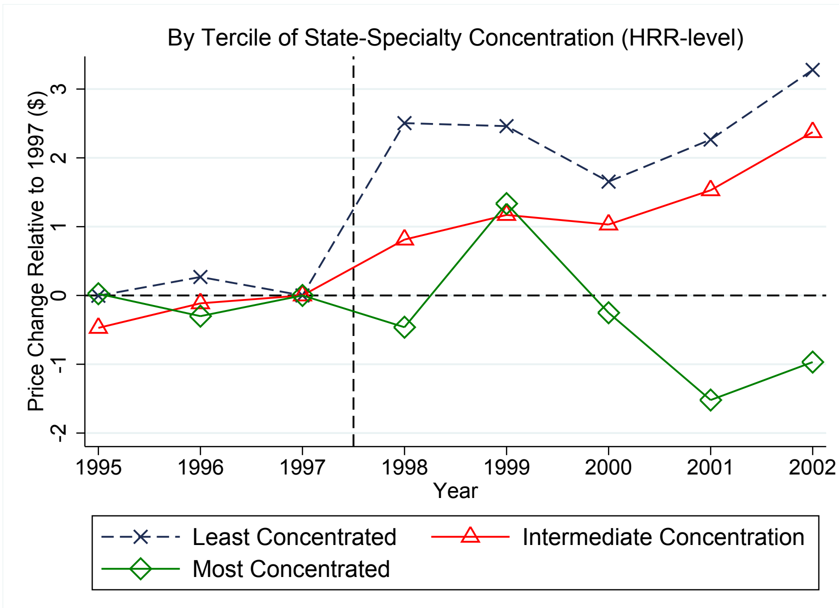
Table C.3 explores the stability of our reduced form results when estimated at different levels of aggregation. For this table, we run reduced-form price-following regressions at levels of aggregation ranging from the individual claim up to national aggregates by service. In between these extremes are state (our baseline geography), the Dartmouth Hospital Referral Region (HRR) and Hospital Service Area (HSA).[9] Specifically, we include estimates from each of the following regressions, where j indexes services, t indexes years, s indexes a geographic unit (state, HRR, or HSA), and i indexes individual claims:

$$\begin{aligned} P^{\text{Private}}_{i,j,s,t} &= \beta \cdot \text{PredChg}^{\text{Medicare}}_{j} \times \text{Post1998}_t + X_{i,j,s,t}\psi + \mu_j \mathbb{1}_j + \mu_s \mathbb{1}_s + \mu_t \mathbb{1}_t \\ &\quad + \mu_{j,s}\mathbb{1}_j \cdot \mathbb{1}_s + \mu_{t,s}\mathbb{1}_t \cdot \mathbb{1}_s + e_{i,j,s,t} \qquad \text{(C.1)} \\ \overline{P^{\text{Private}}}_{j,s,t} &= \beta \cdot \text{PredChg}^{\text{Medicare}}_{j} \times \text{Post1998}_t + X_{j,s,t}\psi + \mu_j \mathbb{1}_j + \mu_s \mathbb{1}_s + \mu_t \mathbb{1}_t \\ &\quad + \mu_{j,s}\mathbb{1}_j \cdot \mathbb{1}_s + \mu_{t,s}\mathbb{1}_t \cdot \mathbb{1}_s + e_{j,s,t} \qquad \text{(C.2)} \\ \overline{P^{\text{Private}}}_{j,t} &= \beta \cdot \text{PredChg}^{\text{Medicare}}_{j} \times \text{Post1998}_t + X_{j,t}\psi + \mu_j \mathbb{1}_j + \mu_t \mathbb{1}_t + e_{j,t}. \qquad \text{(C.3)} \end{aligned}$$

Regression (C.1) is at the individual claim level, regression (C.2) at the geographic area-by-service-by-year level ($j \times s \times t$), and regression (C.3) is at the service-by-year level ($j \times t$)—*i.e.* national aggregation. We use three different geographic levels for s: state (as in the baseline from the main text), HRR, and HSA. Table C.3 below shows very similar coefficients across these different levels of aggregation. Our baseline estimates are thus essentially unchanged by varying the level of aggregation.

[9] These regions are defined by *The Dartmouth Atlas of Health Care* and the definitions are available at http://www.dartmouthatlas.org/tools/downloads.aspx?tab=39 (accessed August 29, 2015).

Appendix Figure C.1: Effect of Surgical Payment Shock on Prices



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 the HHI of physician groups, computed at the specialty-by-HRR level.

Sources: Authors' calculations using Medicare and Thompson Reuters MarketScan data.

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

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent Variable:				Private Payment Level			
Public Payment	1.160**	1.155**	1.160**	0.971**	0.674**	1.321**	1.114**
	(0.225)	(0.212)	(0.226)	(0.107)	(0.112)	(0.261)	(0.250)
Insurance Plan Type Control	-11.493		-11.429	-32.530	22.656**	-2.937	-11.843
	(28.164)		(28.217)	(40.462)	(4.755)	(17.765)	(28.023)
Cost Sharing Fraction	-2.311	-2.253		-2.404	1.767**	-3.091	-2.391
	(6.143)	(6.213)		(6.024)	(0.190)	(5.962)	(6.125)
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

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. 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

	(1)	(2)	(3)
Dependent Variable:	Public Payment	Private Payment	
	1st Stage	Red. Form	IV
Payment Shock × Post 1997	1.201**	1.393**	
	(0.070)	(0.260)	
Public Payment			1.169**
			(0.209)
N	17,552	17,552	17,552
Number of Clusters	2,194	2,194	2,194
Number of Services	2,194	2,194	2,194
Geographic Unit	National	National	National

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. 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 1995 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: The Effect of Aggregation on Reduced Form Price-Following Estimates

	(1)	(2)	(3)	(4)	(5)
Dependent Variable:			Private Payment Level		
Payment Shock × Post-1997	1.290**	1.386**	1.294**	1.261*	1.393**
	(0.291)	(0.258)	(0.389)	(0.434)	(0.260)
Weighted	No	Yes	Yes	Yes	Yes
Level of Aggregation	Claim	State	HRR	HSA	National
N	143,832,536	303,728	579,192	809,040	17,552
Number of Services	2,168	2,194	1,388	1,173	2,194
R^2	0.530	0.900	0.921	0.919	0.989

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 equations (C.1) through (C.3). Column 1 estimates equation (C.1) on the underlying claims data, where the unit of observation is an individual claim. Columns 2 through 4 show estimates of (C.2), with aggregation to the geographic level indicated in each column. Column 5 reports our estimate of equation (C.3), where data are aggregated up to the service-by-year level. Both the payment shock and private payment outcome are expressed in dollar terms. Observations are constructed at the level indicated in each regression. Standard errors are calculated allowing for arbitrary correlation among the errors associated with each service. Sources: Authors' calculations using Medicare claims and Thompson Reuters MarketScan data.

Appendix Table C.4: Jointly Estimating the Effects of Both 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 × Post-1997	1.248**	1.106**		1.249**	1.107**	
	(0.108)	(0.173)		(0.032)	(0.217)	
Geographic Shock × Post-1996				0.926**	1.032*	
				(0.085)	(0.470)	
Instrumented Medicare Payment			0.886**			0.888**
			(0.162)			(0.162)
N	128,712	128,694	128,694	128,712	128,694	128,694
Number of Clusters	156	156	156	199	199	199
Number of Services	156	156	156	156	156	156
Geographic Unit	Pre-Consolidation Payment Area			Pre-Consolidation Payment Area		
Additional Significance Tests—p-value Against the Following Nulls:						
H_0 : coefficient = 1	0.02	0.54	0.48			0.49
H_0 : coefficient = 1.45		0.05	0.001			0.001
H_0: coefficients equal				0.0005	0.89	

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 equation (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 also 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-payment locality-by-year level. The panel is balanced in the sense that each service-by-locality 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.

Appendix Table C.5: Baseline Estimates Without Imposing Panel Balance

	(1)	(2)	(3)
Dependent Variable:	Medicare Payment	Private Payment	
	1st Stage	Reduced Form	IV
Payment Shock $\times$ Post 1997	1.212**	1.232**	
	(0.075)	(0.273)	
Public Payment			1.017**
			(0.212)
N	593,705	593,705	593,705
Number of Clusters	3,386	3,386	3,386

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 5.2. Columns 1 and 2 report estimates of equation (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). Observations are constructed at the service-by-state-by-year level, and are weighted according to the number of times the service is observed in Medicare claims in 1997. Standard errors are calculated allowing for arbitrary correlation among the errors associated with each service. 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.

D Further Analysis of Heterogeneity in Medicare's Influence

The main text presents sub-sample analyses that explore the relationship between market characteristics and the size of Medicare's effects on private payments. The analysis emphasizes the characteristics most tightly linked to parameters in the model from section 3, namely concentration on the physician and insurer sides of the market. An important limitation of the analysis is that these market characteristics are not randomly assigned. This appendix has three objectives. First, we formally test the statistical significance of differences in Medicare's influence in high and low concentration markets by estimating a regression model containing the relevant interactions. Second, we examine heterogeneity along other dimensions of interest, as discussed briefly in section 6.3 of the main text. Third, we show that the estimated differences in the strength of Medicare's effects across markets are robust to controlling quite flexibly for interactions between Medicare's payment changes and a variety of regional economic and demographic characteristics.

D.1 Estimation Framework

This appendix's examination of heterogeneity in the strength of Medicare's effects involves an augmented version of our baseline regression specification. To the baseline specification, we add interactions between the predicted Medicare price change and a vector of market characteristics, denoted $Z_{j,s}$:

$$\begin{aligned} P_{j,s,t}^{\text{Private}} = {} & \beta_1 \cdot \text{PredChg}_j^{\text{Medicare}} \times \text{Post1998}_t + \beta_2 \cdot \text{PredChg}_j^{\text{Medicare}} \times \text{Post1998}_t \times Z_{j,s} \\ & + \mu_j^1 \mathbb{1}_j + \mu_s^1 \mathbb{1}_s + \mu_t^1 \mathbb{1}_t + \mu_j^2 \mathbb{1}_j \cdot Z_{j,s} + \mu_s^2 \mathbb{1}_s \cdot Z_{j,s} + \mu_t^2 \mathbb{1}_s \cdot Z_{j,s} \\ & + \mu_{j,s}^1 \mathbb{1}_j \cdot \mathbb{1}_s + \mu_{t,s}^1 \mathbb{1}_t \cdot \mathbb{1}_s + \mu_{j,s}^2 \mathbb{1}_j \cdot \mathbb{1}_s \cdot Z_{j,s} + X_{j,s,t}\gamma_1 + X_{j,s,t} \cdot Z_{j,s}\gamma_2 + e_{j,s,t} \end{aligned} \tag{D.1}$$

We allow the coefficients on all time-varying controls to vary with the relevant interaction variable.[10] The variables in $Z_{j,s}$ include the market characteristics of interest as well as a variety of potential confounds. The case for causal interpretations in the estimates of sources of heterogeneity is, of course, not as strong as for our estimates of Medicare's average effects. Nonetheless, our exploration of sensitivity to the inclusion of potential confounds sheds light on the robustness of the relationship between the strength of Medicare's influence and the characteristics of primary interest.

D.2 Results

Column 1 of Table D.1 presents an estimate of equation (D.1) in which the market characteristic of interest is physician concentration. The relationship between the price following

[10] In equation (D.1) we have omitted interactions between the heterogeneity dimension and the state-by-service code fixed effects ($\mathbb{1}_j \cdot \mathbb{1}_s \cdot Z_{j,s}$). Some characteristics will vary only at the state level, so the state and state-by-service fixed effects already account for any heterogeneity along these dimensions. Others do vary at the state-by-service level, but we have omitted the three-way interaction ($\mathbb{1}_j \cdot \mathbb{1}_s \cdot Z_{j,s}$) because of the computational burden.

coefficient and physician concentration is statistically significant. Column 8 of Table D.1, along with Table D.2, show that this result is robust to controlling for interactions between the Medicare price shock and numerous area characteristics. This includes controlling for interactions between the price shock and state fixed effects, as reported in column 8.[11]

Column 2 of Table D.1 presents an estimate of equation (D.1) in which the market characteristic of interest is insurer concentration. The relationship between the price following coefficient and insurer concentration is statistically significant. Table D.3 shows that this result is robust to controlling for interactions between the Medicare price shock and numerous area characteristics. An additional aspect of insurance carriers' leverage, in particular during the time period we consider, may be the threat to physicians' payment rates posed by HMOs (Cutler, McClellan, and Newhouse, 2000). In column 3 we add an interaction between the payment shock and the state's HMO market share in 1997. HMO-abundant states indeed have more price-following, and column 7 shows that this is true even after controlling for the insurance market HHI.[12]

Column 4 of Table D.1 considers the share of small physician groups in a market. It shows that a high prevalence of small groups is strongly associated with the strength of Medicare's influence on private prices. This result is robust to varying the definition of small groups and to simultaneously controlling for physician concentration, as shown in columns 7 and 8.

We next consider heterogeneity in price-following according to Medicare's size relative to the private market. We find that the public-private ratio enters significantly, with a coefficient of 1.3. The larger the relative size of the Medicare market, the larger the price-following coefficient. Moving from the first to the 99th percentile of the Medicare Relative Size distribution is associated with moving from a price transmission coefficient of 0.3 to 1.5. Columns 7 and 8 show that this result is robust to simultaneously controlling for the other characteristics we examine, along with additional observable area characteristics.

The final aspect of heterogeneity that we consider is the *ex ante* price dispersion in the market for each service. In column 6, we add an interaction with the coefficient of variation within each state-service cell, computed on data prior to 1998. We find that high *ex ante* price dispersion is associated with weaker price-following. One possible interpretation of this fact is that the price heterogeneity is measuring how closely a market is following Medicare's lead in the cross-section. This would make it natural to expect markets with more dispersed prices *ex ante* to follow Medicare less closely. Like the previous results, this finding is robust to simultaneously controlling for the other characteristics we examine, along with additional observable area characteristics.

[11]In this regression, the identifying variation for the interaction with HHI comes from differences in concentration across specialties within a state.

[12]The insurance market variables are omitted from column 8 because they vary at the state level, and column 8 controls for a state-specific interaction term. So only interactions with variables defined at the state-by-code level, rather than the state level, are included in column 8.

Appendix Table D.1: Heterogeneity in Price-Following By Market Characteristics

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dependent Variable:				Private Payment Level				
Payment Shock × Post-1997	1.312**	1.350**	1.496**	1.171**	1.015**	1.573**	1.123**	0.905*
	(0.199)	(0.217)	(0.294)	(0.204)	(0.144)	(0.331)	(0.195)	(0.418)
Payment Shock × Post-1997	-0.853**						-0.562**	-0.462*
× Specialty HHI	(0.298)						(0.207)	(0.219)
Payment Shock × Post-1997		0.421**					0.823**	
× Insurance HHI		(0.150)					(0.207)	
Payment Shock × Post-1997			0.604**				0.669**	
× HMO Share 1997			(0.177)				(0.173)	
Payment Shock × Post-1997				0.949**			0.319^{+}	0.300^{+}
× Spec. Sole Practice Share				(0.233)			(0.168)	(0.159)
Payment Shock × Post-1997					1.386**		1.382**	1.596**
× Public-Private Ratio					(0.516)		(0.525)	(0.596)
Payment Shock × Post-1997						-1.132**	-1.721**	-1.342**
× *Ex Ante* Price Dispersion						(0.409)	(0.598)	(0.415)
Additional Interactions:								
Payment Shock ×	None	None	None	None	None	None	Density	State FE
N	245,296	245,296	245,296	245,296	245,296	245,296	245,296	245,296
Number of Codes	1,364	1,364	1,364	1,364	1,364	1,364	1,364	1,364

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 (D.1). Observations are constructed at the service-by-state-by-year level. The panel is balanced in the sense that each service-by-state panel is only included if public and private prices are available for each year from 1995 through 2002, plus all of the interaction variables are available. Observations are weighted according to the number of times the service is observed in Medicare claims in 1997. The dependent variable in all columns is the level of the average private payment. Standard errors are calculated allowing for arbitrary correlation among the errors associated with each service. All of the variables interacted with the payment shock have been converted into z-scores, except for the public-private ratio. Further details of the construction of all variables are available in the note to Table 1, Appendix B, and in the main text. Sources: Authors' calculations using Medicare claims, Thompson Reuters MarketScan data, and data obtained from the National Association of Insurance Commissioners, by permission. The NAIC does not endorse any analysis or conclusions based upon the use of its data.

Appendix Table D.2: Robustness Checks on 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	1.293**	1.278**	1.182**	1.374**	1.319**	1.238**
	(0.156)	(0.167)	(0.150)	(0.249)	(0.219)	(0.180)
Payment Shock × Post-1997	-0.909*	-0.818**	-0.969**	-0.734**	-0.829**	-0.988**
× Specialty HHI	(0.364)	(0.293)	(0.323)	(0.240)	(0.259)	(0.352)
N	240,264	240,264	240,264	240,264	240,264	240,264
Number of Codes	1,303	1,303	1,303	1,303	1,303	1,303
Controls Below Fully Interacted?	Yes	Yes	Yes	Yes	Yes	Yes
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 Completion)	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 reduced form specifications based on equation (D.1). Observations are constructed at the service-by-state-year level. The panel is balanced in the sense that each service-by-state panel is only included if public and private prices could be measured 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 weighted average of those interactions. The construction of all variables is further described in the note to Table 1 and in the main text. Sources: Authors' calculations using Medicare claims and Thompson Reuters MarketScan data, and Ruggles et al. (2010).

Appendix Table D.3: **Robustness Checks on Heterogeneity in Surgical CF Shock's Effect by Insurer Concentration**

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable:	Private Payment Level					
Payment Shock × Post-1997	1.370**	1.307**	1.371**	1.416**	1.362**	1.344**
	(0.204)	(0.183)	(0.236)	(0.289)	(0.243)	(0.231)
Payment Shock × Post-1997	0.296*	0.599*	0.537**	0.603*	0.684*	0.431*
× Insurer HHI	(0.149)	(0.274)	(0.168)	(0.241)	(0.313)	(0.207)
N	293,688	293,688	293,688	293,688	293,688	293,688
Number of Codes	2,194	2,194	2,194	2,194	2,194	2,194
Controls Below Fully Interacted?	Yes	Yes	Yes	Yes	Yes	Yes
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 Completion)	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 reduced form specifications based on based on equation (D.1). Observations are constructed at the service-by-state-year level. The panel is balanced in the sense that each service-by-state panel is only included if public and private prices could be measured 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 weighted 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.

E Short-Term Price-Following Estimates

E.1 Background

In our discussion of institutional detail in section 2, we presented anecdotal evidence from practitioners on the nature of contracting between physicians and insurers. Practitioners describe many negotiated payment schedules as being benchmarked directly to Medicare's relative rates, often with a constant markup. Payments associated with contracts of this form would move mechanically with updates to Medicare's relative rates. This contracting institution might thus underlie some of the dynamics we observe. Specifically, it might help to explain the relatively short-run effect of the change in Medicare's Conversion Factors for surgical relative to non-surgical care.

This appendix presents an analysis that, using data from a single large insurer, may shed light on the prevalence of payments linked directly to Medicare's relative rates. We examine physician payments from Blue Cross Blue Shield of Texas (BCBS-TX) for the 2010 calendar year. We exploit a key piece of institutional detail about BCBS-TX's Medicare-benchmarked payments. Each year, CMS announces hundreds of updates to the RVUs that RBRVS assigns to specific services. BCBS-TX in turn issues publicly available newsletters in which it informs its physicians of the date on which it will incorporate the new year's RBRVS values into its physician fee schedule. We therefore know that the switch from 2009 to 2010 RVUs was scheduled for July 1, 2010. Using changes in the RVUs that RBRVS assigned to specific services between these two years, we will estimate the relationship between updates to Medicare's RBRVS and changes in BCBS-TX payments surrounding July 1, 2010.

Two features of this analysis make it plausible to interpret the estimates below as reflecting the role of mechanical benchmarking as opposed to active contract renegotiation. The first key distinction between the analysis in this appendix and that in the main text is the time horizon. In the main text, we estimate the effect of Medicare and private payment changes over a horizon of several years. The analysis presented below occurs entirely within 2010. Second, in comparison with the changes to the surgical and non-surgical Conversion Factors, the updates analyzed below did not dramatically alter physician groups' average Medicare reimbursement rates. Consequently, these RBRVS updates less drastically alter the bargaining positions of any given physician group.

E.2 Estimation

We analyze the RVU changes using the universe of BCBS-TX claims data. These data describe the outpatient care provided to for roughly 3.5 million insured members, primarily through PPO plans. Because the data are from a single insurer and contain unique identifiers for physician groups, they allow us to track the payments associated with identifiable insurer-physician group pairs. We use data on BCBS-TX payments to 76,548 physician groups, for 2,028 unique services in 2010. The year's worth of data on these payments covers 18.37 million underlying claims. RVU changes for the services whose prices we study have a standard deviation of 8.1 percent.

We compute average BCBS-TX payments at the physician group-by-service-by-month level. We then run the following regressions, where j indexes services, t indexes months, and g indexes physician groups:

$$\overline{P^{\text{BCBS}}}_{j,g,t} = \beta \cdot \text{PredChg}_j^{\text{Medicare}} \times \text{PostJuly}_t + \upsilon_j \mathbb{1}_j + \upsilon_t \mathbb{1}_t + \upsilon_{j,g,t} \tag{E.1}$$

$$\overline{P^{\text{BCBS}}}_{j,g,t} = \beta \cdot \text{PredChg}_j^{\text{Medicare}} \times \text{PostJuly}_t + \upsilon_j \mathbb{1}_j + \upsilon_g \mathbb{1}_g + \upsilon_{j,g} \mathbb{1}_j \cdot \mathbb{1}_g + \upsilon_t \mathbb{1}_t + \upsilon_{j,g,t}. \tag{E.2}$$

The variable $\text{PredChg}_j^{\text{Medicare}}$ is constructed as $\overline{P^{\text{Medicare}}}_{j,2009}$, the average 2009 Medicare payment for service j, multiplied by the percent change in the RVUs assigned to service j from 2009 to 2010. We estimate the regression on payments aggregated to the physician group-by-code-by-month level ($j \times g \times t$). Regression (E.2) adds physician group fixed effects ($\mathbb{1}_g$) and group-by-code fixed effects ($\mathbb{1}_j \cdot \mathbb{1}_g$). In both specifications, we cluster standard errors at the service code level, since that is the level at which our main independent variable (the Medicare price shock) varies.

We run each of these regressions using two weighting schemes. The first set of weights is analogous to those we use in the paper; they are set to add up to the pre-shock Medicare quantities (specifically, the 2009 Medicare quantities for these same services). We divide this Medicare quantity for service j by the number of claims for service j in the BCBS-TX sample. We then add up the weights within a group-service-year cell. The second weighting scheme simply uses the number of underlying BCBS-TX claims as our weight.

Under the following assumptions, estimates of β from the above regressions can be interpreted as the short-run mechanical effect of RBRVS updates on BCBS-TX's RBRVS-benchmarked payments. First, the set of services benchmarked to Medicare must not change at the same time as BCBS-TX implements the new RVUs. Second, other contract terms must also remain constant, or at least not change in ways that are correlated with the RVU changes. We would not expect these assumptions to hold over the long run, especially the second one. When RVU changes alter physicians' bargaining positions, we would expect them to alter the outcomes of subsequent round of active contract renegotiation. But over the time span of a few months, as we examine here, this is unlikely—especially since physician contracts tend not to be renegotiated annually. Dunn and Shapiro (2015) report, for example, that "contracts between large health systems and commercial payers are typically negotiated every three years, sometimes up to every five years, whereas contracts between smaller practices and insurers are typically set on an auto-renewing annual basis with provisions that allow either party to terminate the contract." While some renegotiations likely took place during the period we analyze, it seems reasonable to expect their influence to be minimal.

E.3 Results

Table E.1 reports the estimates of equations (E.1) and (E.2) using the two sets of weights. The estimates imply that a \$1 change in Medicare's payments, as driven by updates to the RBRVS, was associated on average with a \$0.40 to \$0.50 change in BCBS-TX's payments. Under the assumptions discussed above, this estimates the contribution of RBRVS-benchmarked payments to the short-run relationship between public and private payments.

If the payments in our main MarketScan data have the same degree of Medicare benchmarking as BCBS-TX, the estimate would thus suggest that \$0.40 to \$0.50 of our year-1 estimate of \$1.20 (from Figure 4) can be attributed to RBRVS-benchmarking. Note that this \$1.20 estimate persists over subsequent years. From this we infer that the insurers and physician groups will generally agree to maintain these mechanical changes in subsequent years' negotiations.

Appendix Table E.1: **Short-Term Responses to Medicare RVU Updates in BCBS-TX**

	(1)	(2)	(3)	(4)
Dependent Variable:	Price or average price in BCBS-TX data			
RVU Shock	0.520**	0.442**	0.402**	0.370**
× Post-Update	(0.125)	(0.116)	(0.088)	(0.089)
Fixed Effect	Code	Group-Code	Code	Group-Code
Number of Codes	2,028	2,028	2,028	2,028
N	2,499,583	2,499,583	2,499,583	2,499,583
Weights	Texas Medicare service quantity, 2009		BCBS-TX service quantity	

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 equations (E.1) and (E.2). Columns 1 and 3 report estimates of equation (E.1), and columns 2 and 4 show equation (E.2). Both the payment shock and private payment outcome are expressed in dollar terms. The unit of observation is the billing code-by-physician group-by-month. In columns 1 and 2, observations are weighted so that the weights for each service add up to the number of times the service is observed in Medicare claims from Texas in 2009. In columns 3 and 4, the weights are based on the Texas data used in the regression. Standard errors are calculated allowing for arbitrary correlation among the errors associated with each service. Sources: Authors' calculations using Medicare claims and Blue Cross/Blue Shield of Texas claims data.

F Analysis of Quantities in the Medicare and MarketScan Data

Section 6.5 notes that the evolution of quantities in the Medicare and MarketScan data does not support strong conclusions regarding the short-run effects of the Conversion Factor change on care provision. This appendix summarizes the relevant data. The Conversion Factor change increased payments for non-surgical care relative to surgical care. We can thus convey the relevant quantity movements concisely by observing the share of total care that falls into the surgical category. A decrease in this share implies a decline in the provision of surgical relative to non-surgical care.

Appendix Figure F.1 displays the evolution of surgical relative to non-surgical care provision in both the Medicare and MarketScan data. Panel A shows the Medicare claims while panel B shows the MarketScan claims. Both panels report two measures. The first, which is the most complete quantity metric, accounts for differences between high and low resource intensive services by weighting each service by its number of RVUs. The second measure is simply the raw fraction of total service claims.

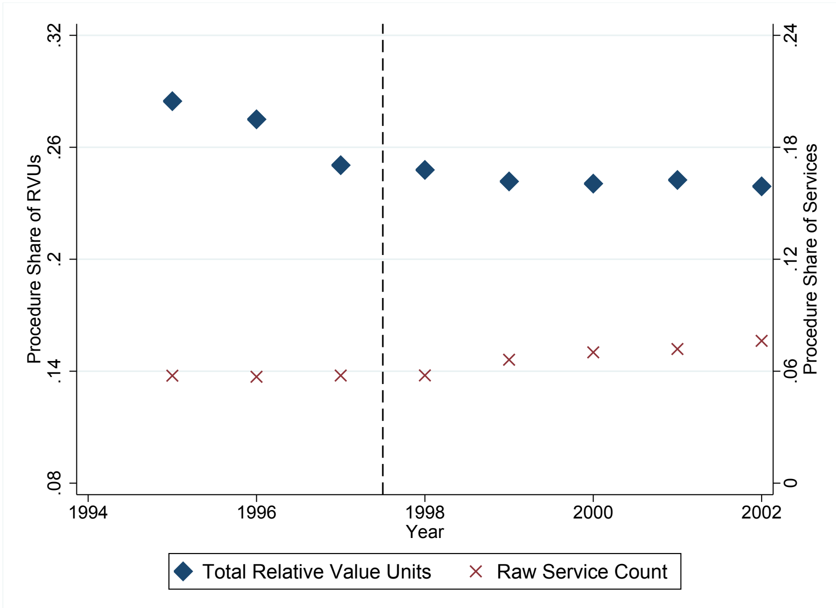
Looking first at the Medicare data, it is apparent that the surgery share declines between the initial and later years of the sample. The decline appears, however, to have largely occurred between 1996 and 1997, which precedes the payment shock. Consequently, regression specifications yield very different results depending on whether or not they control for a baseline trend in surgical service provision.[13] The figure also reveals that the RVU weights are quite relevant for assessing the relative evolution of surgical and nonsurgical service provision. While surgical services' RVU-weighted share declines, the unweighted share is relatively stable throughout the sample.

The evolution of quantities in the MarketScan data exhibit similar sensitivity. The surgery share is again non-trivially lower during the last years of the sample than during the initial years. In these private sector data, the surgery share exhibits significant volatility from year to year. It is at its highest in 1995 and 1999 and at its lowest in 2001 and 2002. Regressions again yield very different results depending on whether or not they control for a baseline trend in surgical service provision. Similarly, one again obtains a quite different picture depending on whether one weights by each service's number of RVUs. These features of the data underly our summary that the data do not support strong conclusions regarding the short-run effects of the Conversion Factor merger on care provision.

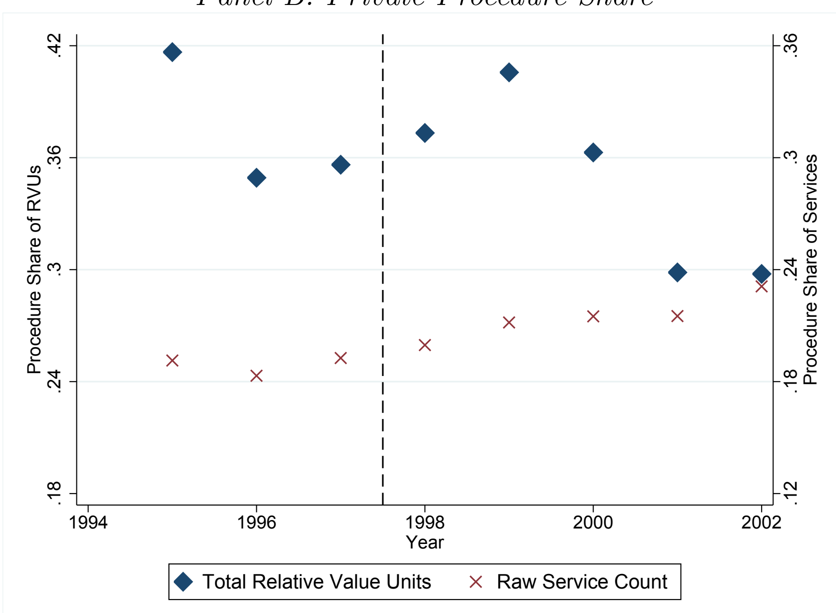
[13]Notably, our pricing results are entirely robust to this modification, plus a variety of additional robustness checks shown in Appendix C.

Appendix Figure F.1: Surgical and Non-Surgical Service Quantities

Panel A: Medicare Procedure Share



Panel B: Private Procedure Share



Note: The figures show the share of Medicare (Panel A) and MarketScan (Panel B) services that are classified as Procedures, calculated both with and without weighting for service intensity using Medicare's Relative Value Units.

G Quantifying the Resource Reallocation Resulting from Our Payment Shocks

G.1 Calculating Specialty Financial Exposure

In section 7.1, we quantify the implications of our results for the incomes of physicians of different specialties. Since different specialties provide different mixes of surgical and nonsurgical care, we use Medicare claims data to determine the empirical split of each specialty's payments into surgical and nonsurgical services. The American Medical Association (Gonzalez and Zhang, 1998) provides data on the average overall reimbursements and take-home pay for a range of different specialties. Combining these two sources, we compute the effects of Medicare's surgical reimbursement cut, Medicare's nonsurgical payment increase, and possible spillovers into private payments. Table 4 shows the results.

G.2 Estimating the Overall Reallocations from the Conversion Factor Change

Medicare spending on surgical care amounted to $15.8 billion (inflation-adjusted to 2013 dollars) after the surgical payment cut (CMS 2001, Table 57). Based on the surgical payment reduction, and assuming no behavioral response, this means that surgical payments were $1.8 billion lower than they would have been under the previous Conversion Factor regime, with nonsurgical payments higher by the same amount. Private payers spent $254 billion on physician and clinical services in 1998 (Centers for Medicare and Medicaid Services (CMS), 2014). Assuming the same medical-surgical split as in Medicare, and using our baseline price-following estimate of 1.16, yields a reallocation of $5.9 billion in private insurance payments due to the medical-surgical pricing change.

This estimate of the private sector spillover is subject to two caveats. The first is a standard concern of our estimates' external validity. As we observed in section 4.1, the MarketScan data represent the universe of claims associated with a selected set of plans provided by large employers. Small employers' plans or individual market insurance may be more or less likely to pay physicians according to fee schedules influenced by Medicare's relative payments.

Second, Medicare may exert more influence over fee-for-service payments than over capitated payments. On this point it is important to keep in mind that, although managed care was pervasive during the period we study, it typically did not translate into capitated payment of physician groups. The Community Tracking Study (CTS) reveals that throughout the period we study, roughly 40 percent of the revenue of physicians' practices was linked to managed care contracts. A much smaller share of revenue was prepaid or capitated. In 1996 this share was 16 percent (CSHSC 1999, 56), while in 2004 it was 13 percent (CSHSC 2006, 4-29). Throughout our sample, capitated payments thus represented less than one-third of the revenues associated with managed care. Additionally, the share of revenue reported in the CTS as being capitated includes payments under both Medicaid and Medicare managed care arrangements, which are sizable. The capitated share of physician revenues from private sector payers is thus not particularly large. In the context of Medicare's payments for surgical procedures relative to other services, we make a conservative adjustment and assume 16 percent of payments are capitated and thus unaffected by payment changes. If these payments are also affected, then our estimate of the ratio of the private sector spillover to Medicare's direct effect on would increase to 3.9 from the 3.3 that we use in section 7.

G.3 Estimating the Overall Reallocations from the Payment Locality Overhaul

We must also consider external validity in the context of our estimates of the effects of changes in Medicare's system of geographic adjustments. Here too, Medicare's influence on payments in small-group and individual market plans may be either greater or smaller than its influence on the payments from plans sponsored by large employers. But the concern about capitation that we raised in the previous paragraph is less relevant here. Recall that these geographic payment changes altered payments across the board rather than differentially across services. In fee-for-service settings, benchmarking to Medicare's relative payment menu can create a strong, mechanical link between public and private payments for one service relative to another, which would not be relevant in capitated contracts. But since this mechanism is less applicable when considering across-the-board payment changes, our results in this context likely reflect Medicare's influence on the physician's opportunity cost, which would be relevant for both capitated and fee-for-service payments. So there is less need to exclude capitated payments when calculating the spillovers from broad-based geographic payment changes.

One way to benchmark the magnitudes of these reallocations is by comparison with similarly motivated but more directly financed programs. For example, the Critical Access Hospital (CAH) program is a prominent means through which Medicare subsidizes rural health care. By obtaining CAH designation, rural hospitals may claim higher payments than they would otherwise be entitled to receive. Unsurprisingly, rural hospitals lobby fiercely to maintain their access to these designations.[14] In 2010, CAH subsidies amounted to approximately $300 million (MedPAC, 2012). Our findings imply that the CAH's magnitude is on par with the locality consolidation's direct effects and less than half the size of its private payment spillovers.[15]

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

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

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