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

Daniel R. Arnold† Christopher M. Whaley‡

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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 market structure, and in particular, mergers, to test the impact of health care prices on wages and benefit design. We find that hospital mergers lead to a $580 increase in hospital prices and a similar, $622 reduction in wages. The reduction in wages is driven by mergers that occur within hospital markets, and we find little wage effect for cross-market hospital mergers. Our results imply that consumers bear the price effects of hospital mergers in the form of reduced wages. We also find evidence of changes in benefit design structure and adoption of high-deductible health plans. 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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†UC Berkeley. Email: danielarnold@berkeley.edu

‡RAND Corporation. Email: cwhaley@rand.org

1 Introduction

Over 150 million Americans receive health insurance benefits from an employer (Kaiser Family Foundation, 2018). These benefits are provided as a non-taxed form of compensation to workers and their dependents. While the use of health benefits as a form of compensation has both tax and risk-pooling advantages, one under-explored economic 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 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). In addition, prices negotiated by insurers on behalf of employers are often substantially higher than prices paid by public payers – largely due to changes in the market structure of the health care delivery system (White and Whaley, 2019; Chernew, Hicks and Shah, 2020; Cooper et al., 2019*b*). High prices paid by private insurance, which mainly consists of employer-sponsored health plans, are a key reason why the U.S. spends considerably more on health care than other developed countries (Anderson et al., 2003; Anderson, Hussey and Petrosyan, 2019; Papanicolas, Woskie and Jha, 2018).

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 how employers substitute between wage and health benefits compensation remains understudied, which is particularly relevant given the size of employer-sponsored insurance market in the United States. From a theoretical perspective, the pass-through between rising health care costs and lower wages depends on employee valuation of higher health care costs (Summers, 1989). For benefits that are highly-valued, workers will supply labor at similar levels as an equivalent wage payment. Several existing studies have used this economic framework to estimate how extensive-margin requirements to provide additional benefits, or additional forms of benefits, change wages and labor market outcomes (e.g., Gruber (1994) and Kolstad and Kowalski (2016)). These studies, summarized below, find that these additional forms of benefits are highly-valued by workers, and thus lead to near-complete pass-through to wages.

However, few studies have examined if the intensive-margin changes in the costs of health care lead to changes in wages and other labor market outcomes. If health benefits become more costly, and this cost increase is due largely to price increases rather than improvements in the quality of benefits, workers may be less willing to accept equivalently lower wages. As suggestive evidence, Figure 1 plots inflation-adjusted trends in wages for workers who receive employer-sponsored insurance and the mean premium costs of a health insurance plan for a family over the 2008 to 2018 time period. Over this time period, inflation-adjusted wages increased by \$2,614, a net increase of 4.5%. Health insurance costs increased by \$4,721 (31.7%), from \$14,895 in 2008 to \$19,616 in 2018, suggesting an intensive-margin pass-through rate of approximately 50%, which is substantially lower than observed in previous studies.

While this suggestive evidence indicates an inverse relationship between health benefits and wage benefits, it does not indicate a clear causal relationship or indicate how else employers might respond to increased health care costs. In this paper, we extend the existing literature on the impacts of health benefits on labor market outcomes to examine how intensive-margin changes in the costs of health benefits impact labor market outcomes, and in particular, worker wages. Empirically, we examine how recent changes in health care provider market structure, hospital mergers, impact both health care costs and wages and other labor market outcomes for workers who provide insurance through their employer. To do so, we combine detailed data on wages and worker demographics from the American Community Survey (ACS) with medical claims data from the Health Care Cost Institute (HCCI), which combines data from three of the largest health insurers in the U.S.[1]. We also use data from the Hospital Cost Report Information Service (HCRIS), which aggregates cost

[1] The three insurers are Aetna, Humana, and UnitedHealth. These three insurers collectively have 25% market share of the private insurance business.

and revenue data from all U.S. hospitals. We use these combined data sources to measure how changing health care market structures leads to changes in both health care costs and labor market outcomes.

As a source of exogenous variation, we use hospital mergers. Between 2010 and 2015, the number of hospital mergers increased by 70% (Ellison, 2016). Examining the impact of provider consolidation on wages is particularly relevant because increases in employer health care costs are largely driven by increases in provider prices (HCCI, 2019), which in turn are driven in part by horizontal consolidation among hospitals. In addition, while substantial evidence links increases in health care prices to consolidation among hospitals (e.g., Gaynor and Town (2011), Gaynor, Ho and Town (2015), Scheffler and Arnold (2017), Scheffler, Arnold and Whaley (2018), Cooper et al. (2019*b*)), how these costs are financed through lower wages and other labor market outcomes is not known.

Consistent with existing studies, we find that over the 2009 to 2018 period, hospital mergers led to an $580 (3.6%) increase in hospital prices. On the reverse side of the market, we find that hospital mergers lead to a $622, approximately 1.0%, reduction in wages. We find minimal impacts on hours worked or employment. Further supporting a causal link between the costs paid by employer health plans and wages, we do not find evidence of health care price or spending increases following cross-market hospital mergers, which other studies have found do not lead to increases in prices (Dafny, Ho and Lee, 2019). We correspondingly do not find a wage impact following these mergers. Instead, the the effect is driven by mergers that occur between hospitals in the same market. We find stronger results in already concentrated hospital markets and for workers with a college degree.

While our results indicate that employers respond to rising health care costs by reducing wages, employers may have other strategic responses to health care costs. In particular, they may change the structure or generosity of their health benefits, but it is unclear how prices influence these employer decisions. As one notable example, the last decade has seen a rapid growth in high-deductible health plans (HDHPs), which require patients to bear a larger upfront share of health care costs. Presently, almost half of U.S. workers are covered by a plan that requires them to incur the first $1,000 in costs before insurance coverage begins (Peterson-KFF, 2018). While the consequences of HDHPs has been studied, how employers make the decision to change benefits, and whether employers strategically respond to supply-side changes in prices or market structure, has not been examined.[2] We extend our results to tests the impacts of local-market changes in health care prices and provider market structure on the growth of HDHPs. We find that these supply-side factors have a small impact on the adoption of CDHPs. Our two combined results suggest that firms respond to higher health care costs cost by adopting the blunt instrument of HDHPs, and by using the even more blunt instrument of reducing wages.

This paper contributes to two relevant literature. First, 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. As first noted by Summers (1989), the trade-off between wages and benefits depends on how workers value health insurance compensation relative to wage compensation. Most notably, Gruber (1994) examines the wage impact 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) examines the impact of employers providing any insurance coverage and find 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. The existing literature on benefits and wages does not directly address this question. Complicating measurement of these trade-offs is the structure of employer benefits in the U.S. Unlike the models outlined

[2]Several recent papers have examined the impacts of HDHPs on both patient price sensitivity and patient utilization of care. In one notable example, Brot-Goldberg et al. (2017) finds adoption of a HDHP within a large firm lead patients to reduce both unnecessary and necessary care, but does not impact patient use of lower-priced providers. Several studies using national data do not find that firm-level adoption of HDHPs leads to increases in reductions in the use of low-value services or price shopping (Haviland et al., 2016; Sinaiko, Mehrotra and Sood, 2016; Beeuwkes Buntin et al., 2011), although some evidence shows increased price shopping for laboratory tests (Zhang et al., 2018).

in Summers (1989), Gruber (1994), and others, where employers are either make the decision to, or are mandated to, provide select forms of benefits, in context we study, the costs of providing benefits becomes more expensive, and does so for all firms within a geographic market.

A more similar framework comes from papers that model the wage impacts of payroll taxes. In Section 2, we develop a model to join these two literatures. We show that if workers fully value increased health care costs, for example, if cost increases represent improvements in value or quality, then the model initially developed by Summers (1989) should hold. However, if increased health care costs are due to price increases that do not improve quality, then workers should respond similar to responses observed from the literature on payroll taxes.

How these dynamics influence how employers respond to changes in health care market structure and prices has not been extensively examined. 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). However, 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, or examined how changing market structures impact wages.

Further, we are not aware of any other studies that examine the impacts of provider consolidation on wages, or other outcomes beyond health care markets. Employer-sponsored insurance is responsible for approximately one third of U.S. health care spending, $1.2 trillion per year (White and Whaley, 2019). How employers fund these costs has not been thoroughly examined. Similarly, many hospital mergers are designed as a mechanism to negotiate higher prices from insurers and the employers who purchase insurance benefits. While existing work has clearly demonstrated that health care market consolidation leads to higher prices, linking changes in market structure to wages is important to understand how these higher prices are paid for.

Examining the effect of health care costs and provider market structure 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 consume 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 the wages of lower-income workers. Thus, increased health care spending may be an important contributor to wage inequality.

This paper also fits into a more recent literature on the effects of health care price variation and price trends among the commercially insured population. This literature has empirically extended key factors first raised by Arrow (1963) that distinguish health care markets from other markets. 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).

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. For example, 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 in 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. Additional work has examined provider market structure and how consolidation strategies are used to increase bargaining leverage and thus prices (Ho, 2009*a*; Gowrisankaran, Nevo and Town, 2015*a*). Beaulieu et al. (2020), in the most comprehensive study to date on the impact of hospital mergers and acquisitions on quality, found hospital acquisition by another hospital or health system was associated with modestly worse patient experiences and no significant changes in readmission and mortality rates. The lack of measurable quality increases suggests that hospital mergers lead to pure price increases, rather than increases in quality that potentially offset price increases, and thus may increase employee valuation of health benefits.

In this paper, we extend the existing literature on health care consolidation by examining the impacts of changes in health care market structure, and in particular, hospital mergers on wages and other labor market outcomes. We also extend these results to examine the broader question of how rising health care costs are passed on to reductions in worker wages. 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 negotiated 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). We supplement this data with detailed information on revenues from private insurers for each U.S. hospital.

This approach raises a second 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 led to 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 use hospital mergers, which we consider to be exogenous following previous studies. We test if the difference in health care prices caused by hospital mergers is reflected in differences in wages. To do so, we use the HCCI and HCRIS 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.

While the impacts of changes in market structure on prices and spending have been widely studied (e.g., see Gaynor, Ho and Town, 2015, for a review), the pass-through impacts on non-health benefits, primarily wages, have not been thoroughly examined. Understanding the incidence of health care cost increases is important for both policy and economic reasons. The labor economics literature has not fully addressed the extent to which health benefit costs are passed to workers. Additionally, while regulators examine potential impacts on provider prices when reviewing health care consolidation events, they have typically not considered impacts on wages and other labor market outcomes. Our results imply that the price effects scrutinized by regulators do not occur in a vacuum, and are instead borne by workers in the form of lower wages. These impacts are of particular importance given the structure of employer-sponsored insurance in the United States. Our results imply that the impacts of rising health care costs are passed through in the form of lower wages and benefits.

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 estimate compensating wage differential of increasing care costs on worker wages (Rosen, 1986). Conceptually, this question is similar to those put forth by Summers (1989), formalized by Gruber and Krueger (1991), and summarized in Baicker and Chandra (2006). In our model, suppose that firms provide health insurance to their employees and 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 care costs on the labor market is the marginal α – the value of the marginal dollar of health insurance spending. Importantly, unlike the models originating with Summers (1989), this α does not measure the worker-level trade-off between wages and receiving health insurance benefits, but rather, the trade off between wages and health insurance costs. Ultimately, the marginal α's value depends on the source of insurance cost increases. If insurance costs are increasing because insurance coverage provides access to additional services (e.g., preventive screenings) or because new technologies are covered by insurance (e.g., new cancer therapies), then the marginal α is likely to be high. However, if costs are rising due to increases in administration costs, rent-seeking, or other cost increases not valued by patients, the marginal α will be close to zero.

In equilibrium, 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 there is no effect on employment. If $\alpha = 0$, then the results are identical to those obtained for the incidence of a payroll tax – a reduction in both wages (but not by as much as in the $\alpha = 1$ case) and employment. 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 analytically 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 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. Given the scope of the HCCI data, using the raw claims data is not computationally feasible. 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 index as follows. First, we use the weighted average ratio of the market-level price for a specific procedure relative to the nationwide average price (Dunn, Shapiro and Liebman, 2013; Dunn et al., 2013; Neprash et al., 2015). 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 procedure, 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 procedure and the denominator measures total spending across all procedures. We then measure the weighted average ratio of the mean procedure-specific price in each market (g) to the average procedure 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 procedures observed nationally and is less than one otherwise.

We focused our measurement of prices to prices charged for hospital-based services. To identify procedures, we used Diagnosis-Related Group (DRG) codes, which are used by Medicare and other private insurers to group hospital-based services into single procedures. As a sensitivity test, we also measured overall prices in a market, and used Current Procedure Terminology (CPT) codes to identify individual procedures.

3.2 Hospital Cost Report Data

We supplement the detailed HCCI data on medical claims with hospital-level data from the Hospital Cost Report Information System (HCRIS). All Medicare-certified hospitals are required to submit annual cost reports to the Centers for Medicare and Medicaid Services (CMS). These cost reports include information on hospital revenues, capacity, discharge volume, and operating costs. Hospital revenues and discharge volume are further dis aggregated into insurance payer-specific measures. Private insurance fields were added to the HCRIS data in 2009 and the data extend through 2018.[3]

We use these measures to construct total hospital revenues for patients with private insurance over our sample period. We calculate both total hospital revenues from private patients and revenues per-privately insured patient discharge, which is similar to the average

[3] We use the HCRIS data provided by the RAND Hospital data (https://www.rand.org/pubs/tools/TL303.html).

price per commercial patient. While the HCRIS data allow us to calculate revenues per private insurance discharge, they do not directly allow us to identify market-level spending. If consolidation improves the efficiency of care, then higher per-discharge costs may actually lead to lower spending by employers and private insurers.

To measure market-level spending, we use data from the InterStudy survey of insurers. The InterStudy data contains zip code-level information on insurance enrollment by insurance company and product type (e.g., employer-sponsored insurance, Medicare Advantage, Medicaid HMO, etc.). We use the privately insured population in each market from the InterStudy data as the denominator population for total spending. We divide hospital specific revenue from

3.3 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 Associations (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. We also include hospital mergers in each market. We use both data provided by Cooper et al. (2019*b*) and a similar approach using AHA data to extend beyond their sample.

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).

3.4 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*b*; Gowrisankaran, Nevo and Town, 2015*b*). 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 ($SpecVI_{gt-1}$). To measure hospital market structure, we use both an indicator for mergers in that market and hospital concentration. 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.

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

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 2 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 **??**, 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%.

Figure 3 presents the number of hospital mergers per year in our sample. In a given year, there are approximately 100 hospital mergers, but merger volume peaked in 2013.

5.2 First Stage Results: Effect of Hospital Mergers on Prices and Spending

Figure 4 presents the impacts of hospital mergers on private insurance hospital prices. Hospital mergers lead to a \$580 increase in the mean price per hospital service. However, as shown in the figure, the magnitude of the price increase grows in each year following merger and acquisition activity. By the fourth year following a merger, prices are \$871 higher, and \$1,241 higher in the fifth and greater years following a merger. Noticeably, we do not observe a pre-trend increase in prices, which helps further the causal argument that hospital mergers lead to an increase in prices. The mean hospital price is \$16,351, and so the increases we observe translate to an approximately 3.6% increase in hospital prices.

5.3 Reduced Form Results

5.3.1 Effects of Hospital Mergers on Wages and Labor Market Outcomes

Table 2 presents results that examine the impacts of hospital mergers on worker wages. Wages for workers who receive employer-sponsored insurance, our primary outcome, decline by $622 following hospital mergers within an MSA in the specification with just the worker controls and MSA fixed effects (Panel A). When adding fixed effects for occupation and industry, the impact on wages is a $610 and $638 reduction, respectively. Relative to the mean wage of $59,979, the $638 reduction in column 3 corresponds to a 1.1% relative reduction in wages.

Panels B and C present similar results, but distinguish between within-MSA hospital mergers (Panel A) and cross-market hospital mergers. Consistent with the previous results, where we find price and spending effects for within-MSA mergers but not cross-market mergers, and the results in Dafny, Ho and Lee (2019), our results are driven by within-MSA mergers. In our preferred specification in column 3 that includes the full set of MSA, industry, and occupation fixed effects, within-MSA mergers lead to a $1,065 reduction in worker wages. The effect of within-MSA hospital mergers on wages does not depend on the regression specification. As shown in Panel C, we do not find that hospital mergers that occur across markets lead to changes in worker wages.

Figure 5 presents an event study version of the effect of hospital mergers on wages. In the four years prior to hospital merger, wages are slightly trending upwards. Following hospital merger activity, there is an immediate reduction in wages, which increases in the first year following mergers. The trend stabilizes in the remaining post merger years. Appendix Figures X and X present similar results for within- and cross-market mergers that are consistent with the main regression result. We find a sharp change in wages following within-market hospital mergers, but do not find any meaningful change in wage trends following cross-market mergers.

5.3.2 Heterogeneous Effects of Hospital Mergers on Labor Market Outcomes

We next examine how these results vary by patient characteristics and demographics. As shown in Column A, we find that the effects largest for workers with a college degree. For overall mergers, we find a $693 wage reduction among college education workers, and find an imprecisely-estimated $170 reduction for workers without a college degree. However, for workers without a college degree, we find a 0.1-hour increase in the number of hours worked. We do not find any change in the probability of employment based on education.

We also find differences by worker race (Panel B), gender (Panel C), and whether the worker is above or below age 40 (Panel D). We find that the wage impacts are largest for white workers. We find similar magnitudes on the number of hours worked, but the results are only statistically significant for white workers. For worker gender, we find that hospital mergers lead to reductions in wages for both men and women, but the wage impact is largest for female workers. This finding is consistent with previous evidence that employer-sponsored insurance contributes to the male-female wage gap (Cowan and Schwab, 2016). We again find small changes in the number of hours worked and employment status. For worker age, we find that the incidence of the wage effects of hospital mergers on wages falls on workers above age 40. Workers below age 40 have a slight increase in the number of hours, but neither age group experiences changes in employment.

The results presented in the Appendix show that the heterogeneous differences observed are driven by within-market, rather than cross-market hospital mergers.

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. To do so, we use the individual-level HCCI data to test if changes in local-market hospital prices and spending at hospitals leads to an increased probability of enrollment in a CDHP.

As shown in the first column of Table 6, we find that a one dollar increase in hospital prices leads to a 0.0024 percentage point increase in the probability of CHDP enrollment. When applying the mean increase observed earlier from an increase in hospital prices, the effect translates to a 1.4 percentage point increase in the likelihood of CDHP enrollment. This effect is similar when instrumenting for hospital prices using hospital mergers (Panel B). The 2SLS coefficient of 0.0047 percentage points translates to a 2.7% increase in the probability of CDHP enrollment.

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 the second column, we find that local-market price increases are reflected in patient cost sharing. A $1 increase in hospital prices leads to a $0.03 increase in patient cost sharing when using OLS and a $0.125 increase when using 2SLS. Applying the magnitude of the hospital merger price increase of $581 results in a patient cost-sharing increase ranging from $17.4 to $73.

Finally, we measure the share of total health care spending in a market that is paid by patients. As shown in column 4, we find that as health care prices increase, patients are responsible for a smaller relative portion of total health care spending after instrumenting for price increases. We estimate that a $1 increase in hospital prices leads to a 0.8% reduction in the share of spending paid by the patient. 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, hospital revenues, and wages, we use plausibly exogenous 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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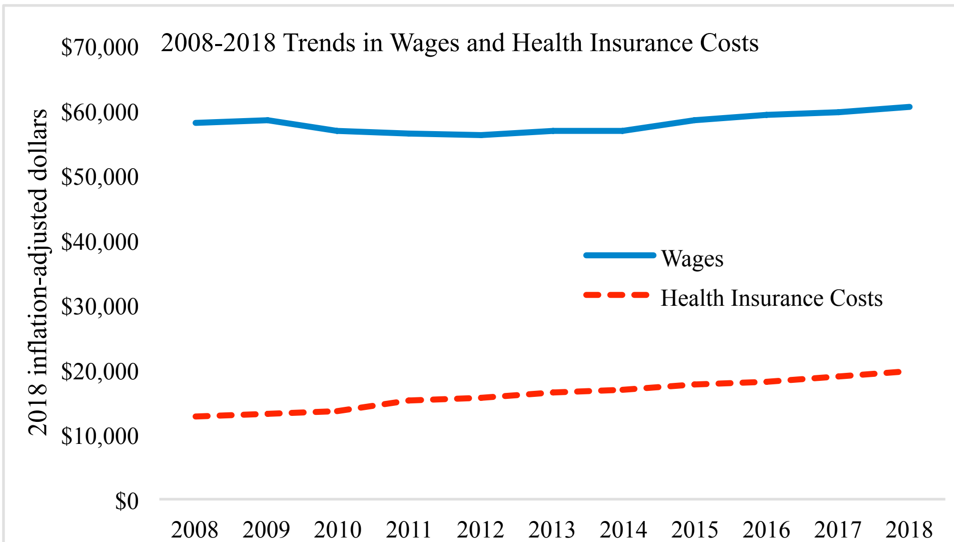
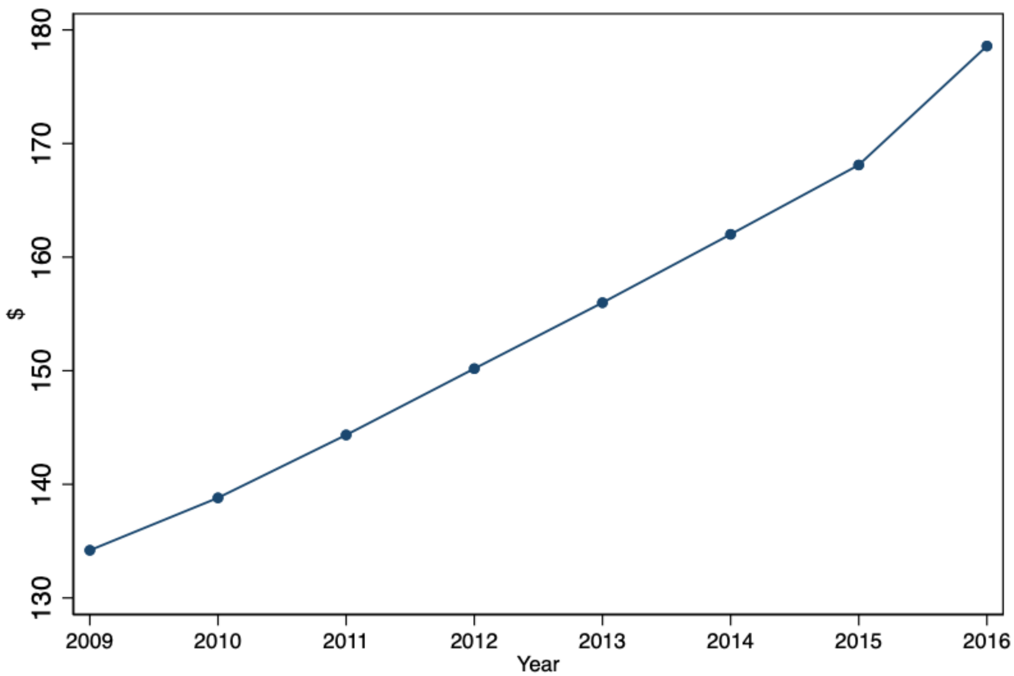
Tables and Figures

Figure 1: 2008 to 2018 Trends in Wages and Employer Health Insurance Costs



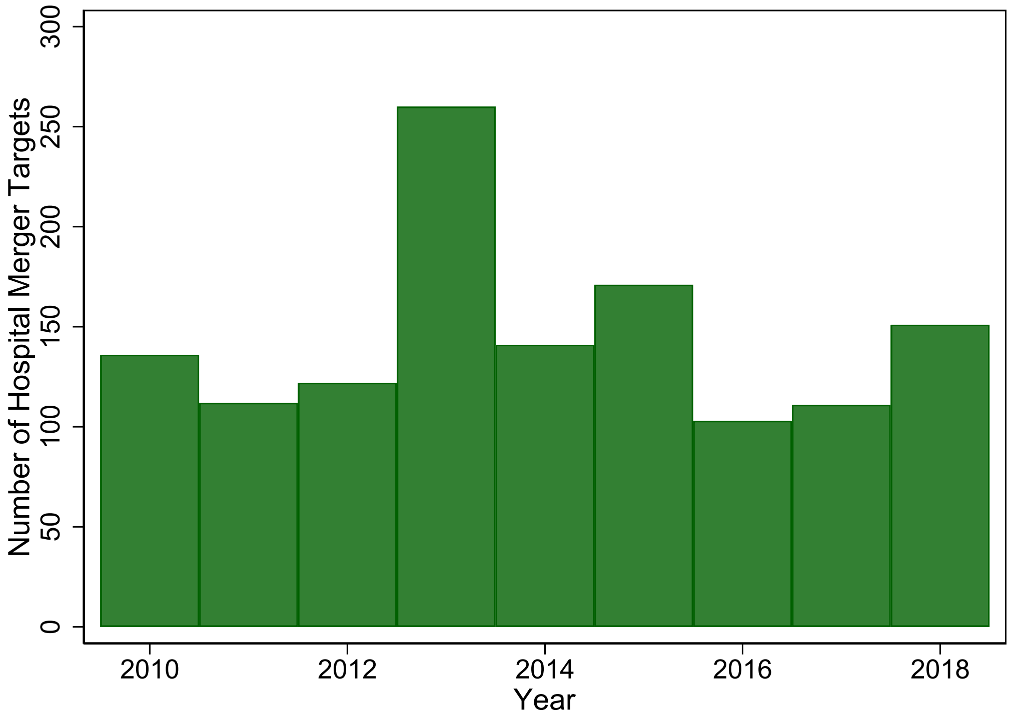
Source Wage income data is derived from the American Community Survey (ACS). The wage sample is limited to ACS respondents who receive health insurance from an employer or union, are between the ages of 20-64, and presently in the labor force. Data on health insurance premiums for a family or group plan is from the Kaiser Family Foundation (Premiums and Worker Contributions Among Workers Covered by Employer-Sponsored Coverage, 1999-2019). **Notes** Wage income and insurance premiums are both inflation-adjusted to 2018 dollars using the Consumer Price Index.

Figure 2: 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 3: Hospital Merger Targets, 2010-2018



Source: Authors' analysis of data from the American Hospital Association's Annual Survey Database.
Notes: Only includes targets in the 290 MSAs included in the ACS wage analysis.

Table 1: Association Between Hospital M&A and Inpatient Prices/Spending per Enrollee

(a) All M&A

	Price	Price (100)	Price (200)	Spending	Spending (100)	Spending (200)
Post M&A	580.072**	278.630	355.218	54.541**	27.264**	35.856**
	(250.958)	(192.844)	(222.586)	(21.854)	(11.313)	(15.749)
Observations	1,963	1,963	1,963	1,963	1,963	1,963
Adjusted R2	0.921	0.905	0.915	0.971	0.943	0.960
Dep. Var. Sample Mean	16,320	12,455	14,368	1,025	554	760

(b) Within M&A

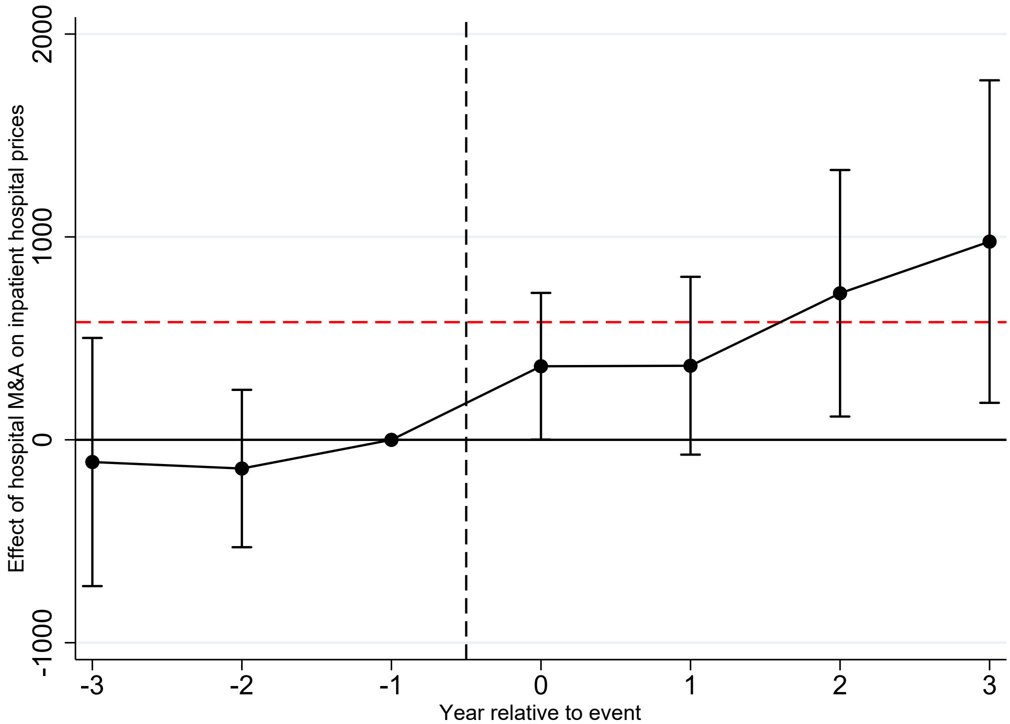
	Price	Price (100)	Price (200)	Spending	Spending (100)	Spending (200)
Post M&A	625.729*	205.928	337.583	53.571*	19.895	32.652
	(330.252)	(241.117)	(306.093)	(30.591)	(15.805)	(24.248)
Observations	1,149	1,149	1,149	1,149	1,149	1,149
Adjusted R2	0.923	0.904	0.914	0.957	0.928	0.947
Dep. Var. Sample Mean	16,103	12,243	14,146	1,011	557	759

(c) Cross M&A

	Price	Price (100)	Price (200)	Spending	Spending (100)	Spending (200)
Post M&A	790.702*	481.815	571.653	85.992*	39.731**	55.108**
	(437.988)	(359.842)	(389.329)	(45.264)	(18.220)	(25.899)
Observations	970	970	970	970	970	970
Adjusted R2	0.913	0.895	0.908	0.973	0.941	0.963
Dep. Var. Sample Mean	16,681	13,053	14,896	1,021	559	762

Notes: Price=inpatient price. Spending=inpatient spending per enrollee. (100) and (200) include only the top 100 and 200 most common DRGs, respectively. All regressions include MSA and year fixed effects. Observations more than three periods before or after treatment are excluded. The estimates shown are weighted averages of the event study's post-treatment coefficients. Standard errors are clustered by MSA. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Figure 4: Association Between Hospital M&A Lags/Leads and Inpatient Prices



Source: Authors' analysis of inpatient price data from the Health Care Cost Institute and hospital merger data from the American Hospital Association (AHA). Study period 2010 to 2016.

Notes: The dotted red line is a weighted average of the post-M&A coefficients and equals $580.

Table 2: Association Between Provider Market Concentration and Inpatient Prices/Spending per Enrollee

	Price	Price (100)	Price (200)	Spending	Spending (100)	Spending (200)
Number of Hospitals	-173.914^{**}	-86.069	-143.556^{**}	-10.735	-4.977	-7.379
	(84.095)	(73.898)	(69.243)	(8.508)	(4.474)	(5.572)
Primary Care HHI	0.316	0.042	0.167	0.019	0.006	0.012
	(0.283)	(0.142)	(0.207)	(0.024)	(0.011)	(0.015)
Specialist HHI	0.017	0.010	0.040	0.008	0.004	0.009
	(0.131)	(0.111)	(0.129)	(0.012)	(0.007)	(0.009)
VI Primary Care	-12.942	-2.430	-6.755	-0.985	-0.181	-0.387
	(12.666)	(9.021)	(10.236)	(1.075)	(0.610)	(0.807)
VI Specialist	-1.041	4.196	2.699	-0.321	0.340	0.111
	(6.640)	(6.967)	(6.574)	(0.687)	(0.357)	(0.438)
Observations	1,837	1,837	1,837	1,837	1,837	1,837
Adjusted R2	0.927	0.907	0.918	0.972	0.942	0.960
Dep. Var. Sample Mean	16,351	12,465	14,383	1,028	555	762

Notes: Price=inpatient price. Spending=inpatient spending per enrollee. (100) and (200) include only the top 100 and 200 most common DRGs, respectively. All regressions include MSA and year fixed effects. Observations more than three periods before or after treatment are excluded. Standard errors are clustered by MSA. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 3: Effect of M&A on Wage Income

(a) Panel A: All M&A

	(1)	(2)	(3)
Post M&A	-621.977**	-610.277**	-637.948***
	(283.632)	(245.654)	(216.483)
Observations	5,960,618	5,960,618	5,960,618
Adjusted R2	0.225	0.284	0.351
FE	MSA	MSA, IND	MSA, IND, OCC
# of MSAs	290	290	290
# of Treated MSAs	228	228	228
# of Control MSAs	62	62	62

(b) Panel B: Within M&A

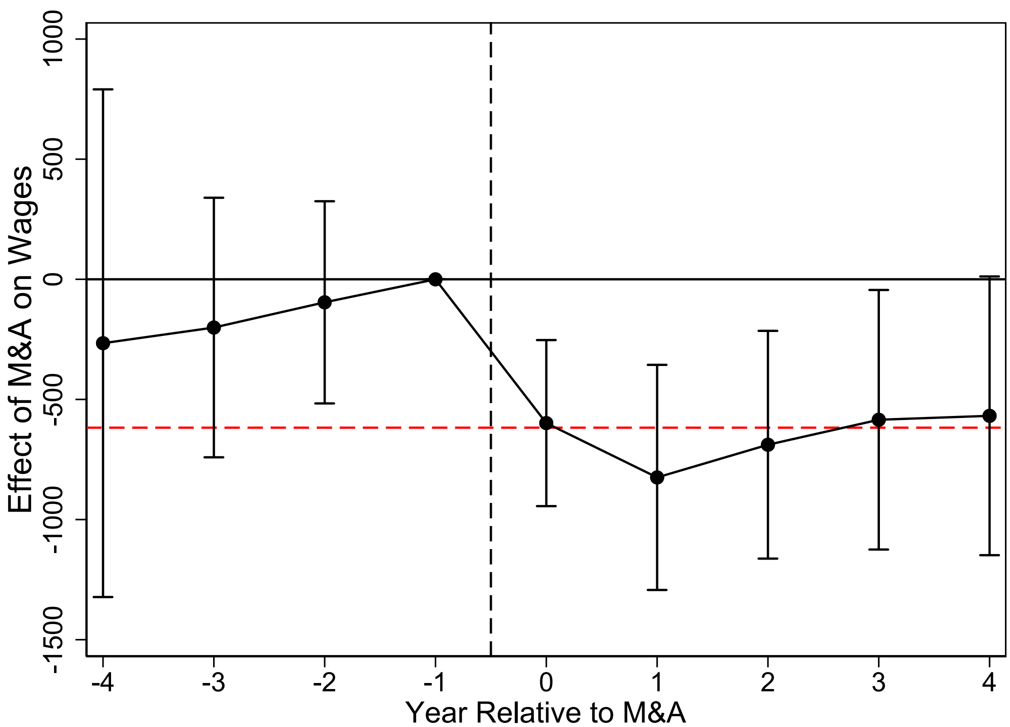
	(1)	(2)	(3)
Post M&A	-1,067.259***	-997.509***	-1,064.563***
	(315.619)	(272.378)	(231.251)
Observations	4,581,968	4,581,968	4,581,968
Adjusted R2	0.222	0.280	0.347
FE	MSA	MSA, IND	MSA, IND, OCC
# of MSAs	169	169	169
# of Treated MSAs	107	107	107
# of Control MSAs	62	62	62

(c) Panel C: Cross M&A

	(1)	(2)	(3)
Post M&A	597.313	479.928	437.896
	(395.472)	(340.003)	(309.218)
Observations	634,877	634,877	634,877
Adjusted R2	0.215	0.274	0.346
FE	MSA	MSA, IND	MSA, IND, OCC
# of MSAs	119	119	119
# of Treated MSAs	57	57	57
# of Control MSAs	62	62	62

Notes: Dependent variable is annual wage income. Health care workers are excluded. All regressions include controls for sex, race/ethnicity, education, and age along with year fixed effects. Regressions use ACS survey weights and standard errors are clustered by strata. Panel B includes as treated only MSAs that had within market mergers over the study period, but no cross market mergers. Panel C includes as treated only MSAs that had cross market mergers over the study period, but no within market mergers. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Figure 5: Association Between Hospital M&A Lags/Leads and Wages



Source: Authors' analysis of wage data from the American Community Survey (ACS) and hospital merger data from the American Hospital Association (AHA). Study period 2010 to 2018.

Notes: The dotted red line is a weighted average of the post-M&A coefficients and equals -$618.

Table 4: Effect of M&A on Hours Worked

(a) Panel A: All M&A

	(1)	(2)	(3)
Post M&A	0.094***	0.081**	0.074**
	(0.036)	(0.033)	(0.031)
Observations	5,960,618	5,960,618	5,960,618
Adjusted R2	0.100	0.154	0.218
FE	MSA	MSA, IND	MSA, IND, OCC
# of MSAs	290	290	290

(b) Panel B: Within M&A

	(1)	(2)	(3)
Post M&A	0.074	0.069	0.056
	(0.048)	(0.043)	(0.041)
Observations	4,581,968	4,581,968	4,581,968
Adjusted R2	0.098	0.150	0.214
FE	MSA	MSA, IND	MSA, IND, OCC
# of MSAs	169	169	169

(c) Panel C: Cross M&A

	(1)	(2)	(3)
Post M&A	0.070	0.048	0.041
	(0.096)	(0.083)	(0.076)
Observations	634,877	634,877	634,877
Adjusted R2	0.115	0.185	0.248
FE	MSA	MSA, IND	MSA, IND, OCC
# of MSAs	119	119	119

Notes: Dependent variable is annual wage income. Health care workers are excluded. All regressions include controls for sex, race/ethnicity, education, and age along with year fixed effects. Regressions use ACS survey weights and standard errors are clustered by strata. Panel B includes as treated only MSAs that had within market mergers over the study period, but no cross market mergers. Panel C includes as treated only MSAs that had cross market mergers over the study period, but no within market mergers. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 5: Subgroup Analyses

(a) No College vs. College

	Wage Income		Hours		Employed	
	No College	College	No College	College	No College	College
Post M&A	-169.866	-693.388^{*}	0.100^{***}	0.028	-0.000	-0.002
	(113.982)	(401.588)	(0.036)	(0.044)	(0.002)	(0.002)
Observations	3,359,388	2,601,225	3,359,388	2,601,225	8,124,332	3,786,803
Adjusted R2	0.303	0.296	0.259	0.146	0.053	0.044
Dep. Var. Sample Mean	41,240	81,976	39.5	42.6	0.63	0.82

(b) White vs. Non-White

	Wage Income		Hours		Employed	
	White	Non-White	White	Non-White	White	Non-White
Post M&A	-915.021^{***}	-193.355	0.062^{*}	0.061	-0.001	0.003
	(223.110)	(363.326)	(0.033)	(0.062)	(0.001)	(0.003)
Observations	4,656,777	1,303,841	4,656,777	1,303,841	8,731,636	3,179,499
Adjusted R2	0.350	0.361	0.232	0.177	0.070	0.078
Dep. Var. Sample Mean	61,276	50,951	41.0	40.0	0.71	0.63

(c) Male vs. Female

	Wage Income		Hours		Employed	
	Male	Female	Male	Female	Male	Female
Post M&A	-496.626^{*}	-777.196^{***}	0.110^{***}	0.027	0.001	-0.002
	(298.216)	(163.152)	(0.039)	(0.042)	(0.002)	(0.002)
Observations	3,209,576	2,751,041	3,209,576	2,751,041	6,071,155	5,839,980
Adjusted R2	0.342	0.338	0.185	0.199	0.075	0.062
Dep. Var. Sample Mean	69,793	46,446	43.0	38.3	0.74	0.64

(d) Under 40 vs. Over 40

	Wage Income		Hours		Employed	
	(1)	(2)	(3)	(4)		
	Under 40	Over 40	Under 40	Over 40	Under 40	Over 40
Post M&A	-224.161	$-1,105.132^{***}$	0.186^{***}	0.003	0.000	-0.002
	(196.674)	(288.582)	(0.047)	(0.034)	(0.002)	(0.002)
Observations	2,423,457	3,537,161	2,423,457	3,537,161	5,286,062	6,625,073
Adjusted R2	0.386	0.337	0.305	0.160	0.110	0.098
Dep. Var. Sample Mean	45,148	68,520	39.2	41.9	0.68	0.70

Notes: Health care workers are excluded. All regressions include controls for sex, race/ethnicity, education, and age along with MSA and year fixed effects. The wage and hours regressions additionally include industry and occupation fixed effects. Regressions use ACS survey weights and standard errors are clustered by strata. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 6: Benefit Design Results

	OLS			
	(1)	(2)	(3)	(4)
	CDHP	Total Cost Sharing	ln(Total Cost Sharing)	Cost Sharing %
Spending per enrollee	2.42e-05***	0.0331***	0.000118***	0.000847***
	(3.46e-07)	(0.00116)	(2.30e-06)	(2.38e-05)
Observations	27,482,473	27,482,473	27,482,473	27,482,473
R2	0.039	0.049	0.086	0.018
MSA FE	X	X	X	X

	2SLS			
	(1)	(2)	(3)	(4)
	CDHP	Total Cost Sharing	ln(Total Cost Sharing)	Cost Sharing %
Spending per enrollee	4.71e-05***	0.125***	0.000162*	-0.00835***
	(1.38e-05)	(0.0463)	(9.14e-05)	(0.000951)
Observations	27,478,643	27,478,643	27,478,643	27,478,643
R2	0.039	0.049	0.086	0.012
MSA FE	X	X	X	X
F-stat	1.7e+04	1.7e+04	1.7e+04	1.7e+04

Notes: CDHP is a dummy variable equal to one if an individual was enrolled in a consumer-driven health plan. Includes at 10% random sample of enrollees under 65 in the HCCI commercial claims database for years 2010-2016. All regressions include controls for sex, age band, and Charlson index along with MSA and year fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Appendix

Additional Tables and Figures

Table A1: Association Between Hospital M&A and Inpatient Prices/Spending per Enrollee – Log Version

(a) All M&A

	ln(Price)	ln(Price (100))	ln(Price (200))	ln(Spending)	ln(Spending (100))	ln(Spending (200))
Post M&A	0.028**	0.022*	0.021	0.058**	0.060**	0.053**
	(0.013)	(0.013)	(0.013)	(0.025)	(0.025)	(0.025)
Observations	1,963	1,963	1,963	1,963	1,963	1,963
Adjusted R2	0.937	0.925	0.930	0.931	0.911	0.921

(b) Within M&A

	ln(Price)	ln(Price (100))	ln(Price (200))	ln(Spending)	ln(Spending (100))	ln(Spending (200))
Post M&A	0.032**	0.015	0.021	0.064**	0.048*	0.051*
	(0.015)	(0.016)	(0.016)	(0.027)	(0.026)	(0.027)
Observations	1,149	1,149	1,149	1,149	1,149	1,149
Adjusted R2	0.939	0.923	0.927	0.923	0.894	0.907

(c) Cross M&A

	ln(Price)	ln(Price (100))	ln(Price (200))	ln(Spending)	ln(Spending (100))	ln(Spending (200))
Post M&A	0.024	0.024	0.021	0.070*	0.079**	0.070*
	(0.022)	(0.023)	(0.021)	(0.038)	(0.038)	(0.038)
Observations	970	970	970	970	970	970
Adjusted R2	0.931	0.918	0.924	0.926	0.902	0.915

Notes: HHI is scaled to range from 0 to 1. Price=inpatient price. Spending=inpatient spending per enrollee. (100) and (200) include only the top 100 and 200 most common DRGs, respectively. All regressions include MSA and year fixed effects. Observations more than three periods before or after treatment are excluded. The estimates shown are weighted averages of the event study's post-treatment coefficients. Standard errors are clustered by MSA. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table A2: Association Between Provider Market Concentration and Inpatient Prices/Spending per Enrollee – Log Version

	ln(Price)	ln(Price (100))	ln(Price (200))	ln(Spending)	ln(Spending (100))	ln(Spending (200))
Number of Hospitals	−0.0120**	−0.0095*	−0.0119***	−0.0193***	−0.0144*	−0.0175**
	(0.0047)	(0.0051)	(0.0044)	(0.0074)	(0.0076)	(0.0075)
Primary Care HHI	0.1549	0.0342	0.0908	0.4920**	0.2662	0.3698*
	(0.1643)	(0.1123)	(0.1429)	(0.2307)	(0.2180)	(0.2248)
Specialist HHI	0.0169	−0.0015	0.0274	0.1658	0.1232	0.1855
	(0.0722)	(0.0791)	(0.0798)	(0.1206)	(0.1234)	(0.1245)
VI Primary Care	−0.0006	−0.0001	−0.0003	−0.0019*	−0.0011	−0.0013
	(0.0007)	(0.0006)	(0.0006)	(0.0011)	(0.0010)	(0.0011)
VI Specialist	−0.0001	0.0002	0.0001	0.0001	0.0005	0.0004
	(0.0004)	(0.0005)	(0.0004)	(0.0008)	(0.0009)	(0.0008)
Observations	1,837	1,837	1,837	1,837	1,837	1,837
Adjusted R2	0.939	0.925	0.930	0.932	0.908	0.920

Notes: HHI is scaled to range from 0 to 1. Price=inpatient price. Spending=inpatient spending per enrollee. (100) and (200) include only the top 100 and 200 most common DRGs, respectively. All regressions include MSA and year fixed effects. Observations more than three periods before or after treatment are excluded. Standard errors are clustered by MSA. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$