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The Effects of Health Care Prices on Wages

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

Daniel R. Arnold[†] Christopher M. Whaley[‡]

July 21, 2020

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 \$521 increase in hospital prices, a \$579 increase in hospital spending among the privately insured population and a similar, \$638 reduction in wages. Both the hospital price and spending increases and the reduction in wages are driven by mergers that occur within hospital markets, rather than 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 caused by provider concentration are passed to workers in the form of lower wages and less generous benefits.

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1 Introduction

The effects of fringe benefits on worker wages is a long-standing economic question. Understanding the effect of benefits on wages is particularly relevant for health insurance, as 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 ([Cooper et al., 2019b](#); [White and Whaley, 2019](#); [Chernew, Hicks and Shah, 2020](#)). 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. Understanding the extent to which rising health care costs are passed-through to lower wages 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. Sev-

eral 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 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.

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, in particular – impact both health care costs and wages and other labor market outcomes for workers who receive 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 ag-

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

gregates cost 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.

We exploit hospital mergers as a source of exogenous variation. 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. (2019b)), 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 2010 to 2016 period, hospital mergers led to an \$521 (2.6%) increase in hospital prices and a \$579 increase in per-privately insured enrollee spending at hospitals. On the reverse side of the market, we find that hospital mergers lead to a \$637, approximately 1.0%, reduction in wages from 2010 to 2018. 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 is consistent with a recent study that found cross-market mergers within a state lead to more modest hospital price increases and cross-market mergers across state lines do not yield significant increases (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 for workers with a college degree and those over 40-years-old.

While our results indicate that employers respond to rising health care costs by reducing wages, employers may also 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 test 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 meaningful effect on the adoption of HDHPs. Our results imply that the level of increased health care spending observed over the 2010 to 2016 period contributed to 15% of the observed increased enrollment in high-deductible plans. Our two combined results suggest that firms respond to higher health care costs by both adopting the blunt instrument of HDHPs, and by using the even more blunt instrument of reducing wages.

This paper contributes to two relevant literatures. 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

²Several 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).

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

A similar framework comes from papers that model the wage impacts of payroll taxes. In [Section 2](#), we develop a model that builds from 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 similarly to responses observed from the literature on payroll taxes. Empirically, [Gruber \(1997\)](#) finds that a large reduction in payroll tax falls fully on worker wages, with no impacts on employment.

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., 2019b](#)). The same authors find prices for hospital services have increased much faster than for other health care services ([Cooper et al., 2019a](#)). Similar work has found that employer-funded health plans reimburse hospitals at 240% of Medicare rates ([White and Whaley, 2019](#)). A common reason for price variation is horizontal consolidation between

providers, and vertical integration among hospitals and 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. (2019b) 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, 2009a; Gowrisankaran, Nevo and Town, 2015a). 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 2010-2016 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., 2019b; 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 be 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 are 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—hospital mergers—as a source of exogenous variation. 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. [Section 6](#) analyzes the impact of rising health care costs on benefit design and [Section 7](#) concludes.

2 Conceptual Framework

Our goal in this paper is to estimate the 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). Suppose that firms provide health insurance to their employees and labor demand (L_d) is given by

$$L_d = f_d(W + C), \quad (1)$$

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

$$L_s = f_s(W + \alpha C), \quad (2)$$

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 may be close to zero.

In equilibrium, it can be shown that

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

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 will be given by

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

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

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

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 for health care services, we used 2010-2016 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 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 Standardized Prices

We focused our measurement of prices to prices paid 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.

To construct an average inpatient price for every MSA-year combination in the data, we first sum the amounts paid of all facility claims associated with an admission. Second,

we average the prices of all admissions (in a given MSA-year combination) that have the same primary diagnosis-related group (DRG) code. This creates a DRG-MSA-year dataset. Finally, we weight observations by the number of admissions associated with each DRG code. This leads to an MSA-year level average inpatient price dataset (hereafter, standardized prices).

3.1.2 Price Index

We construct the price index as follows. First, we use the weighted average ratio of the market-level price for a specific DRG 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 DRG-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 DRG, indexed by k , as

$$w_k = \frac{price_k q_k}{\sum_{k=1}^K price_k q_k} \quad (6)$$

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

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

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

3.2 Hospital Spending

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 disaggregated 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 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. Specifically, we divide hospital revenues from private patients by the number of privately insured enrollees in each market to create a measure of spending per enrollee.

³We use the HCRIS data provided by the RAND Hospital data <https://www.rand.org/pubs/tools/TL303.html>.

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 a census of U.S. hospitals. AHA data is widely used to measure hospital market concentration (Cooper et al., 2019b; 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. (2019b) 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, independent physician association (IPA), and independent.

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 weights 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 2010-2018 ACS data. This sample contains 17.1 million individuals between the ages of 18 and 64, an average of just under 2 million per year. In our main analysis, we restrict the ACS population to those (1) currently employed and who receive insurance from an employer

and (2) are not health care workers. This restriction limits the sample size by 54%, to a total of 7.8 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.

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. Restricting the data to individuals located in the 290 MSAs available in the ACS data leads to an analytic sample of 6.0 million people.

4 Empirical Approach

We estimate the impacts of hospital mergers on both hospital price/spending and labor market outcomes. This approach allows us to estimate the spending and wage impacts of changes in hospital market structure, but does not account for variation that impacts the entire country, such as the introduction of new technologies. To implement our approach we estimate a first stage and a reduced form model. The first stage model estimates the impact of hospital mergers on hospital prices and spending among the privately insured population.

Specifically, we estimate a market-level regression of the form

$$price_{gt} = \alpha + \theta merger_{gt} + \zeta_g + \tau_t + \epsilon_{igt}, \quad (8)$$

where $price_{gt}$ is the standardized price in market g in year t and $merger_{gt}$ is a dummy variable equal to one if market g had experienced a hospital merger during our study period in or before year t . Market (ζ_g , MSA) and year (τ_t) fixed effects are also included to account for time-invariant market differences and temporal trends, respectively. We estimate this regression using OLS for three different outcome measures – $price_{gt}$ (shown in equation 8), $index_{gt}$, and $spending_{gt}$, where $index_{gt}$ and $spending_{gt}$ are the price index and hospital spending per enrollee measures discussed above.

Our reduced form model estimates the impact of hospital mergers on wages. Specifically, for each ACS respondent i in market g during year t , we estimate a regression of the form

$$wage_{igt} = \alpha + \gamma merger_{gt} + \beta X_{igt} + \zeta_g + \tau_t + \epsilon_{igt}, \quad (9)$$

where $wage_{igt}$ is the annual wage income of worker i in market g in year t , $merger_{gt}$ is a dummy variable equal to one if market g had experienced a hospital merger during our study period in or before year t , and X_{igt} is a robust set of controls (age, gender, sex, race, education). Market (ζ_g , MSA) and year (τ_t) fixed effects are also included to 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 while applying ACS sampling weights and clustering standard errors at the level of the ACS' sampling strata. We obtain similar results when clustering at the MSA-level and when not applying the ACS sampling weights.

The θ coefficient on $merger_{gt}$ in 8 measures the effect of hospital mergers on standardized prices, the price index, or hospital spending per enrollee, depending on the outcome measure used. The γ coefficient on $merger_{gt}$ measures the effect of hospital mergers on wages. We

interpret the ratio of the estimated γ and θ coefficients similarly to how they would be interpreted in a standard instrumental variables framework. That is, the ratio of γ and θ measures the pass-through of standardized prices (or the price index or hospital spending per enrollee) to wages, using the local variation in prices/spending caused by hospital mergers.

In both models, a causal interpretation requires the standard difference-in-differences assumptions. A potential threat to a causal interpretation could occur through the non-random and selected nature of hospital mergers. Hospital and other health care providers derive pricing power through internalizing patient willingness to pay for services (Ho, 2009b; Gowrisankaran, Nevo and Town, 2015b). 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. A violation of the validity of our 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. As with other consolidation events, most hospital mergers require regulatory approval and it is not uncommon for delays and hospitals strategically based on wage shocks would have to account for these idiosyncratic delays.

We address these potential threats using several approaches. First, when estimating the effects of mergers on both sets of outcomes, we include MSA fixed effects, and thus remove an time-invariant differences in market incomes. In the labor-market results, we also include fixed effects for industry and occupation. Finding that including these additional controls leads to large differences in our estimated effects would suggest that our results may be driven by market-level changes in the composition of workers or firms, which may be related to hospital decisions to merge. However, we find that including these controls leads to little change in our results. We also test for parallel trends by using an event study approach, where we test for differences in prices/spending and labor market outcomes in the years before and after mergers occur. We find stable and null differences in both sets of outcomes

in the years prior to merger, but changes in prices/spending and wages following mergers. Finally, we examine within- and cross-market hospital mergers. We find that our effects for both outcomes are driven by within-market mergers but if our results are driven by unobserved shocks, then we should find similar effects for cross-market mergers.

5 Results

5.1 Descriptive Characteristics

5.1.1 Price Trends

Figure 2 plots trends in standardized prices over our study period. From 2010 to 2016, standardized prices (weighted by MSA population) increased from \$15,337 to \$20,751, an absolute difference of \$5,414 and a relative difference of 35%. However, as shown in Figure 3, which normalizes prices to each MSA's 2010 price levels and plots the mean, 25th percentile, and 75th percentile price growth, MSAs vary in their price growth. While the mean MSA has experienced a price increase of 30%, the 25th percentile growth is 27% and the 75th percentile growth is 34%. Figure 4 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 Effect of Hospital Mergers on Prices and Spending

Table 1 presents the effects of hospital mergers on standardized prices (column 1), indexed prices (column 2), and hospital spending (column 3). For all hospital mergers (Panel A), mergers are associated with a \$521 increase in mean prices, which translates to a 2.6% relative increase in prices, and a \$579 in per-enrollee spending on hospitals by patients with private insurance. These coefficients were calculated as a weighted average of the post-merger coefficients that we estimated in the event study versions of our models (see Figure 5). We

find stronger effects for mergers that occur within-market (Panel B) than mergers that occur between markets (Panel C). For within-market mergers, we find a \$703 price increase, which translates to a relative 3.9% increase, and a \$513 increase in hospital spending, albeit not precisely estimated. Consistent with the results in [Dafny, Ho and Lee \(2019\)](#), which finds that only cross-market mergers within a state have an impact on prices, we do not find that cross-market mergers lead to higher prices or hospital spending.

Figure 5 presents graphical depictions of the event study versions of our models. The dotted red lines in the figures correspond to the estimates shown in Panel A of Table 1. 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 \$768 higher. 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 lack of a pre-trend also occurs when using indexed prices as the dependent variable. Relative prices are centered around zero prior to a merger occurring, and then steadily increase in each year post-merger to a maximum of 0.038 in the fourth year following a hospital merger. For hospital spending among the privately insured population, spending differences are centered around \$0 pre-merger and steadily increase to \$893 in the fourth year following a hospital merger.

As presented in Figure 6, we find consistent results when separating mergers that happen within market from cross-market mergers. For within-market mergers, we do not find pre-implementation differences for standardized prices, indexed prices, or hospital spending. For each measure, we observe increases in each post-implementation year, with the strongest increases for the standardized and indexed price measures. For cross-market mergers, we do not observe pre-merger differences in the price or spending measures, or trends in the measures. Following cross-market mergers, we observe modest increases in each outcome, but the results are not statistically significant in any post-merger year.

5.3 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). This estimate corresponds to the estimate of γ in equation 9. 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.

Following the price and spending results, Panels B and C present similar results, but distinguish between within-MSA hospital mergers 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 mean changes in worker wages.

Figure 7 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. For this event study, we use the specification in column 3 of Table 2 that includes MSA, industry, and occupation fixed effects. In the four years prior to hospital mergers, wages are not statistically different from \$0 between the treatment and control MSAs. Following hospital merger activity, there is an immediate reduction in wages, which slightly increases in magnitude in the first year following mergers. The trend stabilizes in the remaining post-merger years.

Consistent with both the mean results in Table 2 and the price and spending results

in Figure 6, the results in Figure 8 show meaningful differences in the wage impacts of within- and cross-market mergers. For within-market mergers (Panel A), pre-merger prices are centered around \$0 and have wide confidence intervals. Following within-market mergers, wages drop by \$772, and increase further to \$1,059 in the second post-merger year. By the fourth year following a merger, wages are \$643 lower than the year prior to the merger. Although the 95% confidence interval overlaps zero, the result is still statistically significant at conventional thresholds (p-value = 0.054). For cross-market hospital mergers, we likewise do not find a pre-merger difference in prices. However, following a merger, prices actually increase, albeit the results are not statistically significant.

These wage results mirror the results we find when looking at the effect of hospital mergers on hospital prices and spending. Following mergers, we find price and spending increases among privately-insured patient populations that are similar in magnitude to the reduction in wages among workers who receive private insurance through an employer. Consistent with previous work (Dafny, Ho and Lee, 2019), we do not find that mergers that occur between hospitals in separate markets lead to changes in prices or changes in wages. Instead, both the hospital price and spending increases and the reductions in wages occur following hospital mergers that occur between hospitals in the same market.

In Table 3, we examine the effect of hospital mergers on hours worked. As in our analysis of worker wages, the sample population is limited to workers who are employed and who receive insurance through an employer or union. Thus, our results are interpreted as the intensive-margin change in hours, conditional on working. For all hospital mergers (Panel A), we find a 0.09-hour to 0.07-hour reduction in weekly hours worked. Off of the baseline mean of 41 weekly hours, our results indicate an approximately 0.2% increase in weekly hours. In Panels B and C, we observe similar coefficient magnitudes when separating between within- and cross-market mergers. However, these estimates are not statistically significant.

Our final labor market outcome is the probability of being employed. For this analysis, we expand the sample to include all ACS respondents ages 18 to 64. We do not find any

effect of combined (Panel A) or within-market hospital mergers on employment. We find a small increase in employment following cross-market mergers. However, while precisely estimated, the magnitude of the coefficient indicates a less than 1% relative increase in the probability of employment.

5.3.1 Heterogeneous Effects of Hospital Mergers on Labor Market Outcomes

We next examine how these results vary by patient characteristics and demographics (Table 5). As shown in Panel A, we find that the effects are 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 of mergers are largest for white workers, but baseline mean wages are 16.8% lower for non-whites than for whites. This difference is after adjustment for age, education and MSA, industry, and occupation, and could reflect systematic race-based differences in economic opportunities (Chetty et al., 2020). We find similar magnitude differences on the number of hours worked, but the results are only statistically significant for white workers. For worker gender, we find that mean wages are 33.5% lower for women 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.

6 Impacts on Benefit Design

Finally, we 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 health care spending, which includes spending on inpatient, outpatient, and physician services in what follows, leads to an increased probability of enrollment in a HDHP. For computational reasons, we select a random 10% sample of the HCCI population, which leaves us with 27.5 million patient-year observations.

Among this population, we estimate the impacts of annual health care spending per enrollee on the likelihood of enrollment in a HDHP plan, the annual amount spent by patients in the form of cost-sharing payments, and the share of annual spending attributed to a patient that is paid by the patient. We use both panel variation in spending and instrument for spending using hospital mergers and estimate the following model:

$$\begin{aligned} spending_{gt} &= \alpha + \gamma merger_{gt} + X_{it} + \zeta_g + \tau_t + \mu_{igt} \\ HDHP_{igt} &= \alpha + \delta \widehat{spending}_{gt} + X_{it} + \zeta_g + \tau_t + \varepsilon_{igt}, \end{aligned}$$

where $spending_{gt}$ is average health care spending per enrollee in market g in year t , $merger_{gt}$ is a dummy variable equal to one if market g had experienced a hospital merger during our study period in or before year t , and X_{igt} is a set of controls (age, gender, Charlson index). Market (ζ_g , MSA) and year (τ_t) fixed effects are also included to account for time-invariant market differences and temporal trends, respectively. In the second equation, $HDHP_{igt}$ equals one if individual i is enrolled in a HDHP in year t and zero otherwise and $\widehat{spending}_{gt}$

is the the predicted spending estimated from the first equation. We estimate the second equation for three outcomes measures in addition to $HDHP_{igt} - cs_{igt}$ (the dollar value of cost-sharing paid by enrollee i in year t), $ln cs_{igt}$ (the natural log of the cost-sharing paid by enrollee i in year t), and $pct cs_{igt}$ (the amount of cost sharing paid by enrollee i in year t as a percent of total health care spending by enrollee i in year t).

As shown in the first column of Table 6, we find that a \$1 increase in health care spending per enrollee leads to a 0.0024 percentage point increase in the probability of enrollment in a high-deductible health plan. This effect translates to a slightly larger 0.047 percentage point increase when instrumenting for hospital prices using hospital mergers (Panel B). When applying the mean increase observed earlier from an increase in spending on hospital services, \$579, the effect translates to a 2.7 percentage point increase in the likelihood of HDHP enrollment.

We also estimate similar regressions that test if increases in local-market health care spending lead to changes in patient cost sharing. 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 health care spending 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 \$579 results in a patient cost-sharing increase ranging from \$17 to \$72.

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 spending per enrollee increases, patients are responsible for a smaller relative portion of total health care spending after instrumenting for price increases. We estimate that a \$100 increase in health care spending per enrollee 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.

To place these results in context, inflation-adjusted spending among this population increased by \$501 (18% relative increase) over the 2010 to 2016 period. Our results thus imply that this increase in spending lead to a 0.012 percentage point increase in high-deductible plan enrollment. Over this time period, enrollment in HDHP plans increased by 15.4 percentage points, from 14.1% in 2010 to 29.5% in 2016. Thus, our results imply that rising health care spending accounted for 0.08% of the growth in high-deductible health plans.

7 Conclusion

This paper examines the relationship between changes in provider market structure on both 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, hospital mergers, to estimate the effect of health care prices on wages. We find that hospital mergers lead to a \$521 increase in hospital prices, a \$579 increase in mean hospital spending, and a \$638 reduction in wages. The effects are driven by within-market hospital mergers.

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. Our results imply that health care reforms with mechanisms for lowering prices are likely underestimating their potential savings if they do not include impacts on wages.

Many recent policy proposals seek to constrain health care cost growth in the U.S. 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. Appropriately accounting for the incidence of health care spending growth is critical for evaluating proposed policy reforms.

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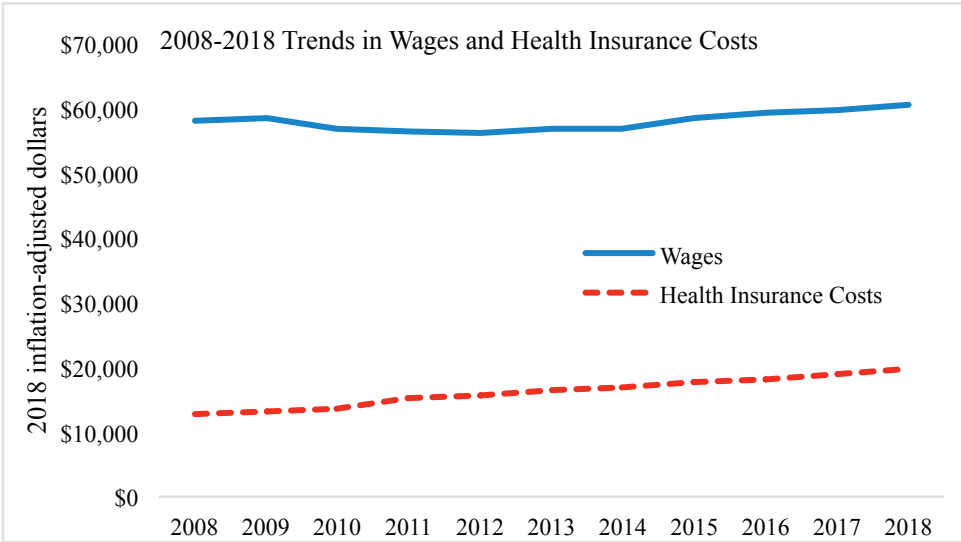
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Tables and Figures

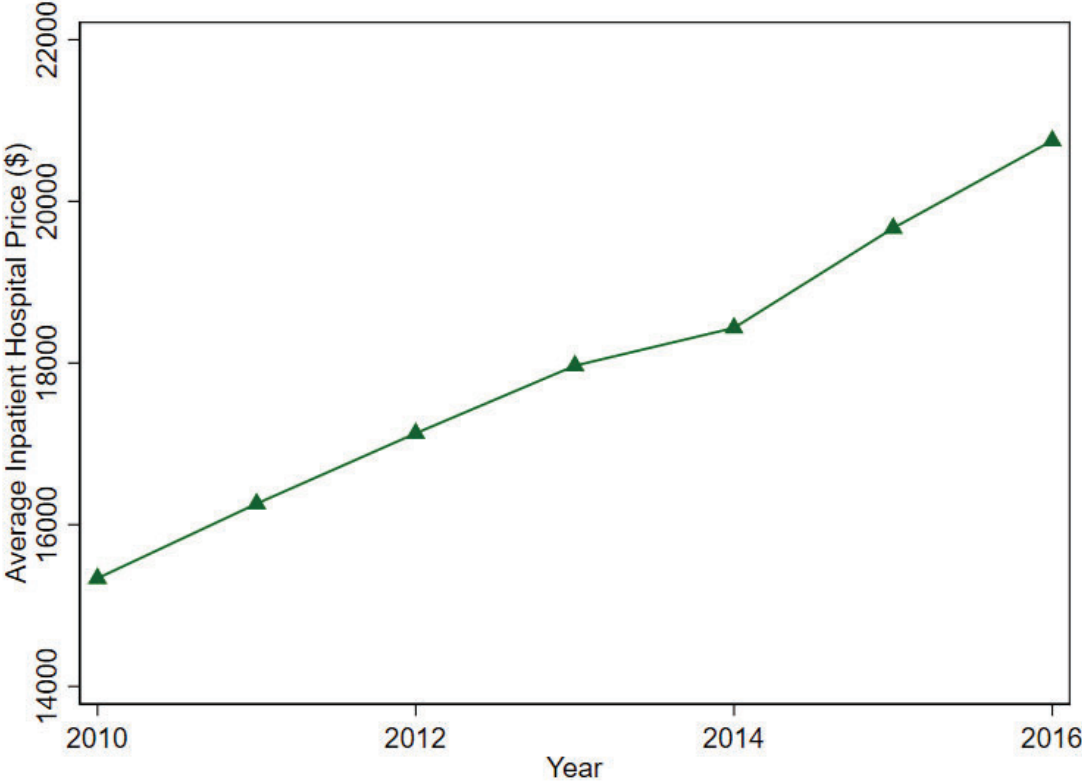
Figure 1: 2008 to 2018 Trends in Wages and Employer Health Insurance Costs



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.

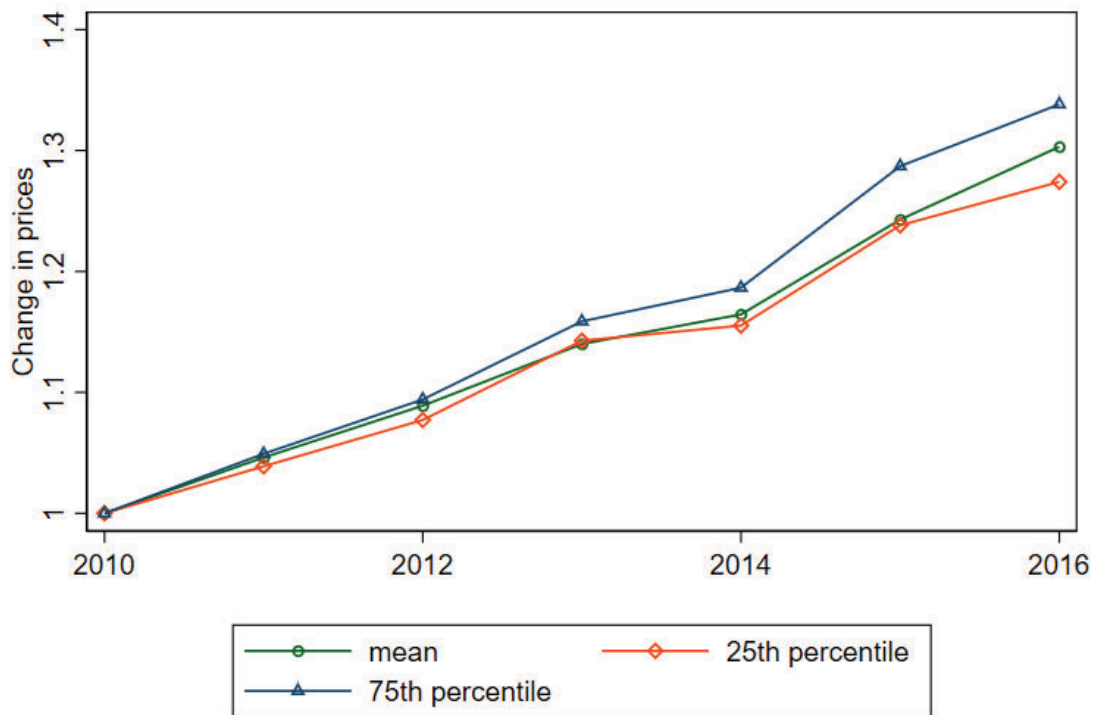
Figure 2: Standardized Prices (weighted by MSA population), 2010-2016



Source: Authors' analysis of commercial claims data from the Health Care Cost Institute (HCCI).

Notes: Standardized prices are first calculated at the DRG level and then weighted by DRG volume to create one average number for each year.

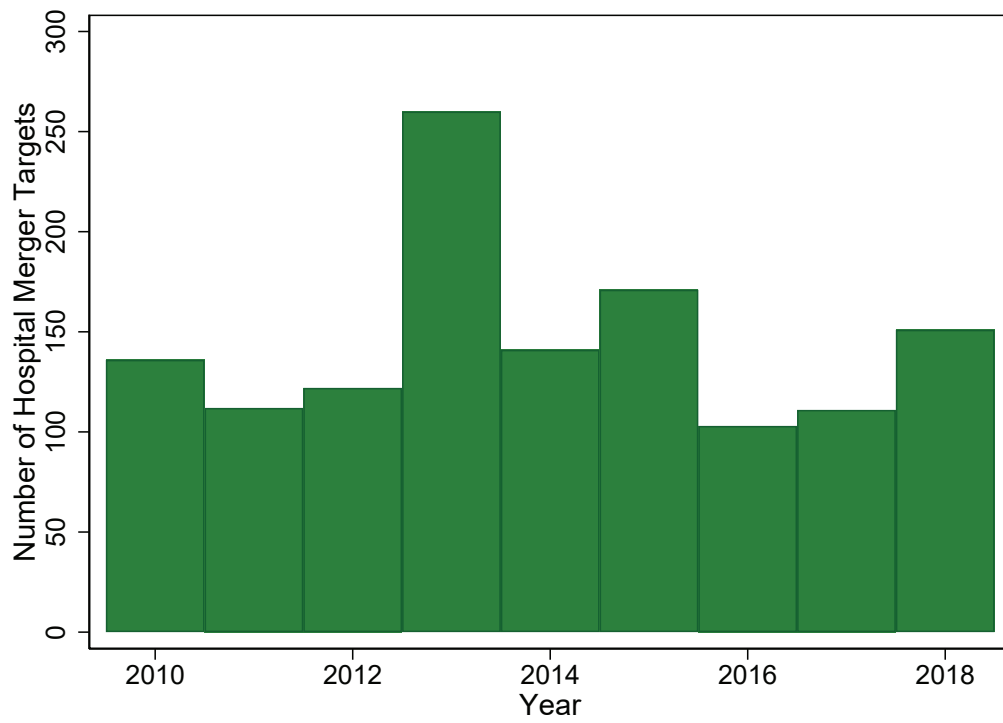
Figure 3: Standardized Price Growth Across MSAs, 2010-2016



Source: Authors' analysis of commercial claims data from the Health Care Cost Institute (HCCI).

Notes: The standardized price level in each MSA was normalized to be 1 in 2010.

Figure 4: 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: First Stage Results: Effect of Mergers on Hospital Prices and Spending per Enrollee

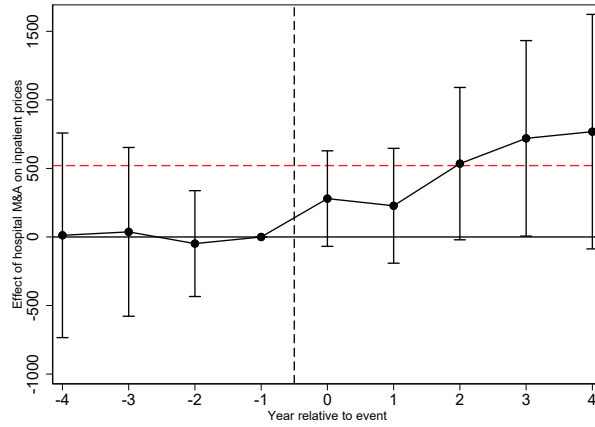
(a) All M&A			
	Price	Price Index	Spending per Enrollee
Post M&A	520.912* (275.244)	0.026* (0.015)	578.731* (309.419)
Observations	2,660	2,660	2,492
No. of MSAs	380	380	356
Adjusted R2	0.908	0.927	0.833
Dep. Var. Sample Mean	16,738	1.01	5,155
(b) Within M&A			
	Price	Price Index	Spending per Enrollee
Post M&A	703.117** (356.267)	0.039** (0.018)	512.723 (406.224)
Observations	1,526	1,526	1,421
No. of MSAs	218	218	203
Adjusted R2	0.918	0.923	0.851
Dep. Var. Sample Mean	16,551	0.99	5,317
(c) Cross M&A			
	Price	Price Index	Spending per Enrollee
Post M&A	742.885 (474.271)	0.016 (0.022)	157.207 (340.669)
Observations	1,239	1,239	1,085
No. of MSAs	177	177	155
Adjusted R2	0.900	0.917	0.799
Dep. Var. Sample Mean	17,025	1.08	4,652

Source: Authors' analysis of inpatient hospital prices from the Health Care Cost Institute (HCCI) and hospital spending from the Healthcare Cost Report Information System (HCRIS).

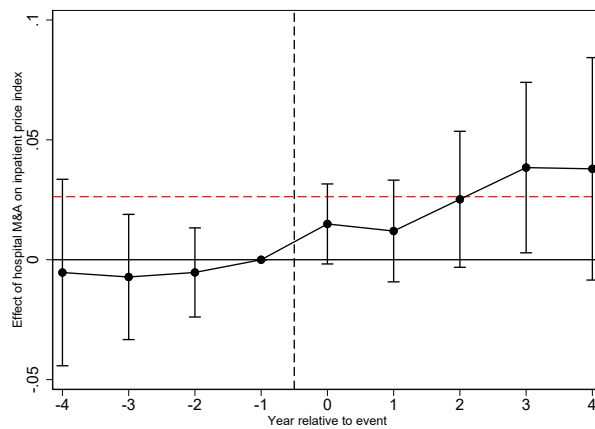
Notes: All regressions include MSA and year fixed effects. The estimates shown are weighted averages of the event study's post-treatment coefficients. Standard errors are clustered by MSA. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Figure 5: Event Study: Association Between Hospital M&A Lags/Leads with Inpatient Prices and Spending

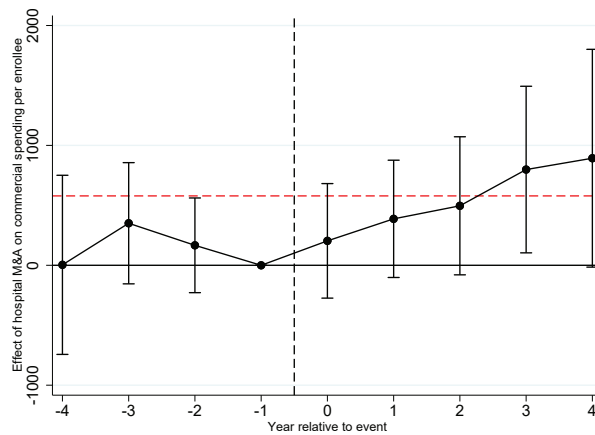
(a) Standardized Prices



(b) Price Index



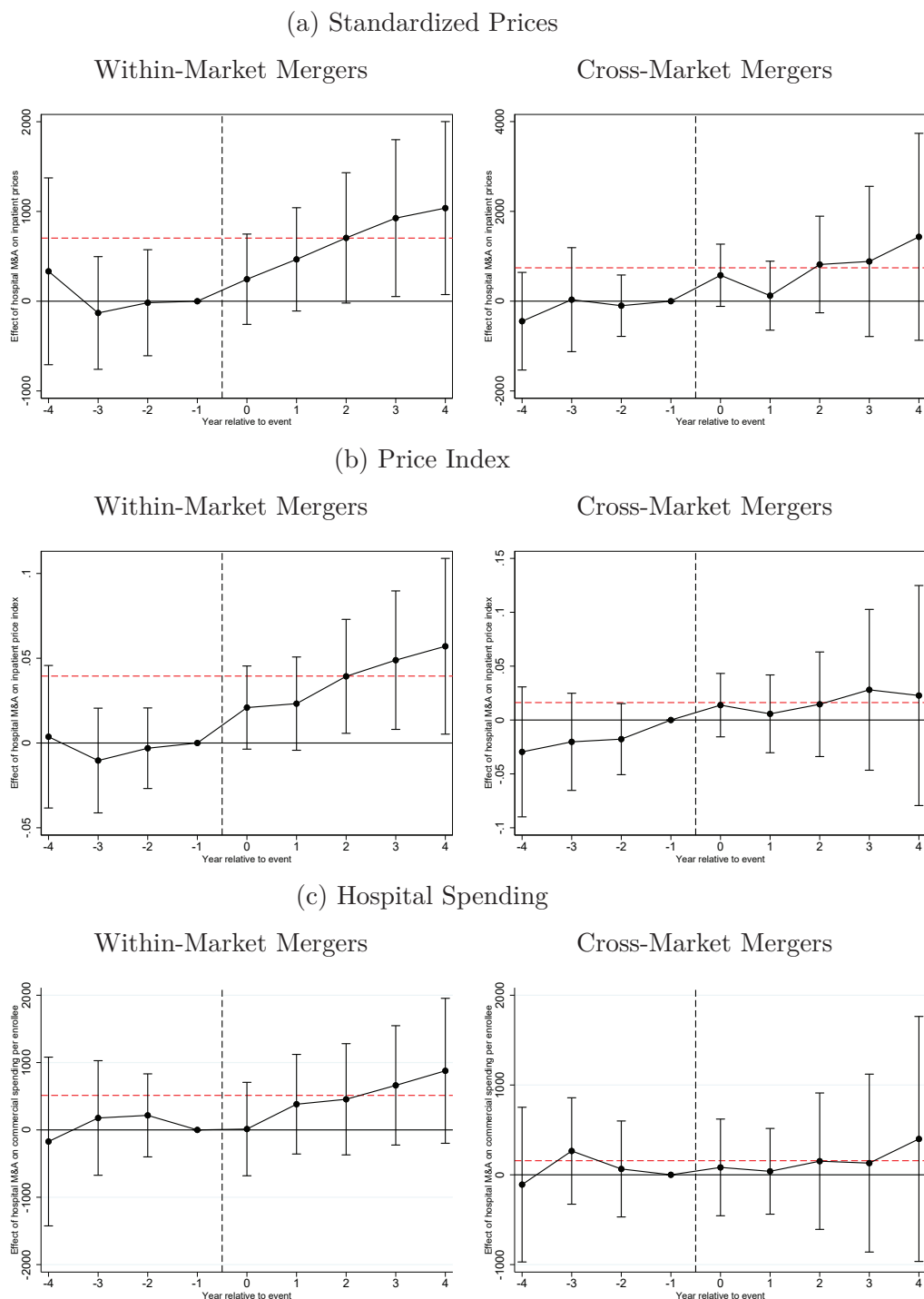
(c) Hospital Spending



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 lines are weighted averages of the post-M&A coefficients shown in Panel A of Table 1.

Figure 6: Event Study: Association Between Within and Cross-Market Hospital M&A Lags/Leads with Inpatient Prices and Spending



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 lines are weighted averages of the post-M&A coefficients shown in Panels B and C of Table 1.

Table 2: Effect of M&A on Wage Income

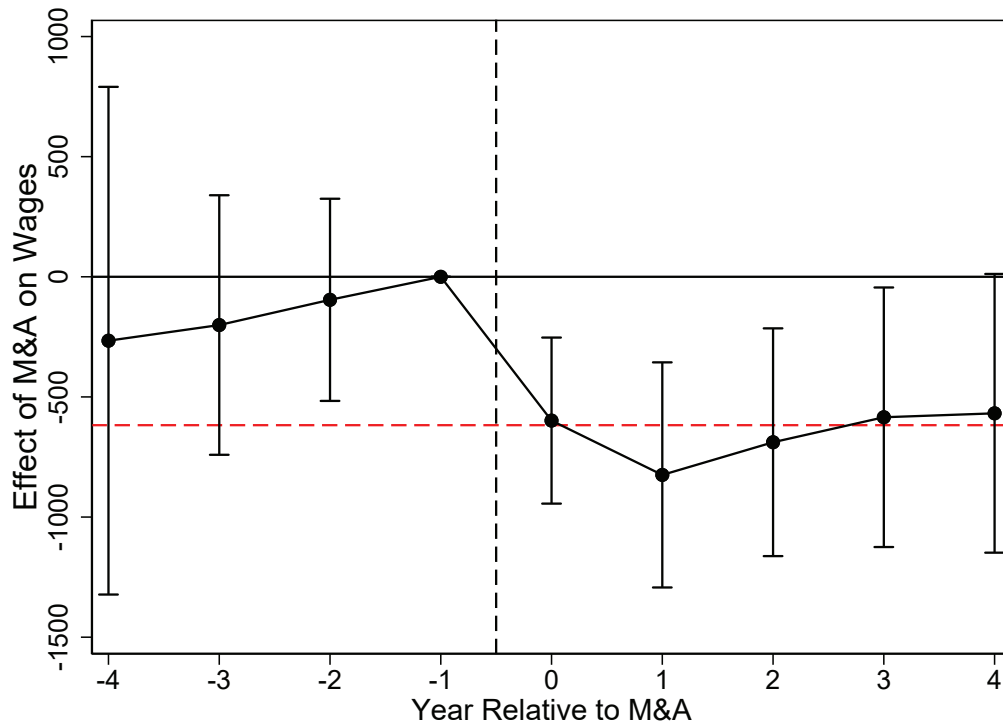
(a) All M&A			
	(1)	(2)	(3)
Post M&A	-621.977** (283.632)	-610.277** (245.654)	-637.948*** (216.483)
Observations	5,960,618	5,960,618	5,960,618
Adjusted R2	0.225	0.284	0.351
FE	MSA	MSA, IND	MSA, IND, OCC
# of MSAs	290	290	290
# of Treated MSAs	228	228	228
# of Control MSAs	62	62	62

(b) Within M&A			
	(1)	(2)	(3)
Post M&A	-1,067.259*** (315.619)	-997.509*** (272.378)	-1,064.563*** (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) Cross M&A			
	(1)	(2)	(3)
Post M&A	597.313 (395.472)	479.928 (340.003)	437.896 (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. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

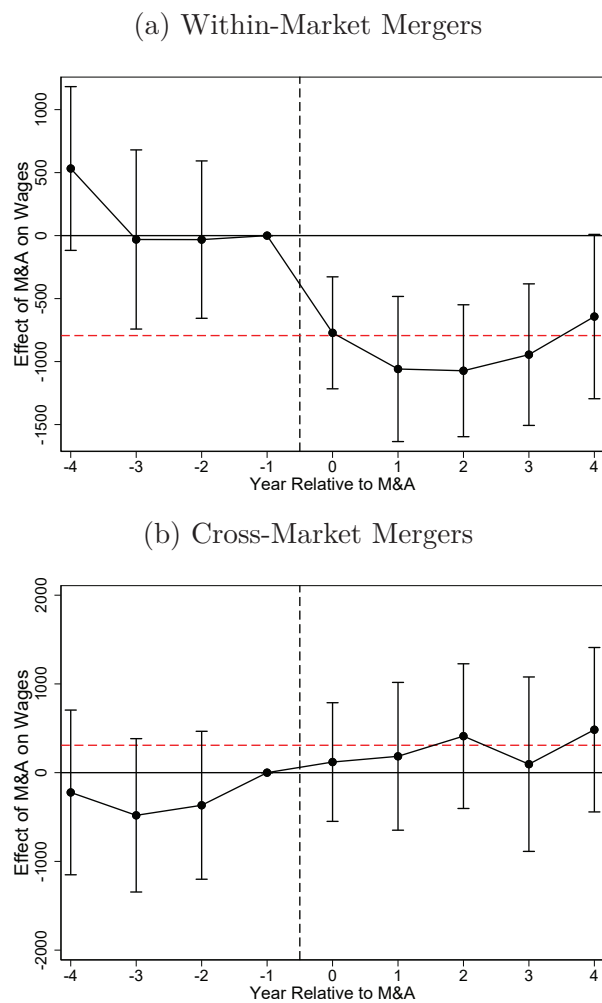
Figure 7: Association Between Hospital M&A Lags/Leads and Wages



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.

Figure 8: Association Between Within and Cross-Market 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 -\$794 and \$308 for Panels A and B, respectively.

Table 3: Effect of M&A on Hours Worked

(a) Panel A: All M&A			
	(1)	(2)	(3)
Post M&A	0.094*** (0.036)	0.081** (0.033)	0.074** (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 (0.048)	0.069 (0.043)	0.056 (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 (0.096)	0.048 (0.083)	0.041 (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 4: Effect of M&A on the Probability of Being Employed

(a) All M&A	
(1)	
Post M&A	-0.001 (0.002)
Observations	11,911,135
Adjusted R2	0.075
FE	MSA
# of MSAs	290

(b) Within M&A	
(1)	
Post M&A	-0.002 (0.002)
Observations	9,154,109
Adjusted R2	0.076
FE	MSA
# of MSAs	169

(c) Cross M&A	
(1)	
Post M&A	0.009*** (0.003)
Observations	1,353,721
Adjusted R2	0.070
FE	MSA
# of MSAs	119

Notes: Dependent variable is a dummy variable equal to one if employed, zero otherwise. Includes all 18-64 year olds in the sample, except for health care workers. 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 (113.982)	-693.388* (401.588)	0.100*** (0.036)	0.028 (0.044)	-0.000 (0.002)	-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*** (223.110)	-193.355 (363.326)	0.062* (0.033)	0.061 (0.062)	-0.001 (0.001)	0.003 (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* (298.216)	-777.196*** (163.152)	0.110*** (0.039)	0.027 (0.042)	0.001 (0.002)	-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) Under 40	(2) Over 40	(3) Under 40	(4) Over 40	Under 40	Over 40
Post M&A	-224.161 (196.674)	-1,105.132*** (288.582)	0.186*** (0.047)	0.003 (0.034)	0.000 (0.002)	-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

	<u>OLS</u>			
	(1)	(2)	(3)	(4)
	CDHP	Total Cost Sharing	ln(Total Cost Sharing)	Cost Sharing %
Spending per enrollee	2.42e-05*** (3.46e-07)	0.0331*** (0.00116)	0.000118*** (2.30e-06)	0.000847*** (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

	<u>2SLS</u>			
	(1)	(2)	(3)	(4)
	CDHP	Total Cost Sharing	ln(Total Cost Sharing)	Cost Sharing %
Spending per enrollee	4.71e-05*** (1.38e-05)	0.125*** (0.0463)	0.000162* (9.14e-05)	-0.00835*** (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. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$