Who Pays for Health Care Costs? The Effects of Health Care Prices on Wages*

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Abstract

Over 150 million Americans receive health insurance benefits from an employer as a form of compensation. In recent years, health care costs have grown rapidly, raising concerns that increased health care spending crowds-out wage increases. We leverage geographic variation in health care price growth caused by changes in hospital and physician market structure to test the impact of health care prices on wages and benefit design. We use changes to hospital and physician market structure as a source of exogenous variation. Our reduced form results find that that hospital concentration is associated with a 2.5% reduction in wages, changes to physician market structure do not have a strong impact on wages. Using this variation, we find markets that experience 10% higher price growth than the national average experience 4.1% slower wage growth. Distribution impacts. This effect is concentrated among workers without a college degree. We also find that a 10% increase in health care prices leads to a 9.5% increase in the the growth of high-deductible health plans and a 6% increase in health care costs paid by patients. Overall, our results show how rising health care costs are passed to workers in the form of lower wages and less generous benefits.

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1 Introduction

Over 150 million Americans receive health insurance benefits from an employer. These benefits are provided as a non-taxed form of compensation to workers and their dependents (Summers, 1989). While the use of health benefits as a form of compensation has advantages, one under-explored consequence is that it exposes worker compensation to increases in health care costs. This exposure is particularly notable, as going back to at least 1980, U.S. health care spending has increased substantially faster than inflation (Kamal and Cox, 2018). According to the Kaiser Family Foundation, average annual family premium contributions and out-of-pocket spending rose by 128% (from $2,061 to $4,706) and 145% ($1,231 to $3,020), respectively, from 2003 to 2018 (Rae, Copeland and Cox, 2019). Increasing health care costs place downward pressure on the ability of employers to compensate employees through wages and other forms of benefits, but the extent of the offset remains understudied. This cost growth also occupies the focus of proposals to reform American health care.

How do employers respond to rising health care costs? Do they suppress wages to cover rising health care costs? Do they shift more of the cost burden to their employees through higher employee premium contributions or with the use of high-deductible plans and other forms of cost-sharing? The potential cost to wages is an important indirect cost of rising health care costs that often gets overlooked in debates on health care reform. Its tendency to be overlooked could be because identifying the extent to which it occurs has alluded researchers. Historically, the commercial claims data necessary to accurately measure commercial health care spending has been unavailable to researchers. In this paper, we seek to fill in the gap in the literature by linking 2009-2016 commercial claims data from the Health Care Cost Institute (HCCI) with data on wages from the American Community Survey (ACS).

While several papers have considered the impacts of changes in insurance generosity, few papers have considered the impacts of changes in health care prices and spending on wages. Most notably, Gruber (1994) examined the wage impacts of requiring employers to provide coverage for specific services and finds health care costs are passed on to employees with little change in employment outcomes. More recently, Kolstad and Kowalski (2016) examined the impacts of employers providing insurance coverage and found close to full pass-through between employer health benefits and wages. However, for many employers, the costs of providing health insurance to their employees has increased, even in the absence of providing additional benefits.

Increases in employer health care costs are driven by increases in provider prices (HCCI, 2019), which in turn are driven in part by horizontal consolidation among hospitals. Between 2010 and 2015, the number of hospital mergers increased by 70% (Ellison, 2016). Substantial evidence links increases in health care prices to consolidation among hospitals. A detailed review of the hospital merger literature found that out of nine studies identified, prices increased (or increased faster relative to trend) for hospitals that consolidated relative to control group hospitals in all but one case (Gaynor and Town, 2011). The observed increase was often quite large. Tenn (2011) found that prices at Sutter hospital increased 28-44% after its merger with Alta-Bates hospital, relative to the control group. More recently, Scheffler and Arnold (2017) found hospital prices were 11% higher in highly concentrated hospital markets than unconcentrated markets and Cooper et al. (2019*b*) found that compared to hospitals with four or more local competitors, monopoly hospitals had prices that were 12% higher. Recent hospital mergers have not been linked to improvements in quality outcomes (Beaulieu et al., 2020). The lack of measurable quality impacts suggests that hospital concentration results in a price increase, rather than an improvement in value.

In addition to horizontal consolidation, many health care markets are integrating vertically (Nikpay, Richards and Penson, 2018; Richards, Nikpay and Graves, 2016). Physician practices are becoming increasingly vertically integrated with hospitals, which allows for both increases in the prices of physician services due to increased negotiation power and hospitals to direct referred services to higher-priced hospital settings. In addition, vertical integration is occurring in specialties that are particularly important for insurance networks, which further provides additional price bargaining leverage to hospitals and health systems.

In a review of the literature on the effects of integration on prices, Post, Buchmueller and Ryan (2018) found that four of the six studies they identified found an unambiguously positive relationship between integration and price. Baker, Bundorf and Kessler (2014) found modest increases of about 3% in hospital prices among more vertically integrated markets and estimated future integration could produce increases of 3-10% (Baker, Bundorf and Kessler, 2015). Neprash et al. (2015*b*) found outpatient price increases of roughly 3% due to vertical integration while Capps, Dranove and Ody (2018) found an average physician price increase of 14%, with notable variation by specialty.

Post, Buchmueller and Ryan (2018) also reviewed ten studies that examined the effects of vertical integration on spending. Among the ten studies, eight concluded vertical integration was associated with higher spending. Magnitudes of the effects on spending varied widely. Among studies using commercial data, estimates ranged from 2-3% (Baker, Bundorf and Kessler, 2014; Neprash et al., 2015*b*) up to 10% (Baker, Bundorf and Kessler, 2015) and 20% (Robinson and Miller, 2014).

While the impacts of changes in market structure on prices and spending have been widely studied (see Gaynor, Ho and Town, 2015, for a review), the impacts on non-health benefits, primarily wages, have not been studied. In this paper, we extend the existing literature on health care consolidation and prices and examine the impacts of increases in both health care spending and prices on wages and other labor market outcomes. Understanding the effect of health care costs on labor market outcomes is particularly relevant for two reasons. First, previous studies have observed wage stagnation, particularly for lower-education workers (e.g., Autor, Katz and Kearney, 2008). The extent to which workers are being paid in health care benefits rather than monetary benefits is not well understood. Health benefits are typically paid for at the firm-level, rather than at the individual-level. Thus, the potential impacts on wages are borne both by employees who consumer health care services and those who do not.

In addition, other recent research has highlighted the reasons behind growing wage inequality (see e.g., Autor, Manning and Smith, 2016; Card, Heining and Kline, 2013; Moretti, 2013; Mueller, Ouimet and Simintzi, 2017). Most employer benefits are set at the same amount across the firm. Increased health care spending is likely to have a disproportionate impact on wages for lower-income workers. Thus, increased health care spending may be an important contributor to wage inequality.

Examining these questions faces several empirical challenges. First, few data sources contain detailed information on health care prices. In this paper, we use 2009-2016 national data from the Health Care Cost Institute (HCCI). The HCCI data contain inpatient, outpatient, physician, and pharmacy claims for over 50 million commercially insured individuals per year. The claims come from UnitedHealth, Aetna, and Humana – the first, third, and fifth largest U.S. health insurers by enrollment in 2018 (Haefner, 2019). The data allow us to calculate actual prices paid for services (rather than charges) and the total annual medical spending of enrollees in the database. HCCI data has been used extensively by researchers to measure health care prices and spending (Cooper et al., 2019*b*; Curto et al., 2019; Pelech and Hayford, 2019).

A second empirical challenge is how to identify the effect of rising health care prices on wages. We focus on local-market variation in health care prices. In effect, we compare regions with higher health care price growth with regions with slower price growth. The variation here comes from mergers, which we consider to be exogeneous following previous studies. We test if the difference in health care prices is reflected in differences in wages. To do so, we use the HCCI data to construct year and market-specific indices of health care prices and spending for each Metropolitan Statistical Area (MSA) in the U.S. We link these local-market measures to data from the American Community Survey on wage compensation and employment status.

This approach raises a final concern – the potential endogeneity between local-market health care price growth and unobserved shocks to wages in that market. Examining the relationship between health care costs and employee compensation is inherently challenging given the fact that unobserved firm and occupation characteristics may by correlated with both health care costs and wages. For instance, many firms and occupations that attract high-skilled workers typically provide both high wages and generous (expensive) health care benefits. It is also possible that this type of endogeneity exists over time when comparing changes in health insurance costs and wages. Most of the prior work in this area has addressed the endogeneity problem by identifying exogenous variation in health insurance costs across individuals in cross-sectional data. For example, Baicker and Chandra (2006) used regional variation in medical malpractice laws as an instrument for health insurance prices and found that a 10% increase in premiums resulted in a 2% decrease in wages for individuals covered by employer-sponsored insurance. Two studies have used panel data to address the endogeneity problem by controlling for time-invariant observed and unobserved firm and occupation characteristics through fixed effects and long-differences specifications (Anand, 2017; Buchmueller and Lettau, 1997). A limitation of this approach is that estimates could be biased if there are unobserved within-firm changes over time that a correlated with both health insurance costs and compensation. For example, an increase in the number of high-skilled workers who are more expensive to insure would result in higher compensation and higher health insurance costs.

We address the endogeneity concern by leveraging changes in health care market structure as a source of exogenous variation. We instrument for changes in health care prices with changes in the market concentration of insurers, hospitals, and physicians in that market. The variation in market structure creates variation in health care prices, which should not be linked to unobserved differences in wages or employment outcomes.

Over the eight year time period of our data sample, prices for health care services increased by an annual rate of 6% while, among the employed population, regression-adjusted wages increased at an average rate of 1.8%. We find that a 10% increase in an MSA's health care prices relative to the rest of the country leads to a 4.1% reduction in wages in that MSA. The entirety of the effect is concentrated among individuals without a college degree, where we find a 5.4% reduction in wages.

This paper fits into a recent literature on the effects of health care price variation and price trends among the commercially insured population. Most notably, recent work has identified the wide degree of price dispersion that exists both across and within many health care markets (Cooper et al., 2019*b*). The same authors find prices for hospital services have increased much faster than for other health care services (Cooper et al., 2019*a*). Similar work has found that privately insured plans reimburse hospitals at 240% of Medicare rates (White and Whaley, 2019). A common reason for price variation is consolidation between providers, and vertical integration among physician practices (Baker, Bundorf and Kessler, 2014; Baker et al., 2014; Gaynor, Ho and Town, 2015; Fulton, 2017; Scheffler, Arnold and Whaley, 2018).

This paper also contributes to a broader literature on the compensating differentials between employer health care costs and other benefits. Several papers have estimated the effects of increased health insurance premiums on labor market outcomes and wages (Baicker and Chandra, 2006; Anand, 2017; Goldman, Sood and Leibowitz, 2005). To the best of our knowledge, no previous study has analyzed the effects of underlying health care costs on labor market outcomes or firm decisions. Further, we are not aware of any other studies that examine the impacts of provider consolidation on wages, or other outcomes beyond health care.

This paper proceeds as follows. Section 2 outlines the conceptual framework for our analysis. Section 3 describes the data used for this study while Section 4 presents the empirical approach used to estimate our main effects. Section 5 presents our regression results and Section 7 concludes.

2 Conceptual Framework

Our goal in this paper is to the effect of increasing health insurance costs effect on the compensating wage differential (Rosen, 1986). Conceptually, we think about this problem using the ideas put forth by Summers (1989), formalized by Gruber and Krueger (1991), and summarized in Baicker and Chandra (2006). Suppose that labor demand (L_d) is given by

$$L_d = f_d(W + C), \tag{1}$$

where W is wages and C is insurance costs. Further suppose that labor supply is given by

$$L_s = f_s(W + \alpha C), \tag{2}$$

where αC is the monetary value that employees put on health insurance. The key to determining the effect of rising health insurance costs on the labor market is the marginal α – the value of the marginal dollar of health insurance spending. Ultimately, the value of marginal α depends on the source of the insurance cost increase. If insurance costs are increasing because the cost of valuable health care services is increasing – marginal α is likely to be high. However, if costs are rising due to increases in administration costs or rent-seeking – marginal α will be close to zero. We suspect that the high level of merger activity during our study period makes the latter scenario more likely.

Assume now that average and marginal α are equal such that increases in the cost of insurance are valued in the same way as the existing level of insurance spending. In this case, it can be shown that

$$\frac{dW}{dC} = \frac{-\eta^d - \alpha\eta^s}{\eta^d - \eta^s}, \tag{3}$$

where η^d and η^s are the elasticities of demand and supply for labor, respectively. If $\alpha = 1$, then wages fall by the full cost of the insurance, and if $\alpha = 0$, then the results are identical to those obtained for the incidence of a payroll tax. Additionally, the proportional change in employment of will be given by

$$\frac{dL}{L} = \frac{\eta^d(W_0 - W_1 - dC)}{W^0}, \tag{4}$$

where W_0 and W_1 represent the initial and final levels of wages, respectively.

Equation 3 implies that reductions in wages will be less than the increase in health insurance costs if $\alpha < 1$. In this scenario, employees value increased insurance at less than the cost to the employer, which implies costs cannot fully be shifted to wages and employment will fall. Thus, the basic model suggests rising health care costs should lead to lower wages with an ambiguous effect on employment.

Suppose now there are two types of workers (H and L). Assuming marginal α and C are the same for both types, equation 3 becomes

$$\frac{dW_H}{dC} = \frac{-\eta_H^d - \alpha\eta_H^s}{\eta_H^d - \eta_H^s} \text{ and } \frac{dW_L}{dC} = \frac{-\eta_L^d - \alpha\eta_L^s}{\eta_L^d - \eta_L^s}, \tag{5}$$

where the group whose wages fall further as health care costs increase depends on relative elasticities of labor demand and supply.

The ambiguity of these analytical predictions makes assessing the labor effects of rising health care costs on labor market outcomes fundamentally an empirical question.

3 Data

3.1 Data on Health Care Prices

To measure local-market prices and spending for health care services, we used 2009-2016 data from the Health Care Cost Institute (HCCI). The HCCI data pools claims data from UnitedHealth, Aetna, and Humana – the first, third, and fifth largest U.S. health insurers by enrollment in 2018 (Haefner, 2019). The HCCI data covers nearly 50 million individuals per year and includes observations from every U.S. state and metropolitan area.

In addition to its wide geographic coverage, an important advantage of the HCCI data is its inclusion of negotiated prices. For each of the 8 billion claims in the database, the HCCI data includes the "allowed amount" that represents the contracted price between a provider and the respective HCCI insurer. The HCCI data includes negotiated prices for specific procedures and providers.

Unfortunately, we are not able to link the HCCI data at the individual-level to information on wages. Instead, we construct market-level measures of health care prices and spending. Given the scope of the HCCI data, using the raw claims data is infeasible. We instead construct price and spending indices for each geographic market. Our primary results use Metropolitan Statistical Areas (MSAs) as the geographic units. We obtain similar results when using other units, including counties, Hospital Referral Regions (HRRs), and Hospital Service Areas (HSAs).

3.1.1 Price Index

We construct the price and spending indices as follows. First, we use the weighted average ratio of the market-level price for a specific CPT code relative to the nationwide average price (Dunn, Shapiro and Liebman, 2013; Dunn et al., 2013; Neprash et al., 2015*a*). This index allows for price differences across markets to be captured in a single metric. Other approaches include estimating procedure-level regressions with fixed effects for each geographic market and recovering the fixed effect for each market. However, recent work finds that the easier implement index approach produces similar results, as the more computationally-burdensome regression approach (Johnson and Kennedy, 2020).

More formally, we define weights for each CPT code, indexed by k, as

$$w_k = \frac{price_k q_k}{\sum_{k=1}^{K} price_k q_k} \tag{6}$$

where $price_k$ represents the nationwide average price for the service and q_k measures the procedure's total volume. Thus, the numerator measures total spending for the specific CPT code and the denominator measures total spending across all CPT codes. We then measure the weighted average ratio of the mean CPT code-specific price in each market (g) to the average CPT code price as

$$index_g = \sum_{k=1}^{K} \frac{price_{kg}}{price_k} \frac{w_k}{\sum w_{kg}} \tag{7}$$

where $\sum w_{kg} = 1$ if the MSA contains prices for all CPT codes observed nationally and is less than one otherwise.

3.2 Data on Health Care Market Characteristics

We use two sources of data to measure the composition of health care markets in each geographic region. For hospitals, we use data from the American Hospital Association's (AHA) Annual Survey. The AHA data contains information on hospital characteristics (e.g. number of beds) and is generally treated as census of U.S. hospitals. AHA data is widely used to measure hospital market concentration (Cooper et al., 2019*b*; Scheffler, Arnold and Whaley, 2018; Fulton, 2017; Moriya, Vogt and Gaynor, 2010). Following other papers that use the AHA data, we construct the hospital-specific Herfindahl-Hirschman Index (HHI) in each geographic market. We treat hospitals in the same geographic market that are owned by the same system as one hospital for the purpose of HHI calculations. We measure market shares using hospital admissions.

For physician markets, we use data from the SK&A Office Based Physicians Database provided by IQVIA. The SK&A data is a census of office-based physicians and provides detailed information on physician practices. The data lists the specialties of all physicians working in a practice along with the non-physician health care professionals (e.g. nurse practitioners, nurses) who work in the practice.

Importantly, the data also provides ownership information for each physician practice. Specifically, the SK&A data has health system, hospital, and medical group identifiers. Physicians often appear in the data with more than one of these three identifiers. Thus, we define physician organization ownership hierarchically as follows: health system, hospital, medical group, site. If physicians do not have one of the three identifiers (health system, hospital, medical group), they are assigned to an organization that includes the physicians operating at their same site.

We use the number of full-time-equivalent physicians in an organization to measure the market shares we use as inputs for our physician HHI calculations. The full-time-equivalent weight we assign to a physician at a particular site is one divided by the number of sites at which the physician works. For instance, if a physician works at three sites, we assign 0.33 FTE to each site. We calculate five physician HHIs: primary care, cardiology, hematology/oncology, orthopedics, and radiology. The primary care HHI includes physicians listed as having one of the following specialties: family practitioner, general practitioner, geriatrician, internist, internal medicine/pediatrics, pediatrician. Only physicians in an organization with the specialty of current interest are included in market share calculations. These specialties were chosen because their numbers in the SK&A data closely match those reported by the AMA Masterfile and they are some of the most highly compensated specialties (see Fulton, 2017, for details). We also calculated a specialist HHI which is a weighted average (using number of full-time-equivalent physicians) of the cardiology, hematology/oncology, orthopedics, and radiology HHIs.

We also measure hospital-physician integration using the SK&A data. Specifically, we measure the percent of full-time-equivalent primary care physicians and specialists in a market that are in practices owned by a hospital or health system. Specialists here include all non-primary care specialties – not just the specialties included in the four specialist HHIs we calculated. The health system and hospital identifiers in the SK&A data were used to calculate these measures. Like the AHA data, the SK&A data has been used by several other studies to measure physician market structure (Scheffler, Arnold and Whaley, 2018; Nikpay, Richards and Penson, 2018; Barnes et al., 2018; Scheffler and Arnold, 2017; Baker, Bundorf and Kessler, 2016; Richards, Nikpay and Graves, 2016; Dunn and Shapiro, 2014).

3.3 Data on Wages

Finally, our individual-level data on wages and employment status comes from the American Community Survey (ACS) (Ruggles et al., 2019). To be consistent with the pricing data, we use 2009-2016 ACS data. This sample contains 8.34 million individuals between the ages of 19 and 64, an average of just over 1 million per year. In our main analysis, we restrict the ACS population to those currently employed and who receive insurance from an employer. We do not require ACS respondents to have insurance through their own employer, and include spouses and other family members who are dependents on another family member's employer-sponsored health insurance. This restriction limits the sample size by 32%, to a total of 5.7 million people.

From the ACS data, we identify individual-level information on demographics (age, gender, race, education), industry (NACIS codes), and occupation. The ACS data also contains sampling weights, which are designed to weight the ACS sample to be nationally representative.

The ACS data contains multiple questions on income, including total income, wage and salary income, and other forms of income. We use wage and salary income as our primary measure of wages because compared to other forms of income (e.g. investment or rental income), wage income is most directly linked to employer benefit decisions. As a placebo test, we measure the impacts of health care market structure and concentration on non-wage forms of income. Local-market shocks to health care spending should not impact broader economic returns (e.g. stock market investments).

We use the publicly available ACS data, which does not include respondent zip code and limits identifiable counties to those with at least 100,000 individuals. Thus, we use

Metropolitan Statistical Areas (MSAs) as our primary geographic unit. Other studies have used Dartmouth Atlas-constructed Hospital Referral Regions (HRRs) to measure health care markets. HRRs are similarly broad as MSAs. For example, the US has 306 HRRs and 384 MSAs.

4 Empirical Approach

We leverage the large literature on geographic variations in health care prices to estimate the impact on wages. The primary disadvantage is that our variation is driven by local-market variations in prices. We are not able to account for variation that impacts the entire country, such as the introduction of new technologies.

To implement this approach, for each ACS respondent i in market g during year t, we start by estimating a regression of the form

$$\ln(wage_{igt}) = \alpha + \gamma price_{gt} + \beta X_{igt} + \zeta_g + \tau_t + \epsilon_{igt}. \tag{8}$$

This regression regresses log wages on our local-market price measure ($price_{gt}$) and a robust set of controls (X_{igt} = consumer age, gender, sex, race, education). Market (ζ_g, MSA) and year (τ_t) fixed effects account for time-invariant market differences and temporal trends, respectively. We iteratively add fixed effects for worker occupation and industry codes. We estimate this regression using OLS and cluster standard errors at the level of the ACS' sampling strata. We similarly weight this model using the ACS sampling weights. We obtain similar results when clustering at the MSA-level and not weighting.

The γ coefficient on $price_{gt}$ measures the effects of changes in local-market health care prices on wages. Under the assumption that conditional on the controls and fixed effects, any unexplained variation in ϵ_{igt} is not correlated with changes in local-market prices, then this OLS regression can be interpreted causally.

However, there are several reasons to think that this assumption may not be valid. For one, hospital and other health care providers derive pricing power through internalizing patient willingness to pay for services (Ho, 2009; Gowrisankaran, Nevo and Town, 2015). Patient willingness to pay is a function of income. Thus, any unobserved local-market productivity or income shocks may influence patient willingness to pay for health care services. Providers may respond to this increase in willingness to pay by increasing prices. Thus, there is the possibility that omitted variable bias will lead to bias in the OLS regression.

As a solution to this potential bias, we leverage consolidation trends that have substantially changed the health care industry in recent years. For each market, we use the above data on health care horizontal and vertical integration to construct measures of local-market competitiveness. The higher prices that occur from the increased bargaining power following horizontal and vertical integration creates a price shifter for our local-market health care price indices.

We use these measures as an instrument for local-market prices and estimate the following regression:

$$\begin{aligned}\textbf{First stage: } price_{gt} = \alpha + \zeta_1 Hosp_{gt-1} + \zeta_2 PCPVI_{gt-1} + \zeta_3 SpecVI_{gt-1} \\ + \beta X_{gt} + \gamma market_g + \tau year_t + \epsilon_{igt}\end{aligned} \tag{9}$$

The first stage model regresses prices on lagged hospital market structure ($Hosp_{gt-1}$), the share of primary care physicians vertically integrated with a hospital or health system ($PCPVI_{gt-1}$), and the share of specialists vertically integrated with a hospital or health system ($SpecVI_{gt-1}$). For hospital market concentration, we follow the thresholds used by the DOJ/FTC Horizontal Merger Guidelines and categorize HHIs as less than 1,500 (competitive), between 1,500 and 2,500 (moderately concentrated), and above 2,500 (concentrated). We include the same set of controls as in equation 8.

The second stage model uses predicted prices from the first stage regression to measure the effect of health care prices on log wages.

$$\textbf{Second stage:} \quad \ln(wage_{igt}) = \alpha + \eta \widehat{price}_{gt} + \beta X_{gt} + \gamma market_g + \tau year_t + \epsilon_{igt}. \qquad (10)$$

We estimate this model using two-stage least squares (2SLS) and again use the ACS sampling weights and variance clusters.

A causal interpretation of the η coefficient requires the standard instrumental variables assumptions. First, our market concentration measures must have predictive power on local-market prices. As shown in Table 1, our first stage regressions indicate that, consistent with the several previous papers, increases in horizontal and vertical provider consolidation increases health care prices. Changes to insurer concentration have minimal influence on prices. Our F-statistics are above conventional thresholds.

The second assumption is that our set of instruments, changes in health care market structure, are not correlated with unobserved differences in local-market wages, ϵ_{igt}. Following the omitted variables bias example, one potential violation of this assumption is if providers consolidate in part due to unobserved shocks to local-market wages. While this assumption is not testable, we believe it is reasonable for several reasons. First, using changes in market structure as an IV relies on both the existence and timing of local-market changes in price. A violation of the validity of this approach requires that the timing of shocks that create both unobserved variation in wages and changes in prices occur simultaneously with changes in market structure. However, the timing of changes in market structure, is unlikely to occur with much precision. Many consolidation events, for example hospital or insurance mergers, require regulatory approval. The decision to vertically integrate varies by physician practice, and precise coordination of vertical integration is unlikely to occur in markets with many physician groups.

5 Results

5.1 Descriptive Characteristics

5.1.1 Price Trends

Figure 1 plots trends in prices over our study period. From 2009 to 2016, average non-indexed prices (weighted by MSA population) increased from \$134 to \$179, an absolute difference of \$45 and a relative difference of 34%. However, as shown in Figure 2, which normalizes prices to each MSA's 2009 price levels and plots the mean, 25th percentile, and 75th percentile price growth, MSAs vary considerably in their price growth. While the mean MSA has experienced a price increase of 32%, the 25th percentile growth is 19% and the 75th percentile growth is 41%.

5.1.2 Market Structure Trends

Figure 3 plots lagged hospital, primary care, and specialist MSA-level HHIs from 2009 to 2016. MSA-level HHIs were weighted by MSA population to create Figure 3. Average hospital HHI increased by 289 points (or 11%) over the period from 2,670 to 2,959. Both specialist and primary care HHI increased. Specialist HHI increased by 224 points from 1,666 to 1,890 and primary care HHI increased by 365 points from 410 to 775.

Figure 4 shows the percentage of primary care physicians and specialists in practices owned by hospitals/health systems from 2009 to 2016. The measures were again calculated at the MSA-level and then population weighted to create the figure. Hospital-physician integration has increased dramatically since 2009. The percentage of primary care physicians in practices owned by hospitals/health systems increased by 18 percentage points from 27% in 2009 to 45% by 2016. Simultaneously, the percentage of specialists in practices owned by hospitals/health systems increased by 27 percentage points from 22% in 2009 to 49% by 2016.

5.2 First Stage Results

Table 1 presents regression results that test the association between changes in MSA-specific health care market structure and MSA-specific prices (market-level version of equation 9). Column 2 shows a statistically significant positive relationship between hospital market concentration and the price index. The coefficients for moderately concentrated and concentrated hospital markets are similar in magnitude at 0.0134 (p-value < 0.01) and 0.0128 (p-value $= 0.045$), respectively. Additionally, the coefficient for the percentage of primary care physicians in small in magnitude, but also statistically significant (0.000377, p-value $= 0.019$). The coefficient for the percentage of specialists is positive, but not statistically significant. The F-stat for the model is 23.2.

5.3 Reduced Form Results: Effects of Hospital Market Structure on Wages

Table 2 presents the main reduced form results that examine the effects of changes in hospital market structure on wages. The first three columns use level wages as the regression outcome, while columns 4-6 use log-transformed wages. For each wage measurement, we iteratively add fixed effects for MSA, industry, and occupation. The preferred specification that includes the full set of fixed effects implies that moving from a competitive (e.g. HHI below 2,500) to a concentrated (HHI above 2,500) hospital market is associated with a \$511 (2.5%) reduction in wages. The results in columns 4-6 that use log-wages as the dependent variable are more precisely estimated than when using level wages in columns 1-3. For each wage specification, adding the additional industry and worker occupation fixed effects have little impact on the main coefficients of interest. The coefficients on the physician market structure variables indicate that changes in physician market structure has a negligible impact on wages. The year coefficients indicate steady wage growth over the study period of 2010 to 2016.

5.3.1 Event Study Results: Hospital Market Structure Lags and Leads and Wages

We next examine the impacts of lags and leads in hospital market structure on wages. Given the lack of an impact in the previous results for changes in physician markets on wages, we focus on changes in hospital concentration. These event study regressions are important for three reasons. First, our main specification uses a one-year lag of hospital concentration to examine changes in prices and wages. It is unclear if a one-year lag is the appropriate specification, or if given the multi-year nature of many employer contracts, if a longer lag-period is more appropriate. Examining lagged effects can also inform the persistence of changes in hospital market structure on wages. Finally, examining the lead effects (e.g. how do current wages respond to future changes in hospital market concentration) can help inform the validity of our empirical approach.

More formally, we estimate a regression of the form that includes three lag and lead periods in the following specification:

$$wage_{igt} = \alpha + \sum_{l=-3}^{l=3} \gamma_l hosp_{gl} + \beta X_{igt} + \zeta_g + \tau_t + \epsilon_i. \tag{11}$$

In this model, the γ_l coefficients measure the association between hospital market structure and wages in the three years prior and following. To measure hospital market structure, we use a binary non-concentrated (HHI below 2,500) and concentrated (HHI above 2,500) indicator. The lag coefficients, which measure the impacts of previous changes in hospital market structure on current wages, are informative because they help identify the time duration between changes in market structure and changes in wages. They also test if changes in market structure lead to persistent changes in wages, or if wage shocks recover. At the same time, the lead coefficients identify the impact of future market structure changes on current wages. In this sense, they serve as a parallel trends test that can help identify potential bias in our main estimates.

Figure 5 presents these results using level wages as the wage measure. Relative to markets with competitive hospital markets, markets with hospital markets that become concentrated three years ago have \$573 lower wages. For markets that became concentrated in the last two and one years, the coefficient magnitudes are similar,\$573 and \$470, respectively, but are less precisely estimated. We do not find a difference in wages in markets that become concentrated in the current year. Importantly for our estimation approach, we do not find current wage differences in markets that become concentrated in the future.

Similar results are presented in Figure 6, which uses log-transformed wage as the dependent variable. When using log-wages, the largest and most precisely-estimated effect, -2.5%, occurs for markets that become concentrated in the prior year. The two-year and three-year lag coefficients are -1.2% and 0.9%, although both are less precisely estimated. We similarly do not observe meaningful effects for the lead coefficients.

These event study results, combined with the previous reduced form results, indicate that changes in hospital market structure have meaningful impacts on wages for workers who receive employer-sponsored insurance. To our knowledge, these impacts have not been previously quantified, but have meaningful impacts on US workers. The mechanism underlying these effects is the increase in prices following hospital concentration observed in several other studies. We next model how these price increases impact wages.

5.4 Effect of Health Care Prices on Wages

Table 3 presents our primary results on the effects of local-market health care prices on wages. Columns 1-3 present OLS results and columns 4-6 present 2SLS results. For each model, the first column includes MSA fixed effects, the second column adds industry fixed effects, and the third column adds occupation fixed effects. The OLS results show a small and not statistically significant relationship between local-market health care prices and log-transformed wages. The coefficient magnitudes range between -0.048 and -0.053. However, the instrumented results show more meaningful effects. The magnitude of the effect ranges from -0.44 (p-value = 0.16) with just MSA fixed effects, to -0.55 (p-value = 0.055) when adding industry fixed effects, and -0.54 when adding occupation fixed effects (p-value = 0.04).

The preferred specification in column 6 implies that an annual 10% increase in health care prices, relative to the rest the country, leads to a 0.54 decrease in log-transformed wages. Re-transforming the coefficient $(\exp(\beta) - 1)$ implies a 4.1% reduction in local-market wages. The fixed effects hold stable industry and occupation, and so these wage differences do not account for compositional changes between markets. To place these results in context, the upper quartile of markets experienced a 11.1% increase in relative health care prices over our study period (2009-2016). Applying our estimate of the offsetting effect on wages implies a 4.6% reduction in wage income.

5.5 Robustness

To test the validity of these results, we estimate the same model, but use alternative price index constructions. First, to limit concerns that outlier services drive market-level prices, we restrict the procedures used to construct the price index to the 200 and 100 most common procedures in each calendar year. As shown in Table 4, these results are almost identical to our main results.

Second, we redefined the index in equation 7 such that w_k was no longer divided by $\sum w_{kg}$. This change forces CPT code prices to receive equal weight in every MSA, and CPT codes that do not appear in an MSA to receive zero weight. If an MSA had a lot of missing CPT codes, it would receive an artificially low price index under this construction. The results using this construction of the price index, also in Table 4, are nearly identical to the main results.

5.6 Heterogeneity in Effects

5.6.1 College Degree

We first examine differential effects between individuals with and without a college degree. Conceptually, individuals without a college degree may be more constrained in their labor market opportunities and may be less able to obtain higher paying jobs in the event of wage stagnation due to increased health care costs. In addition, many firms set employee premium contribution amounts at the firm level, and do not consider different contribution levels by income. Thus, wage offsets in the event of premium increases are more likely to impact lower-income workers.

To avoid testing the effects health care prices on income by income status, we estimate our main model separately for individuals with and without a college degree. In the ACS data, 56% of workers in our sample do not have a college degree.

Table 5 presents these results for workers with (Panel A) and without (Panel B) a college degree. In the OLS and IV results for college-educated workers, we do not find that increased health care prices have an effect on wages. The OLS coefficients are negative in columns 1-3, but small in magnitude and not statistically significant. The IV results in columns 4-6 are also small in magnitude and not statistically significant, but the sign changes to positive. In contrast, the IV results for workers without a college degree are negative and larger than the main results. When re-transformed, the -0.76 coefficient in column 6 implies that off of the mean price index of 1.0, a 10% increase in a MSA's health care prices leads to a 5.3% reduction in wages for workers without a college degree.

These results imply that all of the pass-through effect of health care prices on wages is borne by workers without a college education.

6 Impacts on Benefit Design

Finally, we also consider potential responses by employers besides passing health care costs through as decreased wages. In particular, the period we analyze coincides with the rapid growth in high-deductible health plans (HDHPs). While the effects of HDHPs have been extensively studied (Sood et al., 2013; Haviland et al., 2016; Brot-Goldberg et al., 2017; Zhang et al., 2018), what factors lead to the adoption of HDHPs has received less attention.

We use the market-level data and test if an increase in the MSA-specific price index leads to an increase in the share of the population enrolled in a consumer-directed health plan (CDHP). The CDHP share for each MSA is constructed using HCCI's enrollment data. Importantly, we observe CDHP, rather than HDHP enrollment. CDHPs typically consist of HDHP combined with savings account, such as Health Savings Account (HSA) or Flexible Spending Account (FSA). Unfortunately, we cannot identify employer contributions to health savings accounts.

As shown in Table 6, we find that price increases lead to a corresponding increase in the share of the population enrolled in a CDHP. When not weighting (column 1), a 10% increase in a MSA's prices leads to a 1.2 percentage point increase in CDHP enrollment. When weighting by population in the HCCI data, column 2 shows a 2.1 percentage point increase. Based on the population-weighted mean CDHP enrollment rate of 22.1%, these results translate into relative increases of 5.4% and 9.5%, respectively. Consistent with existing studies, the year coefficients show a steady increase in the growth of CDHPs of approximately 2.7 percentage points per year.

The results are even larger when examining individual-level enrollment in CDHPs (Table 7). Across the 110.8 million patient-year observations in the HCCI data, a 10% increase in local-market health care prices leads to a 140 percentage-point increase in the probability that an individual is enrolled in a CDHP. Thus, when incorporating individual-level controls and enrollment decisions, we observe a substantially larger increase in the effect of health care prices on benefit design decisions.

We also estimate similar regressions that test if increases in local-market prices lead to changes in patient cost sharing. We first construct a similar index as our price index, but use patient cost sharing as the primary measure of interest. For patient cost sharing we include all forms of cost-sharing payments (e.g. coinsurance, copay, and deductible payments). As shown in columns 3-4, we find that local-market price increases are reflected in patient cost sharing. The unweighted and weighted results are nearly identical and show that a 10% increase in local-market prices leads to a 6% increase in patient cost-sharing payments in that MSA.

Finally, we measure the mean share of total health care spending in a market that is paid by patients. As shown in columns 5-6, we find that as health care prices increase, patients are responsible for a smaller relative portion of total health care spending. We estimate that a 10% increase in an MSA's health care prices leads to an approximately 0.5% reduction in the portion of health care costs paid by patients. This result implies that while increasing health care prices lead to increased spending, patients are not responsible for the full increase in the form of cost-sharing payments. Intuitively, insurance limits patient exposure to cost sharing increases, but does not limit exposure to health care prices in the form of reduced wages or other forms of compensation.

7 Conclusion

This paper examines the relationship between rising health care costs and wages. Using detailed data on market structure, health care prices, and wages, we use plausibly exogeneous variation in health care market structure to estimate the effect of health care prices on wages. We find that markets which experience 10% higher price growth than the national average experience 4.1% slower wage growth. Additionally, we find the effect is concentrated among workers without a college degree.

Due to the unique way in which health care is financed for many Americans, recent changes to health care markets have broad-reaching impacts. Our results suggest Americans doubly feel the effects of rising health care costs – through higher health care prices and slower wage growth. This means that health care reforms with mechanisms for lowering prices are likely underestimating their potential savings if they do not include impacts on wages.

This has important implications for both health and social policy. From the perspective of health policy, it has long been known that the U.S. has a health care cost problem. The U.S. is much higher than other countries in terms of the percent of GDP occupied by health care and U.S. health care price growth frequently outpaces growth in the overall CPI. The list of strategies for containing health care cost growth is too large to discuss at length here. Among the options frequently discussed are vigorous antitrust enforcement with respect to health care mergers, reducing waste in terms of over and improper use of services, and Medicare-for-All. Importantly though, stated savings from any of these measures are *understated* if they do not include the impact on wages. An as we have shown in Section 5 the indirect cost to wages can be magnitudes greater than the direct cost of medical care. Thus making it critical that it be included when assessing proposed health care reforms.

While controlling health care costs would alleviate the reduced wage growth we identify in this paper, other interim measures should be considered. For instance, if rising health care costs continue, is there a way to redistribute the burden so that it is not exclusively felt by workers without a college degree? Strategies to redistribute this burden are likely to be of particular interest to policymakers at the current time given the host of other factors, such as technological change, that are already pushing in the direction of increasing wage inequality.

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Tables and Figures

Table 1: First Stage Regression Results

	(1)
Dependent Variable: price index	
Hospital - Competitive	REF
Hospital - Moderately Concentrated	0.0134***
	(0.00410)
Hospital - Concentrated	0.0128**
	(0.00638)
% Primary Care Hosp/HS	0.000377**
	(0.000161)
% Specialists Hosp/HS	8.21e-06
	(0.000110)
Observations	3,044
R-squared	0.971
FE	MSA, Year
F-stat	23.2

Source Authors' analysis of data from the Health Care Cost Institute (HCCI), American Hospital Association (AHA)'s Annual Survey Database, Managed Market Surveyor File from HealthLeaders-InterStudy (now Decision Resources Group), and SK&A's Office Based Physicians Database. **Notes** All market structure measures are lagged by one year. Regressions are weighted by MSA population. *** p<0.01, ** p<0.05, * p<0.1

Table 2: Reduced Form Results

	(1)	(2)	(3)	(4)	(5)	(6)
		Level Wages			Log Wages	
Concentrated Hospital Market	-571.1	-506.7	-511.5	-0.0244**	-0.0238***	-0.0251***
	(459.9)	(416.3)	(366.3)	(0.00977)	(0.00870)	(0.00836)
% Primary Care Hosp/HS	28.62*	25.56*	23.68**	0.000466	0.000337	0.000157
	(16.29)	(14.25)	(11.98)	(0.000368)	(0.000319)	(0.000298)
% Specialist Hosp/HS	-7.900	-9.093	-6.939	8.41e-05	-3.97e-05	-3.79e-05
	(11.22)	(9.929)	(8.217)	(0.000235)	(0.000213)	(0.000196)
2011	168.5*	436.8***	683.5***	-0.0186***	0.000393	0.00801**
	(95.78)	(91.06)	(84.30)	(0.00396)	(0.00367)	(0.00357)
2012	1,393***	1,632***	1,887***	0.0201***	0.0443***	0.0538***
	(271.9)	(239.3)	(196.8)	(0.00579)	(0.00517)	(0.00488)
2013	3,066***	3,249***	3,279***	0.0676***	0.0954***	0.101***
	(327.6)	(286.9)	(236.4)	(0.00644)	(0.00577)	(0.00543)
2014	4,054***	4,302***	4,279***	0.0997***	0.129***	0.135***
	(368.8)	(324.5)	(268.0)	(0.00706)	(0.00633)	(0.00596)
2015	5,791***	6,012***	5,956***	0.146***	0.175***	0.181***
	(394.7)	(347.5)	(286.8)	(0.00746)	(0.00670)	(0.00629)
2016	7,501***	7,787***	7,637***	0.185***	0.215***	0.219***
	(460.4)	(405.6)	(337.6)	(0.00864)	(0.00779)	(0.00731)
Observations	5,108,939	5,108,939	5,108,939	5,108,939	5,108,939	5,108,939
MSA FE	X	X	X	X	X	X
Industry FE		X	X		X	X
Occupation FE			X			X
R-squared	0.230	0.287	0.362	0.053	0.198	0.240

Source Authors' analysis of data from the American Community Survey (ACS), American Hospital Association (AHA)'s Annual Survey Database, and SK&A's Office Based Physicians Database.
Notes This table presents the reduced form results that examine the effect of market structure (lagged by one year) on wages. The dependent variable is log-transformed wages. Column 1 includes MSA fixed effects, column 2 adds industry fixed effects, and column adds fixed effects for occupation. All columns include controls for age, gender, race, and education. All regressions are weighted using the ACS' sampling weights and standard errors are clustered at the ACS sampling strata level. *** p<0.01, ** p<0.05, * p<0.1.

Table 3: Effect of Health Care Prices on Wages

	(1)	(2)	(3)	(4)	(5)	(6)
		OLS			2SLS	
Dependent Variable: ln(wage income)						
Price Index	-0.0475	-0.0513	-0.0531	-0.440	-0.545*	-0.544**
	(0.0412)	(0.0380)	(0.0365)	(0.313)	(0.284)	(0.270)
2010	-0.00836**	-0.00874***	-0.00394	-0.00758**	-0.00775**	-0.00293
	(0.00357)	(0.00339)	(0.00346)	(0.00365)	(0.00347)	(0.00354)
2011	-0.0111***	-0.00545	0.00368	-0.0110***	-0.00523	0.00392
	(0.00400)	(0.00379)	(0.00387)	(0.00404)	(0.00384)	(0.00392)
2012	-0.000627	0.00944**	0.0199***	-0.000662	0.00939**	0.0199***
	(0.00459)	(0.00415)	(0.00414)	(0.00460)	(0.00417)	(0.00417)
2013	0.0211***	0.0358***	0.0443***	0.0192***	0.0335***	0.0420***
	(0.00460)	(0.00424)	(0.00422)	(0.00469)	(0.00434)	(0.00432)
2014	0.0272***	0.0461***	0.0560***	0.0237***	0.0416***	0.0516***
	(0.00462)	(0.00426)	(0.00428)	(0.00522)	(0.00482)	(0.00478)
2015	0.0539***	0.0760***	0.0851***	0.0504***	0.0716***	0.0808***
	(0.00454)	(0.00415)	(0.00414)	(0.00505)	(0.00470)	(0.00463)
2016	0.0829***	0.109***	0.116***	0.0801***	0.105***	0.113***
	(0.00455)	(0.00422)	(0.00419)	(0.00495)	(0.00466)	(0.00459)
Observations	5,662,018	5,662,018	5,662,018	5,662,018	5,662,018	5,662,018
MSA FE	X	X	X	X	X	X
Industry FE		X	X		X	X
Occupation FE			X			X
R-squared	0.054	0.151	0.201	0.049	0.030	0.011

Source Authors' analysis of data from the American Community Survey (ACS), American Hospital Association (AHA)'s Annual Survey Database, and SK&A's Office Based Physicians Database.

Notes This table presents the main regression results that examine the effect of MSA-level health care prices on wages. The dependent variable is log-transformed wages. Columns 1-3 present OLS results and columns 4-6 present 2SLS results, which use hospital, insurance, and physician market structure as instruments for prices. Columns 1 and 4 include MSA fixed effects, columns 2 and 5 add industry fixed effects, and columns 3 and 6 add fixed effects for occupation. All columns include controls for age, gender, race, and education. All regressions are weighted using the ACS' sampling weights and standard errors are clustered at the ACS sampling strata level. ** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 4: Alternative Price Index Specifications

	(1)	(2)	(3)	(4)	(5)	(6)
		OLS			2SLS	
Dependent Variable: ln(wage income)						
Price Index (200 services)	-0.0368	-0.0351	-0.0303	-0.526*	-0.615**	-0.575**
	(0.0398)	(0.0348)	(0.0330)	(0.285)	(0.258)	(0.241)
Price Index (100 services)	-0.0306	-0.0309	-0.0239	-0.505*	-0.589**	-0.530**
	(0.0370)	(0.0320)	(0.0302)	(0.284)	(0.255)	(0.239)
Alternative Price Index	-0.0646	-0.0603	-0.0599	-0.432	-0.514*	-0.521*
	(0.0439)	(0.0406)	(0.0391)	(0.318)	(0.288)	(0.273)
Alternative price index (200 services)	-0.0366	-0.0348	-0.0296	-0.532*	-0.622**	-0.582**
	(0.0399)	(0.0349)	(0.0330)	(0.288)	(0.261)	(0.244)
Alternative price index (100 services)	-0.0310	-0.0313	-0.0241	-0.504*	-0.589**	-0.530**
	(0.0370)	(0.0320)	(0.0302)	(0.285)	(0.256)	(0.239)
Observations	5,662,018	5,662,018	5,662,018	5,662,018	5,662,018	5,662,018
MSA FE	X	X	X	X	X	X
Industry FE		X	X		X	X
Occupation FE			X			X

Source Authors' analysis of data from the American Community Survey (ACS), American Hospital Association (AHA)'s Annual Survey Database, and SK&A's Office Based Physicians Database.
Notes This table presents the sensitivity test regression results that examine the effect of MSA-level health care prices on wages, but use different health care price index constructions. The dependent variable is log-transformed wages. Columns 1-3 present OLS results and columns 4-6 present 2SLS results, which use hospital, insurance, and physician market structure as instruments for prices. Columns 1 and 4 include MSA fixed effects, columns 2 and 5 add industry fixed effects, and columns 3 and 6 add fixed effects for occupation. All columns include controls for age, gender, race, and education. All regressions are weighted using the ACS' sampling weights and standard errors are clustered at the ACS sampling strata level. *** p<0.01, ** p<0.05, * p<0.1.

Table 5: Effects by Education

	(1)	(2)	(3)	(4)	(5)	(6)
		OLS			2SLS	
Dependent Variable: ln(wage income)						
Panel A: College Degree						
Price Index	-0.0309	-0.0511	-0.0447	0.0693	-0.124	-0.0672
	(0.0631)	(0.0585)	(0.0556)	(0.378)	(0.354)	(0.339)
Observations	2,492,778	2,492,778	2,492,777	2,492,778	2,492,778	2,492,777
F-statistic				23.04	23.07	23.06
Panel B: No College Degree						
Price Index	-0.0586	-0.0545	-0.0569	-0.755*	-0.772**	-0.771**
	(0.0470)	(0.0424)	(0.0403)	(0.417)	(0.378)	(0.378)
Observations	3,169,240	3,169,240	3,169,240	3,169,240	3,169,240	3,169,240
F-statistic				17.81	17.814	17.82

Source Authors' analysis of data from the American Community Survey (ACS), American Hospital Association (AHA)'s Annual Survey Database, and SK&A's Office Based Physicians Database.
Notes This table presents the main regression results that examine the effect of MSA-level health care prices on wages. The dependent variable is log-transformed wages. Panel A includes individuals with a college degree and Panel B includes people without a college degree. Columns 1-3 present OLS results and columns 4-6 present 2SLS results, which use hospital, insurance, and physician market structure as instruments for prices. Columns 1 and 4 include MSA fixed effects, columns 2 and 5 add industry fixed effects, and columns 3 and 6 add fixed effects for occupation. All columns include controls for age, gender, race, and education. All regressions are weighted using the ACS' sampling weights and standard errors are clustered at the ACS sampling strata level. ** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 6: Effect of Prices on Benefit Design and Patient Cost Sharing

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable	CDHP Share		Patient cost sharing index		Percent cost sharing	
Price Index	0.115***	0.210***	0.600***	0.609***	-0.0538***	-0.0490***
	(0.0284)	(0.0535)	(0.0733)	(0.0941)	(0.0103)	(0.0128)
2010	0.0240***	0.0264***	-0.0313***	-0.00459	0.00915***	0.0113***
	(0.00198)	(0.00129)	(0.00666)	(0.00375)	(0.000760)	(0.000531)
2011	0.0631***	0.0632***	-0.0263***	-0.00381	0.0149***	0.0159***
	(0.00296)	(0.00284)	(0.00865)	(0.00425)	(0.00106)	(0.00104)
2012	0.101***	0.102***	-0.0345***	-0.0101**	0.0196***	0.0212***
	(0.00398)	(0.00441)	(0.00957)	(0.00502)	(0.00125)	(0.00106)
2013	0.128***	0.127***	-0.0362***	-0.00829	0.0253***	0.0265***
	(0.00416)	(0.00545)	(0.0102)	(0.00649)	(0.00139)	(0.00131)
2014	0.147***	0.154***	-0.0568***	-0.0177*	0.0289***	0.0306***
	(0.00461)	(0.00698)	(0.0109)	(0.00902)	(0.00162)	(0.00150)
2015	0.168***	0.178***	-0.0690***	-0.0193*	0.0287***	0.0316***
	(0.00504)	(0.00709)	(0.0105)	(0.0115)	(0.00160)	(0.00165)
2016	0.167***	0.187***	-0.0875***	-0.0193*	0.0248***	0.0294***
	(0.00511)	(0.00579)	(0.0117)	(0.0109)	(0.00180)	(0.00187)
Observations	3,048	3,048	3,048	3,048	3,048	3,048
Weighted		X		X		X
R-squared	0.822	0.888	0.803	0.871	0.818	0.872

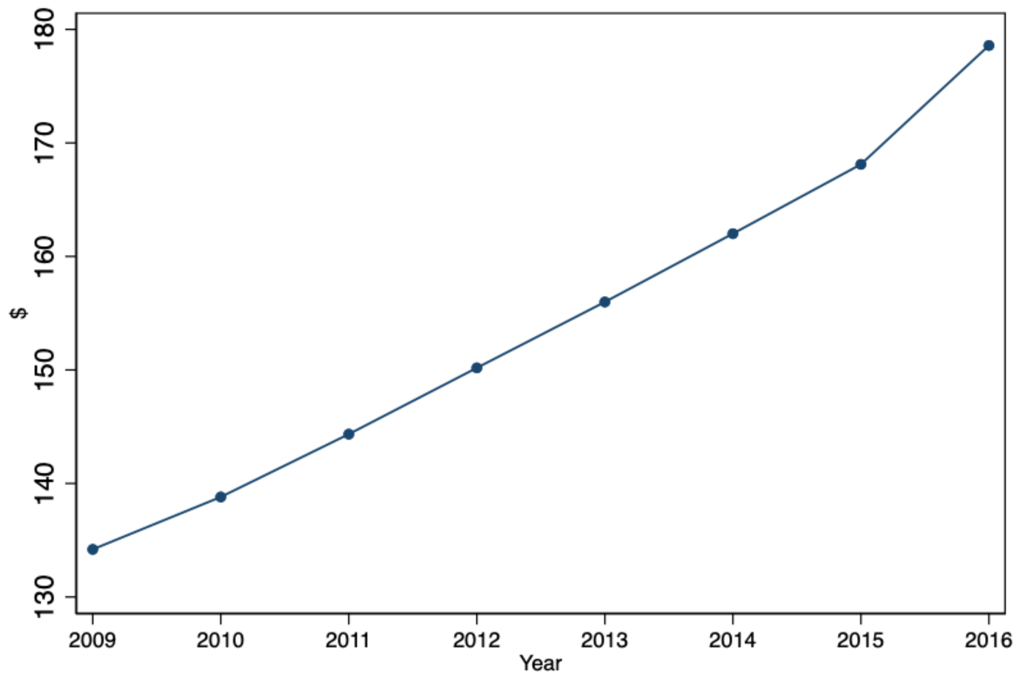
Source Authors' analysis of data from the Health Care Cost Institute (HCCI), American Hospital Association (AHA)'s Annual Survey Database, Managed Market Surveyor File from HealthLeaders-InterStudy (now Decision Resources Group), and SK&A's Office Based Physicians Database.
Notes This table presents regression results that examine the effect of MSA-level health care prices on benefit design and patient cost sharing. In columns 1-2, the dependent variable is the share of the population enrolled in a CDHP. In columns 3-4, the dependent variable is an index of the amount patients pay out-of-pocket for health care services. The dependent variable in columns 5-6 is the percentage of health care costs paid by patients out of pocket. Columns 2, 4, and 6 are weighted by the number of patients in each MSA. Standard errors are clustered at the MSA level and all regressions include MSA fixed effects. ** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 7: Effect of Prices on the Likelihood of Enrollment in a CDHP Plan

	(1)
Dependent variable	Enrolled in a CDHP Plan (0/1)
Price Index	1.409***
	(0.00717)
2010	0.338***
	(0.00110)
2011	0.710***
	(0.00106)
2012	0.875***
	(0.00105)
2013	1.066***
	(0.00102)
2014	1.308***
	(0.000972)
2015	1.361***
	(0.000944)
2016	1.265***
	(0.000888)
Observations	110,868,570
R-squared	0.841

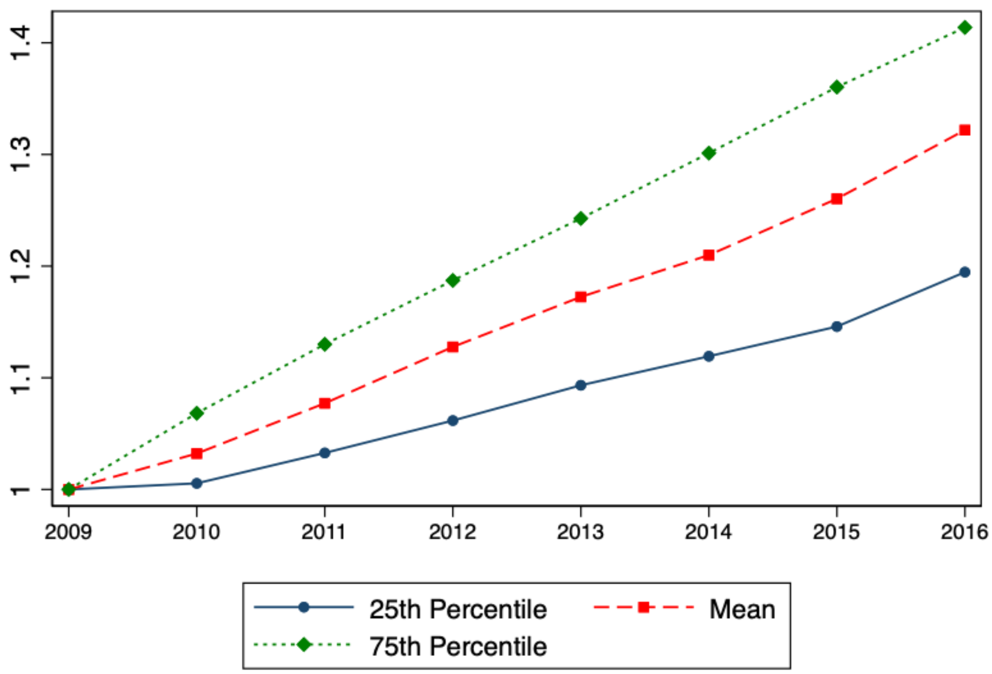
Source Authors' analysis of data from the Health Care Cost Institute (HCCI).
Notes This table presents the results of an individual-level logistic regression that examines the effect of prices on the likelihood of enrollment in a CDHP plan. The dependent variable equals 1 if an individual is enrolled in a CDHP plan and 0 otherwise. Age, gender, and the Charlson Index are included as controls along with MSA fixed effects. *** p<0.01, ** p<0.05, * p<0.1.

Figure 1: Average Prices (weighted by MSA population), 2009-2016



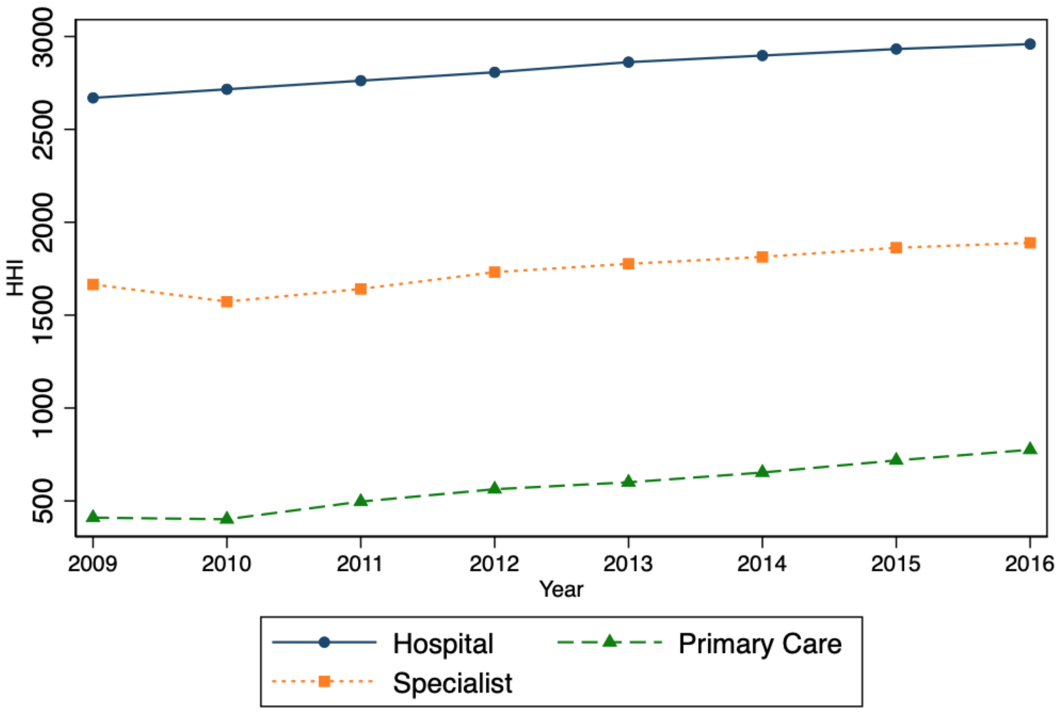
Source Authors' analysis of commercial claims data from the Health Care Cost Institute (HCCI). **Notes** Price is calculated by dividing total medical spending by the number of claims.

Figure 2: Price Growth, 2009-2016



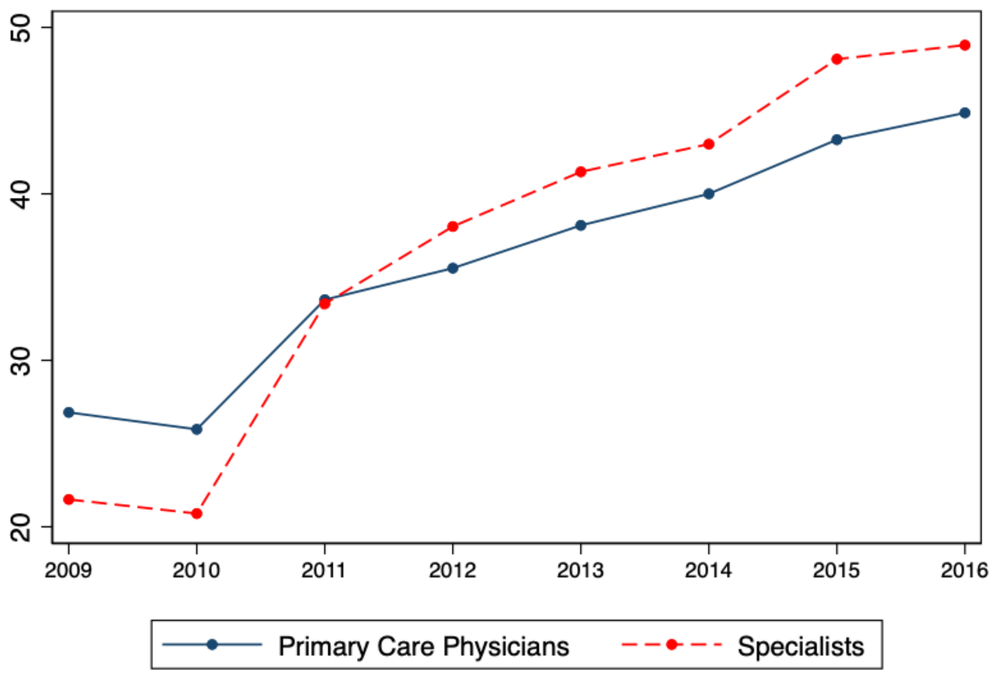
Source Authors' analysis of commercial claims data from the Health Care Cost Institute (HCCI). **Notes** MSA prices were normalized to 2009 levels.

Figure 3: Horizontal Market Concentration (MSA-level), 2009-2016



Source Authors' analysis of data from the American Hospital Association (AHA)'s Annual Survey Database, Managed Market Surveyor File from HealthLeaders-InterStudy (now Decision Resources Group), and SK&A's Office Based Physicians Database. **Notes** HHI=Herfindahl-Hirschman Index. All HHIs were calculated at the MSA-level and then population weighted to create a yearly national average.

Figure 4: Percentage of Physicians in Practices Owned by Hospitals/Health Systems, 2009-2016



Source Authors' analysis of data from SK&A's Office Based Physicians Database. **Notes** The percentages of primary care physicians and specialists in practices owned by hospitals/health systems were calculated at the MSA-level and then population weighted to create yearly national averages.

Figure 5: Association Between Hospital Market Concentration Lags/Leads and Wages

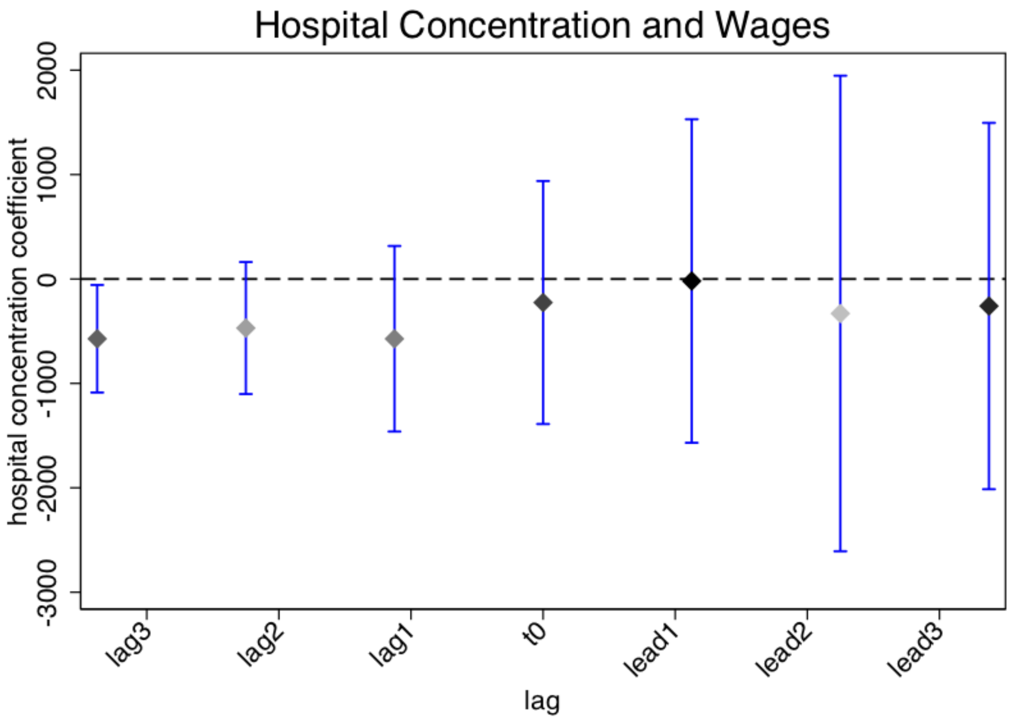


Figure 6: Association Between Hospital Market Concentration Lags/Leads and Log Wages