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ABSTRACT

We study the effect of mandates requiring COVID-19 vaccination among healthcare industry workers adopted in 2021 in the United States. There are long-standing worker shortages in the U.S. healthcare industry, pre-dating the COVID-19 pandemic. The impact of COVID-19 vaccine mandates on shortages is ex ante ambiguous. If mandates increase perceived safety of the healthcare industry, marginal workers may be drawn to healthcare, relaxing shortages. On the other hand, if marginal workers are vaccine hesitant or averse, then mandates may push workers away from the industry and exacerbate shortages. We combine monthly data from the Current Population Survey 2021 to 2022 with difference-in-differences methods to study the effects of state vaccine mandates on the probability of working in healthcare, and of employment transitions into and out of the industry. Our findings suggest that vaccine mandates may have worsened healthcare workforce shortages: following adoption of a state-level mandate, the probability of working in the healthcare industry declines by 6%. Effects are larger among workers in healthcare-specific occupations, who leave the industry at higher rates in response to mandates and are slower to be replaced than workers in non-healthcare occupations. Findings suggest trade-offs faced by health policymakers seeking to achieve multiple health objectives.

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1 Introduction

The COVID-19 pandemic - like other influenza, plague, and cholera pandemics during the past several centuries - was a massive shock to the global economy (Jordà et al., 2022). Unlike wars that impact economies primarily by destroying physical capital, pandemics do so predominantly by affecting labor supply and demand, and through general disruptions in supply chains that can distort labor markets and lead to changes in consumer demand for goods and services. The emergence of healthcare innovations that effectively reduce the spread of diseases can curtail the adverse impacts of pandemics on economies. Effective vaccines were a major innovation globally in response to the COVID-19 pandemic. However, in the United States, and despite high vaccine effectiveness and government campaigns that included removing financial barriers and information provision, many Americans opted against vaccination. As of September 2021, only 70% of American adults had received the primary vaccine that had been widely available since the spring (Hamel et al., 2021). The non-trivial share of unvaccinated Americans may have reduced the benefits of vaccines to the U.S.

To increase vaccination rates, many states adopted COVID-19 vaccine mandates for workers in specific industries, including healthcare, in 2021. The response of the American public to these mandates was decidedly mixed. Proponents of the healthcare industry mandate argued that mandates were required to reduce disease contraction in healthcare facilities - patients in these facilities are generally vulnerable to infectious diseases such as COVID-19 and thus require protection, while critics were concerned about negative impacts on the economy, individual rights, vaccine safety, and so forth (Canning et al., 2022). In terms of the response of healthcare workers to the mandates, the impact is ex ante unclear. On the one hand, vaccines may have promoted a safe workplace, drawn workers to the healthcare sector, and reduced ongoing workforce shortages. Alternatively, workers who preferred to remain unvaccinated may have decided to exit, or not enter, the healthcare industry, exacerbating pre-pandemic shortages. Shekhar et al. (2021) and Caiazzo and Stimpfel (2022) provide evidence demonstrating vaccine hesitancy among healthcare workers, but to date, the impact on employment of healthcare sector workers has not been evaluated, and we are the first to study this question. How healthcare workers respond to vaccine mandates is important for policymakers to understand for future pandemics and for planning the staffing of healthcare facilities.

We estimate the impact of state COVID-19 vaccine mandates on employment in the healthcare industry. We draw data from the monthly Current Population Survey (CPS) from January 2021 to December 2022 – the period in which state COVID-19 employment vaccine mandates were in place. We match vaccine mandates to the CPS data on state and month, and isolate healthcare workers based on detailed industry information available in the CPS. We apply difference-in-differences methods and control for state vaccine mandates in other industries, as these mandates may impact workers' next-best option if they choose to leave, or not enter, healthcare for another industry.

Due to pre-COVID-19 shortages or working under constant pressure since the beginning of the COVID-19 pandemic, there are elevated labor market shortage concerns about some specific types of workers such as healthcare workers in long-term care facilities, registered nurses, and mental healthcare providers. To investigate whether state vaccine mandates exacerbated the shortages in these particular industries, we specifically analyze the effect of the mandates on the likelihood of employment in these sectors. In part of the analysis, we separately consider workers in the healthcare industry whose occupations are specific to healthcare and those that are more general. Exit of workers in healthcare-specific occupations may be particularly concerning as these workers are directly involved in patient care and are more likely to have industry-specific skills and to be more difficult to replace if they quit. For example, a dental or medical assistant plausibly requires much more industry-specific training than does a landscaper or cook.

Beyond examining 'level' effects, that is whether or not a worker is employed in the healthcare sector, we also explore the extent to which healthcare COVID-19 vaccine mandates impact 'transitions,' entry and exit by workers to and from the healthcare sector. To estimate transitions, we exploit the panel nature of the CPS and link respondents over time, tracking the industries in which they work to observe such transitions. Exploring entry and exit allows us to assess whether changes in levels are driven by incumbent workers leaving the industry or potential new workers electing not to enter, behaviors which could have differential impacts on the experience of the healthcare workforce, and potentially quality of care. Thus, we are able to study whether COVID-19 vaccine mandates impact the stock and flow of healthcare industry workers.

Our findings suggest that state mandates requiring COVID-19 vaccines among employees reduce the probability of working in the healthcare industry by 6%. The probability declines to a somewhat larger degree among healthcare occupations (7%) than non-healthcare occupations (5%).[1] We explore heterogeneity across state policies (e.g., social insurance) and worker demographics (e.g., sex, race, education), and demonstrate broadly similar effects of the mandates across the groups we consider. We show that states adopting and not adopting a COVID-19 vaccination mandate follow similar trends in healthcare employment pre-mandate and that results are robust to a range of sensi-

[1] These estimates are not statistically distinguishable from one another (see Section 4).

tivity analyses. Studying employment transitions in and out of the healthcare industry, we find a large increase in outflows that is not offset by a small increase in inflows. An interpretation of the increased job churning is that quitting workers who are vaccine-averse are partially replaced by new hires who are either pro-vaccine or neutral on the issue. We also find that the increased outflow rate resulting from the mandate is much higher for workers in healthcare-specific occupations than in other occupations. Despite this gap in apparent quitting across these two types of occupations, the hiring rates are similar, which may reflect higher costs of replacing the workers with healthcare-specific skills compared to non-healthcare workers. In sum, our results suggest that efforts to increase vaccination rates through employment-based mandates in specific industries may have had the unintended consequence of exacerbating the ongoing workforce shortages in many healthcare markets across the U.S.

The paper is organized as follows. Section 2 discusses background on vaccines and the U.S. healthcare workforce, and outlines our contributions. Data, variables, and methods are reported in Section 3, while the main results and robustness checking and extensions are listed in Sections 4 and 5 respectively. Section 6 offers a discussion and conclusion.

2 Background and contribution

2.1 Background

Vaccine mandates are not a new phenomenon and have been used as a policy tool to control the spread of infectious diseases since the 19th century in the U.S. and other countries. In 1855, the state of Massachusetts mandated smallpox vaccinations for children before going to school to prevent the spread of this disease. Since then, several other states have mandated vaccination against different diseases for school-aged children, with all states requiring some vaccination for school-entry at the time of writing (National Council of State Legislatures, 2023). In 1902, due to a smallpox outbreak in Massachusetts, the state required a smallpox vaccine for all residents.[2] Smallpox vaccine requirements began much earlier in the U.S. Army during the American Revolutionary War in 1775. Since then, different vaccines, including but not limited to influenza, cholera, and hepatitis A, have been administered in the army depending on the type of infectious disease soldiers might be exposed to in each battleground (Grabenstein et al., 2006). In the case of the COVID-19 pandemic, arguably the worst pandemic in recorded history which has led to 6.8 million hospitalizations and 1.2 million deaths in the U.S.

[2] Please see Chervinsky (2021).

(COVID Data Tracker, 2024),[3] vaccines with over 85% effectiveness were developed in less than one year after the beginning of the pandemic in 2020. These vaccines substantially reduce the likelihood of becoming infected with COVID-19, and the probability of severe disease and death among those infected. Governments in the U.S. used mandates for specific groups to increase vaccination rates and reduce the disease burden.

Healthcare workers were among the initial groups eligible to receive the COVID-19 vaccine in the U.S. when the vaccine was released in December 2020.[4] However, due to vaccine hesitancy, which has historical roots dating back to the pre-COVID-19 era (Paterson et al., 2016), a substantial portion of healthcare workers were not comfortable receiving COVID-19 vaccines because of their high perceived risk or low perceived effectiveness of the vaccines (Gagneux-Brunon et al., 2021; Kwok et al., 2021; Gu et al., 2022). One study based on nearly 3,500 healthcare workers from several healthcare systems in five U.S. states over the period October to November 2020 suggests that 8% of respondents did not have any plans to receive a COVID-19 vaccine (Shekhar et al., 2021). Other studies in the U.S. report hesitancy rates between 8 and 16.4% (Biswas et al., 2021). Immunization is important for healthcare workers for at least two reasons. First, remaining unvaccinated while working in a healthcare facility could increase the risk of patients becoming infected with the virus. Second, working unvaccinated in the healthcare setting could increase the risk of infection among the workers (unvaccinated and vaccinated), which could exacerbate the healthcare worker shortage.

COVID-19 vaccine mandates have proven to be an effective policy for increasing vaccine take-up in the general population outside the U.S. Cross-country studies find that these mandates increase vaccination rates in Canada, Germany, France, and Italy (Karaivanov et al., 2022; Fitzpatrick et al., 2023). COVID-19 vaccine mandates in the U.S. also show a positive impact on the general population's vaccination rates, with effects appearing as early as three weeks following the mandates' implementation dates (Howard-Williams et al., 2022; Okpani et al., 2024). McGarry et al. (2022) and Reses et al. (2023) study the effect of these mandates in 15 states using the National Healthcare Safety Network and suggest that these mandates increase the COVID-19 vaccination rate among healthcare workers in nursing home facilities, but with no observed impact on staff shortages. Earlier studies also show that state influenza vaccination requirements for hospital workers 1995-2017 increased the vaccination rate among healthcare workers by 4.2% (Lindley et al., 2019), and reduced pneumonia and influenza mortality rates by

[3] These numbers are current as of February 20, 2024.

[4] Military personnel were the first professional group in the U.S. mandated to receive COVID-19 vaccines after the Food & Drug Administration approved the Pfizer vaccine (Elliott and Chambers, 2022).

2.5% (Carrera et al., 2021).

Although there are no unified state-representative databases tracking the vaccination rates among healthcare workers in the U.S. covering our analysis period,[5] some states report these statistics for a limited time period in which many state-level healthcare worker COVID-19 vaccination mandates took effect. For example, the state of Maine reports monthly statistics for May-October 2021.[6] Using a pre-post design, these data suggest a 24% increase in COVID-19 vaccination rates among healthcare workers in October 2021, one month after the mandate became effective in September 2021, compared to August 2021, showing an increase from 78% to 97%.

The healthcare industry labor shortage is a multifaceted problem including heightened demand for healthcare services in the baby boomer generation and increases in longevity due to advances in medical care. Based on pre-COVID-19 research in the U.S., healthcare workers' burnout during COVID-19 pandemic likely exacerbated ongoing healthcare worker shortages (Willard-Grace et al., 2019) and increased medical errors (Shanafelt et al., 2010). Staff shortages may delay admitting patients to healthcare facilities and prevent patients from receiving quality care due to the increased medical errors and staff burnout. A report by the Joint Commission showed that sentinel events[7] rose by 20% in 2022 compared to 2021 (The Joint Commission, 2023).[8] McGarry et al. (2020) report that 20% of nursing home facilities experienced staffing and protective equipment shortages in the first few months of the COVID-19 pandemic in 2020. However, a survey of facilities in four federal healthcare programs reveals that these facilities were understaffed prior to the pandemic, with the pandemic exacerbating, rather than originating, this staffing problem (Pandemic Response Accountability Committee, 2023).

Recent healthcare industry reports show that more healthcare providers are leaving the industry than in earlier time periods: over 145,000 healthcare providers left the industry 2021-2022 (Popowitz, 2023). Half of the industry leavers were physicians, 25% were nurses, and the remainder were physician assistants, physical therapists, and licensed clinical social workers. Academic research also suggests that the supply of nurses plummeted by 100,000 only in 2021 (McGarry et al., 2020; Auerbach et al., 2022).

[5] The National Healthcare Safety Network data is one exception. However, to the best of our knowledge these data are available only to state Health Departments.

[6] Authors' analysis of data reported by the state of Maine.

[7] Sentinel event is defined as an event that results in patient's death, permanent harm, or severe harm.

[8] The Joint Commission is a non-profit organization in the U.S. that accredits hospitals and healthcare systems.

2.2 Contribution

Our work connects to at least three strands of economic literature. First, we study the impacts of the policies designed in the wake of the pandemic on the U.S. labor market. A multitude of studies examine the direct impact of the pandemic on economic outcomes, generally suggesting strong and negative effects (Alon et al., 2020; Polyakova et al., 2020; Brodeur et al., 2021; Lee et al., 2021; Montenovo et al., 2022). Other studies have investigated state and federal social-distancing policies (e.g., school closures, shelter-in-place mandates, and mask mandates) adopted during the pandemic prior to development of pharmaceutical interventions when the primary public health response relied on behavioral changes (Courtemanche et al., 2020; Lyu and Wehby, 2020; Abouk and Heydari, 2021), again generally finding that these policies negatively impacted labor markets. Our work examines a policy implemented after the height of the pandemic in 2020 and targeting the use of effective vaccines to curb disease spread, and also documents potentially negative impacts on labor markets.

Second, we add to the literature on potential distortions of public health policies on labor markets. For example, a long-standing question in labor economics is the extent to which the system of employer-sponsored health insurance leads to 'job-lock' in the U.S. labor market (Gruber and Madrian, 1994; Garthwaite et al., 2014; Maclean and Webber, 2022), or the extent to which mandating health insurance may reduce employment or wages (Summers, 1989).[9] Further, due to earnings caps, and social insurance programs (e.g., Supplemental Security Income and Medicaid), some recipients may reduce their labor supply to remain eligible (Yelowitz, 1995; Neumark and Powers, 2005).

Third, we contribute to the literature that examines factors influencing the size and composition of the healthcare workforce. Studies have examined the impact of reimbursement, job strain, economic conditions, insurance design, and so forth (Chen et al., 2018; Dillender et al., 2021; Forsythe et al., 2020; Buerhaus et al., 2022; Cantor et al., 2022; Cortes and Forsythe, 2023; Barnes et al., 2023; Shen et al., 2024). We focus on state actions that compel employees to receive a pharmaceutical intervention that healthcare workers may have reservations against.

[9] A unique feature of the U.S. labor market is the tight link between employment and insurance. Even after the Affordable Care Act (ACA) created options for health insurance outside employment, the Act also mandated that employers provide insurance, and the vast majority of insured working age adults receive coverage from their employer (Keisler-Stankey et al., 2023). This system creates the potential that employees will remain in jobs they would otherwise leave to retain health insurance coverage.

3 Data and methods

To study the impact of COVID-19 vaccine mandates on healthcare industry employment, we draw data from two primary sources. First, we use the monthly Current Population Survey (CPS) to construct a sample of individual workers, including those working in healthcare. Second, we utilize policy data on healthcare worker vaccination mandates from the Centers for Disease Control and Prevention (CDC). Here we describe variable construction and samples from these data sources.

3.1 Current Population Survey

We draw data on workers from the monthly CPS between January 2021 and December 2022, available through the University of Minnesota's IPUMS system (Flood et al., 2023). We begin the study period in 2021 as, prior to this year, COVID-19 vaccines were not readily available to the general American public. Indeed, the Food & Drug Administration (FDA) approved the first vaccines under emergency use authorization on December 11th, 2020 for most people 16 years and older (Food & Drug Administration, 2021). In 2021 the general population of adults, in large numbers, began to receive the vaccine (Kates et al., 2022). We close the study period in December 2022 as the COVID-19 pandemic wound down and many states repealed their vaccine mandates in late 2022 or soon thereafter (Howard-Williams et al., 2022).

The CPS reference week is the week that includes the 12th day of the month (U.S. Census Bureau, 2022). The exception to this rule is December, for which the reference week is the week that includes the 5th day of the month, if that week is fully in December. We match mandates to the CPS data based on reference week, that is we require that the reference week proceed the effective date for the mandate. Of the approximately 150,000 observations in each monthly sample, for most of the analysis we retain all civilians aged 21-64 who report being employed at the survey. We use the full civilian population in that age range when we analyze employment and labor force participation in Section 4.1.

To identify workers in the healthcase industry, we use the IPUMS variable *IND*. Table A1 lists the specific healthcare industries and associated IPUMS codes included in our analysis. We construct an indicator variable coded as one if the respondent reports working in the healthcare industry and zero otherwise.

To estimate if the impact on healthcare industry employment varies by occupation, we identify (non-) healthcare occupations using the IPUMS variable *OCC* among workers within the healthcare industry. Healthcare occupation (HCO) is defined to equal one for healthcare practitioners and technical occupations and healthcare support occupations as listed in Table A2, and equals zero otherwise. Non-healthcare occupation ($nHCO$) is defined as workers in other occupations within the healthcare industry.

The CPS includes basic demographic information for respondents that we include in our regressions to reduce residual variation and increase precision, though as we show in Section 4 our results are not sensitive to including or excluding these variables. We select the following demographics: age (21-29 years, 30-39 years, 40-49 years, 50-59 years, and 60-64 years, with 21-29 years as the omitted category), sex (male and female, with male as the omitted category), race (White, Black, and other, with White as the omitted category), ethnicity (Hispanic and non-Hispanic, with non-Hispanic as the omitted category), born outside the U.S., and education (less than high school, high school, some college, and college or higher, with less than high school as the omitted).

In Table A3, we report demographics of CPS respondents stratified by healthcare industry, here we include both the employed and those not employed as the objective of this exercise is to compare individuals across groups in terms of relevant variables. We categorize respondents as i) non-healthcare industry, ii) healthcare industry, iii) healthcare industry and non-healthcare occupation, and iv) healthcare industry and healthcare occupation. In comparing those in the healthcare industry and not, those working in the healthcare industry are more likely to be female (77% vs. 48%) and are more likely to hold a college degree (50% vs. 37%). Other demographics, while not identical, are more similar across the two groups. When we strategy respondents in the healthcare industry by healthcare occupation, we that the two groups are largely similar across the outcomes we consider, though those working in healthcare occupations have somewhat higher educational attainment, are more likely to be born outside of the U.S., and are more likely to work part-time.

A limitation of using CPS data, and all surveys of which we are aware, to study the COVID-19 period is that response rates were lower and non-responders may be a non-random sample of all CPS respondents. Previous research has documented this issue with the CPS (Heffetz and Reeves, 2021; Ward and Edwards, 2021), though much of this work focuses on 2020, the height of the pandemic, and we study years 2021 and 2022.

3.2 Vaccination mandate data

We use policy data compiled by the CDC to measure mandates that compel workers to have COVID-19 vaccines as a requirement for employment in the healthcare sector (Howard-Williams, 2022). The CDC data are available through the end of September 2022. To supplement these data, we conducted a review of state health departments and news media to identify mandate expirations from September 2022, the end of the time period covered by the CDC database (Howard-Williams, 2022), to December 2022, the final month included in our study period. Full details on our search to locate effective dates in Q4 2022 are available on request from the corresponding author. However, as we show in Section 5, our results are robust to excluding Q4 2022 from the sample.

We construct an indicator coded one if a state mandates that employees in the healthcare industry must be fully vaccinated. Fully vaccinated is defined by the mandates during this period as having two COVID-19 shots from Moderna or Pfizer, or one shot from Johnson & Johnson. Most mandates allow for some exemptions (e.g., health or religious exemptions); workers who do not receive the vaccination must meet other requirements including regular COVID-19 testing and wearing multiple masks at all times. We do not separate out mandates that do and do not allow for exemptions as both types of mandates create the benefits of a safer workplace and impose costs on vaccine-averse employees, and thus affect the utility of working in the healthcare sector.

We include healthcare worker and long-term care worker mandates in our analysis because the long-term care industries are a subset of overall healthcare, and the two mandates are highly correlated. Only two states (Delaware and Massachusetts) have a long-term care mandate, but not a healthcare worker mandate. As we show in Section 4, our results are very similar if we use only the healthcare worker mandates (i.e., we code states with only a long-term care mandate as a part of the comparison group). We also conduct a 'leave-one-out' analysis in Section 5 in which we sequentially drop each treated state and find that results are very similar across the leave-one-out samples. When a state has both a healthcare and long-term care mandate, and the effective or expiration dates differ, we use the earlier date for the effective date and the later date for the expiration. For brevity, we refer to both types of mandates as 'healthcare' except when explicitly studying only healthcare industry mandates.

These mandates - implemented by 17 states - were adopted between September and November 2021, hence we have very little staggering in mandate adoption. See Figure 1 for the geographic distribution of healthcare industry COVID-19 mandates, with the effective date for each adopting state listed in the figure notes. In Figure 2, we show the timing of mandate adoption. Some states let their mandates expire between March 2022 and November 2022. We incorporate expirations into our coding of the mandates, though we will show in Section 5 that our results are robust to ignoring expirations, and treating these mandates as an 'absorbing state' variable. Some mandates include specific language that non-compliant workers will face adverse employment penalties (e.g., termination). Figure 3 shows which mandates have such language, and their expiration dates. We test the impact of these 'strong' mandates that impose specific employment penalties on non-compliant workers in robustness checking (Section 5).

An important question is the extent to which mandates are enforced. To the best of our knowledge, there are no data on how well states enforced their mandates, and thus we cannot study this aspect of the mandates. However, we will use insight from our event-study findings, in particular the dynamics we observe (reported in Section 4), to provide suggestive evidence on employer compliance.

The CDC data do not include the exact jobs impacted by the mandates, thus we include all employees of healthcare industries in our analysis. A potential limitation of the CDC policy data is that they do not include sub-state (e.g., city or county) vaccine mandates, and thus we do not incorporate this information. The geography of relevance for these mandates is the location of employment, but the CPS data include the location of residence. In analysis of the 2019 (i.e., prior to the COVID-19 pandemic) American Community Survey available through IPUMSUSA (Ruggles et al., 2023), we find that 94% of employed adults 21 to 64 years of age live and work in the same state, but just 77% live and work in the same county.[10] These numbers suggest that measurement error will be lessened by focusing on state (vs. sub-state) mandates. Additionally, the CPS suppresses sub-state geographic information (e.g., 30% of respondents do not have county of residence information), furthering our concerns regarding measurement error involved in studying sub-state mandates.

We control for vaccine mandates targeting other groups of workers in regressions: congregate facility (e.g., jails or prisons, we refer to these as 'jail'), government, and schools.[11] The rationale for including these mandates is that they may impact a worker's next-best option. For example, a worker contemplating entering (or exiting) the health-care industry following adoption of a vaccine mandate in this industry may consider the extent to which other feasible jobs also require vaccination. However, as we will show in Section 4, our results are not sensitive to excluding these additional mandates.

3.3 Methods

We estimate the difference-in-differences (DID) regression in Equation 1:

$$H_{i,s,t} = \beta_0 + \beta_1 Hvaccine_{s,t-1} + Ovacine_{s,t}\beta_2 + \beta_3 Pop_{s,y(t)} + \beta_4 X_i + \alpha_s + \alpha_t + \epsilon_{i,s,t} \quad (1)$$

[10] The CPS does not include this information, hence we use the American Community Survey.

[11] One state, Hawaii, mandated vaccination among private workers, but we do not include this mandate given the limited variation.

where $H_{i,s,t}$ represents an indicator for working in the healthcare industry for respondent i in state s in time (month-year period) t. $Hvaccine_{s,t-1}$ is an indicator variable taking on a value of one if the state has a healthcare vaccine mandate in place and zero otherwise. We lag this variable by one month as the time required to become fully vaccinated was approximately one month for most vaccines available in 2021 and 2022 in the U.S.[12] $Ovacine_{s,t}$ is a vector of other non-healthcare industry COVID-19 vaccinations (jail, government, and school workers). $Pop_{s,y(t)}$ is the annual (this variable does not vary at the month-year level unlike other variables in the analysis) population using data on from the U.S. Census Bureau (University of Kentucky Center for Poverty Research, 2023).[13] X_i is vector of individual-level variables (age, sex, race, ethnicity, and education). α_s is a vector of state fixed effects and α_t is a vector of time fixed effects (i.e., separate indicators for each month-year pair in our data). $\epsilon_{i,s,t}$ is the error term.

Data are weighted by survey weights provided by the CPS. We cluster standard errors at the level of the state. We estimate OLS in our main specifications. As we show in Section 5, our results are not sensitive to numerous alternative samples and specifications. Our methods for analyzing transition of workers into and out of the healthcare industry are similar, and described in Subsection 4.4.

3.4 Trends and summary statistics

Figure 4 reports trends in the probability of working in the healthcare industry for states that do and do not adopt mandates requiring healthcare workers to be vaccinated for COVID-19. Data are aggregated to the time (month-year) level. While the monthly data are somewhat noisy, prior to the adoption of the initial mandates (September 2021), the two groups of states trend similarly in terms of the share of workers in the healthcare industry. After September 2021, there is a divergence between the two groups. States adopting a mandate experience a sharp decline in the share of workers in the healthcare industry, but this decline is not apparent in other states. In mid-2022, after many states allowed their mandates to expire, the two trend lines reconverge. While these trends do not adjust for myriad factors that were changing over this time period, Figure 4 offers suggestive evidence that mandates requiring employees in the healthcare industry to be

[12] Both Pfizer and Moderna required four weeks between the first and second shot. As described above, full vaccination required two doses of Pfizer and Moderna, and one dose for Johnson & Johnson. The vast majority of Americans received Pfizer and Moderna vaccinations (Our World in Data, 2023). For example, as of June 1st, 2021 of the 296.81 million COVID-19 vaccinations administered in the U.S., 160.31 million where Pfizer and 125.02 million were Moderna, reflecting 96.1% of all vaccines administered (Our World in Data, 2023).

[13] We use U.S. Census Bureau Population data maintained by the University of Kentucky Poverty Research Center.

vaccinated for COVID-19 may have deterred some people from working in this industry.

We report summary statistics for states that adopt a vaccine mandate (measured prior to mandate adoption) and for states that do not (measured over the full sample period) in Table 1. 12% of employed people work in the healthcare industry in both samples. In terms of respondent characteristics, the two groups are quite similar for most variables included in our analysis. However, states that adopt mandates are more racially and ethnically diverse than other states; we suspect that these differences are driven by states such as California, New Mexico, and New York adopting healthcare worker COVID-19 mandates and being more diverse than the average U.S. state (Kaiser Family Foundation, ND). We also observe that workers in mandate adopting states tend to have higher education and are more likely to have been born outside the U.S. than people in non-adopting states. For example, the share of workers with a college degree or higher is 47% in adopting states and just 39% in non-adopting states.

The prevalence of vaccine mandates for government and school workers is greater in states that adopt such mandates for healthcare workers. However, the reverse is true for those people who work in jails. States that adopt healthcare worker COVID-19 vaccine mandates are larger in terms of population than states that do not. As we show in Section 4, our results are robust to controlling for these mandates.

4 Results

4.1 Probability of working in the healthcare industry

Table 2 reports the main results of our study. We 'build up' the regression specification by increasing adding control variables, this approach allows us to observe the extent to which our findings change with different sets of controls. Specification (1) includes only state and period (month-year) fixed effects while specifications (2) and (3) add in respondent and state characteristics respectively. In specification (4), we include the same covariates in specification (3) but define a mandate for healthcare workers only (that is, we do not code the states with only long-term care worker mandates as having a healthcare worker mandate, this change results in just two states - Delaware and Massachusetts - be recoded as part of the comparison group).

Reassuringly, our findings are stable across the four specifications. In particular, we find that post-mandate, the probability that a worker is employed in the healthcare sector declines by 0.49 to 0.72 percentage points ('ppts'). In our preferred specification - which includes state and period fixed effects, respondent characteristics, and state characteristics - we find that the probability of working in the healthcare sector among employees declines by 0.72 ppts. Comparing that with the pre-treatment mean in mandate-adopting states implies a decline of 5.8% (= -0.0072/0.1232 $*$ 100%).

17 states adopted healthcare worker COVID-19 vaccine mandates, which potentially creates a small treated cluster setting. Given this number of treated states, we report the t-statistic generated by testing the null of no treatment effect using a score bootstrap approach that has been shown to have better properties in settings with a small number of treated clusters than the state-level clustering we use in our main analyses (Brewer et al., 2017; Kline and Santos, 2012; Roodman et al., 2019). The t-statistic for our preferred specification - column (3) in Table 2 - is -2.8775 (p=0.0040), and thus the results are not appreciably different from those obtained with state-level clustering.[14]

4.2 Identification

<u>Parallel trends</u>: The primary assumption of our DID approach is 'parallel trends:' had adopting states not implemented COVID-19 vaccine mandates for healthcare workers, then adopting and non-adopting states would have followed the same trends in the propensity to work in the healthcare industry in the post-period. This assumption is untestable as we cannot observe counterfactual outcomes for treated states. To provide suggestive evidence on the probability that our data satisfy parallel trends, we estimate an event-study. In particular, we decompose our static DID parameter estimate ($Post \times Treat$) into a series of single event-time indicators from eight months pre-mandate to eight months post-event (more specifically, we interact the treatment indicator with our series of time-to-event indicators). The omitted event-time period is the month prior

[14] In unreported analyses, we have examined the impact of COVID-19 healthcare worker mandates on vaccination rates. In particular, we use data from the CDC on overall vaccination rates among those 18 years and older (Centers for Disease Control and Prevention, 2023) per state-time (month-year) period. CDC works with clinics, retail pharmacies, long-term care facilities, dialysis centers, and other sites across the country to collect vaccine data. We convert vaccination counts to the rate per 100,000 state residents 18+ years using population data from the U.S. Census Bureau maintained by the University of Kentucky Center for Poverty Research (2023) and the share of the population 18+ years that we calcuated from the CPS data. If the mandates we study are inducing some workers to become vaccinated, then we would expect vaccination rates should increase post-mandate. On the other hand, if workers elect to leave (or not enter) the healthcare sector following an mandate to avoid vaccination, then we should observe no change in vaccination rates. Our findings are in line with the latter hypothesis; we do not observe evidence of an increase in vaccination rates post-mandate. We interpret these findings as suggestive only and we choose not to report these results as we suspect that we are under-powered to detect effects as healthcare workers make up a small share of the overall workforce, just 12%, and our vaccination rate data is for the population (working and not working) 18 years and older. Ideally, we would have data measured before and after mandate adoption for all states for the healthcare industry specifically, but to the best of our knowledge such data are not available. Results are available on request from the corresponding author.

to mandate adoption. We code non-mandate adopting states as zero for all event-time indicators. Mandate expirations are ignored in the event-study. We trim the data in event-time for mandate adopting states, that is we exclude time periods more than eight months pre- and post-mandate from the event-study analysis sample.[15] Otherwise, the event-study specification is identical to that outlined in Equation 1.

We report event-study results in Figure 5. Coefficient estimates on the policy lead variables are small and statistically indistinguishable from zero, suggesting that adopting and non-adopting states followed similar trends in our outcome prior to mandate adoption. This pattern of results offers suggestive evidence that our data satisfy the parallel trends assumption, so that the two groups would have followed similar trends in healthcare worker employment post-treatment had the mandate states not been treated. Examination of the policy lags suggests that effects emerge the month of mandate adoption and increase over the second and third months post-mandate, stabilizing after that point and remaining constant through the end of the eight-month study period.

In Figure 6, we report results from alternative event-study specifications and samples. The objective of this exercise is to ensure that our event-study findings are not driven by a specific sample or specification. Reassuringly, results are not sensitive to the different specifications and samples. First, we exclude time-varying covariates, including only state and period fixed effects in the regression. Second, we omit the period -8 (rather than -1); recall that we trim the data in event-time, therefore the -8 indicator is homogeneous in event-time and includes only observations observed eight months prior to mandate adoption. Third, we do not trim the data in event time, thus the most distal leads and lags are heterogeneous in terms of time-to-event.[16] Fourth, we drop periods after mandates expire. Fifth, we exclude the fourth quarter of 2022 as our main policy database only includes mandate effective and expiration dates through September 2022 (though we reviewed health department websites and other sources to locate changes occurring in the final quarter of 2022; see Section 3.2). Finally, we include as many leads and lags as permitted by our data.

While only suggestive, our event-study results are in line with the hypothesis that

[15] We choose the eight-month event window for two primary reasons. First, we observe all adopting states in each event-time period, and thus we are not concerned about changes in the composition of states used to identify time-to-event coefficient estimates. Second, in terms of the lags, most states have their mandate in place eight months post-mandate, thus by trimming the event-study sample in this manner we are not concerned about mandates that expire; as noted earlier in this section, we ignore expirations in the main event-study, though we will show robustness to various alternatives samples and specifications, all of which lead to the same conclusion.

[16] The -8 lead includes period ten to eight months pre-mandate while the +8 lag includes periods eight to 15 months post-mandate.

employers enforce mandates. As noted in Section 3, there is no information on enforcement of vaccine mandates available at the national level to the best of our knowledge. However, we show (see Figure 5) that COVID-19 vaccine mandate effects are observed immediately post-policy, grow quickly, and stabilize after two months. This pattern of results is what we would expect if employers enforce mandates and employees who do not wish to receive the COVID-19 vaccine quit within the first two months of the mandate in response, with the remaining employees being willing to accept the vaccine. In Figure 6, where we include as many lags as permitted by our data, mandate effects appear to dissipate over time, which aligns with employers gradually hiring to replace vaccine-averse workers who quit in response to the mandate. Of course, the more distant lags are more complicated to interpret than those reported in Figure 5 due to changes in the mandates and the composition of states used to identify each lag coefficient estimate.

Staggered treatment adoption: Recent econometric work suggests that using TWFE regressions with a staggered policy roll-out can lead to bias from heterogeneous treatment effects (Goodman-Bacon, 2021; Borusyak et al., 2021). TWFE regressions can be viewed as an weighted average of all possible two-by-two DID contrasts in the data. In particular, with a staggered policy roll-out and dynamics in treatment effects (e.g., COVID-19 vaccine mandates may increase or decrease over time), some of the comparisons will be reasonable (i.e., comparing units that take treatment to untreated units) while others will be forbidden (i.e., comparing units that take treatment to those that were tread earlier in the study period). These forbidden comparisons may lead to bias in regression coefficients and, in some cases, sign reversals. As shown in Figure 2, there is limited staggering in when states adopt their healthcare mandate. However, some states allow their mandates to expire by the end of our study period (December 2022). Thus, our setting is more complex and our treatment variable is not always an 'absorbing state.' This treatment regime creates a comparison in Equation (1) between states that turn off treatment to i) untreated states that do not adopt a mandate by December 2022 and ii) treated states observed during treatment.[17]

Given that some of the comparisons in our data may lead to bias when using DID with Equation (1), we next use an estimator that can accommodate the non-absorbing and staggered treatment regime that state healthcare worker vaccine mandates follow (de Chaisemartin and d'Haultfoeuille, 2020).[18]

[17] In comparisons using the former group of states, we must assume that untreated states can offer a valid counterfactual for states that allow their mandates to expire, had these states retained their mandate. In comparisons using the latter group of states, we assume that treated states serve as a valid counterfactual for states allowing their mandate to expire.

[18] If there is heterogeneity across treated units, even absent dynamics in treatment effects, TWFE

We report results based on the de Chaisemartin and d'Haultfoeuille (2020) procedure in specification (5) of Table 2. We find that the probability of working in the healthcare industry declines by 0.71 ppts or 5.8%. Reassuringly, the effect size is nearly identical to that estimated by our TWFE regression (0.72 ppts or 5.8% in our preferred specification).

In Figure 7, we present an event-study using the de Chaisemartin and d'Haultfoeuille (2020) estimator. We again observe no evidence of differential pre-trends. While the post-treatment coefficient estimates follow the same pattern as our TWFE regression event-study, we lose some precision. The loss of precision is not unexpected as de Chaisemartin and d'Haultfoeuille (2020) note their estimator is less efficient than TWFE.[19]

Balance: Next, we examine the extent to which healthcare worker COVID-19 vaccine mandates are associated with control variables included in Equation 1. To this end, we aggregate the CPS data to the state-period level, we apply CPS-provided survey weights in the aggregation of the data. We then regress each of the (aggregated) included characteristics on the mandate (lagged one month), state fixed effects, and month fixed effects. Standard errors are clustered by state.

We report results graphically in Figure 8. The findings largely mirror summary statistics reported in Table 1. Healthcare worker vaccine mandates are largely uncorrelated, after adjusting for state and time fixed effects, with the time-varying variables we include in our regressions. However, we do find evidence that these mandates are correlated with COVID-19 vaccine mandates for both government and school workers. While ideally we would observe no correlation between the healthcare worker vaccine mandates and any controls, the fact that our findings are not contingent on any specific control variable (see Table 2) is re-assuring.

Sample selection: We focus on a sample of employed people in our main analysis as this is the group that is arguably most directly impacted by healthcare worker COVID-19 vaccine mandates. A possible concern is that the mandates may impact the choice to participate in the labor market and work post-mandate. To examine the importance of such changes, we regress the probability of i) participating in the labor force, ii) employed (vs. not employed and not in the labor force), iii) full-time employment (vs.

may also recover a poor estimate of the ATT. TWFE regression 'variance weights' observations. All else equal, states whose treatment 'turns on' in the middle of the panel will be up-weighted relative to other states. However, as described earlier, the effective date for all states with a mandate falls within a three month period, leaving limited scope for differential weighting to lead to a poor estimate of the ATT when using TWFE regressions. However, because some states allow their mandates to expire, our context is more complex than absorbing state policies.

[19]The de Chaisemartin and d'Haultfoeuille (2020) event-study does not include an omitted period. Thus, we estimate a treatment effect for -1, the month prior to treatment, which is omitted in the TWFE event-study.

part-time employment, not employed, and not in the labor force), and iv) part-time employment (vs. full-time employment, not employed, and not in the labor force) among CPS respondents 21-64 years of age on healthcare worker COVID-19 mandates and other controls using Equation 1. We report results in Table 3.

The coefficient estimates are generally small in magnitude and not statistically different from zero in these regressions. One exception to this pattern of null findings is that the probability of employment post-mandate increases modestly by 0.5 ppts or 0.7%. This finding suggests that healthcare worker COVID-19 vaccine mandates could potentially draw some workers into employment, which may lead to some changes in the composition of the employed sample. However, in Figure A1, we report event-studies for all four employment outcomes that we consider in Table 3. Increases in employment propensity do not emerge until approximately seven months post-mandate, while our healthcare worker findings appear immediately following mandate adoption - we observe a decrease in the probability of working in this industry in the month of mandate that roughly doubles in size by month two, and remains relatively stable through month eight (see Figure 5). Thus, given the differences in the timing of effects, any mandate-induced changes in employment (which we find to be very small in magnitude) cannot explain our main results for the probability of working in the healthcare sector.

4.3 Heterogeneity in mandate effects

In our main analyses, we consider healthcare worker COVID-19 vaccine mandate effects across all workers, thus capturing the average effect across all workers. The impact of the mandates potentially differ based on the worker preferences, the extent to which the state policy environment promotes (or perhaps hinders) the ability to obtain a vaccine, and sector of the healthcare market. In this section, we explicitly examine heterogeneity in healthcare worker COVID-19 vaccine mandate effects across state policies, respondent characteristics, and healthcare industry sector. This analysis allows us to shed light on the distributional impacts of the COVID-19 vaccine mandates or, put differently, policy equity. While we do observe some differences in effect sizes across the groups we consider, coefficient estimates are negative in all cases (with the exception of our analysis of specific healthcare sectors) and 95% confidence intervals generally overlap, which prevents us from drawing strong conclusions regarding heterogeneity in mandate effects.

In Table 4, we explore the importance of two state-level policies that may support workers to become vaccinated post-mandate: paid sick leave (National Partnership for Women & Families, 2023) and Medicaid insurance expansions to non-traditional populations that occurred in conjunction with the Affordable Care Act (Kaiser Family Foundation, 2023b). Paid sick leave offers financially protected time for healthcare as workers do not have to forego wages to receive care available during normal work hours. Insurance can reduce the out-of-pocket cost of healthcare services, including vaccinations. Indeed, previous economic studies show that both paid sick leave (Pichler and Ziebarth, 2017; Pichler et al., 2021; Maclean et al., 2020; Andersen et al., 2023) and ACA Medicaid[20] expansion (Tummalapalli and Keyhani, 2020) increases vaccination rates, though these studies pre-date the COVID-19 pandemic and generally focus on the influenza vaccine - with Pichler et al. (2021) and Andersen et al. (2023) as exceptions that examine the COVID-19 pandemic period. While we lose precision, we find similar effect sizes among respondents in states that do and do not adopt paid sick leave policies. Findings for COVID-19 healthcare worker vaccine mandates are stronger in states that expand Medicaid with the ACA, which suggests that insurance facilitates use of vaccines.

In Table 5, we examine heterogeneity in mandate effects across respondent characteristics: sex (male and female), race (White and non-White), ethnicity (Hispanic and non-Hispanic), education (college degree and less than a college degree), residence (rural and non-rural),[21] and healthcare vs. non-healthcare occupations (see Section 3.1). Mandate effects appear to be stronger among female, White, non-Hispanic, and rural workers, and for healthcare rather than non-healthcare occupations.

The fact that findings are stronger for those working in healthcare occupations, which are occupations for which we expect the highest specific training and human capital for healthcare work, is potentially concerning for patients and healthcare systems as these jobs may be most closely linked to the provision of healthcare. We observe that, following a COVID-19 vaccine mandate, the probability of working in $nHCO$ occupation declines by 0.21 ppts (5%) while the probability of working in a HCO occupation declines by 0.51 ppts (7%), but only the HCO occupation is statistically distinguishable from zero.

We have boot-strapped the differences in coefficient estimates across samples reported in Tables 4 and 5 using a non-parametric bootstrap (1,000s repetitions). The only difference that rises to the level of statistical significance is that for men and women.[22]

While there are shortages across the healthcare industry, several specific industries face elevated challenges, due to pre-existing shortages and increased need for services

[20] One concern with studying ACA Medicaid expansion heterogeneity is the extent to which enrollees work, but (Guth et al., 2023), for example, reports that in 2021, 63% of non-elderly adult Medicaid enrollees worked. The ACA Medicaid expansions largely impacted non-elderly and non-disabled childless adults in terms of those most likely to gain coverage.

[21] We use the IPUMS harmonized variable *METRO* (Ruggles et al., 2023).

[22] We have also used nested models to explore statistical differences. Results, which are qualitatively the same, are available on request.

during the COVID-19 pandemic. In particular, there are severe shortages of mental healthcare workers, with 77% of U.S. counties experiencing a shortage (Thomas et al., 2009) and the Kaiser Family Foundation reports that in 2023, just 27.2% of the U.S. mental healthcare needs are met (Kaiser Family Foundation, 2023a). During the COVID-19 pandemic, mental health conditions sky-rocketed (Cullen et al., 2020), increasing demand for services. Given these factors, we separately examine the impact of employment vaccine mandates on the probability of working in the long-term care industry,[23] working as a registered nurse,[24] and working in a mental healthcare occupation.[25] Table 6 reports results. Only the coefficient estimate for work as a registered nurse is statistically distinguishable from zero. Post-mandate, the probability decreases by 0.24 ppts (6.9%).

4.4 Healthcare job transitions

The effect of vaccine mandates on healthcare industry employment could occur through changes in the rate at which workers exit the industry, enter, or both. Our finding that vaccination mandates lead to decreases in healthcare industry employment imply that exits must increase more than entries. To the extent that both exit and entry increase under the mandate, the higher employee turnover may be associated with exit by those averse to vaccination and entry by those who either find vaccination attractive or at least are not averse to it. The difference between healthcare occupations and non-healthcare occupations within the healthcare industry is also of interest here, because it is likely much harder for the industry to replace workers who have industry-specific skills.

We construct $Exit_t$ and $Enter_t$ variables from the longitudinal version of the CPS, based on the sample rotation scheme of four months in sample, eight months out, and four more months in, as follows. $Exit_t$ is defined for workers in the healthcare industry at time $t-1$; this variable equals one if the worker is no longer in the healthcare industry at t (including non-employment), and equals zero if they remain employed in the healthcare industry. $Enter_t$ is defined for workers in the healthcare industry at time t; this variable equals one if the worker is not in the healthcare industry at $t-1$ (including non-employed), and equals zero if they remain employed in the healthcare industry.[26]

[23] We use the following industry codes from the IPUMS variable *IND*: 8270 and 8290.

[24] We use the following occupation code from the IPUMS variable *OCC*: 3603.

[25] Industry codes available in the CPS do not allow us to separately identify the mental healthcare industry, thus we use the following occupation codes from the IPUMS variable *OCC*: 2001, 2004, 2013, 3422, and 3605.

[26] We identify possibly false match of IPUMS-provided respondent identifiers using information on race, ethnicity, age, and sex and drop about 1% of the original transition sample based on apparent measurement errors: inconsistencies in race, ethnicity, or sex; or decrease or increase more than two years in age across CPS interviews.

To analyze healthcare industry transitions by occupation, we constrain the samples. For healthcare occupations we include only healthcare industry workers in healthcare occupations (HCO=1) at time $t-1$ for exit, and t for entry, respectively. The procedure is analogous for non-healthcare occupations.

Table A4 shows summary statistics for the exit and entry samples by mandate states and non-mandate states, separately. In the mandate states before the mandates and in non-mandate states, about 5.0% of workers exit and enter the healthcare industry each month. The transition rates are slightly higher in mandate states during the pre-mandate period than non-mandate states. The transition rate is higher among non-healthcare than healthcare occupation workers, consistent with more industry-specific skills for the latter than the former.[27] Respondent characteristics are well-balanced except for other races, Hispanics, and those born outside the U.S., where mandate states have higher shares than non-mandate states. As described in Section 3, California, New Mexico, and New York adopted healthcare mandates and these states are some of the most diverse in the U.S. (Kaiser Family Foundation, ND).

We estimate a regression analogous to Equation 1, but replace the dependent variable with either $Exit_{i,s,t}$ or $Enter_{i,s,t}$, using the samples as defined above. We analyze the contemporaneous impact of the mandate $Hvaccine_{s,t}$ on $Exit_{i,s,t}$ and a lagged mandate impact $Hvaccine_{s,t-1}$ on $Enter_{i,s,t}$, based on the reasoning that entry takes time, as establishments must undertake a search process for new employees. We control for respondent characteristics at time $t-1$. One-month lagged CPS weights are used for $Exit$ analysis and contemporaneous weights for $Enter$. We examine the same four specifications used in previous results: i) has only state- and month-fixed effects; ii) adds individual characteristics; iii) adds other state-level vaccine mandates; and iv) defines the healthcare mandate excluding the long-term care mandate.

Table 7 contains regression results. Starting with the top panel, for exit, in all four specifications the mandate is estimated to raise the exit rate from the healthcare industry by close to one ppt. Relative to the pre-treatment mean of 0.0535, this coefficient estimate implies a nearly 20% increase in the rate of leaving the healthcare industry. The bottom panel of Table 7 contains the results for the impact of the healthcare vaccine mandate on entry. The coefficient estimates are generally smaller than for the corresponding specification for exit, although the difference is sometimes small, as in specification (3).

Results for transitions that distinguish by healthcare and non-healthcare occupations

[27] Using the pre-COVID-19 period from 2014 to 2019, we find that HCO workers have a lower probability of exiting from the healthcare industry$_{t-1}$ to other industry$_t$ conditioning on individual characteristics and state and year fixed effects, compared to $nHCO$.

are presented in Table 8. We show results only for specification (2) and (3) from Table 7 because the others are similar. Although, as noted above, turnover (both exit and entry) in healthcare occupations tends to be lower than in non-healthcare occupations before treatment, the results imply a much bigger impact of the mandate on exit from healthcare occupations,although the difference in the estimated impacts is not statistically different from zero. The coefficient of 0.014 implies an approximate 30% increase in the exit rate. The impact on exit of workers in non-healthcare occupations is negligible, implying the workers in non-healthcare occupations do not react negatively to the mandate.

Moreover, turning to the bottom panel with results for entry, the vaccine mandate is estimated to have a much lower impact on hiring in healthcare occupations than either the exit rate among those occupations or compared to entry into non-healthcare occupations. Consistent with the results in Table 5, these findings suggest that the vaccine mandate may have resulted in a relative fall in healthcare occupations, implying a substitution of workers in non-healthcare for those in healthcare occupations.

5 Robustness and extensions

In this section, we explore the robustness of our main findings to alternative specifications and samples. For brevity, we focus on the probability of working in the healthcare industry. We display results graphically in Figure A2. We report the main estimated effect from Table 2, Column (3) on the left. We then sequentially add controls for the COVID-19 vaccination and death rates, and region-by-time fixed effects.[28] The latter captures the effect of region-specific trends in the probability of being employed in the healthcare sector, assuming each U.S. region to be on its own healthcare sector employment trajectory. After that, we tailor the analytical sample by excluding the data from the last quarter of 2022, drop 2022 (when a federal mandate for select healthcare workers was in place), dropping observations for states that removed the mandate during our study period (January 2021 to December 2022) after the expiration date, and dropping divisions that include no state that adopted a mandate, and using the event-study sample (i.e, we trim the data in event time, dropping observations for treated states more than eight months before or after mandate adoption). Next, we ignore any expiration of the mandates, assuming they were in effect until the end of our study period. The next specification uses a 'strong' mandate (one in which the employee will face adverse consequences, such as a firing, also controlling for other mandates) and the contemporaneous, instead of lagged, mandate. Finally, we report estimates from unweighted OLS,

[28] We use the four U.S. Census regions.

probit, and logit regressions. Overall, results are very stable and are not sensitive to the selection of covariates, samples, and the regression models.

The somewhat larger coefficient in the sample dropping 2022, although confidence intervals are overlapping, may deserve a bit more discussion. First, the event-study for the probability of healthcare industry employment (Figure 5) shows that the impact of the mandate increases for two months, reaching a maximum magnitude at two months after the effective date and becomes slightly weaker thereafter, a pattern that might be explained by quick resignations of vaccine-averse employees and slow replacement hiring. Also, the situation changed at the federal level, as the Department of Health and Human Services required full vaccination in facilities receiving Medicare and Medicaid by January 2022. The effectiveness of this rule is unclear, but it could have reduced the impact in 2022 relative to 2021.[29]

We also conduct a 'leave-one-out' analysis in which we sequentially exclude each state that adopts a COVID-19 vaccine mandate and re-estimate Equation 1. The objective of this exercise is to assess the extent to which our findings are driven by the experiences of specific states. Results are reported in Figure A3. Findings are very similar across the leave-one-out samples. In Figure A4, we conduct a parallel analysis in that we estimate effects separately by cohort-of-adoption, that is in each sample we include states that adopt the mandate in the same period (month-year) and non-adopting states. Coefficient estimates are negative for all cohorts, but effects are more pronounced among those states that adopted in September and October 2021 vs. November 2021.

Our study mainly focuses on the impact of COVID-19 employment vaccines on the probability of working in the healthcare industry and our results suggest that these mandates reduce this probability. However, from a welfare perspective, understanding whether these changes in employment have implications for the quality of healthcare delivered is also important. To provide some exploratory evidence on such impacts, we investigate whether there are changes in mortality among adults 65 years and older post-mandate. We select older adult mortality to proxy healthcare quality following previous work (Sloan et al., 2001; Picone et al., 2002; Amiri and Solankallio-Vahteri, 2019; Lasater et al., 2021).[30] The rationale for choosing this metric is that older adults are a vulnerable

[29]Enforcement of the federal requirements was postponed by a lawsuit of 26 states, until the Supreme Court narrowly voted in favor of the requirements in March 2022.

[30]In unreported analyses, available on request, we have examined the impact of these mandates on measures of healthcare shortages, the primary channel through which we expect the mandates to impact health, using data from Kaiser Family Foundation (2023a). We examined the impact of the mandates on the percent of physical healthcare needs met in each state and the number of healthcare professionals required to meet healthcare needs. Analyses suggested that both metrics worsened in states that adopted COVID-19 healthcare industry mandates, but the relationship was only statistically distinguishable from

population and their health, including mortality, is likely more responsive to changes in healthcare quality than younger adults. We acknowledge that this measure is at best a rough proxy for quality and encourage future work on this question. Our intent here is to offer exploratory and suggestive evidence only.

We draw monthly data 2021-2022 from the CDC Wide-ranging ONline Data for Epidemiologic Research ('WONDER') database (Centers for Disease Control and Prevention, 2024)[31] on total deaths, internal deaths,[32] and deaths in an inpatient or outpatient setting among adults 65 years or older. We convert these data to the rate per 100,000 adults 65 years and older in the state. We estimate a state-level version of Equation 1 to study the effects of vaccine mandates for healthcare workers on these outcomes, DID results are reported in Table 9. Overall, our findings suggest that the mandates worsen healthcare quality, as proxied by mortality rates among older adults. For example, in Table 9, column (1), we observe an increase of approximately 10 deaths per 100,000 adults 65 years and older, which implies a 3.0% increase in mortality rates. Event-studies, reported in Figure A5, suggest that all three mortality rates among older adults were trending down in states that would later adopt a mandate, relative to those that did not adopt a mandate, but that pattern was reversed after the mandate was adopted, with increases in mortality rates observed after approximately four months. The observed pre-trends likely work against our ability to detect mandate effects, but to further probe the empirical importance of such differential trends, in unreported analyses we have de-trended the data and results are nearly identical to those reported in Table 9.[33] While

zero for the latter metric only. We also explored comparable metrics for mental healthcare, but findings were inconclusive. We do not present these results as the data are only available at the annual level (we use years 2019, to avoid confounding from 2020, and 2021), thus we cannot exploit the same granularity in mandate timing (month-year) as we do in our main analyses.

[31] We use the provisional mortality statistics given our focus on a very recent time period. In the WONDER data, cells with less than eleven deaths are suppressed. However, there are no suppressed cells in our WONDER extraction.

[32] Internal deaths have no external reason for the death, such as a traumatic injury, and are viewed as 'natural' deaths (Fitzthum, 2022). We classify deaths using ICD-10 codes and include the following deaths as internal deaths: ICD-10 Codes: A00-B99 (certain infectious and parasitic diseases); C00-D48 (neoplasms); D50-D89 (diseases of the blood and blood-forming organs and certain disorders involving the immune mechanism); E00-E88 (endocrine, nutritional and metabolic diseases); F01-F99 (mental and behavioural disorders); G00-G98 (diseases of the nervous system); H00-H57 (diseases of the eye and adnexa); H60-H93 (diseases of the ear and mastoid process); I00-I99 (diseases of the circulatory system); J00-J98 (diseases of the respiratory system); K00-K92 (diseases of the digestive system); L00-L98 (diseases of the skin and subcutaneous tissue); M00-M99 (diseases of the musculoskeletal system and connective tissue); N00-N98 (diseases of the genitourinary system); Q00-Q99 (congenital malformations, deformations and chromosomal abnormalities); R00-R99 (symptoms, signs and abnormal clinical and laboratory findings, not elsewhere classified); U00-U99 (codes for special purposes). We also exclude deaths related to child birth given our focus on adults 65 years and older.

[33] We de-trend the data in the following manner. First, for each state that adopts a mandate by 2022, we estimate a linear trend for each outcome in the pre-mandate years, while for each state that does

this analysis has many caveats, the suggestive results imply that vaccine mandates for healthcare workers may have worsened healthcare quality.

6 Discussion

We document a 6% decline in the probability of working in the healthcare industry after states mandated COVID-19 vaccines for healthcare workers over the period 2021 to 2022. These results are not driven by pre-existing trends and are not sensitive to changes in the estimation method, analytical sample, or covariates included in the analysis. Our sub-sample analysis indicates that the decline in employment is observed in all groups and does not statistically differ by policy environment, sex, education, race, and ethnicity, education level, and place of residence. Dismayingly, we find the reduction is greater for workers in healthcare occupations compared with those in non-healthcare occupations. Further, the estimated effects are sustained several months after the mandate repeal.

Our results highlight the unintended consequences of policies designed to create a safer environment for healthcare workers and patients during the COVID-19 pandemic. We can compare our effect sizes with the impact of the COVID-19 pandemic on the U.S. healthcare workforce. The magnitude of the decline in healthcare worker employment due to vaccine mandates is comparable to the magnitude of the decline in employment in the healthcare industry during the first few months of the pandemic, beginning in March 2020, before experiencing a gain in employment rebounding to the pre-COVID-19 employment levels (Cantor et al., 2022; Buerhaus et al., 2022; Forsythe et al., 2020; Cortes and Forsythe, 2023; Shen et al., 2024).

Cantor et al. (2022) estimate that the employment level in the healthcare industry declined from 22.2 million in 2019 to 21.1 million by the second quarter of 2020 followed by a 700,000 increase after a year. Therefore, the employment level by the end of June 2021 (two months before the first COVID-19 vaccine mandate took effect) was 400,000 lower than the pre-COVID-19 level. Given that healthcare worker COVID-19 mandates were adopted in states that covered nearly 41% of the U.S. population (University of Kentucky Center for Poverty Research, 2023),[34] and considering employment proportional to the population level, our estimates suggest over a 536,000 decline in the healthcare worker employment level in the U.S. that is attributable to state vaccine mandates.

not adopt such a mandate by 2022, we estimate the linear trend over the full study period. Second, we subtract the estimated trends (separately for each outcome) off the outcome variable. We then estimate Equation 1 on these de-trended outcomes.

[34] We use U.S. Census population data prepared by the University of Kentucky Poverty Research Center.

We also find that the increased outflow rate resulting from the mandate is much higher for workers in healthcare-specific occupations than in other occupations, while the hiring rates are lower in the healthcare-specific occupations. The estimated results likely reflects higher costs of replacing the workers with healthcare-specific skills compared to non-healthcare workers in the industry. They also suggest that the industry may have reacted to the loss of workers in healthcare occupations by hiring workers with more general skills. This substitution may have implications for the quality of patient care, but that question is beyond the scope of this paper.

Our study has limitations. First, we lack data on mandate enforcement by employers, preventing us from exploring this margin of the mandates. Second, we have no information on the effect of healthcare worker vaccine mandates on healthcare workers' COVID-19 vaccination rates. This limitation is due to the unavailability of the vaccination statistics for this particular group at the national level (see Section 2). Although interesting, we argue that reporting these results is not essential for our analysis as we investigate how these mandates affect willingness to work in the healthcare industry. Furthermore, compelling evidence of vaccine hesitancy among healthcare workers makes our results plausible. Finally, we lack data on compliance costs to firms, which may be important. Earlier work from the pre-COVID-19 period shows that firms do not fully comply with health mandates (Maclean et al., 2020).

Our findings suggest that although vaccine mandates have shown to be successful in increasing vaccine uptake in non-U.S. settings, they appear to exacerbate on-going labor shortages in the healthcare industry. Therefore, to be able to benefit from the public health impact of these mandates, without harming the healthcare workforce, policymakers could consider combining mandates with other interventions that promote vaccination such as direct financial incentives; providing vaccines at the worksite and without costs; and education campaigns designed to increase information (and decrease dis-information) about vaccines in terms of effectiveness, positive externalities, and safety. Another possible approach - if elevating vaccination rates among healthcare workers comes at too high a cost in terms of workforce shortages - is to focus such efforts on other groups, for example patients in healthcare facilities and their family members, and society more broadly defined. Further, again returning to policy efforts targeting healthcare workers, providing generous paid sick leave to such workers and encouraging the use of this leave when showing signs of COVID-19 may be another method to reduce disease spread within healthcare facilities. In sum, future waves of COVID-19 and other viruses are likely to emerge and there could be another pandemic, and thus learning from the most recent experience could help develop policies for future events.

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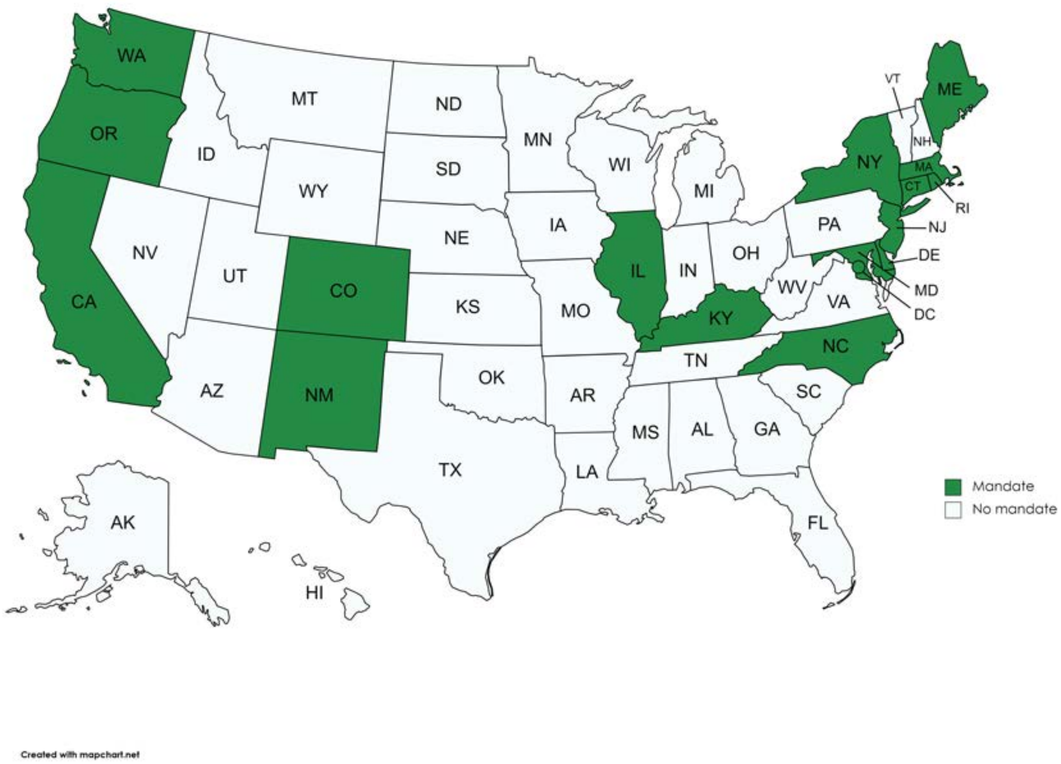
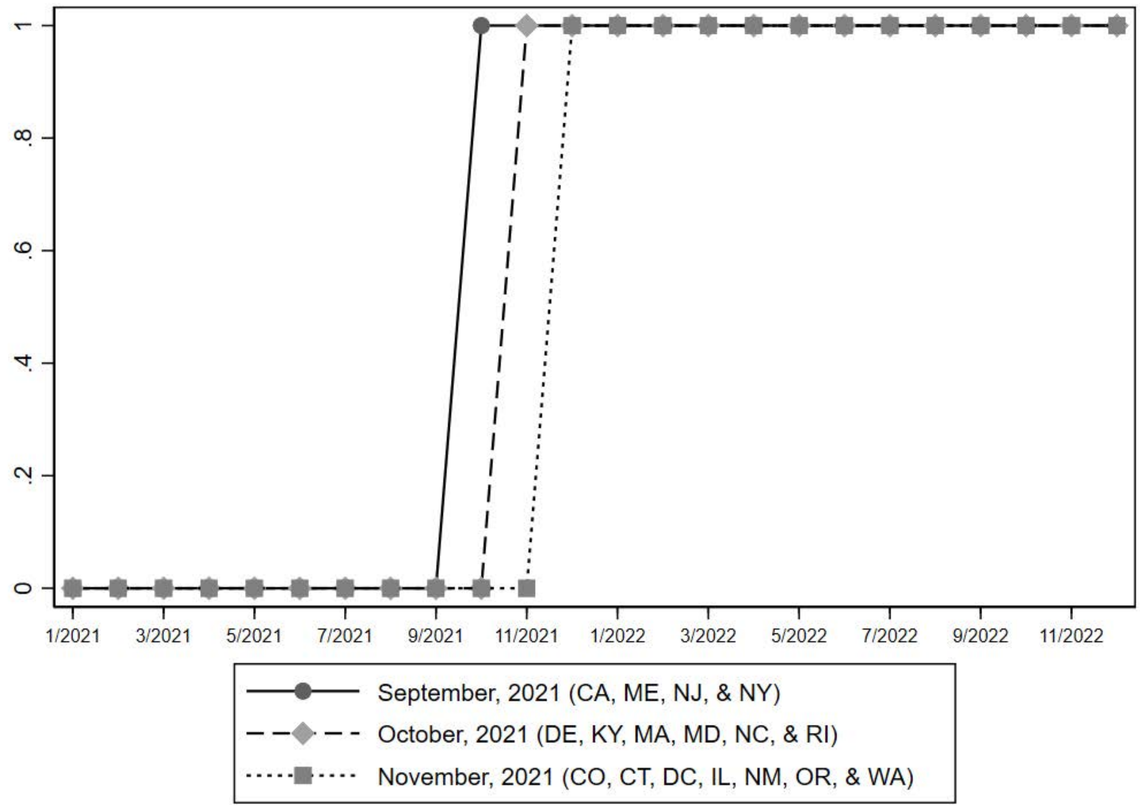
7 Figures and tables

Figure 1: Geographic distribution of state COVID-19 vaccine mandates for healthcare workers in the U.S.



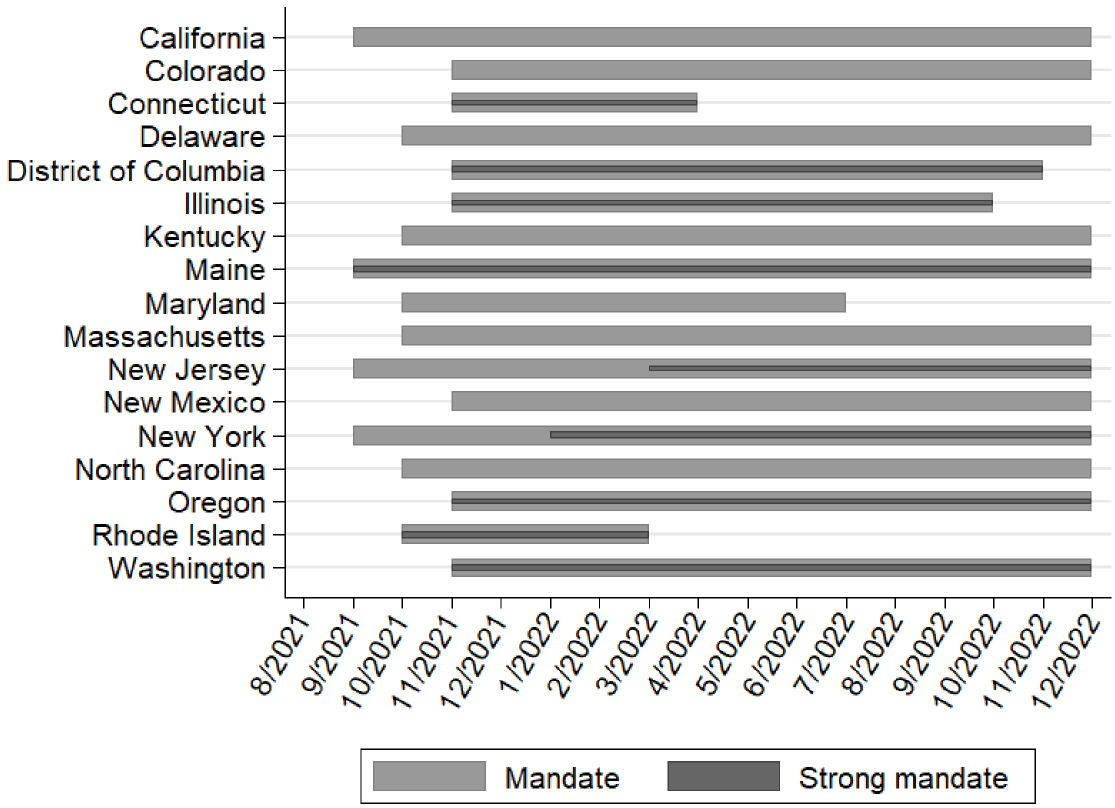
Notes: Data source is Howard-Williams (2022). Treatment states (effective periods) are as follows: September 2021 (CA, ME, NJ, and NY), October 2021 (DE, KY, MA, MD, NC, and RI), and November 2021 (CO, CT, DC, IL, NM, OR, and WA).

Figure 2: Temporal distribution of state COVID-19 vaccine mandates for healthcare workers in the U.S.



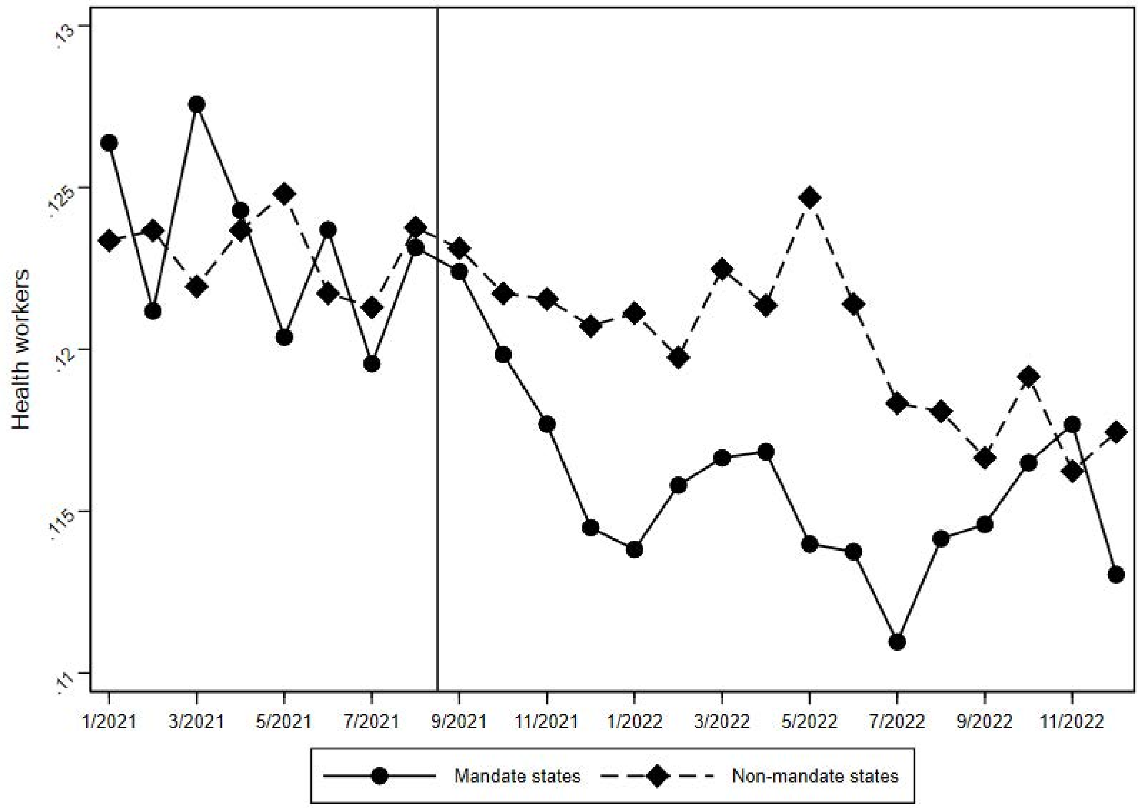
Notes: Data source is Howard-Williams (2022). Treatment states (effective periods) are as follows: September 2021 (CA, ME, NJ, and NY), October 2021 (DE, KY, MA, MD, NC, and RI), and November 2021 (CO, CT, DC, IL, NM, OR, and WA).

Figure 3: Temporal distribution of state COVID-19 vaccine mandates for healthcare workers in the U.S.



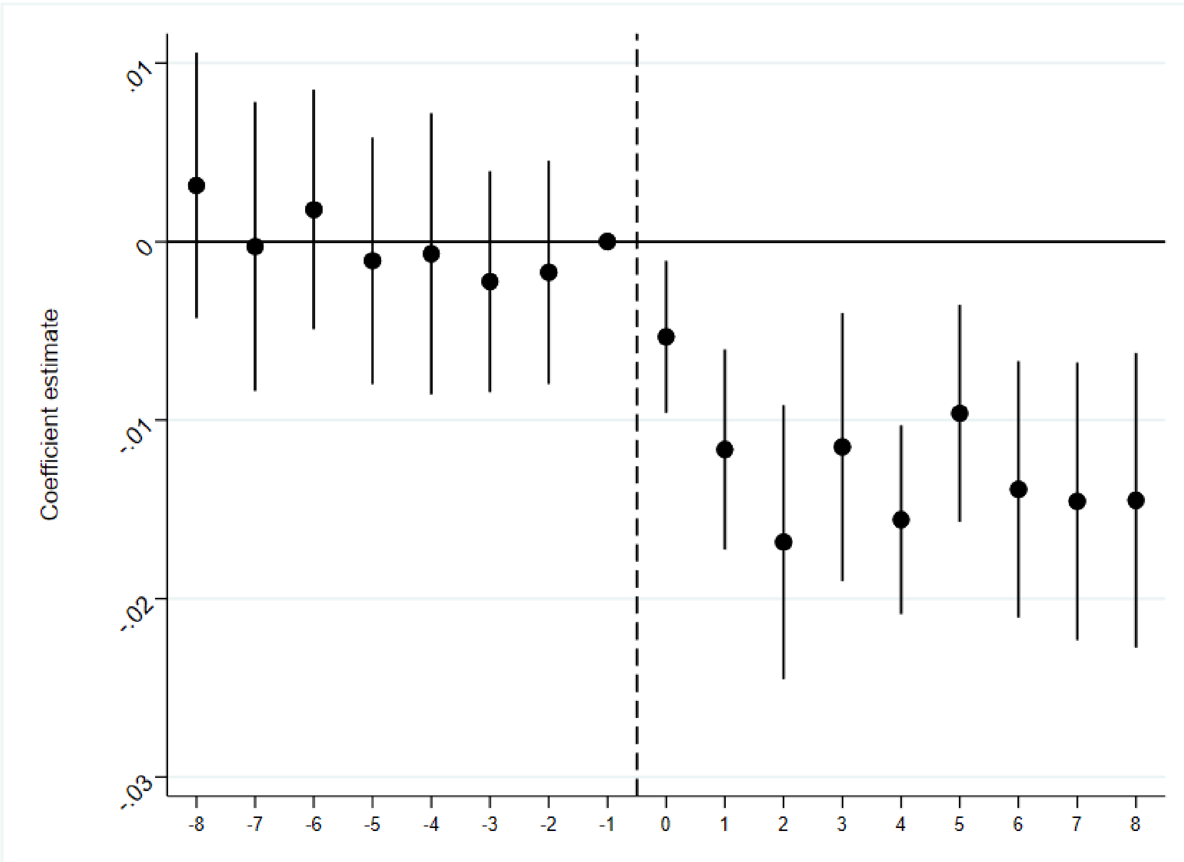
Notes: Data source is Howard-Williams (2022). Dates when the mandates became effective (states) are as follows: September 2021 (CA, ME, NJ, and NY), October 2021 (DE, KY, MA, MD, NC, and RI), and November 2021 (CO, CT, DC, IL, NM, OR, and WA) Mandate expiration dates (state) are as follows: March 2022 (RI), April 2022 (CT), July 2022 (MD), October 2022 (IL), and November 2022 (DC). CA, CO, DE, MA, ME, NC, NJ, NM, NY, OR, KY, and WA either allow their mandate after our study period or do not have an expiration date.

Figure 4: Trends in the probability of being a healthcare worker: CPS 2021-2022



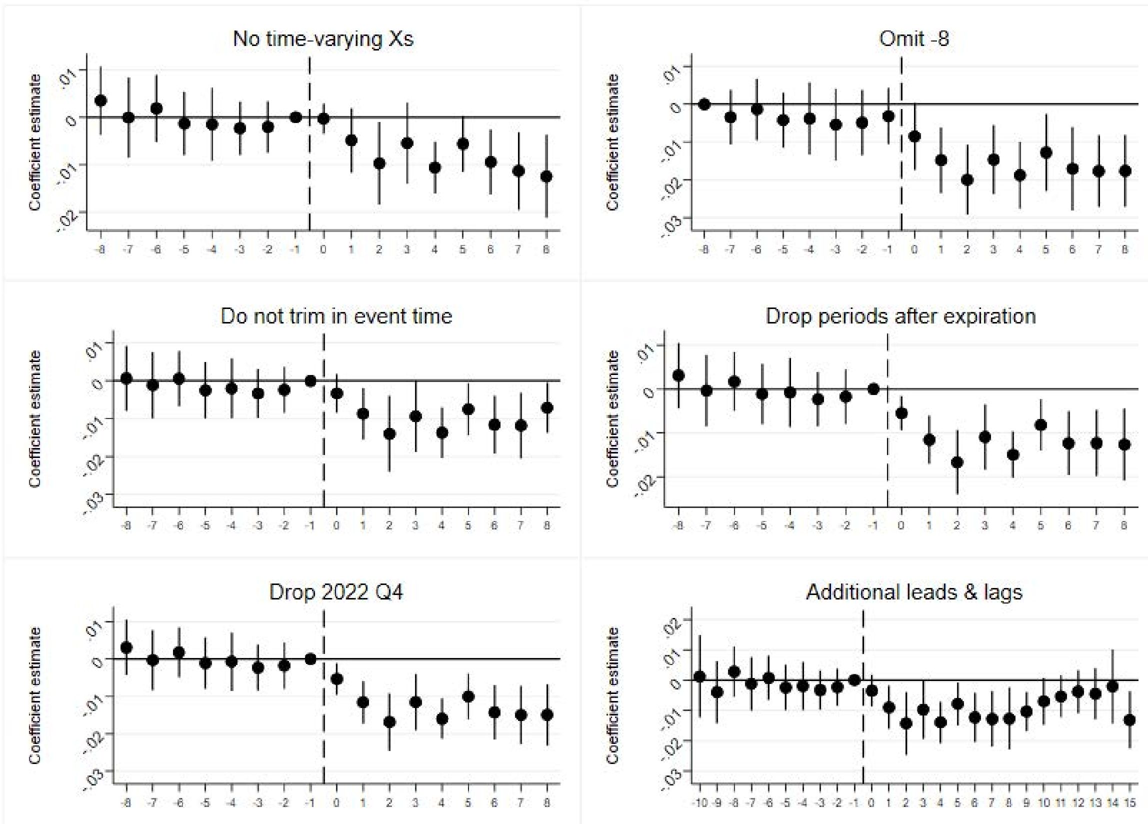
Notes: CPS = Current Population Survey. States with a healthcare worker vaccine mandate include: CA, CO, CT, DC, DE, IL, KY, MA, MD, ME, NC, NJ, NM, NY, OR, RI, and WA. Data are aggregated to the treatment group (treatment = have mandate, comparison = do not have mandate) year level. Data are weighted by CPS-provided weights prior to aggregation.

Figure 5: Effect of a state COVID-19 vaccine mandate for healthcare workers on the probability of working in the healthcare sector using an event-study: CPS 2021-2022



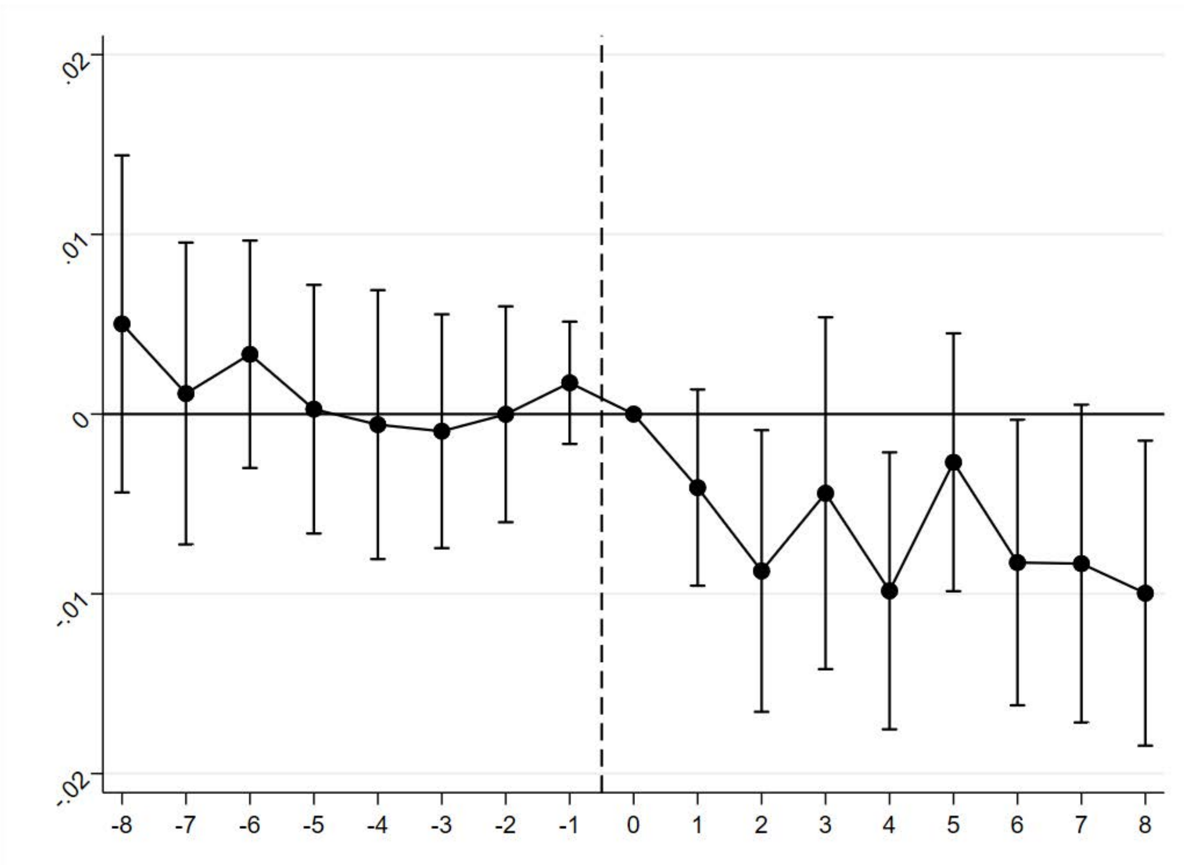
Notes: CPS = Current Population Survey. The regression includes a state COVID-19 vaccine mandate for healthcare workers (lagged one month), respondent characteristics, state characteristics, state fixed effects, and month-year fixed effects. The unit of observation is a respondent in a state in a month-year. Data are weighted by the CPS-provided survey weights. Regressions estimated with OLS. -1 is the omitted category. Data are trimmed in event-time for the treatment group (states that adopt a mandate): observations more than eight months before or after mandate adoption are excluded. Data are not trimmed in event-time for the comparison group (states that dot no adopt a mandate). Coefficient estimates are reported in circles and 95% confidence intervals that account for within-state clustering are reported with vertical lines.

Figure 6: Effect of a state COVID-19 vaccine mandate for healthcare workers on the probability of working in the healthcare sector using alternative event-study specifications and samples: CPS 2021-2022



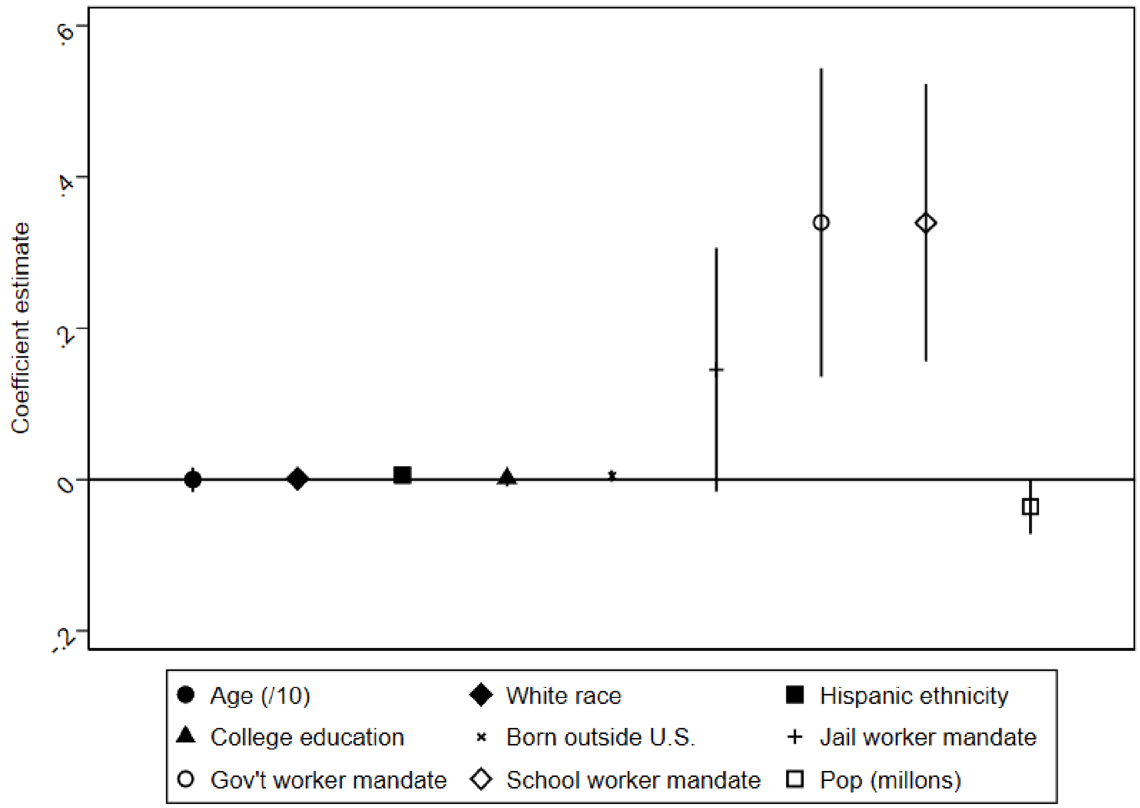
Notes: CPS = Current Population Survey. The regression includes a state COVID-19 vaccine mandate for healthcare workers (lagged one month), respondent characteristics, state characteristics, state fixed effects, and month-year fixed effects unless otherwise noted. The unit of observation is a respondent in a state in a month-year. Data are weighted by the CPS-provided survey weights. Regressions estimated with OLS. -1 is the omitted category unless otherwise noted. Data are trimmed in event-time for the treatment group (states that adopt a mandate): observations more than eight months before or after mandate adoption are excluded unless otherwise noted. Data are not trimmed in event-time for the comparison group (states that dot no adopt a mandate). Coefficient estimates are reported in circles and 95% confidence intervals that account for within-state clustering are reported with vertical lines.

Figure 7: Effect of a state COVID-19 vaccine mandate for healthcare workers on the probability of working in the healthcare sector using an event-study specification that is robust to bias from dynamic and heterogeneous treatment effects: CPS 2021-2022



Notes: CPS = Current Population Survey. The regression includes respondent characteristics, state characteristics, state fixed effects, and month-year fixed effects. The unit of observation is a respondent in a state in a month-year. Data are weighted by the CPS-provided survey weight. Regressions estimated with a method proposed by de Chaisemartin and d'Haultfoeuille (2020). Coefficient estimates are reported in circles and 95% confidence intervals that account for within-state clustering are reported with vertical lines.

Figure 8: Test of covariate balance: 2011-2022



Notes: See Section 3 for data sources. The outcome variable is reported in the figure legend. The regression includes a state COVID-19 vaccine mandate for healthcare workers (lagged one month), state fixed effects, and month-year fixed effects unless otherwise noted. The unit of observation is a state in a month-year. Data are weighted by the CPS-provided survey weights prior to aggregating to the state-month-year level. Regressions estimated with OLS. Coefficient estimates are reported in circles and 95% confidence intervals that account for within-state clustering are reported with vertical lines.

Table 1: Summary statistics: CPS 2021-2022

Sample:	Mandate states, pre-mandate	Non-mandate states
Outcome		
Healthcare worker	0.12	0.12
Respondent characteristics		
21-29 years	0.20	0.21
30-39 years	0.26	0.25
40-49 years	0.23	0.23
50-59 years	0.22	0.22
60-64 years	0.09	0.08
Male	0.53	0.53
Female	0.47	0.47
White	0.75	0.79
African American	0.11	0.14
Other race	0.14	0.08
Non-Hispanic	0.79	0.84
Hispanic	0.21	0.16
Hispanic	0.21	0.16
Less than high school	0.07	0.06
High school	0.22	0.27
Some college	0.24	0.28
College degree	0.47	0.39
Born inside the U.S.	0.76	0.84
Borth outside the U.S.	0.24	0.16
State characteristics		
Jail worker COVID-19 vaccine mandate	0.000	0.012
Gov't worker COVID-19 vaccine mandate	0.054	0.045
School worker COVID-19 vaccine mandate	0.003	0.002
Population	17,249,385	12,241,176
Observations	138,945	648,714

Notes: CPS = Current Population Survey. The unit of observation is a respondent in state in a month-year. Data are weighted by CPS-provided survey weights.

Table 2: Effect of a state COVID-19 vaccine mandate for healthcare workers on the probability of working in the healthcare sector: CPS 2021-2022

Specification:	(1)	(2)	(3)	(4)	(5)
Healthcare worker COVID-19 mandate (lag one month)	-0.0049**	-0.0049**	-0.0072***	-0.0051**	-0.0071**
	(0.0023)	(0.0022)	(0.0025)	(0.0025)	(0.0034)
State fixed effects	Y	Y	Y	Y	Y
Period fixed effects	Y	Y	Y	Y	Y
Respondent characteristics	N	Y	Y	Y	Y
State characteristics	N	N	Y	Y	Y
No LTC worker mandate	N	N	N	Y	N
DCDH	N	N	N	N	Y
Pre-treatment mean	0.1232	0.1232	0.1232	0.1232	0.1232
Observations	1,018,333	1,018,333	1,018,333	1,018,333	1,018,333

Notes: CPS = Current Population Survey. LTC = long-term care worker. DCDH = de Chaisemartin and d'Haultfoeuille (2020). The unit of observation is a respondent in a state in a month-year. Data are weighted by the CPS-provided survey weights unless otherwise noted. Regressions estimated with OLS. Standard errors are clustered at the state level and are reported in parentheses.
***, **, and * = statistically different from zero at the 1%, 5%, and 10% level.

Table 3: Effect of a state COVID-19 vaccine mandate for healthcare workers on the probability of employment and being in the labor market: CPS 2021-2022

Outcome:	In labor force	Employed	Fulltime	Parttime
Healthcare worker COVID-19	0.0028	0.0050**	0.0007	0.0024
mandate (lag one month)	(0.0025)	(0.0025)	(0.0034)	(0.0023)
Pre-treatment mean	0.7694	0.7192	0.5592	0.1316
Observations	1,372,596	1,372,596	1,372,596	1,372,596

Notes: CPS = Current Population Survey. The regression includes a state COVID-19 vaccine mandate for healthcare workers (lagged one month), respondent characteristics, state characteristics, state fixed effects, and month-year fixed effects. The unit of observation is a respondent in a state in a month-year. Data are weighted by the CPS-provided survey weights. Regressions estimated with OLS. Standard errors are clustered at the state level and are reported in parentheses.
***, **, and * = statistically different from zero at the 1%, 5%, and 10% level.

Table 4: Heterogeneity by state characteristics in the effect of a state COVID-19 vaccine mandate for healthcare workers on the probability of working in the healthcare sector: CPS 2021-2022

Stratify by PSL mandate:	Mandate	No mandate
Healthcare worker COVID-19 mandate (lagged one month)	-0.0054	-0.0048
	(0.0036)	(0.0032)
Pre-treatment mean	0.1219	0.1272
Observations	309,692	708,641
Stratify by ACA Medicaid expansion:	Expansion	Non-expansion
Healthcare worker COVID-19 mandate (lagged one month)	-0.0084***	-0.0033
	(0.0028)	(0.0032)
Pre-treatment mean	0.1231	0.1248
Observations	738,003	280,330

Notes: CPS = Current Population Survey. PSL = paid sick leave mandate. ACA = Affordable Care Act. The regression includes a state COVID-19 vaccine mandate for healthcare workers (lagged one month), respondent characteristics, state characteristics, state fixed effects, and month-year fixed effects. The unit of observation is a respondent in a state in a month-year. Data are weighted by the CPS-provided survey weights. Regressions estimated with OLS. Standard errors are clustered at the state level and are reported in parentheses.
***, **, and * = statistically different from zero at the 1%, 5%, and 10% level.

Table 5: Heterogeneity by respondent characteristics in the effect of a state COVID-19 vaccine mandate for healthcare workers on the probability of working in the healthcare sector: CPS 2021-2022

Stratify by sex:	Men	Women
Healthcare worker COVID-19 mandate (lagged one month)	-0.0024	-0.0127**
	(0.0026)	(0.0049)
Pre-treatment mean	0.0578	0.1963
Observations	531,812	486,521
Stratify by race:	White	Non-white
Healthcare worker COVID-19 mandate (lagged one month)	-0.0060*	-0.0102*
	(0.0030)	(0.0059)
Pre-treatment mean	0.1117	0.1577
Observations	817,500	200,833
Stratify by ethnicity:	Hispanic	Non-Hispanic
Healthcare worker COVID-19 mandate (lagged one month)	-0.0041	-0.0077***
	(0.0053)	(0.0026)
Pre-treatment mean	0.0928	0.1314
Observations	153,782	864,551
Stratify by education:	College	No college
Healthcare worker COVID-19 mandate (lagged one month)	-0.0085*	-0.0058**
	(0.0044)	(0.0027)
Pre-treatment mean	0.1390	0.1090
Observations	432,255	586,078
Stratify by rural residence:	Rural	Non-rural
Healthcare worker COVID-19 mandate (lagged one month)	-0.0046	-0.0083***
	(0.0041)	(0.0030)
Pre-treatment mean	0.1280	0.1222
Observations	343,990	674,343
Stratify by occupation:	Non-healthcare	Healthcare
Healthcare worker COVID-19 mandate (lagged one month)	-0.0021	-0.0051**
	(0.0022)	(0.0021)
Pre-treatment mean	0.0458	0.0774
Observations	1,018,333	1,018,333

Notes: CPS = Current Population Survey. The regression includes a state COVID-19 vaccine mandate for healthcare workers (lagged one month), respondent characteristics, state characteristics, state fixed effects, and month-year fixed effects. The unit of observation is a respondent in a state in a month-year. Data are weighted by the CPS-provided survey weights. Regressions estimated with OLS. Standard errors are clustered at the state level and are reported in parentheses.
***, **, and * = statistically different from zero at the 1%, 5%, and 10% level.

Table 6: Heterogeneity by sector in the effect of a state COVID-19 vaccine mandate for healthcare workers on the probability of working in the healthcare sector: CPS 2021-2022

Outcome:	Long-term care	Nurses	MH/SUD
Healthcare worker COVID-19	-0.0002	-0.0024*	0.0006
mandate (lagged one month)	(0.0010)	(0.0014)	(0.0005)
Pre-treatment mean	0.0125	0.0290	0.0033
Observations	1,018,333	1,018,333	1,018,333

Notes: CPS = Current Population Survey. MH = mental health. SUD = substance use disorder. The regressions includes a state COVID-19 vaccine mandate for healthcare workers (lagged one month), respondent characteristics, state characteristics, state fixed effects, and month-year fixed effects. The unit of observation is a respondent in a state in a month-year. Data are weighted by the CPS-provided survey weights. Regressions estimated with OLS. Standard errors are clustered at the state level and are reported in parentheses.
***, **, and * = statistically different from zero at the 1%, 5%, and 10% level.

Table 7: Effect of a state COVID-19 vaccine mandate for healthcare workers on the probability of exiting and entering the healthcare sector: CPS 2021-2022

Specification:	(1)	(2)	(3)	(4)
Exit:				
Healthcare worker COVID-19	0.0095**	0.0098**	0.0088**	0.0087**
mandate	(0.0041)	(0.0042)	(0.0039)	(0.0034)
Pre-treatment mean	0.0535	0.0535	0.0535	0.0535
Observations	82,211	82,211	82,211	82,211
Entry:				
Healthcare worker COVID-19	0.0059*	0.0064**	0.0087*	0.0067
mandate (lagged one month)	(0.0031)	(0.0031)	(0.0046)	(0.0042)
Pre-treatment mean	0.0505	0.0505	0.0505	0.0505
Observations	81,926	81,926	81,926	81,926
State fixed effects	Y	Y	Y	Y
Period fixed effects	Y	Y	Y	Y
Respondent characteristics	N	Y	Y	Y
State characteristics	N	N	Y	Y
No LTC worker mandate	N	N	N	Y

Notes: CPS = Current Population Survey. The variable of interest = healthcare worker COVID-19 mandate. The variable is contemporaneous for exit and lagged one month for entry. The unit of observation is a respondent in a state in a month-year. Data are weighted by the CPS-provided survey weights at $t-1$ for exit and t for entry. Standard errors are clustered at the state level and are reported in parentheses.
***, **, and * = statistically different from zero at the 1%, 5%, and 10% level.

Table 8: Heterogeneity by occupation in the effect of a state COVID-19 vaccine mandate for healthcare workers on the probability of exit and enter the healthcare sector: CPS 2021-2022

Specification:	(1)	(2)	(3)	(4)
Exit:	Healthcare occupation		Non-healthcare occupation	
Healthcare worker COVID-19 mandate	0.0143***	0.0142**	0.0019	0.0007
	(0.0041)	(0.0061)	(0.0082)	(0.0105)
Pre-treatment mean	0.0501	0.0501	0.0595	0.0595
Observations	52,194	52,194	30,017	30,017
Entry:	Healthcare occupation		Non-healthcare occupation	
Healthcare worker COVID-19 mandate	0.0049	0.0078	0.0098**	0.0126*
(lagged one month)	(0.0040)	(0.0060)	(0.0049)	(0.0073)
Pre-treatment mean	0.0491	0.0491	0.0530	0.0530
Observations	52,065	52,065	29,861	29,861
State characteristics	N	Y	N	Y

Notes: CPS = Current Population Survey. The variable of interest = healthcare or long-term care worker mandate. The variable is contemporaneous for exit and lagged one month for entry. The regressions includes a state COVID-19 vaccine mandate for healthcare workers, respondent characteristics, state fixed effects, and month-year fixed effects. The unit of observation is a respondent in a state in a month-year. Data are weighted by the CPS-provided survey weights at $t-1$ for exit and t for entry. Standard errors are clustered at the state level and are reported in parentheses.

***, **, and * = statistically different from zero at the 1%, 5%, and 10% level.

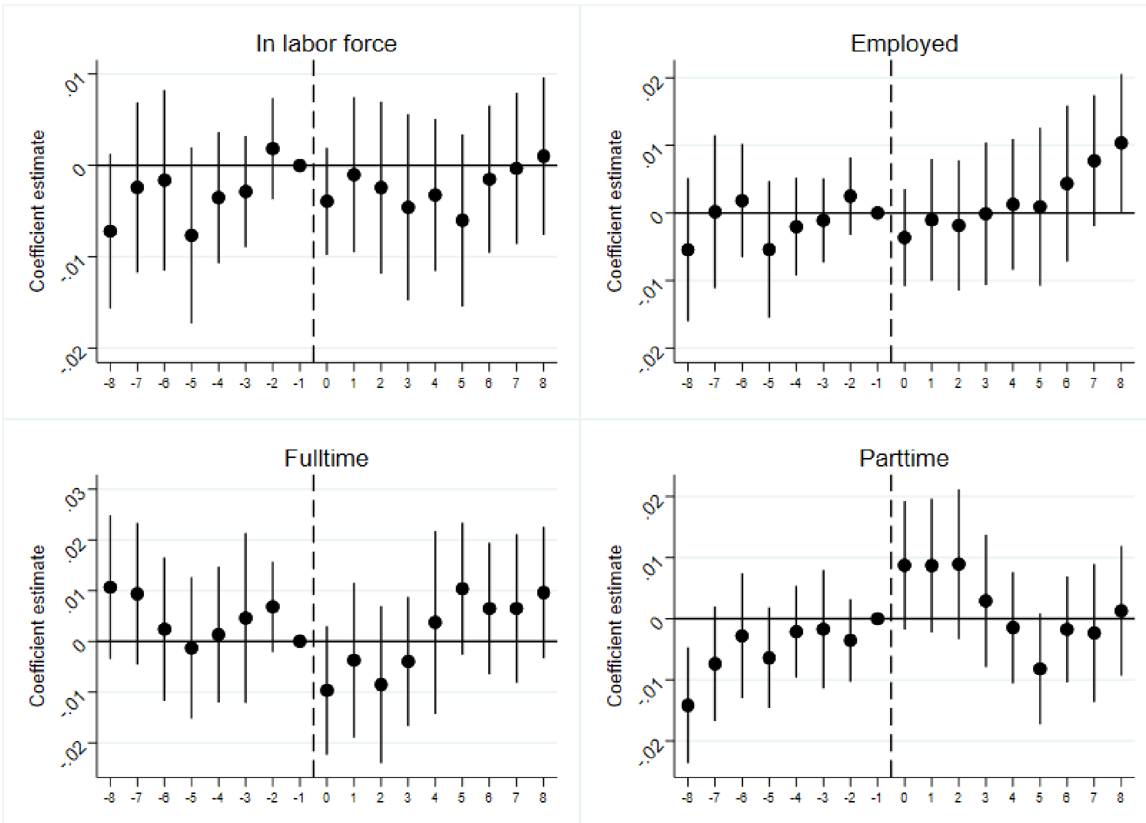
Table 9: Effect of a state COVID-19 vaccine mandate for healthcare workers on adults 65 years and older mortality rates: NVSS 2021-2022

Outcome:	All deaths	Internal detahs	Deaths in medical facility
Healthcare worker COVID-19 mandate	10.4207*	10.2769*	8.8203**
(lagged one month)	(5.3332)	(5.2901)	(3.7825)
Pre-treatment mean	343.5314	332.3248	124.0883
Observations	1,224	1,224	1,224

Notes: Notes: The regression includes state policies, state fixed effects, and month-year fixed effects. The unit of observation is a state in a month in a year. Data are weighted by the state population 65 years and older. Regressions estimated with OLS. Standard errors are clustered around the state and reported in parentheses. ***,**,* = statistically different from zero at the 1%, 5%, and 10% level.

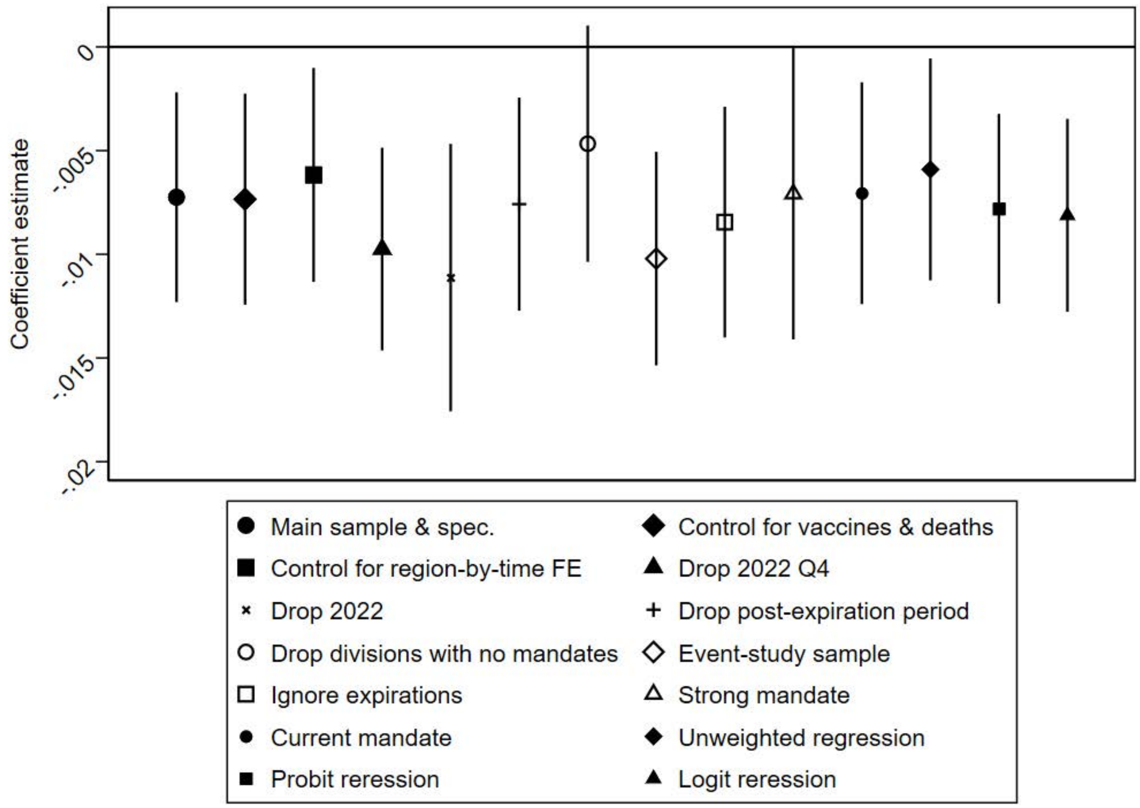
8 Appendix

Figure A1: Effect of a state COVID-19 vaccine mandate for healthcare workers on the probability of employment outcomes using an event-study: CPS 2021-2022



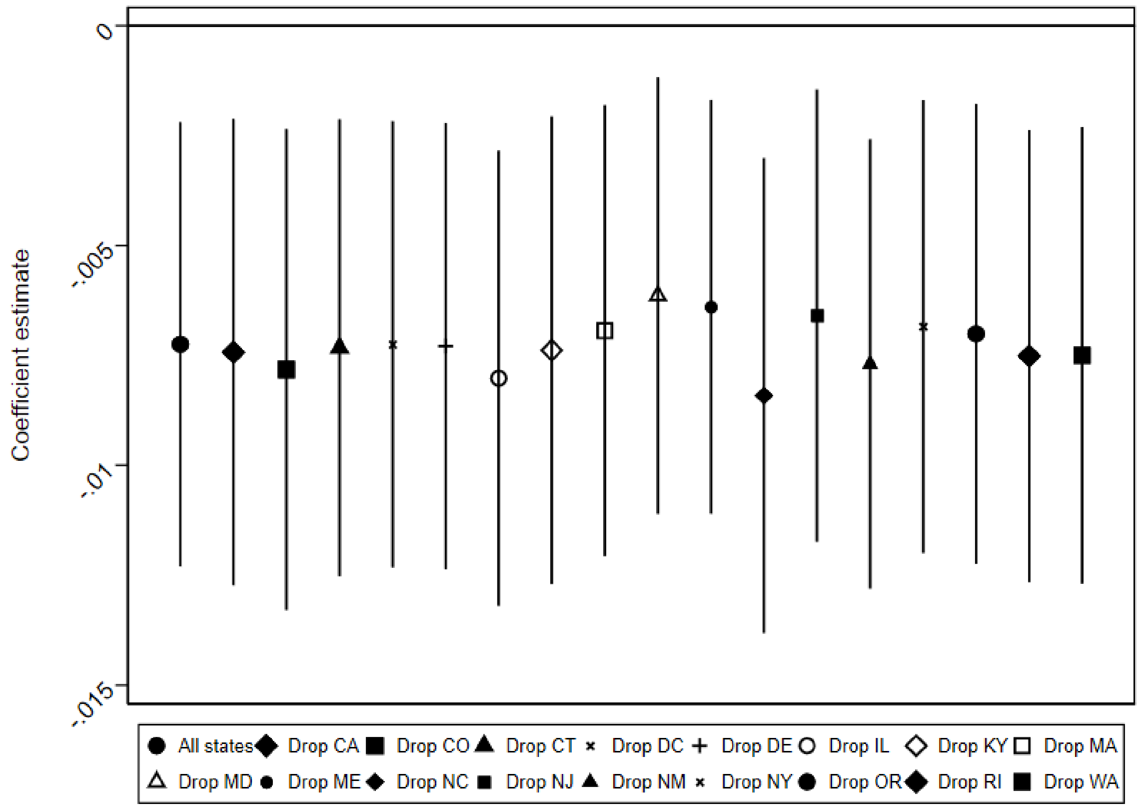
Notes: CPS = Current Population Survey. The regression includes a state COVID-19 vaccine mandate for healthcare workers (lagged one month), respondent characteristics, state characteristics, state fixed effects, and month-year fixed effects. The unit of observation is a respondent in a state in a month-year. Data are weighted by the CPS-provided survey weights. Regressions estimated with OLS. -1 is the omitted category. Data are trimmed in event-time for the treatment group (states that adopt a mandate): observations more than eight months before or after mandate adoption are excluded. Data are not trimmed in event-time for the comparison group (states that dot no adopt a mandate). Coefficient estimates are reported in circles and 95% confidence intervals that account for within-state clustering are reported with vertical lines.

Figure A2: Effect of a state COVID-19 vaccine mandate for healthcare workers on the probability of working in the healthcare sector using different specifications and samples: CPS 2021-2022



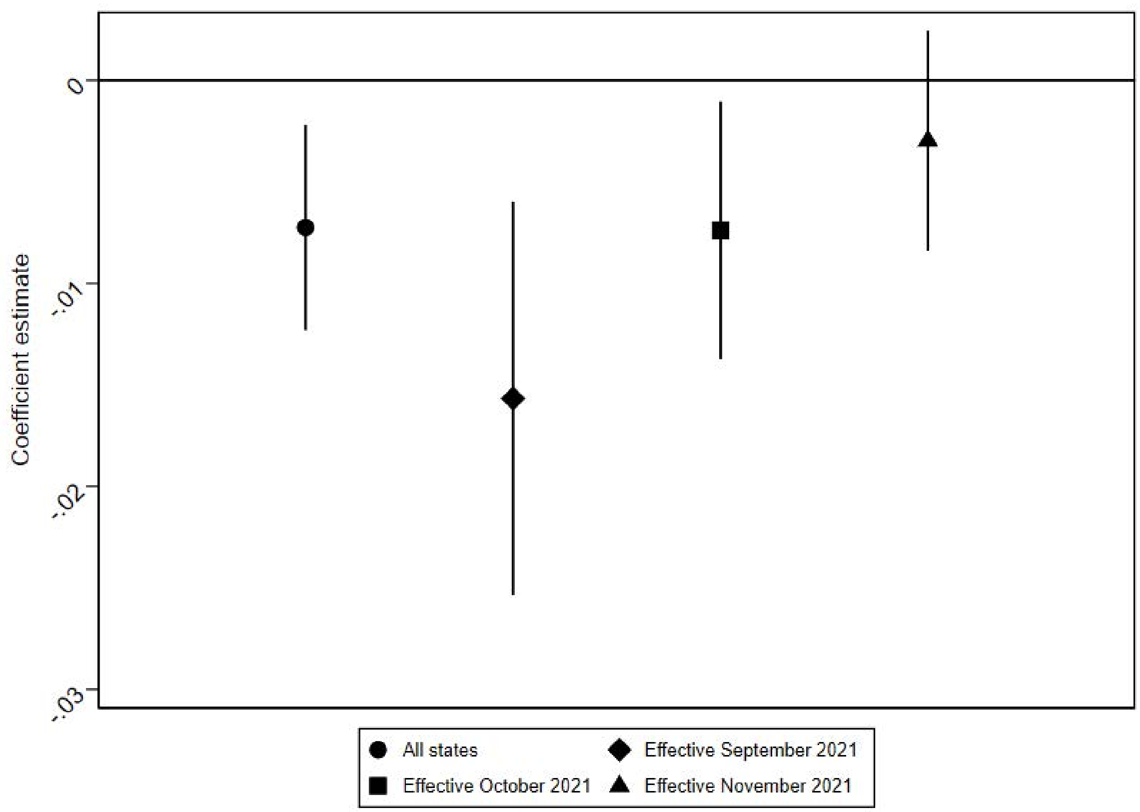
Notes: CPS = Current Population Survey. The regression includes a state COVID-19 vaccine mandate for healthcare workers (lagged one month), respondent characteristics, state characteristics, state fixed effects, and month-year fixed effects unless otherwise noted. The unit of observation is a respondent in a state in a month-year. Data are weighted by the CPS-provided survey weights unless otherwise noted. Regressions estimated with OLS unless otherwise noted. Coefficient estimates are reported in circles and 95% confidence intervals that account for within-state clustering are reported with vertical lines.

Figure A3: Effect of a state COVID-19 vaccine mandate for healthcare workers on the probability of working in the healthcare sector sequentially excluding each treatment state: CPS 2021-2022



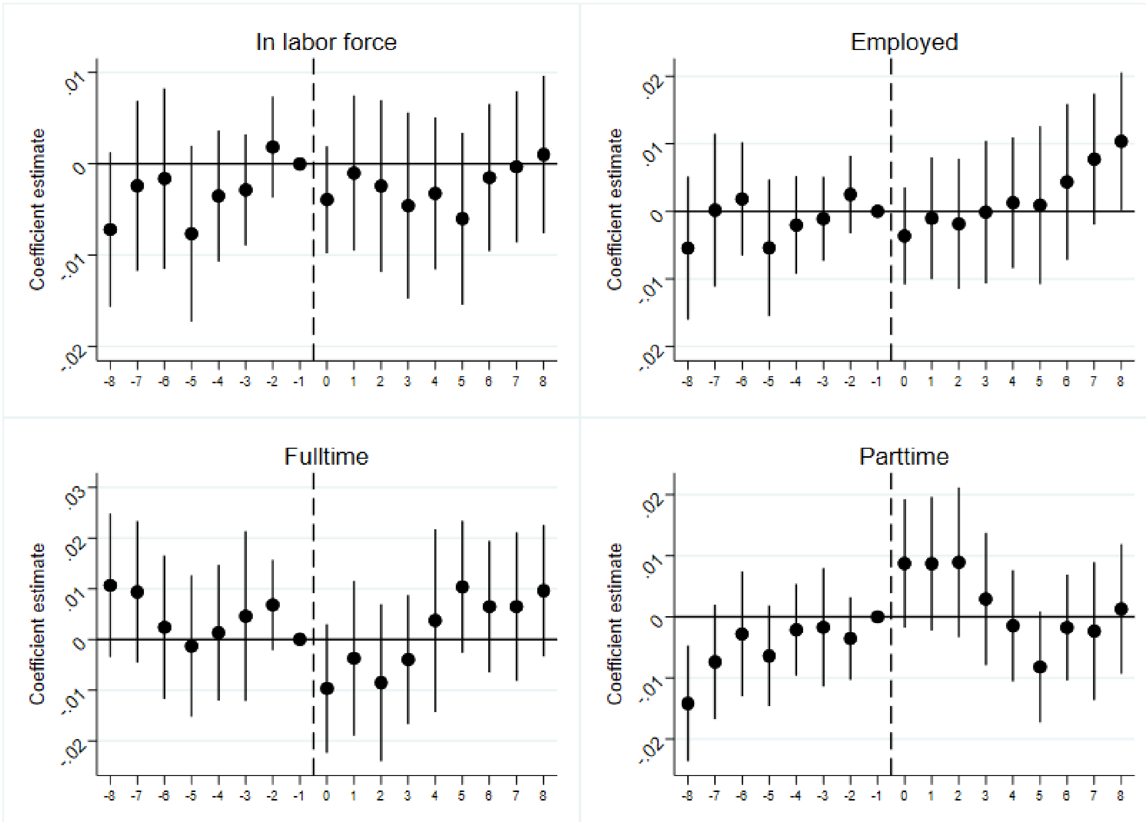
Notes: CPS = Current Population Survey. The excluded state is listed in the figure legend. The regression includes a state COVID-19 vaccine mandate for healthcare workers (lagged one month), respondent characteristics, state characteristics, state fixed effects, and month-year fixed effects. The unit of observation is a respondent in a state in a month-year. Data are weighted by the CPS-provided survey weights. Regressions estimated with OLS. Coefficient estimates are reported in circles and 95% confidence intervals that account for within-state clustering are reported with vertical lines.

Figure A4: Effect of a state COVID-19 vaccine mandate for healthcare workers on the probability of working in the healthcare sector by cohort of mandate adoption: CPS 2021-2022



Notes: Cohort of adoption = groups of states that adopted a mandate in the same period (month-year). CPS = Current Population Survey. The excluded state is listed in the figure legend. The regression includes a state COVID-19 vaccine mandate for healthcare workers (lagged one month), respondent characteristics, state characteristics, state fixed effects, and month-year fixed effects. The unit of observation is a respondent in a state in a month-year. Data are weighted by the CPS-provided survey weights. Regressions estimated with OLS. Coefficient estimates are reported in circles and 95% confidence intervals that account for within-state clustering are reported with vertical lines.

Figure A5: Effect of a state COVID-19 vaccine mandate for healthcare workers on adults 65 years and older mortality rates using and event-study: CDC WONDER 2021-20222



Notes: CDC WONDER = Centers for Disease Control and Prevention Wide-ranging ONline Data for Epidemiologic Research. The regression includes a state COVID-19 vaccine mandate for healthcare workers (lagged one month), state characteristics, state fixed effects, and month-year fixed effects. The unit of observation is a state in a month-year. Data are weighted by the state population 65 years and older. Regressions estimated with OLS. -1 is the omitted category. Data are trimmed in event-time for the treatment group (states that adopt a mandate): observations more than eight months before or after mandate adoption are excluded. Data are not trimmed in event-time for the comparison group (states that dot no adopt a mandate). Coefficient estimates are reported in circles and 95% confidence intervals that account for within-state clustering are reported with vertical lines.

Table A1: Healthcare industry codes: CPS 2021-2022

Industry label	Industry code
Offices of physicians	7970
Offices of dentists	7980
Offices of chiropractors	7990
Offices of optometrists	8070
Offices of other health practitioners	8080
Outpatient care centers	8090
Home health care services	8170
Other health care services	8180
General medical and surgical hospitals, and specialty	8191
Psychiatric and substance abuse hospitals	8192
Nursing care facilities	8270
Residential care facilities, without nursing	8290

Notes: CPS = Current Population Survey. We use the variable *IND* included in the IPUMS harmonized CPS data set (Flood et al., 2023).

Table A2: Healthcare Occupation codes: CPS 2021-2022

Healthcare practitioners and technical occupations			
3000	Chiropractors	3261	Acupuncturist
3010	Dentists	3270	Healthcare diagnosing or treating practitioners, all other
3030	Dietitians and nutritionists	3300	Clinical laboratory technologists and technician
3040	Optometrist	3310	Dental hygienist
3050	Pharmacist	3321	Cardiovascular technologists and technician
3090	Other physician	3322	Diagnostic medical sonographer
3100	Surgeon	3323	Radiologic technologists and technician
3110	Physician assistant	3324	Magnetic resonance imaging technologist
3140	Audiologist	3330	Nuclear medicine technologists and medical dosimetrist
3150	Occupational therapist	3401	Emergency medical technician
3160	Physical therapist	3402	Paramedic
3200	Radiation therapist	3421	Pharmacy technician
3210	Recreational therapist	3422	Psychiatric technician
3220	Respiratory therapist	3423	Surgical technologist
3230	Speech-language pathologists	3424	Veterinary technologists and technician
3245	Exercise physiologists and therapists, all other	3430	Dietetic technicians and ophthalmic medical technician
		3500	Licensed practical and licensed vocational nurse
3250	Veterinarian	3515	Medical records specialist
3255	Registered nurse	3520	Opticians, dispensing
3256	Nurse anesthetist	3545	Miscellaneous health technologists and technicians
3258	Nurse practitioner	3550	Other healthcare practitioners and technical occupations
3261	Acupuncturist	3550	Other healthcare practitioners and technical occupations
Healthcare Support Occupations			
3601	Home health aides	3640	Dental assistants
3602	Personal care aides	3645	Medical assistants
3603	Nursing assistants	3646	Medical transcriptionists
3605	Orderlies and psychiatric aides	3647	Pharmacy aides
3610	Occupational therapist assistants and aides	3648	Veterinary assistants and laboratory animal caretakers
3620	Physical therapist assistants and aides	3649	Phlebotomists
3630	Massage therapists	3655	Other healthcare support workers

Notes: CPS = Current Population Survey. We use the variable *OCC* included in the IPUMS harmonized CPS data set (Flood et al., 2023).

Table A3: Demographics of non-healthcare and healthcare industry respondents: CPS 2021-2022

Column	(1)	(2)	(3)	(4)
Sample:	Non-healthcare	Healthcare industry	Non-healthcare occupation	Healthcare occupation
21-29 years	0.21	0.20	0.18	0.21
30-39 years	0.24	0.27	0.26	0.28
40-49 years	0.21	0.23	0.23	0.23
50-59 years	0.22	0.22	0.24	0.20
60-64 years	0.12	0.086	0.093	0.081
Male	0.52	0.23	0.26	0.21
Female	0.48	0.77	0.74	0.79
White	0.76	0.71	0.73	0.71
African American	0.13	0.17	0.18	0.17
Other race	0.11	0.11	0.099	0.12
Non-Hispanic	0.81	0.86	0.85	0.87
Hispanic	0.19	0.14	0.15	0.13
Less than high school	0.086	0.025	0.025	0.024
High school	0.29	0.17	0.22	0.14
Some college	0.26	0.31	0.30	0.32
College degree	0.37	0.50	0.46	0.52
Born inside the U.S.	0.79	0.82	0.85	0.80
Born outside the U.S.	0.21	0.18	0.15	0.20
Rural	0.25	0.25	0.24	0.25
Non-rural	0.75	0.75	0.76	0.75
Northeast	0.17	0.20	0.19	0.20
Midest	0.20	0.23	0.22	0.23
South	0.39	0.36	0.36	0.36
West	0.24	0.22	0.23	0.21
In the labor force	0.75	0.99	0.99	0.99
Employed	0.72	0.97	0.97	0.97
Full-time work	0.56	0.74	0.78	0.72
Part-time work	0.13	0.19	0.16	0.21
Observations	1,244,651	127,945	46,724	81,221

Notes: CPS = Current Population Survey. The sample includes civilians, regardless of current employment status, ages 21-64, with additional restrictions indicated by column. The unit of observation is a respondent in a state in a month-year. Data are weighted by the CPS-provided survey weight. We use the variables *IND* and *OCC*, which provide information of the most recent work for non-employed and of the current work for employed, included in the IPUMS harmonized CPS data set (Flood et al., 2023). Columns (3) and (4) include respondents in the healthcare industry, but stratified by non-healthcare and healthcare occupations.

Table A4: Summary statistics for transition analysis sample: CPS 2021-2022

	$Exit_t$		$Enter_t$	
	Mandate states, pre-mandate	Non-mandate states	Mandate states, pre-mandate	Non-mandate states
$Exit_t$ or $Enter_t$	0.0568	0.0526	0.0547	0.0494
Among $nHCO$	0.0618	0.0589	0.0575	0.0518
	(obs: 4126)	(obs: 19418)	(obs: 4120)	(obs: 19304)
Among HCO	0.0539	0.0492	0.0531	0.0481
	(obs: 7,034)	(obs: 34,206)	(obs: 7,028)	(obs: 34,145)
21-29 years	0.19	0.20	0.18	0.20
30-39 years	0.28	0.27	0.29	0.27
40-49 years	0.22	0.23	0.22	0.23
50-59 years	0.22	0.22	0.22	0.22
60-64 years	0.08	0.09	0.08	0.09
Female	0.75	0.78	0.75	0.78
White	0.69	0.76	0.69	0.76
African American	0.15	0.17	0.15	0.16
Other race	0.16	0.08	0.16	0.08
Hispanic	0.16	0.11	0.15	0.11
Less than high school	0.02	0.02	0.02	0.02
High school	0.15	0.16	0.15	0.16
Some college	0.29	0.33	0.29	0.33
College degree	0.51	0.49	0.51	0.49
Born outside the U.S.	0.22	0.13	0.21	0.13
State characteristics				
Jail worker COVID-19 vaccine mandate	0.00	0.01	0.00	0.01
Gov't worker COVID-19 vaccine mandate	0.05	0.05	0.05	0.05
School worker COVID-19 vaccine mandate	0.00	0.00	0.00	0.00
Population	16324728	11849589	16155885	11793499
Observations	11,160	53,624	11,148	53,449

Notes: CPS = Current Population Survey. The unit of observation is a respondent in state in a month-year. All variables except for population are binary; standard deviation of population is in parentheses. Data are weighted by CPS-provided survey weights at $t-1$ for exit and t for entrance. Individual characteristics are based on the values at $t-1$.