The Effects of Minimum Wage Policy on Self-Employment: Evidence from the Current Population Survey

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Abstract

This paper studies the effects of minimum wages on unincorporated self-employed workers in the US using the 1988-2020 Current Population Survey. Standard state-level difference-in-differences estimates of employment and earnings elasticities find that increasing the minimum wage tends to decrease employment but has little effect on hours worked or earnings in the year following the change. Instrumental variable estimates are consistent with these findings but show large earnings gains for these workers. Using a simple model of labor market search, I show that minimum wage increases are potentially welfare improving and the welfare effects can be identified by changes in self-employment. Given this model, I show that between 1988-2020 minimum wage changes have been welfare improving on average.

Keywords: Minimum Wage, Self-Employment, Gig Work, Spillovers
JEL Codes: J08, J21, J42

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1 Introduction

Research on the effects of minimum wages is vast and a topic of contentious debate.[1] However, little effort has been made in examining how these policies affect workers in alternative forms of employment, particularly those workers in jobs not covered by the minimum wage.[2] This, however, belies the fact that minimum wage changes may have spillover effects on these workers. To address this, in this paper I examine the effects of minimum wage policy on a subset of self-employed workers, the unincorporated self-employed. I do this by estimating standard minimum wage elasticities of employment and earnings for these workers using data from the Current Population Survey's (CPS) Annual Social and Economic Supplement (ASEC). Given these estimates, I then present a simple model of labor market search that provides insight into the welfare effects of the minimum wage given this self-employment option.

Using a standard difference-in-differences identification strategy, I find that increasing the minimum wage tends to decrease employment for these workers but has little effect on hours worked or earnings. For a 10 percent increase in the minimum wage, unincorporated self-employment declines by about 1.7 percent. Using an instrumental variable strategy, I find similar results for employment but a large and highly significant effect on earnings. For the same increase, employment declines by about 2.5 percent (though marginally insignificant with a P-value ≈ 0.11) while earnings increase by almost 6 percent (significant at the 1 percent level). Through the lens of a simple labor search model, these results are likely welfare improving, as raising the minimum wage increases the value of working in traditional labor sector. This reduces employment in the alternative sector, pushing workers into searching for traditional jobs, and raises the average wage of those who remain in self-employment.

[1]See for example: Neumark, Salas and Wascher (2014); Neumark and Shirley (2021); Allegretto, Dube and Reich (2011); Allegretto et al. (2017); Manning (2021)

[2]There is, however, much research on the effects of minimum wages in informal work in developing countries. See for example Comola and de Mello (2011); Magruder (2013); Hohberg and Lay (2015) who examine Indonesia and Pérez Pérez (2020); Alaniz, Gindling and Terrell (2011); Maloney and Mendez (2004); Gindling and Terrell (2007) who examine Latin America.

This work relates primarily to estimates of minimum wage effects on the self-employed in the US. Recent work by Glasner (2022) examines the effects of minimum wages on nonemployers—workers who file self-employment income but have no employees—in an effort to estimate spillover effects on workers not covered by the minimum wage. Examining the 2000-2018 period, the author finds that minimum wage policy had differential effects before and after the 2007-2009 Great Recession. Results for the effects on the level of self-employment and self-employment income receipts prior to the Great Recession are negative and significant, though not for the employment to population ratio, which is in contrast to the finding presented here. The author finds that post-2009 minimum wage changes had negligible to positive effects on nonemployers, with strong positive effects in the transportation industry, which they attribute to the expansion of Uber and Lyft during this period. These results indicate that increasing the minimum wage pushes workers into self-employment, particularly in areas where these ride-sharing platforms are available. However, the Nonemployer Statistics (NES)—the data used in their study—suffers from some complex drawbacks. Data from the NES may be biased due to spurious self-employment take-up associated with the Earned Income Tax Credit (EITC). Specifically, incentive effects of the tax credits from the EITC may induce tax filers to file self-employed income when they may not have otherwise.[3] If EITC expansion (2009-2013) was associated with higher misreporting, then estimates from these periods may show erroneous minimum wage related nonemployer growth if minimum wage changes were correlated with the expansion of the EITC. Given this, survey data may more accurately reflect the choices workers are making in response to minimum wage changes.

Comparatively, Bono-Lunn, Moulton and Scott (2021), using data from the US Census Bureau's American Community Survey (ACS), examine self-employed workers over the 2000-2017 period. The authors in this study focus on the flexibility of the self-employed occupation (as measured in the CPS's Work Schedules Supplement) and find that minimum wage

[3]See Collins et al. (2019); Garin, Jackson and Koustas (2022) for more details.

changes during this period pushed workers in highly flexible jobs out of self-employment. The authors also record estimates of the effects of the minimum wage on unincorporated self-employment in general and show negative (but insignificant) effects. The estimates in Bono-Lunn, Moulton and Scott (2021) largely support those produced here, with the additional 15 years of policy variation likely necessary to identify these effects.

The remainder of the paper proceeds as follows. Section 2 presents a simple labor search model where workers choose whether to work in the traditional labor market or in a self-employment or "gig"-sector to fix ideas about the structure of the setting. Section 3 describes the data employed for the empirical estimations. Section 4 outlines the empirical strategy and the robustness techniques used in the estimation procedure. Section 5 presents the results from the procedures laid out in Section 4. Section 6 concludes.

2 Theoretical Framework

In this section I present a simple model of labor market search from Rocheteau and Tasci (2008)[4] to develop expectations about how the minimum wage impacts employment in an alternative employment sector not covered by the minimum wage and how that can impact total welfare. In this framework workers make decisions about which sector to work in and this decision hinges crucially on their innate productivity in the alternative employment sector. The model presented, while simple, is of particular use because it allows for a welfare interpretation of the reduced form estimates of the effects of minimum wage changes on this alternative employment sector.

Consider a setting where unemployed workers and firms search to fill vacant jobs and match with endogenous rate $p(\theta)$ and $q(\theta)$, respectively. Here $\theta = v/u$ is labor market tightness, where v is the ratio of vacant job postings to workers and u is the the unemployment rate. So the rate at which workers and firms match with each other is directly related to the relative size of those two groups. In general, the "tighter" the labor market is—that is the

[4]See Pissarides (2000) for the seminal continuous time version of this type of model.

bigger θ—the easier it is for workers to find jobs and the harder it is for firms to find workers. This observation is important because increasing the wage, or the binding minimum wage in this case, reduces equilibrium market tightness.[5] This results provides a tension between increasing the value of a job to the worker through the wage and reducing the rate at which workers meet employers and receive that wage.

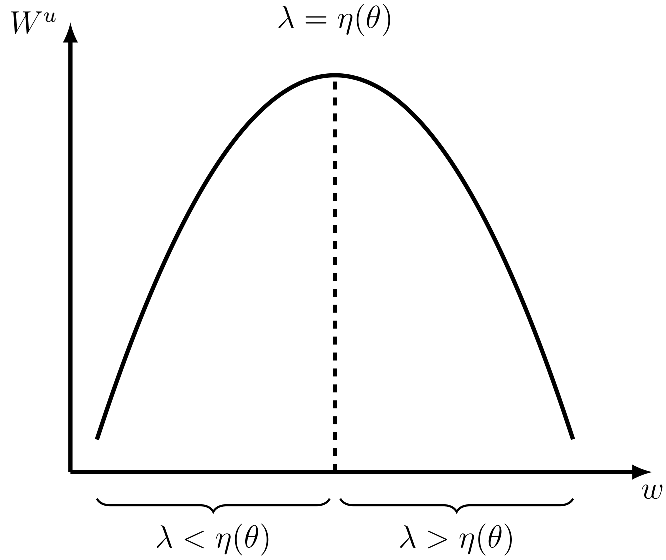
Now suppose that workers have the option to work in an alternative self-employment or "gig" employment sector. This work is characterized by a simple flow utility, e, that is drawn from a distribution, $G(e)$.[6] All workers have some innate e and so choose whether to work in the traditional labor market where they must search for work and face the possibility of losing their job and searching again, or working in this gig-sector. This decision generates a cutoff level of productivity in the gig sector that dictates whether workers choose to work there or not and is tied directly to this tension between the wage in the traditional sector and its effect on labor market tightness. Generally when wages are low, increases in the wage correspond to increases in the value of working in the traditional labor market. This continues until the benefits of the increasing wage are completely offset by the reduction in matches for unemployed workers. If workers observe some level of bargaining power over their wage (e.g. when wages are decided through Nash bargaining) then this point coincides with the point where the elasticity of the matching function with respect to unemployment is equal to workers' bargaining power. From Hosios (1990) we know that this condition, commonly referred to as the "Hosios condition", is welfare maximizing.

This relationship is outlined in Figure 1. Here the y-axis is the value of being unemployed in the traditional sector, W^u, the x-axis is the wage, λ represents the bargaining power of the worker, and $\eta(\theta)$ is the elasticity of the matching function with respect to unemployment. If $\lambda < \eta(\theta)$ then increases in the minimum wage dominate the reduction in matches due to

[5] See Proposition 7 in Rocheteau and Tasci (2008) for details.

[6] This setup is identical to the participation decision in Rocheteau and Tasci (2008). This structure is quite natural if we consider the distribution of gig-sector productivities equivalent to outside productivity in the participation decision. When leaving traditional employment workers have the option to choose the most productive use of their time, in this case it simply may be a gig-sector job rather than retirement.

Figure 1: Value of Unemployment and Minimum Wage Changes



the decline in θ, increasing W^u and as a result total welfare. This continues until $\lambda = \eta(\theta)$ where welfare is maximized, after which the relationship reverses. Since increasing the value of working in the traditional sector, through the value of unemployment, W^u, increases the cut-off level of gig-sector productivity for workers, the employment level for the gig-sector is declining in the left side of this figure and increasing in the right. Thus, if increases in the minimum wage reduces gig-sector employment then, through the lens of this model, we are on the left side of this graph and those reductions indicate that the economy is better off overall. This result also has implications for earnings in the gig-sector. If the cutoff level of gig-sector productivity is increasing, then in equilibrium those who remain in the gig-sector will have higher productivity, increasing the average earnings for the group.

3 Data

In order to analyze the effects of the minimum wage on the unincorporated self-employed this paper uses two primary data sources: the 1988 to 2020 Current Population Survey's (CPS) Annual Social and Economic Supplement (ASEC) and data on changes to minimum wage laws at the state level, including information on when states index their minimum wage to inflation. The Consumer Price Index retroactive series (CPI-U-RS, formerly known as the "research series"), from the US Bureau of Labor Statistics (BLS) inflation estimates, is also used to transform the minimum wage to 2018 dollars. Data used for estimating the propensity for a state to be bound by the federal minimum wage—used in the IV estimation procedure—also includes income per capita in 1969 from the Census's Historical Income Tables and the state ideological index from Berry et al. (2010).

3.1 Current Population Survey

The CPS is a monthly household survey conducted jointly by the Bureau of Labor Statistics (BLS) and the US Census. The original incarnation of the survey began shortly after the Great Depression as a way to measure various employment metrics for the US. It has been administered as the "Current Population Survey" since 1948 and is used in many national economic estimates, for example the monthly US unemployment estimates. The data used in this study begin in 1988, which is the earliest that unincorporated self-employed workers are identifiable in the ASEC, and end in 2020. The ASEC, which is added to the "basic monthly" CPS every March, provides information on the previous year's income for the self-employed, not otherwise available (for example in the Outgoing Rotation or "Earners Study"), and restricts the analysis in this paper to yearly effects (March-to-March).

Data are extracted from the Integrated Public Use Microdata Series (IPUMS), which organizes and harmonizes the data across years (Flood et al. (2021)). These data include information on the respondents age, hours worked, the type of employment the respondent engages in in the reference period, their education, current employment status, labor force status, industry and occupation, marital status, race, sex, the state of the household's residence, as well as information on survey characteristics, such as the individual's current month in the sample rotation and individual identifiers. The general framework for the survey provides that respondents appear in the survey for four consecutive months, are not sampled for the next eight months, and then are sampled again for four more consecutive months.

Table 1: Unincorporated Self-Employed Worker Characteristics, 1988-2020

Observations (total)		326,086
Population (yearly average)		9,328,977
Age (average)		44
	16-25	0.05
	26-35	0.19
	36-45	0.28
	46-55	0.28
	56-65	0.20
Education (average years)		13.73
	<High school	0.15
	≥High school	0.85
Male		0.62
Married		0.68
Wages ($2018)		
	Mean	$23.09
	Median	$13.55

Source: Current Population Survey, Annual Social and Economic Supplement (1988-2020).
Note: The table reports summary statistics for unincorporated self-employed workers from the 1988-2020 CPS ASEC. Numbers represents the proportion of workers with the row characteristic unless noted otherwise.

Put simply, this means a respondent is observed in the same four consecutive months across two consecutive years. Given this framework, I only observe individuals twice and only those individuals with a month-in-sample that overlaps the month of March (though these individuals may be from any of the four month-in-sample cohort pairs, e.g., 1 and 5, 2 and 6, etc.). Table 1 displays worker characteristics for the unincorporated self-employed over this period. These workers represent approximately 9 million workers each year on average and are typically older, married, men, with at least a high school education.

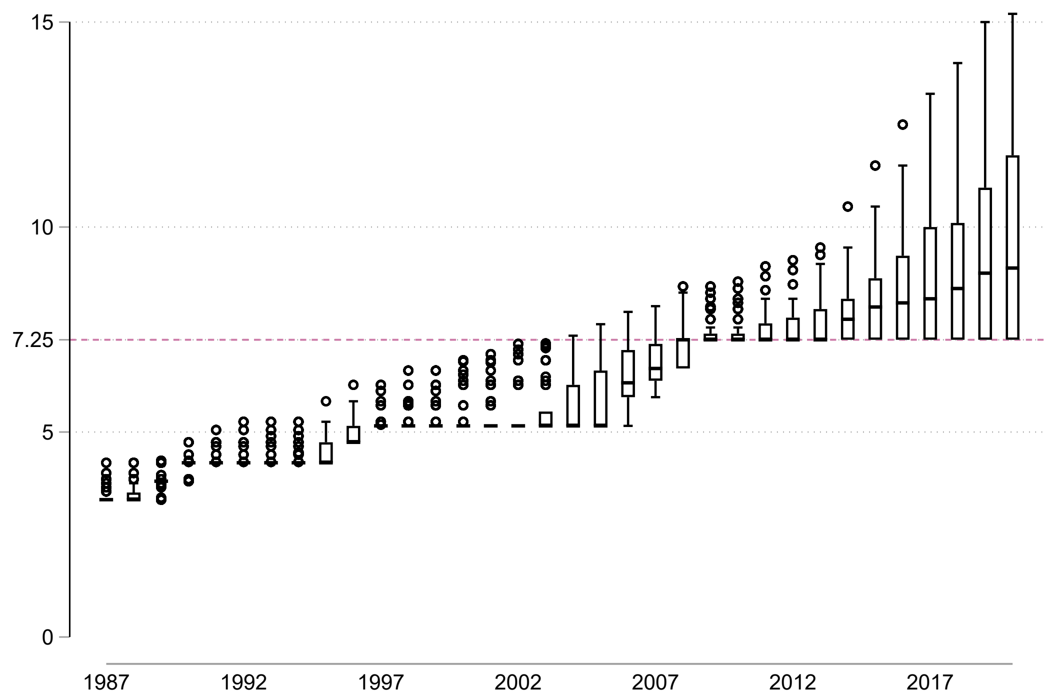
Use of the CPS data for examining unincorporated self-employed workers does have several drawbacks. First, county-level data is only available for about 45 percent of respondents due to Census confidentiality constraints (Flood et al., 2021). This makes a county-level analysis similar to those in Allegretto, Dube and Reich (2011); Jha, Neumark and Rodriguez-Lopez (2022) infeasible, restricting the analysis to state-level variation. Similarly, since individuals are not identifiable below the state level individuals that are treated by sub-state minimum wage policies are also not identifiable. This may bias estimates of minimum wage policies if individuals within cities receive treatment but are categorized as untreated at the state level.

3.2 Minimum Wage Laws

Minimum wage changes recorded at the monthly level are derived from Neumark and Yen (2021). When states have more than one minimum wage for varying firm characteristics (e.g., a separate lower minimum wage for firms that supply health insurance) the larger of the alternatives is used. As mentioned above, these data are limited in that many workers in the self-employment or alternative labor market work in cities that have separate, and often larger, minimum wages. Due to the sampling level of the public use data these individuals are not distinguishable from those outside of major cities and so the prevailing state minimum wage is used. This may limit the measurement of minimum wage effects if changes in employment are due to unobserved city minimum wage changes. For example, we may see disemployment effects for states where nominal state minimum wages are constant, but city minimum wages are increasing. In this case, if minimum wages are declining in real terms then estimates of the effect of minimum wage changes on employment in either type of work would show a positive relationship when in fact the relationship is negative.

Figure 2 displays the range of effective nominal state-level minimum wages by year from 1988 to 2020. This figure highlights the large variation in minimum wages over the period of this study. Figure 3 gives these changes in real terms ($2018). Figure 4 displays the number of effective state-level minimum wage changes over the same period. Together these figures highlight both the frequency and magnitude of variation that allows for identification of minimum wage effects in this study. Of the 51 sub-federal jurisdictions covered in these data (50 states and Washington DC), all experience an effective minimum wage change at least seven times due to federal minimum wage increases. A further 36 have jurisdiction-level

Figure 2: The distribution of effective nominal minimum wages by year



Source: Data is from Neumark and Yen (2021) and the NCSL.

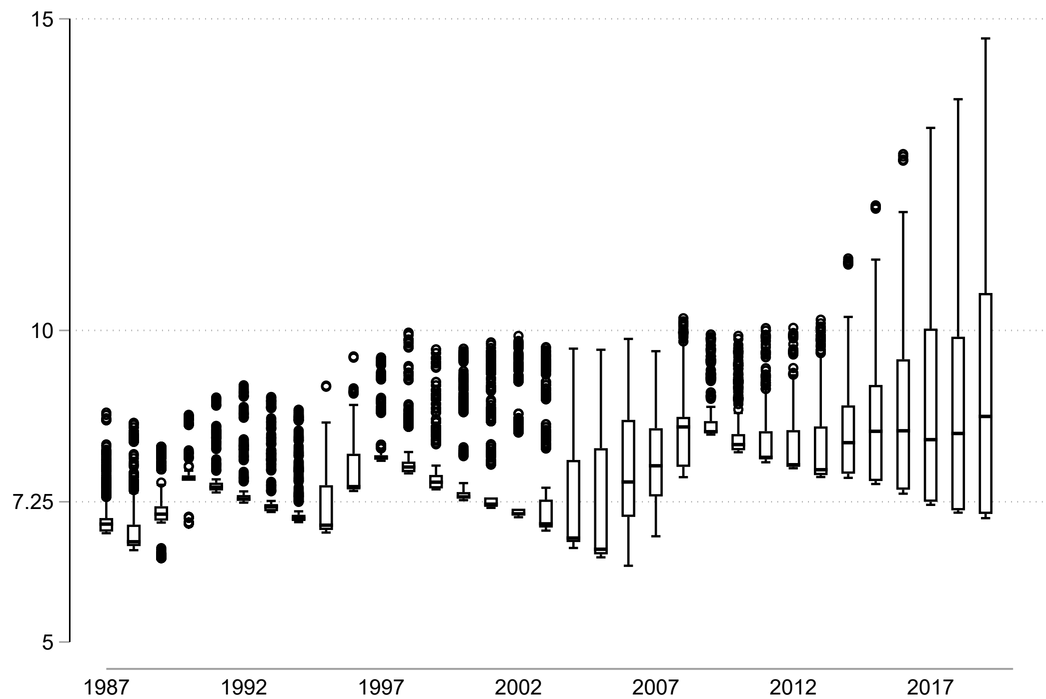
Note: Boxes represent the interquartile range of effective nominal minimum wages for all states in a given year. Whisker ends represent the next largest observation over 1.5 times the difference between the maximum and minimum value of the interquartile range. Subsequent empty circles are observations outside of these values. The lowest effective minimum wage is the federal minimum wage for any given year. Represented by a red dashed line, 7.25 is the effective lowest nominal minimum wage during the period of this study and has been the federal minimum wage since July of 2009.

minimum wage increases over this period, with 15 states only observing federal increases.[7] Overall there are 706 effective minimum wage changes over this period, with yearly variation comprised of 350 individual state-by-year events. This difference is due to the annual nature of the observations, so years with multiple effective minimum wage changes are only counted as a single increase.

Information on laws that index the minimum wage to inflation or commit states to multi-

[7]States that only obverse the federal minimum include: Alabama, Georgia, Idaho, Indiana, Kansas, Louisiana, Mississippi, Oklahoma, South Carolina, Tennessee, Texas, Utah, Virginia, Wisconsin, and Wyoming.

Figure 3: The distribution of effective real minimum wages (2018$) by year

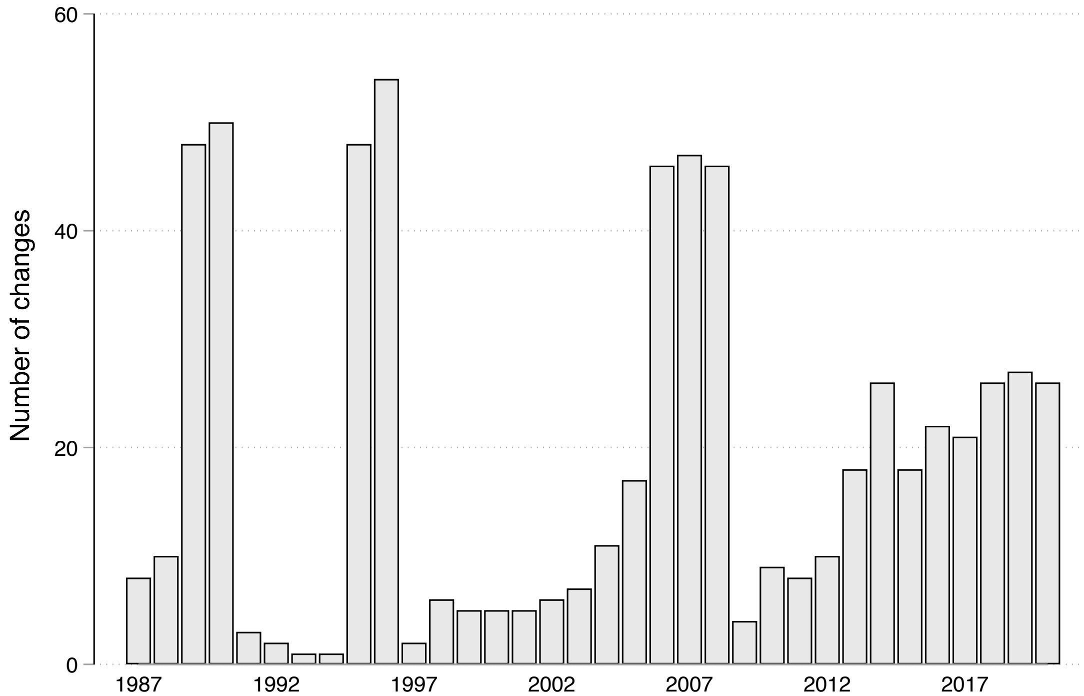


Source: Data is from Neumark and Yen (2021) and the NCSL.

Note: Boxes represent the interquartile range of effective real minimum wages (2018$) for all states in a given year. Whisker ends represent the next largest observation over 1.5 times the difference between the maximum and minimum value of the interquartile range. Subsequent empty circles are observations outside of these values.

year minimum wage increases are from Brummund and Strain (2020) and are extended to present day with the National Conference of State Legislatures (NCSL) and official state legislative data. Table 2 displays the month and year for states that implement laws that either: index the minimum wage to inflation outright or increase the minimum wage by some multi-year scheme and then subsequently index it to inflation after.

Figure 4: Number of Effective Minimum Wage Changes by Year



Source: Data is from Neumark and Yen (2021) and the National Conference of State Legislatures.

4 Empirical Strategy

4.1 Difference-in-Differences

Baseline estimates implement a standard two-way fixed effects difference-in-differences design. Across-state variation in minimum wage policies is used to estimate the spillover effects of the minimum wage on unincorporated self-employed workers. Individual-level data from the CPS ASEC requires that these observations are only observed in March of each year and so estimates are these effects year-over-year. The sample is restricted to those 16-64 years old in all estimates, and to those with recorded business income for wage estimates. The identifying assumption of this approach is that, in the absence of a minimum wage increase, both treated and untreated states would have evolved similarly in their outcomes. This

Table 2: State inflation-indexing minimum wage policies

State	Month	Year
Alaska	February	2015
Arizona	January	2007
California	January	2017
Colorado	January	2007
Connecticut	October	2019
District of Columbia	July	2014
Florida	May	2005
Maine	November	2016
Michigan	December	2018
Missouri	January	2007
Montana	January	2007
Nevada	July	2007
New Jersey	January	2014
Ohio	January	2007
Oregon	January	2003
South Dakota	January	2015
Vermont	January	2007
Washington	January	1999

Source: Brummund and Strain (2020), NCSL, and official state legislative data.
Note: The table reports the month and year in which states experiences an increase in the minimum wage due to a law that includes an indexing provision regardless of whether the indexing provision takes effect immediately or in later years.

assumption has raised some concerns by economists (e.g., Dube, Lester and Reich (2010); Allegretto, Dube and Reich (2011); Allegretto et al. (2017); Meer and West (2016)) and so procedures that include jurisdiction time trends are also presented.

Equation (1) represents the two-way fixed effects difference-in-differences design:

$$y_{ist} = \alpha + \beta \ln(\text{mw})_{st} + \mathbf{X}_{ist}\gamma + \tau_t + \delta_s + \varepsilon_{ist}, \quad (1)$$

where y_{ist} is the outcome of interest for individual i, in year t, in state s; $\ln(\text{mw})_{st}$ is the natural log of the minimum wage for state s, in year t; $\mathbf{X}_{ist}$ are controls for age, age squared, number of years of schooling, and marital status; and τ_t, and δ_s are time and state fixed effects, respectively. Outcomes captured in y include an individual's employment status represented by an indicator, the natural log of the average hourly business income reported by the individual for the reference period, and the natural log of average hours worked over the reference period. Here the parameter of interest, β, represents the average effect of the treatment on the treated (ATT) of an percent change in the minimum wage. The individual employment indicator is defined as one if individual i is classified as unincorporated self-employed during the reference period and is zero if the individual is either employed in traditional employment (responded that they "worked for wages or salary"), is unemployed, is incorporated self-employed, or if the individual is not in the labor force.

As has become common with the standard two-way fixed effects estimation methods in the minimum wage literature (Dube, Lester and Reich, 2010; Allegretto, Dube and Reich, 2011; Meer and West, 2016, etc.), specifications that control for region and state linear time trends are included. Equation (2) represents the two-way fixed effects difference-in-differences design with a census-division-by-year linear time trend:

$$y_{ist} = \alpha + \beta \ln(\text{mw})_{st} + \mathbf{X}_{ist}\gamma + \psi_d \times t + \tau_t + \delta_s + \varepsilon_{ist}, \tag{2}$$

where ψ_d is an indicator for census division d. Equation (3) represents the two-way fixed effects difference-in-differences design with state-linear time trends included:

$$y_{ist} = \alpha + \beta \ln(\text{mw})_{st} + \mathbf{X}_{ist}\gamma + \delta_s \times t + \tau_t + \delta_s + \varepsilon_{ist}. \tag{3}$$

The idea motivating Equations 2 and 3 is to control for unobserved differences in the trends in the outcomes of interest that are potentially correlated with the treatment for those areas. If treatment is correlated with the outcome of interest, say for instance policy makers increase the minimum wage when the economy is doing well and employment is rising, then controlling for region- or state-linear time trends should help control for these differential time paths. However, as Meer and West (2016) points out, if the effect of the policy is to change the growth rate of the outcome, then these controls may attenuate estimates of the true effect. In that regard I also include both a single "lead" model, where I include the log of the minimum wage for the next year in the current year's estimation, and a "lead-lag" model where one lead and two lags of the log of the minimum wage are included.[8] If states are differentially increasing their minimum wage due to trends in the outcome variable then these leads would be correlated with those outcomes.

Equation (4) represents this lead-lag specification. Model differences are in the restriction of T:

$$y_{ist} = \alpha + \sum_{r}^{T} \beta_r \ln(\text{mw})_{st-r} + \mathbf{X}_{ist}\gamma + \tau_t + \delta_s + \varepsilon_{ist}, \tag{4}$$

where $r = -1$ and $T = 0$ in the single lead specification and $r = -1$ and $T = 2$ in the lead-lag specification.[9]

All models also include controls for states that index the minimum wage to inflation, for whether an individual is in full or part-time work, and for the state-level unemployment rate for the month of March in the specified year.

4.2 Instrumental Variable Approach

Despite the large literature, standard estimation methods still suffer from endogeneity in this setting. States often, and sometimes mechanically, make minimum wage policy choices in direct response to economic conditions. For instance, states that index their minimum wages often have mechanism in their policies to cancel an increase if inflation or unemployment is

[8]The authors in Meer and West (2016) use a specification that includes two leads and three lags while in Allegretto, Dube and Reich (2011) the authors use 16 leads and eight lags in their respective lead-lag specification, however the data used in these studies is at, or aggregated to, the quarterly level, so the one-lead-two-lag model presented here should be relatively similar to both in terms of lags, and closer to Meer and West (2016) in terms of leads, given the annual nature of the CPS ASEC.

[9]This specification is comparable to Equation (2) in the working paper version of Meer and West (2016) and Equation (5) in Allegretto, Dube and Reich (2011).

too high.[10] Given that a states decision to increase its minimum wage is endogenous to the current conditions in that state I employ the instrumental variable (IV) approach developed in Baskaya and Rubinstein (2012) to estimate the effect of changes in the minimum wage on employment and earnings. This approach utilizes the differential conditions, such as income per capita, average political ideology, and previous minimum wage policy, in states before the period observed to estimate a state's propensity to be bound by the federal minimum wage in the current period. This propensity, multiplied by the natural log of the real federal minimum wage, is then used as an instrument for the current log minimum wage at the state level.

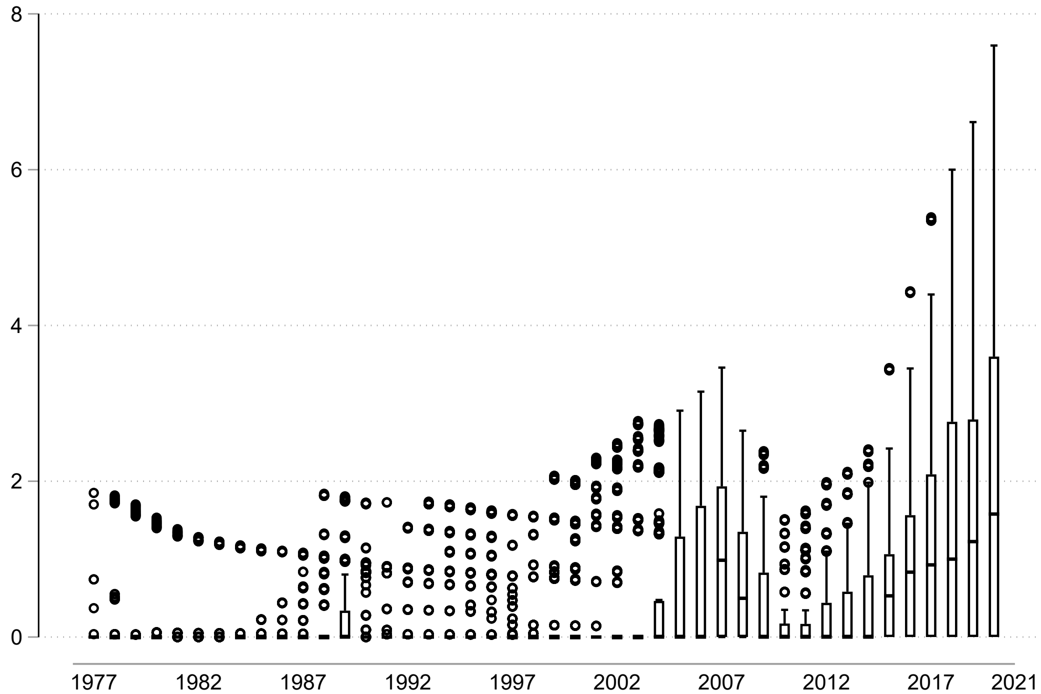
The idea here is that states are differentially affected by changes in the federal minimum wage based on the pre-existing state characteristics that induce states to increase their minimum wages over the federal level. Federal minimum wage changes, compared to state-level changes, are plausibly exogenous to those same states (Card, 1992). Then changes in federal minimum wages, for states where the federal minimum wage binds, do not suffer from this endogeneity issue. Additionally, these pre-characteristics likely do not directly impact states contemporaneous choices to increase their minimum wage. So by instrumenting for the minimum wage with the relative bindingness of the federal minimum wage we can isolate states' own likelihood of increasing their minimum wage, separate from contemporaneous shocks to the location. Figure 5 displays the relative scale of these federal-to-state minimum wage differences over the study while including a 10 year pre-period.

Equation (5) represents this the first stage of the two-stage least-squares (2SLS) approach.

$$MW_{st} = \zeta + \theta F_s FMW_t + \mathbf{X}_{st}\mu + \phi_s + \phi_s t + \phi_t + \epsilon_{st}, \tag{5}$$

[10]For instance, Michigan's "Improved Workforce Opportunity Wage Act 337 of 2018" reads: "An increase in the minimum hourly wage rate as prescribed in subsection (1) does not take effect if the unemployment rate for this state, as determined by the Bureau of Labor Statistics, United States Department of Labor, is 8.5% or greater for the calendar year preceding the calendar year of the prescribed increase." (Hildenbrand, 2019).

Figure 5: The Distribution of Differences in Effective Minimum Wages and the Federal Minimum Wage (2018$) by Year



Source: Data is from Neumark and Yen (2021) and the NCSL.

Note: Boxes represent the interquartile range of the difference between the effective real minimum wages (2018$) for all states in a given year and the federal minimum wage (also in 2018$). Whisker ends represent the next largest observation over 1.5 times the difference between the maximum and minimum value of the interquartile range. Subsequent empty circles are observations outside of these values.

where ζ is a constant, F_s is the time-invariant state propensity (Baskaya and Rubinstein (2012) refer to this as the "traditional propensity") to be bound by the federal minimum wage, FMW_t is the federal minimum wage in year t, $\mathbf{X}_{st}$ are state specific controls included in Equation (1), ϕ_s and ϕ_t are state and year specific control, and ϵ_{st} is the error term. In this specification the state-linear time trend is also included. The second stage is equivalent to Equation (3), where $\ln(mw)$ is the endogenous regressor. One thing to note here is that data from Berry et al. (2010) for the state ideology index do not include the District of Columbia and so the jurisdiction is omitted from all IV estimates.

The propensity for a state to be bound by the federal minimum wage is estimated through a logistic regression. Equation (6) represents this approach:

$$F_{st} = \Lambda \left(\mathbf{Z}_s \pi + \xi_{st} \right), \tag{6}$$

where Λ is the logistic cumulative distribution function, F_{st} is an indicator for if state s is bound by the federal minimum wage in year t, $\mathbf{Z}_s$ are time-invariant characteristics such as income per capita in 1969, the average value over the 1960-1968 state-ideology index from Berry et al. (2010), the proportion of years between 1968-1976 where the federal minimum wage was binding within the state, and ξ_{st} is an error term. Equation (6) is then used to predict the values of F_s for each state in the sample.

5 Results

From Section 2 we know that the effects of the minimum wage on employment and wages may differ depending on where we reside in Figure 1. If wages are sufficiently low in the traditional sector, that increases in the binding minimum wage fully compensate for the reduction in the worker matching rate, then employment should decline and average wages in the self-employment or gig-sector should increase. On the other hand if wages are comparatively higher in the traditional sector, then we may see that minimum wage changes reduce the value of working in these jobs and workers shift to working in self-employment. To shed light on this uncertainty I estimate a battery of tests using specifications outlined in Section 4.

5.1 Employment

Table 3 provides results for estimates of Equations (1)-(4) for extensive margin employment. Coefficients represent the effect of a percent increase in the minimum wage on the probability an individual is an unincorporated self-employed worker. The elasticity of the minimum wage with respect to unincorporated self-employment, η, along with standard errors are below.

All specifications control for individual demographics including age, age squared, marital status, education level, as well as state and year fixed effects, the state unemployment rate for March in the year reported, full or part time work status, and whether the state indexes the minimum wage to inflation. Column 1 represents a baseline estimate of the difference-in-differences design. Results indicate no spillover effects of the minimum wage on the unincorporated self-employed. However, these results rely on the assumption that states' employment rates, absent the policy change, would have evolved similarly over time. As previous studies have argued this may in fact not be the case. Columns 2 and 3 address this issue by including census-division and state linear time trends, respectively. In column 2 we see that this flips the sign of the estimate to negative and increases it in magnitude dramatically. Estimates in column 3 go even further, with effects now significant at the five percent level.

These results imply that, when appropriately controlling for differences in the evolution of state employment outcomes, minimum wage increases decrease the prevalence of unincorporated self-employment in the following year. These results are consistent with being on the left side of Figure 1. In this scenario, increasing the minimum wage pushes workers to traditional employment. It's important to note that this does not necessarily imply that these workers increase traditional employment, they may simply be searching while unemployed.

One way to test for the presence of differences in underlying paths for our outcomes is by including a leading value for the variable of interest. If policy changes are tied to the evolution of our outcome variable, then leading values of our variable of interest should be correlated with this outcome (Allegretto, Dube and Reich, 2011; Meer and West, 2016). Column 4 display these results. As we can see, the coefficient on the leading value for the log of the minimum wage is positive but insignificant (t-statistic of 1.15). Furthermore, the estimated coefficient on the contemporaneous effect of the minimum wage (effect of changes over one year) is still negative, but much smaller and now insignificant. This result is peculiar because if the state-linear time trend is in fact controlling for differences in the evolution of

Table 3: Effects of Minimum Wages on Unincorporated Self-Employment, 1988–2020

	(OLS)	(OLS)	(OLS)	(OLS)	(OLS)	(IV)
$\ln(\text{MW}_t)$	0.0009	-0.0029	-0.0082**	-0.0022	-0.0086*	-0.0118
	(0.0050)	(0.0044)	(0.0040)	(0.0038)	(0.0048)	(0.0073)
$\ln(\text{MW}_{t+1})$				0.0059	0.0071	
				(0.0052)	(0.0053)	
$\ln(\text{MW}_{t-1})$					0.0033	
					(0.0055)	
$\ln(\text{MW}_{t-2})$					0.0045	
					(0.0053)	
Implied η	0.018	-0.059	-0.168**	-0.046	-0.178*	-0.245
η SE	(0.103)	(0.091)	(0.082)	(0.077)	(0.100)	(0.151)
Observations	2,692,499	2,692,499	2,692,499	2,627,021	2,627,021	2,660,656
Controls	Y	Y	Y	Y	Y	Y
Year FE	Y	Y	Y	Y	Y	Y
State FE	Y	Y	Y	Y	Y	Y
Census × Year	N	Y	N	N	N	N
State × Year	N	N	Y	N	N	Y

Standard errors in parentheses

* $p < .1$, ** $p < .05$, *** $p < .01$

Note: The table reports the effect of a minimum wage increase on unincorporated self-employment based on Equations (1)-(4). Observations are at the individual level and so all specification control for individual race, marital status, age, age squared, years of schooling, year, state, the state unemployment rate, and whether the state indexes the minimum wage to inflation. Additionally, estimate standard errors are clustered at the state level. Estimates are weighted with ASEC weights. The first-stage F-statistic for the IV estimates in column 6 is 56.12.

our outcomes in column 3, then we would expect this leading value to be correlated with the outcome and so statistically different from zero. Regardless, an important distinction here (and in column 5) is that we have similarly negative coefficients (even more so for column 5) without the need for the state-linear trends, implying these leads and lags incorporate important information for the evolution of the outcome that is captured in the state trends. Column 5 highlights this idea, where now the cofficient on the contemporaneous minimum wage is almost identical to that from the state linear time trend model in column 3 (-0.0086) and significant at the 10 percent level.

Column 6 reports the IV estimates where the first-stage equation is represented by Equation (5) and the second stage is equivalent to Equation (3). Here we see estimates similar to those in column 3, with an implied elasticity of -0.274, significant at the 10 percent level. The unreported F-stat for first-stage estimate is 56.12, implying the instrument is highly relevant. These estimates are almost twice as large as those in column 3, indicating that these simple state-linear time trends may not be sufficient to correctly control for differences in state outcomes due to contemporaneous shocks.

5.2 Hours

Table 4 displays the same battery of estimates as before, but where the log of annual hours worked is the outcome and so represent elasticities of the minimum wage with respect to intensive margin employment for unincorporated self-employed workers. Controls follow the specifications above, but the sample is restricted to those with positive hours in the survey. Across baseline specification and those including jurisdiction-specific trends, estimates are small and insignificant, and all estimates are positive. However, when leads and lags are included in the specification the coefficients turn negative, though insignificant. Estimates using the IV specification in column 6 show negative results with a P-value of 0.101, marginally insignificant. These results imply that contemporary state-specific shocks might be more important for hours specifically. However, the results in column 6 may be less plausible if we think that self-employed workers, conditional on having positive hours of employment, have more autonomy in choosing the number of hours they work. Since we're already controlling for whether a worker is part-time or full-time, it seems unlikely that firms hiring these workers would be able to adjust over this margin and so changes to average hours are more likely indicative of workers' choices.

5.3 Earnings

Table 5 displays results for the effect of changes in the minimum wage on earnings, where the outcome of interest is the log of the hourly self-employment income. Baseline estimates

Table 4: Effects of Minimum Wages on Annual Hours Worked, 1988–2021

	(OLS)	(OLS)	(OLS)	(OLS)	(OLS)	(IV)
$\ln(\mathrm{MW}_t)$	0.0122	0.0184	0.0097	-0.0175	-0.0144	-0.1124
	(0.0249)	(0.0269)	(0.0291)	(0.0359)	(0.0661)	(0.0671)
$\ln(\mathrm{MW}_{t+1})$				0.0577	0.0571	
				(0.0402)	(0.0461)	
$\ln(\mathrm{MW}_{t-1})$					-0.0363	
					(0.0691)	
$\ln(\mathrm{MW}_{t-2})$					0.0416	
					(0.0408)	
Observations	114,236	114,236	114,236	112,075	112,075	113,290
Controls	Y	Y	Y	Y	Y	Y
Year FE	Y	Y	Y	Y	Y	Y
State FE	Y	Y	Y	Y	Y	Y
Census × Year	N	Y	N	N	N	N
State × Year	N	N	Y	N	N	Y

Standard errors in parentheses

* $p < .1$, ** $p < .05$, *** $p < .01$

Note: The table reports the effect of a minimum wage increase on the log of annual hours worked for unincorporated self-employed workers, based on Equations (1)-(4). Observations are at the individual level and so all specification control for individual race, marital status, age, age squared, years of schooling, year, state, the state unemployment rate, and whether the state indexes the minimum wage to inflation. Additionally, estimate standard errors are clustered at the state level. Estimates are weighted with ASEC weights.

in column 1 indicate that minimum wage changes decrease income for unincorporated self-employed workers, with an elasticity of -0.3084 and significant at the five percent level. Estimates with the inclusion of state-linear time trends see the effects decline to some degree with the elasticity dropping to -0.1058 and no longer significant. However, in the specification with an included lead (column 4), there is little indication that the evolution of state outcomes differ dramatically. Evidence from the specification in column 5 further highlights the delayed effects of minimum wage changes on the wages for these workers. Coefficient on the two-year lag for the specification in column 5 shows large and significant effects with an elasticity of the minimum wage to earnings for these workers of -0.4793, and significant at the one percent level. Results from the IV estimates in column 6 show large, positive, and

Table 5: Effects of Minimum wages on Unincorporated Self-Employment Wages, 1988-2021

	(OLS)	(OLS)	(OLS)	(OLS)	(OLS)	(IV)
$\ln(\text{MW}_t)$	-0.3084**	-0.0859	-0.1058	-0.2791	0.0122	0.5914***
	(0.1338)	(0.1362)	(0.1559)	(0.2249)	(0.2450)	(0.2126)
$\ln(\text{MW}_{t+1})$				-0.1415	-0.2098	
				(0.1696)	(0.1693)	
$\ln(\text{MW}_{t-1})$					-0.0329	
					(0.1586)	
$\ln(\text{MW}_{t-2})$					-0.4793***	
					(0.1512)	
Observations	94,016	94,016	94,016	90,590	90,590	93,151
Controls	Y	Y	Y	Y	Y	Y
Year FE	Y	Y	Y	Y	Y	Y
State FE	Y	Y	Y	Y	Y	Y
Census × Year	N	Y	N	N	N	N
State × Year	N	N	Y	N	N	Y

Standard errors in parentheses

* $p < .1$, ** $p < .05$, *** $p < .01$

Note: The table reports the effect of an increase in the log of the minimum wage on log hourly earnings for unincorporated self-employed workers, based on Equations (1)-(4). Observations are at the individual level and so all specification control for individual race, marital status, age, age squared, years of schooling, year, state, the state unemployment rate, and whether the state indexes the minimum wage to inflation. Additionally, estimate standard errors are clustered at the state level. Estimates are weighted with ASEC weights.

significant effects on earnings, with an elasticity of 0.5914. This result implies that state-specific contemporary shocks likely bias estimates of the effects on earnings in the previous specifications. The IV results here, coupled with those for employment, are also consistent with the model presented in Section 2.

6 Conclusion

This paper provides estimates of minimum wage elasticities for employment, hours, and earnings for unincorporated self-employed workers. These results are derived from standard difference-in-differences estimation procedures and are supported by a battery of robustness estimates. These estimates imply that increasing the minimum wage by 10 percent leads to a 1.7 to 2.5 percent decline in unincorporated self-employment. Estimates of the effects of the minimum wage on earnings, for the same increase, imply and effect from a negligible amount to almost 6 percent increase. These results are consistent with the theme of the pre-Great Recession results from Glasner (2022) and the of those in Bono-Lunn, Moulton and Scott (2021), though do not speak to their estimates of the post recession period or for workers in the transportation industry or for the relative flexibility of the worker's occupation.

Through the lens of the labor search model presented in Section 2, these results imply that between 1988 and 2020 state-level minimum wage policy has been welfare improving. That is, increases in the binding minimum wage have not been completely offset by a decline in the worker matching rate in the traditional labor sector. This effect leads to workers finding opportunities in the traditional labor sector more attractive, inducing them to switch and search, and leading to a decline in the prevalence of unincorporated self-employment. Furthermore, as workers shift from the self-employment or gig-sector to the traditional sector, the average productivity of those who remain increase, which is reflected in an increase in average earnings. Overall these results make it clear that further research on the spillover effects of these types of policies on self-employed workers and alternative workers is warranted.

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