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EFFECTS OF THE MINIMUM WAGE ON EMPLOYMENT DYNAMICS

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Effects of the Minimum Wage on Employment Dynamics  
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**ABSTRACT**

The voluminous literature on minimum wages offers little consensus on the extent to which a wage floor impacts employment. For both theoretical and econometric reasons, we argue that the effect of the minimum wage should be more apparent in new employment growth than in employment levels. In addition, we conduct a simulation showing that the common practice of including state-specific time trends will attenuate the measured effects of the minimum wage on employment if the true effect is in fact on the rate of job growth. Using three separate state panels of administrative employment data, we find that the minimum wage reduces net job growth, primarily through its effect on job creation by expanding establishments. These effects are most pronounced for younger workers and in industries with a higher proportion of low-wage workers.

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# 1 Introduction

The question of how a minimum wage affects employment remains one of the most widely studied – and most controversial – topics in labor economics. During the recent recession, the employment rate for younger or low-skilled workers (who are more likely to be paid wages at or close to the minimum) worsened disproportionately, and following the recession the unemployment gap based on education remains large (Hoynes et al., 2012; United States Bureau of Labor Statistics, 2013). A more conclusive understanding is needed of the effects of the minimum wage if it is to be evaluated alongside alternative policy instruments as a method of increasing the standard of living for low-income households. Moreover, in recent years a number of states have indexed their minimum wages to adjust for inflation, and – despite growing state and federal pressure to continue this practice – there is little evidence on how inflation indexing might alter any effect of the minimum wage.

To date, nearly all studies of the minimum wage and employment have focused on how a legal wage floor affects the employment *level*, either for the entire labor force or a specific employee subgroup (e.g. teenagers or food service workers). We argue that, in a Diamond (1981)-type worker search and matching framework, an effect of the minimum wage should be more apparent in employment *dynamics* – that is, in the actual creation of new jobs by expanding establishments and the destruction of existing jobs by contracting establishments. Diamond argues that transitions to a new employment steady state may be slow, such that it may take some time for any effect of the policy to be visible in the employment level.

In addition to this theoretical foundation, there are several empirical reasons for why effects of the minimum wage should be detected more clearly in job growth than in employment levels. A critical factor is that, unlike many treatments studied in the program evaluation literature, the identifying variation consists of relatively small and temporary changes in a state’s real minimum wage, which are soon dissipated by inflation and increases by other states; we confirm this empirically in Section 2.2. As a result, there is often insufficient time for even sizable effects on the rate of job growth to be reflected in the level of employment.

We also demonstrate that a common practice in this literature – the inclusion of state-specific time trends as a control – will attenuate estimates of how the minimum wage affects the employment level. Specifically, we perform a simulation exercise which shows that if the true effect of the minimum wage is indeed in the growth rate of new employment, then even real causal effects on the level of employment can be attenuated to be statistically indistinguishable from zero. In contrast, the inclusion of state time trends does not induce a similar attenuation problem for estimates of the direct effect on growth.

To implement our analysis, we use a state panel difference-in-differences approach in which we allow for state fixed effects, region-by-time period effects, and state-specific time trends. We examine effects of the minimum wage on employment dynamics and levels using three administrative data sets: the Business Dynamics Statistics (BDS), the Quarterly Census of Employment and Wages (QCEW), and the Quarterly Workforce Indicators (QWI). These data sets vary in their strengths and weaknesses (discussed at length below), but together they encompass a long (1975-2012) panel of aggregate employment metrics for the population of employers in the United States.

Our findings are consistent across all three data sets, indicating that job growth declines significantly in response to increases in the minimum wage. However, we do not find a corresponding reduction in the *level* of employment, particularly in specifications that include state-specific time trends. For reasons discussed below and illustrated in our simulation exercise, we view this null effect for the employment level as neither surprising nor likely to be an accurate reflection of the effect of the minimum wage. Additionally, we decompose the negative net effect on job growth and find that it is primarily driven by a reduction in job creation by expanding establishments, rather than by an increase in job destruction by contracting establishments. These are among the more policy-relevant outcomes related to employment: the change in the number of jobs in the economy, rather than, for instance, the turnover of individuals within existing jobs.

We perform numerous robustness checks to test the validity of our identification strategy, which requires that the pre-existing time-paths of outcomes for states which increase their minimum wages do not differ in an off-trend manner relative to states that do not see an increase. We evaluate this possibility in Section 5.4 by including leads of the minimum wage into our specifications; if increases in the minimum wage showed a negative effect on employment dynamics *before* their implementation, this would suggest that the results are being driven by unobserved trends. This is not the case. Indeed, for our results to be driven by confounders, one would have to believe that increases in the minimum wage were systematically correlated with unobserved shocks to that state in the same time period, but not other states in that region, unrelated to existing state-specific time trends, and that these shocks are not reflected in measures of state-specific demographics or business cycles. Our results are additionally robust to varying the specifications to account for finer spacial and temporal controls, the recent financial crisis, and inflation indexing of state minimum wages, as well as across different panel lengths and time periods.

Finally, we find that the effect on job growth is concentrated in lower-wage industries

and among younger workers. Much of the existing literature focuses on these groups, though it is important to note that the minimum wage could affect other industries or elsewhere in the age distribution (e.g. [Neumark et al., 2004](#)). We report disaggregated results for all two-digit industry codes and for all age groups available in our data, and find that the effects of the minimum wage tend to be concentrated in the more likely groups.

The primary implication of our study is that the minimum wage does affect employment through a particular mechanism. This is important for normative analysis in theoretical models (e.g. [Lee and Saez, 2012](#)) and for policymakers weighing the tradeoffs between the increased wage for minimum wage earners and the potential reduction in hiring and employment. Moreover, we reconcile the tension between the expected theoretical effect of the minimum wage and the estimated null effect found by some researchers. We show that – given the nature of minimum wage policies – finding no effect on the level of employment should be unsurprising, particularly when state-specific time trends are included. In contrast, the minimum wage significantly reduces net job growth, and, in particular, new job creation.

This article proceeds as follows: in [Section 2](#) we provide a brief review of the literature on the employment effects of the minimum wage and build our case for examining job growth directly rather than the employment level. [Section 3](#) presents our econometric model and [Section 4](#) describes the data used in our study. We discuss our empirical results and their robustness in [Section 5](#), and conclude in [Section 6](#).

## 2 Theoretical and Empirical Framework

### 2.1 Challenges in Estimating Employment Effects

The economic literature on minimum wages is longstanding and vast. [Neumark and Wascher \(2008\)](#) provide an in-depth review of the field, which continues to be characterized by disagreement on how a minimum wage affects employment. The majority of recent studies, following [Card and Krueger \(1994\)](#), use difference-in-differences comparisons to evaluate the effect of these policies on employment levels. It is important to note that these models test whether there is a discrete change in the level of employment before and after a state changes its minimum wage, relative to the counterfactual change in other state(s)’ employment.<sup>1</sup> Two issues arise with this approach: first, the minimum wage is often set at the state

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<sup>1</sup>Using within-state variation still leverages changes to the federal minimum wage: if the federal wage floor increases, this effectively acts as a “negative” treatment to the wage differential between states that already use a super-federal minimum wage (and leave it unchanged) and those for whom the federal minimum binds.

level, and different states may not serve as useful counterfactuals; second, the true effect on employment outcomes may not be discrete in levels.

The first issue has received a great deal of attention. A large literature builds upon the basic difference-in-differences framework by modifying the specification to improve the quality of the counterfactuals, and recent empirical specifications often include covariates that capture non-linear differences in local economic climates. For example, [Orrenius and Zavodny \(2008\)](#) use a broad set of business cycle controls (in addition to time period dummies) to account for differing economic environments across states and over time; they find no adverse effects of the minimum wage on the employment of less-educated adults. Taking this a step further, [Allegretto et al. \(2011\)](#) demonstrate substantial heterogeneity in employment patterns across regions of the U.S. and control for this by allowing time period effects to vary by Census Division; they similarly find no effect of the minimum wage on teenage employment. State- or county-specific time trends are often included (e.g. [Page et al., 2005](#); [Addison et al., 2009](#); [Allegretto et al., 2011](#)) to control for pre-existing time-paths in how labor markets evolve within different areas. Recently, even more creative approaches have been employed in efforts to improve the counterfactual: [Dube et al. \(2010\)](#) compare contiguous counties across state lines and find no detrimental effect from minimum wage differentials; [Sabia et al. \(2012\)](#) conduct a synthetic control study and find that a state's increased minimum wage had a large and significantly negative employment elasticity for low-wage teenagers and young adults.

These more complex specifications arguably provide for better counterfactual employment variation, but this improvement comes at a cost: increasing the number of econometric controls reduces the amount of variation available to identify an effect. A recent paper by [Neumark et al. \(2013\)](#) is highly critical of some of these approaches, arguing that these studies have “thrown out so much useful and potentially valid identifying information that their estimates are uninformative or invalid.” In a parallel vein, we discuss how the temporally staggered nature of increases in nominal state minimum wages affects estimation – especially for specifications which control for state time trends.

Unlike many treatments assessed in the program evaluation literature, nominal state minimum wages change fairly often (see [Figure 1](#)). Moreover, minimum wage increases tend

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[Baskaya and Rubinstein \(2012\)](#) employ an interesting hybrid approach. They allow state-driven variation in minimum wage levels to determine the potential impact to each state of a change to the federal minimum wage, and then instrument for these “minimum wage gaps” in a structural model using factors such as political ideology. They find that minimum wages have significant disemployment effects for teenagers, but only when accounting for this variation across states.

to be staggered across states over a span of only a few years. This means that, even with a long state panel, there are mostly short periods which have substantial variation in minimum wages from which to identify employment effects. This is illustrated in Figure 2, which plots the standard deviation of residuals for a regression of real minimum wage on state fixed effects and quarterly (quarter-by-year) fixed effects. Generally, variation in minimum wages increases while some states raise their minimum wages, followed by reductions in variation as the federal minimum wage is increased shortly thereafter.

Despite this type of treatment dynamic, difference-in-differences identification strategies can still find an effect on the level of employment if there is a sufficiently rapid drop in the number of jobs relative to the counterfactual. Given the small margin of net job expansion relative to total employment, this effectively necessitates a (temporary) reduction in the absolute size of the labor force. However, if the minimum wage instead affects the rate of net job growth, then it will take some time for the effect to be reflected in the level of total employment to a degree which would be statistically detectable. In Appendix A, we provide some simple illustrations of how staggered treatment complicates estimation of non-discrete changes in outcomes.

Further complicating matters, if the true effect of a policy is to change the *slope* for the outcome variable, rather than its *level*, then including time trends as controls will attenuate estimates of the policy’s effect; that is, the inclusion of state-specific time trends leads to biased estimation.<sup>2</sup> The basic intuition is that including state-specific time trends as controls will adjust for two sources of variation. First, if there is any *pre-treatment* deviation in outcomes that is correlated with treatment – e.g. if states that exhibit stronger employment growth are also more likely to increase their minimum wage – then this confounding variation may be appropriately controlled for by including state-specific time trends. The potential cost of this added control is that if the actual treatment effect, the *post-treatment* employment variation, acts upon the trend itself, then inclusion of state time trends will attenuate estimates of the treatment effect and often leads to estimating (statistical) null employment effects.<sup>3</sup>

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<sup>2</sup>We are grateful to Cheng Cheng and Mark Hoekstra, as well as Justin Wolfers (Wolfers, 2006) for this insight.

<sup>3</sup>In a large set of models for how the minimum wage affects employment, Sabia (2009a) notes that the estimated elasticity is non-negative only when state-specific time trends are included. He offers two explanations for this peculiarity: “First, it may be that the included state-level time-varying controls failed to capture important differences in retail employment trends associated with minimum wage hikes, and that the negative minimum wage effects found in models 1-9 can be explained by state-specific time-varying unmeasured heterogeneity. However, a second explanation is that the inclusion of state linear time trends reduces potentially important identifying variation, thus rendering minimum wage effects small and insignificant.”

We illustrate this attenuation problem using both a simple example and a Monte Carlo simulation in Appendix B. The toy example depicts employment in two hypothetical states which exhibit identical employment growth rates prior to period  $t = 0$ . After period  $t = 0$ , the employment growth rate in one state falls relative to that in the other state, but there is no discrete change in the level of employment. We then compute the difference-in-differences of state employment, with and without adjustment for state time trends. The computed employment effect is large and negative when state trends are omitted, but shrinks nearly to zero with the inclusion of state time trends. This occurs despite identical pre-treatment employment trends.

In the simulation, we extend this example. Within each Monte Carlo repetition, we simulate a panel of minimum wages and employment in 51 hypothetical states. We impose a treatment effect which relates the minimum wage to employment growth, but not discretely to the level of employment. Then, we use the simulated state-year values to estimate specifications for the effect of the minimum wage on net job growth and on employment, separately with and without including state time trends. In a set of 10,000 Monte Carlo replications, we show that the estimated effect for the minimum wage on job growth is identical to the true effect, regardless of whether state-specific time trends are included as a control. In contrast, although the estimated effect on employment is large and of the correct sign when state linear trends are omitted, the inclusion of state-specific trends into this estimation manages to attenuate the estimated employment effect to (a small and statistically insignificant) zero, despite the true effect being large and negative by construction. As in the simple example, this attenuation occurs despite: (1) a large true effect on job growth by construction; and (2) no systematic correlation of changes to state minimum wages with state effects, year effects, or state time trends (which were all randomly assigned).

If the minimum wage changes the rate of job creation rather than leading to an immediate drop in the number of jobs, then the general lack of significant effects of the minimum wage on employment levels in specifications including time trends is not surprising. This is not necessarily an argument against the inclusion of state time trends as controls, however, because of the aforementioned potential for pre-treatment deviations in employment outcomes. Instead, we argue that using net job growth as an outcome can provide for a more compelling estimate of the employment effect of the minimum wage. As discussed in the following sub-section, the justification for our approach is two-pronged, motivated both by theoretical arguments and econometric concerns.

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Our discussion above – that the treatment effect is in the trend itself – forms a third explanation.



## 2.2 The Case for Examining Employment Dynamics

The basic analysis of the effects of the minimum wage argues for rapid adjustments to a new equilibrium employment level (e.g. [Stigler, 1946](#)). However, transitions to a new employment equilibrium may not be smooth ([Hamermesh, 1989](#)) or may be relatively slow ([Diamond, 1981](#); [Acemoglu, 2001](#)). In this case, the effects of the policy may be more evident in net job creation. The basis for our framework is the role of the minimum wage in a worker search and matching model (e.g. [Van den Berg and Ridder, 1998](#); [Acemoglu, 2001](#); [Flinn, 2006, 2011](#)), summarized concisely in [Cahuc and Zylberberg \(2004\)](#). In this class of models, the minimum wage has opposing effects on job creation. Although it reduces demand for labor by raising the marginal cost of employing a new worker, a higher minimum wage increases the gap between the expected returns to employment relative to unemployment, inducing additional search effort from unemployed workers. By increasing the pool of searching workers (and the intensity of their searching), the minimum wage improves the quality of matches between employers and employees, generating surplus. The theory thus has ambiguous predictions for the effect of a minimum wage on job creation. If workers' additional search effort sufficiently improves the worker-firm match quality, then job creation should not be adversely affected and may even increase. However, if the demand-side effect dominates, then increasing the minimum wage will cause declines in hiring.

The effect of the minimum wage on worker separations is ambiguous as well, but there are compelling reasons to expect a smaller effect on job destruction. For employers, the non-trivial fixed cost associated with hiring a new employee (e.g. screening, interviewing, training) likely encourages reductions in hiring rather than increased layoffs ([Oi, 1962](#)). Psychological factors may additionally lead to a tempered effect of the minimum wage on job destruction. There is a growing management literature on the negative feelings managers have when terminating employees, sometimes called “firing aversion” ([Folger and Skarlicki, 1998](#); [Molinsky and Margolis, 2005](#); [Dubinsky et al., 2011](#)). These studies posit that managers are disinclined to terminate an employee – even if a layoff is justifiable on economic grounds – because “the decision often produces sorrow or guilt, or both” ([Gilbert, 2000](#)). Further offsetting any disemployment effect, the minimum wage increases the compensation for a subset of employees; these workers may be less likely to voluntarily exit employment, especially if they possess firm-specific human capital ([Hamermesh, 1987](#)).

Furthermore, among workers paid hourly in the United States, additional tenure is highly associated with increased pay. It follows that minimum wage employees are likely to be relatively recent hires, a finding documented by [Even and Macpherson \(2003\)](#) and [Dube](#)

et al. (2011). An implication is that minimum wage increases are most likely to affect workers who are (or would be) recent hires.

To test this, we use the Current Population Survey’s Merged Outgoing Rotation Groups (CPS-MORG) to determine the proportion of newly-hired employees who are paid their state’s minimum wage.<sup>4</sup> Performing calculations similar to those in [Even and Macpherson \(2003\)](#) – though expanded to include another decade of data – we use the weights provided in the CPS-MORG for 3.56 million individuals between 1979 and 2012 who are observed in both sets of interviews (spaced twelve months apart) in which questions on earnings were asked. We find that among all employees (in Period 2), 3.25 percent earn the minimum wage. But among those who are newly employed (that is, not employed in the first wave of interviews but employed in the second), 11.46 percent earn the minimum wage; fully 30.6 percent of minimum wage workers are recently employed. Thus, minimum wage compensation is three-and-a-half times more prevalent among new workers than in the entire labor force.<sup>5</sup> Moreover, transitions out of the minimum wage are common. Following [Even and Macpherson \(2003\)](#) again, we find that among those who were paid the minimum wage in the first wave, 16.55% leave the labor force and 5.85% become unemployed. Among those who remain employed, 65.85% are paid in excess of the minimum wage in the following year. Those individuals are paid a median amount of \$1.00 per hour above that year’s minimum wage, or 24% of their previous wage; the 75th percentile of this figure is \$2.75 per hour.

Although there is support for employee hiring to be relatively more affected, the effect of the minimum wage on both job creation and job destruction is ultimately an empirical question. And, because the effects on these gross margins are theoretically ambiguous and potentially opposing, the net employment effect of the minimum wage could take several forms that are not mutually exclusive. First, the minimum wage could affect (positively or negatively) the total employment level. Second, by encouraging a longer duration of worker-firm matches, the minimum wage could reduce turnover of employees within existing jobs.

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<sup>4</sup>The data are made available by the [National Bureau of Economic Research](#). In our evaluation, we consider an employee to be paid minimum wage only when the employee is paid an amount at or below the state’s minimum wage by the hour, treating all salaried workers as being paid above the minimum wage. Although it is possible to impute the hourly wage for non-hourly employees in the CPS, there is reason to be skeptical of such imputed values in this type of analysis ([Card, 1992](#)). Inclusion of workers who self-report hourly earnings below the minimum wage into the “percent at the minimum” is common in this literature (e.g. [Lang and Kahn, 1998](#); [Pedace and Rohn, 2011](#)).

<sup>5</sup>There is also a non-trivial subset of workers paid only somewhat more than the minimum wage who may be affected by an increase. 5.6% of all workers and 19.3% of newly-hired workers are paid within ten percent of the minimum wage. It is further possible that an even larger share of workers than this are affected by minimum wages in light of recent research on “last-place aversion” by [Kuziemko et al. \(2012\)](#).

Finally, a minimum wage could change the net *flow* of workers into employment by altering the job growth rate. Any of these outcomes are consistent with the theoretical relationships discussed above, but the bulk of the literature has focused on the first relationship, investigating how a minimum wage affects the employment level.

Several recent studies offer exceptions to the focus on employment levels. [Dube et al. \(2011\)](#) examine the relationship between the minimum wage and employee turnover using the 2001-2008 Quarterly Workforce Indicators (QWI). They focus on teenagers and restaurant employees employed in contiguous counties across state lines, and find that the minimum wage reduced both new hiring and separations despite having little effect on contemporaneous employment levels. [Brochu and Green \(2012\)](#) assess firing, quit, and hiring rates in Canadian survey data. They find that workers hired within the previous six months are less likely to separate from their jobs in the presence of a higher minimum wage, a result driven in part by a reduction in firings; they find no effect on workers with longer tenure. Similar to [Dube et al. \(2011\)](#), they find a reduction in hiring rates but do not estimate the net effect on job growth.<sup>6</sup>

It is important to note that even if it were the case that minimum wages just reduce employee turnover, this outcome might be undesirable. [Lazear and Spletzer \(2012\)](#) argue that employee churn is an important component of the labor market because it indicates reallocation of workers to jobs in which they are more productive. They link declines in employee turnover to reduced economic output. More to the point, the total employment effects of the minimum wage are of primary interest for policy-making. It is uncertain what policy goals are served by increasing the tenure of voluntary employment through labor market regulations. We believe that the relevant outcome is employers' creation of new jobs or destruction of existing jobs – the *net* movement of workers into employment. In light of the issues discussed above, this effect may follow a slow process.

[Sorkin \(2013\)](#) builds a model that formalizes this potentially slow adjustment of labor demand and applies it to minimum wage increases. He argues that “the ability to adjust labor demand is limited in the short run” and that this “provide[s] an explanation for the small employment effects found in the minimum wage literature.” Fundamentally, this identification problem stems from the “sawtooth pattern” exhibited in states’ real minimum wages.

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<sup>6</sup>Two other recent papers examine the minimum wage and employment dynamics by exploiting institutional features of the minimum wage in Portugal ([Portugal and Cardoso, 2006](#)) and Germany ([Bachmann et al., 2012](#)). These studies also do not focus on job growth. [Hirsch et al. \(2011\)](#) and [Schmitt \(2013\)](#) focus on other channels of adjustment in response to increases in the minimum wage, such as wage compression, reductions in hours worked, and investments in training.

Sorkin argues that “difference-in-difference faces challenges in measuring the treatment effect of interest, which in this case is the effect of a permanent minimum wage increase, whenever there are dynamic responses to the treatment and the treatment itself is time-varying.”

Historically, minimum wages have been set in nominal dollars and not adjusted for inflation, so any nominal wage differential between two states will become economically less meaningful over time.<sup>7</sup> Furthermore, sooner or later every state experiences a nominal increase in its minimum wage, either due to a revision to a state law or because the federal minimum wage increases. Unlike the slow erosion of nominal minimum wage gaps brought about by inflation, an increase in the counterfactual’s minimum wage may quickly close or even reverse this gap. To put it another way, there is no consistent control group in the long run.

This assertion is supported by the graphs of monthly-frequency state real minimum wage data in Appendix C. Looking first only within-state, Figure 10 shows that the mean real state minimum wage increase during 1976-2012 was 55 cents (the median was also 55 cents). By the time the same state next increased its real minimum wage, which took 54 months on average, the previous increase in minimum wage had eroded – via inflation – to an average cumulative real *decrease* of 11 cents (median -12 cents, see Figure 11). In fact, Figure 12 shows that the 62 percent of state-year real minimum wage increases that were eventually fully eroded by inflation did so in, on average, twenty-two months, and the median time elapsed was only sixteen months. Turning instead to comparisons within Census Region, the mean *relative* real increase in state minimum wage was 25 cents (median 13 cents, Figure 13). By the time of the next within-state increase, the prior increase had eroded – both via inflation and from other regional neighbors changing their minimum wages – to an average decrease of 1 cent (median +2 cents, Figure 14). For those 47 percent of state-year increases which fully eroded relative to regional states, this took only 17 months on average (median 12 months, Figure 15).

Granted, there remain numerous state-year increases in the minimum wage that were never fully eroded by inflation or, in a relative sense, by neighboring states later boosting their minimum wages. However, this exercise demonstrates that there is a relatively short duration of time during which a state difference-in-differences estimation can identify the effects of the minimum wage on employment levels. This situation would not be problematic if the

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<sup>7</sup>Ten states now use regional CPI measures to index their minimum wages for inflation, but this is a relatively recent practice (Allegretto et al., 2011). The Federal minimum wage increase proposed by President Obama in 2013 included a provision for annual increases based on inflation, but little is known about how inflation indexing may alter the effects of a minimum wage on employment.

minimum wage affected employment in an abrupt, discrete manner. But if the minimum wage predominantly affects job creation, then it may take years to observe a statistically significant difference in total employment.<sup>8</sup> Thus, while it is true that any reduction in job growth should be reflected eventually in total employment, the empirical challenges discussed in this section may preclude identifying the net effect of the minimum wage by examining employment levels directly. As a result, even though the employment level is the outcome predominantly considered in the empirical literature on the minimum wage, we focus instead on the job *growth* rate. We implement this approach in the following sections.

## 3 Empirical Model

### 3.1 Econometric Specifications

We estimate four reduced-form specifications for the set of employment outcomes at time  $t$ . As discussed in Section 1, we begin by assessing the overall rate of net job growth, which we calculate, following the Census Bureau, as employment at time  $t$  minus employment at time  $t - 1$ , divided by the average of employment at times  $t$  and  $t - 1$  (United States Census Bureau, 2012).<sup>9</sup> We argued in Section 2 that this outcome might serve as a better measure of the true employment effect of the minimum wage. We additionally estimate the effect on the natural log of employment, contrasting this result with that for job growth. Finally, we decompose the effect on net job growth into its respective margins of (log) job creation and (log) job destruction. Our specifications take the following forms:

$$\text{Employment Outcome}_{st} = \beta \cdot \ln(\text{MW})_{st} + \phi_s + \tau_t + \epsilon_{st} \quad (1)$$

$$\text{Employment Outcome}_{st} = \beta \cdot \ln(\text{MW})_{st} + \phi_s + \tau_{rt} + \epsilon_{st} \quad (2)$$

$$\text{Employment Outcome}_{st} = \beta \cdot \ln(\text{MW})_{st} + \phi_s + \tau_{rt} + \psi_s \cdot t + \epsilon_{st} \quad (3)$$

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<sup>8</sup>An additional econometric explanation for statistical insignificance is over-saturating a model beyond the power of the data – a non-trivial concern, especially for studies including both state time trends and very geographically granular time period fixed effects.

<sup>9</sup>Results are virtually unaffected by instead defining the job growth rate as employment at time  $t$  minus employment at time  $t - 1$ , divided by employment at time  $t - 1$ . Similar results are also obtained from using just the log of employment at time  $t$  minus the log of employment at time  $t - 1$ .

$$\text{Employment Outcome}_{st} = \beta \cdot \ln(\text{MW})_{st} + \phi_s + \tau_{rt} + \psi_s \cdot t + \text{Controls}_{st}\Gamma + \epsilon_{st} \quad (4)$$

Specification 1 estimates the classic panel difference-in-differences, examining the impact of the minimum wage after controlling for state ( $\phi_s$ ) and year ( $\tau_t$ ) fixed effects. Because different regions of the country may face heterogeneous economic shocks that are correlated with changes in the minimum wage, we adapt Specification 2 to allow the time fixed effects to vary across Census Regions ( $\tau_{rt}$ ). However, these approaches fail to capture variation in state employment *trends* over time that may be caused by factors that are correlated with changes in the minimum wage, so we add state linear time trends ( $\psi_s \cdot t$ ) in Specification 3. To capture residual *non-linear* variation in state economic environments, several socioeconomic controls (discussed in Section 4.3) are added in Specification 4.

As we discuss in detail above, the inclusion of state-specific time trends acts as a double-edged sword: on one hand, this controls for pre-treatment deviations in employment outcomes that may be correlated with changes to a state’s minimum wage, but on the other hand, this practice will attenuate estimates of employment effects if the true effect acts on the growth rate rather than the employment level. We argue for examining the effect of the minimum wage on net job growth, with the inclusion of state time trends as a control, an approach supported by our Monte Carlo exercise. Thus, this final specification – specifically, examining the effects on net job growth – is our preferred model, as it most thoroughly accounts for potential confounders while overcoming the potential attenuation problem induced via inclusion of state time trends.<sup>10</sup>

## 3.2 Identification Concerns

Although we include time trends and other time-varying state characteristics, it remains possible that unobserved systematic off-trend deviations drive the correlation between employment outcomes and changes in the minimum wage. Moreover, because there is often a delay between minimum wage legislation and enactment (see [Murphy \(2005\)](#) for some ex-

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<sup>10</sup>Many papers in the literature that examine employment outcomes for a specific subgroup (see, for example, [Sabia, 2009b](#)) also include the unemployment rate for *different* subgroups in an attempt to further control for business-cycle effects. Since we are examining the entire labor market, the overall unemployment rate would be highly endogenous, as discussed in [Neumark and Nizalova \(2007\)](#). Nevertheless, when we add the unemployment rate to this specification, our results are unchanged both qualitatively and quantitatively, suggesting that we are already controlling for any relevant business cycle effects. In Section 5.5 We also show that our results are not sensitive to using Census-Division-by-time effects, nor to the inclusion of quadratic time trends.

amples), employers may react in anticipation of a future change to the local minimum wage. In either of these cases, estimated coefficients for leads of the minimum wage will be economically significant. Results from robustness checks including leads (described in Section 5.4) indicate that this is not the case. These results suggest little employment response in advance of a change. But, to the extent that firms do adapt in anticipation of actual adjustments to the minimum wage, this practice reduces the magnitude of the estimated minimum wage effects.

A related concern is that minimum wage legislation passed during expansionary economic climates may take effect in contractionary periods, resulting in a spurious negative correlation between minimum wages and employment (Reich, 2009; Allegretto et al., 2011). This concern seems unwarranted. In our data from 1976-2012, the average state-level unemployment rate is 5.96% in months during which a state increased its minimum wage, compared to 6.10% in months with no minimum wage increase; the difference is statistically insignificant ( $p = 0.12$ ).<sup>11</sup>

Another issue is employer noncompliance. Predictions vary for the theoretical employment effects from employer noncompliance with minimum wage laws, ranging from higher employment levels than with full compliance (e.g. Ashenfelter and Smith, 1979) to the same employment outcome as with compliance (e.g. Yaniv, 2001). Of course, noncompliance may be part of the response to an increased minimum wage, attenuating the effects of these laws. To the extent that increases in minimum wages are positively correlated with the rate of noncompliance, our results may understate their effect relative to that if compliance was complete.

To assess the importance of the functional form of the minimum wage term, we also specified the minimum wage in real levels instead of as a real natural log, addressing concerns raised by Baker et al. (1999). Although the results are robust to this alternate specification, we prefer our natural log specification both because of its elasticity-like interpretation and for consistency with the broader literature.

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<sup>11</sup>Baskaya and Rubinstein (2012) demonstrate pro-cyclical timing in state increases to minimum wages for states that typically had a super-federal minimum wage.

## 4 Data

### 4.1 Employment Data

In our study, we examine employment outcomes in the United States across three administrative data sets, that vary in their strengths and weaknesses. In particular, we study the *Business Dynamics Statistics* (BDS) from the Bureau of the Census, the *Quarterly Census of Employment and Wages* (QCEW) from the Bureau of Labor Statistics (BLS), and the *Quarterly Workforce Indicators* (QWI), also from the Bureau of the Census.<sup>12</sup> The QCEW and QWI report quarterly employment data for each state, while the BDS is annual as of March 12<sup>th</sup>.

Each of these data sets is administrative in nature: the QCEW and QWI programs collect employment data from county unemployment insurance commissions, while the BDS reports employment data furnished to the U.S. Internal Revenue Service. The employer-sourced administrative nature of these data is important for our research question. As discussed in Section 2, a higher minimum wage may induce additional searching effort on the part of the currently unemployed. [Mincer \(1976\)](#) shows that this positive supply elasticity often leads to an increase in the number of *unemployed* that differs substantially from the change in employment. Because employment is the policy-relevant outcome, measuring job counts using employer-sourced data provides a better identification of any disemployment effects of the minimum wage than do surveys of individuals, such as the Current Population Survey. Moreover, as demonstrated by [Abraham et al. \(2009\)](#), employment data directly reported by firms to maintain legal compliance are more accurate than responses to individual-level surveys such as the CPS.

Additionally, each of the data sets we study accounts for virtually the population of non-farm employment. Population-level data provide for a cleaner assessment of the overall policy impact of minimum wages by avoiding sampling error and enabling us to obtain more precise estimates.<sup>13</sup> These gains do not come without cost. In particular, we are limited in assessing any heterogeneity of labor market effects across demographic groups. Nor can we analyze individuals' life-cycle wage dynamics. We view these costs as unfortunate but

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<sup>12</sup>All of the data and code used in this study are publicly-available and provided by the authors online.

<sup>13</sup>These data still contain nonsampling errors such as typographical errors made by businesses when providing information. The BLS and Census Bureau act to minimize erroneous values, such as excluding from deaths establishments that “exit” the employer universe only to re-enter at some later time. Regardless, as noted by [Haltiwanger et al. \(2009\)](#), nonsampling errors are likely to be distributed randomly throughout the data.



worthwhile, justified by the advantages of obtaining a compelling answer to the question of a minimum wage’s employment effects.

To the best of our knowledge, we are the first to estimate effects of the minimum wage using the BDS. However, researchers working on related questions have used some of the employment data available in the QWI (e.g. [Thompson, 2009](#); [Dube et al., 2011](#)) and the QCEW (e.g. [Addison et al., 2009](#); [Dube et al., 2010](#)). We discuss strengths and weaknesses of each of these data sets below.

#### **4.1.1 Business Dynamics Statistics (BDS)**

The BDS data are provided by the U.S. Census Bureau and cover all non-agriculture private employer businesses in the U.S. that report payroll or income taxes to the IRS. The heart of the BDS is the Census Bureau’s internal Business Register, which is sourced from mandatory employer tax filings and augmented using the Economic Census and other data to compile annual linked establishment-level snapshots of employment statistics (on March 12<sup>th</sup>). The Census Bureau releases the BDS as a state-year panel of employment dynamics, currently covering 1977 to 2011.

One weakness of the BDS is that because the data are annual, rather than quarterly, the BDS does not allow us to match as closely any possible changes in employment to changes in the minimum wage as allowed for by the QCEW. Additionally, the BDS does not disaggregate state employment statistics by industry. However, the BDS provides the longest U.S. panel of gross job creation and destruction, thereby allowing us to deconstruct the effect of the minimum wage into these respective employment margins.

#### **4.1.2 Quarterly Census of Employment and Wages (QCEW)**

The *Quarterly Census of Employment and Wages* (QCEW), housed at the Bureau of Labor Statistics, is a program which originated in the 1930s to tabulate employment and wages of establishments which report to the Unemployment Insurance (UI) programs of the United States. Per the BLS, employment covered by these UI programs today represents about 99.7% of all wage and salary civilian employment in the country (including public sector employment). The BLS currently reports QCEW data by state for each quarter during 1975-2012, a span slightly longer than that of the BDS.<sup>14</sup> The data are disaggregated by

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<sup>14</sup>Employment levels – and therefore also quarterly job growth rates – are not available in the QCEW for Alaska and the District of Columbia for any quarters during 1978-1980. Employment data is not missing for any other states or periods. For each of our regression outcomes, we include observations only if the job

NAICS industry codes for 1990-2012. The QCEW form the longest and most temporally granular panel of data included in our study, making these data ideal for contrasting the effect of the minimum wage on employment from that on job growth rates. However, a weakness is that the QCEW does not report measures of gross job creation or destruction, such that the QCEW cannot be used to decompose the net effect on job creation into its respective margins.

### 4.1.3 Quarterly Workforce Indicators (QWI)

The *Quarterly Workforce Indicators* are data provided as part of the Longitudinal Employer-Household Dynamics (LEHD) program by the Bureau of the Census. Similar to the QCEW, these data originate from county employment insurance filings.<sup>15</sup> As noted above, for our study a major weakness of the QCEW is the lack of data on gross job creation and destruction, and a weakness of the BDS is that the data are only annual. At a first glance, the QWI data seem to bring the best of both worlds, including quarterly data on all non-agricultural employment as well as gross job creation and destruction. Additionally, compared to the QCEW or BDS, a main advantage of the QWI is that these data offer finer measures of employees demographics such as age.

Yet, for our research design, a major shortcoming of the QWI is the substantially shorter – and highly unbalanced – length of the panel. At its onset in 1990, only four states opted into the QWI program, and additional states gradually joined each year (except 1992) through 2004. From 2004 on, the QWI includes forty-nine states (Massachusetts and Washington, D.C. are never included). Thus, the starting date for QWI participation varies considerably across states, and many are relatively recent.

In light of the issues (especially real minimum wage erosion) discussed in Section 2, a longer span of data is preferable to study the employment effects of the minimum wage. Additionally, the years spanned by the QWI panel exhibit fairly high frequency variation in effective state minimum wages (see Figure 1). Baker et al. (1999) demonstrate that the frequency of changes to state minimum wages within a panel has major implications for estimated employment effects, and that estimates based on shorter panels are especially sensitive to high frequency variation of minimum wages.

Ultimately, we feel that – their shared strengths aside – each of these administrative data

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growth rate is available for the state that period.

<sup>15</sup>In fact, the QWI and QCEW originate identically from the same county unemployment insurance records. Thus, differences in the data stem from either the periods during which each state or county is included, or differing imputation methods employed by BLS versus Census (Abowd and Vilhuber, 2013).

sets brings various strengths and weaknesses towards answering our research question. In the results tables, we report estimates separately for each data set, and we find that the effect of the minimum wage is consistent across these data sets.

## 4.2 State Minimum Wages

We draw historical data on state minimum wages from state-level sources.<sup>16</sup> For the QCEW and QWI, we use the minimum wage value as of the first of each quarter. For the BDS, we use the value as of the previous March 12th each year, directly corresponding to the panel years in the BDS data. Some states have used a multiple-track minimum wage system, with a menu of wages that differ within a year across firms of different sizes or industries; we therefore use the maximum of the federal minimum wage and the set of possible state minimum wages for the year. To the extent that there is firm-level heterogeneity in the applicable wage level, our definition allows the minimum wage term to serve as an upper bound for the minimum wage a firm would actually face. We transform minimum wages into constant 2011 dollars using the (monthly) CPI-U from the Bureau of Labor Statistics.<sup>17</sup>

## 4.3 Other Control Variables

Although our econometric specifications include an extensive set of time period controls, precision may be gained by accounting for additional state-specific time-varying covariates. We merge the BDS and minimum wage data with state-level controls from several sources.

The Census Bureau's Population Distribution Branch provides annual state-level population counts, including estimates for intercensal values. Total state population represents a determinant of both demand for (indirectly by way of demand for goods and services) and supply of employees. Because states differ non-linearly in their population changes, controlling directly for population may be important. The range in population between states and across time is enormous, so we use the natural log of state population in our specifications.

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<sup>16</sup>Although historical state minimum wage data are available from sources such as the U.S Department of Labor (<http://www.dol.gov/whd/state/stateMinWageHis.htm>), these data suffer several limitations. For one, minimum wage values are only reported only as of January first each year, whereas the panel used in our study necessitates values as of other dates. Additionally these DOL data incompletely characterize changes to state minimum wages, especially during the early years of our panel. This DOL table is frequently used as the source of historical state minimum wage values for recent studies in this literature, and we caution future researchers to be careful not to inadvertently attribute minimum wage changes to years in which they did not occur. The full set of data we use is available online.

<sup>17</sup>Because we use a national-level deflator, specifying the log minimum wage term as real versus nominal does not affect our results. Time period fixed effects incorporate this added variation.

We additionally include the share of this population aged 15-59, which provides a rough weight for how population might affect demand for versus supply of labor. Demographic controls such as these are commonly used in this literature (e.g. Burkhauser et al., 2000; Dube et al., 2010). Following Orrenius and Zavodny (2008), we also include the natural log of real gross state product per capita.<sup>18</sup> After controlling for state population, this term can be thought of as a rough proxy for average employee productivity as well as a measure of state-level fluctuations in business cycles (Carlino and Voith, 1992; Orrenius and Zavodny, 2008). We include these controls at an annual level – the most granular temporal level available; these covariates are not available quarterly.

Our study uses data on the fifty states and Washington, D.C., when available. Both the QCEW and BDS include these 51 entities, the QCEW from 1975-2012 and the BDS from 1977-2011.<sup>19</sup> The QWI includes at most 49 states (Massachusetts and Washington, D.C. are never present), but the state panel is highly unbalanced, with included states entering at any point from 1990-2004. Rather than impose some arbitrary cutoff, we include all available periods for each data set in our analysis.

We present summary statistics for state minimum wages and employment variables for each data set in Table 1. As discussed in Section 2.2, a more detailed presentation of state minimum wage levels and changes is available in Appendix C.

## 5 Results

### 5.1 Net Job Growth

We begin by examining the effects of the minimum wage on the net job growth rate in Row [1] of Table 2. The results are fairly consistent across the columns for the three data sets: there is a negative and statistically significant effect on job growth. They are also similar in magnitude in each data set. Recall that the BDS is an annual panel whereas the QCEW

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<sup>18</sup>We compute the log of the real value of total GSP per capita using all industry codes, including government. Results are virtually unaffected by using  $\ln(\text{real private sector GSP/capita})$  instead, but we view total GSP as the more appropriate definition given that the population term reflects total state population.

<sup>19</sup>In the BDS, there is a data quality issue for observations for Alaska in 1989 and 1990 and Oregon for 1993 and 1994. For instance, the annual change in employment is about 35 percent from 1993 to 1994 in Oregon, an implausibly large magnitude that dwarfs any annual change seen in the data, including those during the recent recession. Discussions with authorities in the Census Bureau revealed that data corruption necessitated coarsely imputing values. We elected to include these four observations anyway, none of which were associated with a change in the nominal minimum wage. Dropping them yields virtually identical results.

and QWI are quarterly panels. Annualizing the effects in Column (4) of Panels (B) and (C) – which include state-specific time trends, region-by-time effects, and additional controls – to better compare them to Panel (A) yields effects on the growth rate of -0.0586 and -0.042 for the QCEW and QWI, respectively. These are quite close to the effect of -0.0587 for the BDS.

Because the outcome is defined as a growth rate, the result in Column (4) of Panels (A) and (B) indicates that a real minimum wage increase of ten percent reduces job growth in the state by around 0.5 percentage points (during these years, the average state employment growth rate was 2.0 percent annually). In other words, a ten percent increase to the minimum wage results in a reduction of approximately one-quarter of the net job growth rate.<sup>20</sup>

Recall also that we are examining the entire labor market. To the extent that not all workers are affected by increases in the minimum wage, the effect is likely to be more concentrated on these portions of the wage distribution. Therefore, this result implies a large reduction in the rate of net new positions created, one that may appear implausible on first inspection. Keep in mind, as discussed earlier, that hourly earnings near the minimum wage are especially prevalent for newly-employed workers – indeed, 22.1% of these individuals are paid within one dollar (real 2011 values) of the prevailing minimum wage.

In particular, there is a temptation to extrapolate this effect by exponentiating the lower growth rate over a long period. This extrapolation assumes a policy in which a state *permanently* raises its real minimum wage by a constant amount relative to its counterfactual. As discussed in Section 2, there are frequent changes and thus no long-term comparison group. For this reason, calculations that extend this change in the growth rate for many years are strongly out-of-sample and therefore unreliable. It is important to note that this situation does not imply any problem with the estimation, but is a critical factor for *inference* about the estimated result for job growth. To obtain more direct inference about the effect as measured in counts of jobs, we note that this decrease in the net job growth rate implies an elasticity of -0.05 in total employment after one year.<sup>21</sup> This hypothetical increase will continue to have an effect in future years, though as discussed in Section 2, it will be eroded both by inflation and by the changes in the state’s comparison group. The effective elasticity

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<sup>20</sup>The baseline annualized growth rate in the QWI differs substantially from that in the BDS and QCEW. Although this is partly because the QWI covers a only a more recent time-span, much of this difference is due to the highly skewed sampling of states into the QWI over time. For this reason, we are reluctant to draw inference based on *average* growth rates in the QWI.

<sup>21</sup>A reduction in the growth rate by 0.5 percentage points, evaluated at the mean, results in approximately 10,000 fewer jobs relative to the counterfactual in the following year on a baseline of approximately 1.9 million jobs, yielding an elasticity of -0.05.

over the typical relevant time frame is -0.12.<sup>22</sup> That is, each ten percent increase in a state’s real minimum wage, relative to its regional neighbors, causes a 1.2 percent reduction in total employment relative to the counterfactual by the end of five years. This slow effect is in the spirit of the results found by [Neumark and Nizalova \(2007\)](#), who show that young adults exposed to higher minimum wages at the time they were likely to have been entering the labor force have worse longer-term labor market outcomes.

## 5.2 Total Employment

We directly examine the elasticity of employment with respect to the minimum wage in Row [2] of Table 2. The estimated elasticities tend to be fairly large and negative, though generally statistically insignificant, in Columns (1) and (2) of each panel, which do not include state trends. However, the specifications in Columns (3) and (4), which do include these trends, yield effects very close to zero. Thus, we can replicate a common – though certainly not universal – result in the literature of no measured effect of the minimum wage on the level of employment; this is unsurprising in light of our discussion in Section 2.

Recall that the results in Columns (3) and (4) identify changes relative to each state’s linear time trend. Essentially, then, our result is that of no *discrete* change in employment levels relative to trend. This is consistent with theories of relatively slow transitions to a new employment steady state, as in [Diamond \(1981\)](#) and [Acemoglu \(2001\)](#), as well as our discussion of the econometric issues involved in estimating the effects on employment levels in Section 2. Likewise, there is strong evidence of an attenuation problem induced by the state time trend controls. The estimated employment elasticity is large and negative in specifications (1) and (2), but the coefficient moves close to zero when state time trends are included in Columns (3) and (4). In contrast, point estimates for the effect on the job growth rate in Row [1] are fairly similar across specifications. As we discussed in Section 2, there could be *pre-treatment* variation in employment outcomes that is correlated with changes to state minimum wages, so we do not advocate the exclusion of state time trends altogether. Rather, we argue that the results for net job growth provide for more compelling inference about the true disemployment effect of the minimum wage.

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<sup>22</sup>Suppose that in year one a state increases its real minimum wage by 10%, relative to other states within its Census region. The average erosion rate in our panel predicts a remaining effective difference of 6.64% in year two. This relative difference shrinks to 3.87% by year three, to 2.31% by year four, and to 0.84% by year five, before fully eroding. This suggests a cumulative five-year effect that is 2.37 times that observed in year one.

### 5.3 Job Creation by Expanding Firms and Job Destruction by Contracting Firms

Although the results for net job growth provide an overall measure of the policy’s effect, it may be helpful to better understand how the minimum wage effects each component piece of net job growth. In this section, we separately estimate the gross effects of the minimum wage, beginning with the creation of jobs by expanding establishments. These measures are only available in the BDS and the QWI, reported in Panels (A) and (C). Column (4) of Row [3] shows an estimated elasticity of -0.23, with a standard error of 0.064 from the BDS and a smaller but still statistically significant elasticity of -0.14 from the QWI. That is, a ten percent increase in the minimum wage reduces the gross creation of new jobs in expanding firms by about 1.4 to 2.0 percent.

The effect of the minimum wage on job destruction, in Row [4], is somewhat less clear. The point estimate is statistically significant and positive only in Column (4) of Panel (A) and only at the 10% level; this estimate is also not robust to the tests in Section 5.4. The estimate in Column (4) of Panel (C) is negative, though noisy. Furthermore, because we cannot say whether these reductions are the result of establishments choosing not to fill vacancies created by voluntary separations, or a proactive decision to engage in labor force reduction, we do not make very much of this result.<sup>23</sup> We therefore conclude that the changes in the net job growth rate are primarily due to the decrease in job creation by expanding establishments, rather than an increase in job destruction by contracting establishments.

### 5.4 Robustness Checks

In this section, we present a number of alternative specifications to assess the robustness of our results. In particular, we perform a common falsification test, show that our results are consistent for different time periods in our sample, and demonstrate invariance of our results to allowing for finer spatial and time controls, accounting for minimum wage inflation indexing, and dropping the years of the recent financial crisis

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<sup>23</sup>We also examine establishment entry and exits in the BDS, in part motivated by [Burdett and Mortensen \(1998\)](#) and [Van den Berg and Ridder \(1998\)](#), who posit that increasing a minimum wage may eliminate firms which are unprofitable at the higher wage level. We find no evidence of effects on these outcomes.

### 5.4.1 Leading Values

If increases in the minimum wage appear to have an effect on employment dynamics *before* their implementation – especially if contemporaneous changes lose their effect – then our results might be driven by unobserved trends. We investigate this possibility by including several variations of a leading value of the minimum wage as a check of pre-existing deviations in the trend between treated states and their counterfactual. Table 3 presents these results for the job growth rate in each data set. Column (1) reproduces the baseline results from Column (4) of Table 2 as a reference. In Column (2), we include an indicator which equals one if the nominal minimum wage changes the following period. In Columns (3)-(5), we include the leading value of the log of the minimum wage either two, three, or four periods in advance.<sup>24</sup>

For both the BDS and the QCEW, in Column (2) the coefficient for the current period minimum wage term remains close to the baseline in Column (1), while the coefficient for the the leading indicator term is very small. Similarly, including two-, three-, or four-period leads in Columns (3)-(5) does not substantially change the result for the current period minimum wage in general, while the leading values are fairly small and are only statistically significant for the two-period lead in the QCEW. However, note the coefficient on leading value is actually *positive*, suggesting that confounding trends leading to both lower job growth and higher minimum wages are not at work. Results are mostly similar for the QWI, although the standard errors increase from the baseline enough to lose statistical significance in Columns (2) and (3), and the leading indicator term in Column (2) is significant at  $p = 0.08$ . These less-robust results in the QWI are not particularly surprising given the issues with the QWI discussed in Section 4. Finally, as with the QCEW – to the extent the leading indicator term for the QWI *does* indicate some deviation in job growth prior to the minimum wage actually increasing – the sign is positive, not suggestive of an unaccounted for negative shock.

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<sup>24</sup>Note that we do not include a one-period leading value nor include multiple leads simultaneously. This is because there is explicit collinearity between the current and the lag of the minimum wage term. For simplicity, suppose that the true data-generating process is  $Y_t = \beta_1 \ln(MW_t) + \epsilon_t$ . Ordinarily, including  $\ln(MW_{t+1})$  would show no effect in this regression. However, since  $Y_t = \frac{(emp_t - emp_{t-1})}{\frac{1}{2}(emp_t + emp_{t-1})}$  and  $\ln(MW_{t+1})$  is related to  $Y_{t+1}$ , which includes  $emp_t$ , adding a single-period lead introduces substantial endogeneity. This is not an issue for leads of at least two periods difference from each other. If the pre-trend identification assumption is violated, it is difficult to believe that it would not be apparent two periods prior as well. Moreover, including a binary variable for whether there is a change in the following period (as opposed to the actual minimum wage value) yields little indication that there is some negative shock that is correlated with both increases in the minimum wage and reductions in job growth.



## 5.5 Additional Geographic and Time Controls

It is possible that unobserved heterogeneity that drives both changes in the minimum wage and changes in job growth remains despite our extensive set of controls. In Column (6), we add Census *Division* by time period effects (there are four Census Regions containing nine Divisions). In Column (7), we add quadratic state-specific time trends. While, as noted by [Neumark et al. \(2013\)](#), these sorts of controls have the potential to oversaturate the model, neither Column (6) nor Column (7) is meaningfully different from the main specification for any of the three data sets. It is therefore difficult to believe that some other source of unobserved heterogeneity is driving our results.

### 5.5.1 Inflation-Indexing

Next, in order to assess whether states that have shifted to indexing their minimum wage for inflation affect our results, we drop these observations. Using data on wage indexing from [Allegretto et al. \(2011\)](#), this reduces the samples only slightly. The results, found in Column (8) of Table 3, are very similar to our main results. It may not be surprising that automatic increases are not driving our results – not necessarily because they are more predictable, but because these policies are relatively recent and affect few observations.

### 5.5.2 Financial Crisis

The 2008-2009 recession saw striking changes in employment. Because we include time period fixed effects (often by region), the recent recession should not unduly affect our results. However, these two years of our panel additionally experienced several large and high frequency changes in real minimum wage levels, primarily resulting from the federal increases during these years (see Figure 1). As a check that these particular years are not overly influencing identification of the minimum wage term, we estimate specifications using only pre-2008 data. As seen in Column (9) of Table 3, the coefficients are not meaningfully different from our main results.

### 5.5.3 Sensitivity to Time Period

For difference-in-differences estimates, there is nearly always a concern that results could be particular to the time-span included in the study. Generally, it seems most appropriate to use all available periods within a data set unless given a compelling reason to do otherwise. However, such an approach cannot guarantee that estimated effects are not particular to the

time period used. In this section, we evaluate results obtained from estimating the same specification using an extensive set of subsamples of time periods from the BDS and QCEW.

We use specification 2 from the text, because many of the time-spans are short, e.g. five years, undermining the effectiveness – and importance – of state time trends. For this model, we iteratively estimate results using all possible subsamples of spans of five or more consecutive years in the BDS (1977-2011) and QCEW (1975-2012). In total, this approach yields 496 point estimates using the BDS and 595 for the QCEW.<sup>25</sup>

For these 496 and 595 regressions in the BDS and QCEW, respectively, we retain the point estimates for the log real minimum wage term only, discarding the within-model standard errors. Sorting the coefficients by magnitude, the first point estimate to have a value greater than zero in the BDS is at the 95.16th percentile. The median coefficient is -0.0413, and the mean is -0.0369. Of those coefficients that have a positive value, all are for spans of ten or fewer years (the maximum span is 35 years), and more than half (14/25) occur for spans of only five or six years. Figure 3 depicts a histogram of point estimates for these time-spans of the BDS. For the QCEW, the first point estimate to have a value greater than zero is at the 94.62th percentile. The median coefficient is -0.011 (recall that this is for quarterly growth rates, whereas the BDS is annual). The mean is -0.011 as well. Of those coefficients that have a positive value, all are for periods which span thirteen or fewer years (the maximum span is 38 years), and more than half (17/33) occur for spans of only five, six, or seven years. Figure 4 depicts a histogram of point estimates for these time-spans of the QCEW.<sup>26</sup>

This yields several conclusions. First, the coefficients obtained from using the full panels of available data are quite similar to those obtained in this iterative estimation exercise. If anything, the coefficients we estimate using the full BDS and QCEW indicate a *less negative* effect of the minimum wage than those found in this exercise. Put another way, if we were to draw a consecutive span of years *at random* from the BDS and QCEW, then at least half of the time that estimated effect of the minimum wage on job growth would be at least

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<sup>25</sup>For instance, for an initial year of 1977, the BDS has 31 possible spans of at least five years: 1977-1981, 1977-1982, 1977-1983, ..., 1977-2011. For an initial year of 1978, the BDS has 30 possible such spans. An initial year of 1979 has 29 possible spans, etc. The latest initial year for the BDS is 2007, which has a single span of five years from 2007-2011. Note that we consider spans by calendar year. For the QCEW, we could instead consider all possible combinations of quarters spanning 20 quarters, but this would not add much in the way of inference and would greatly increase the computational time necessary to run the full set of regressions (going from 595 spans of at least five calendar years to 6441 possible spans of at least 20 quarters).

<sup>26</sup>Estimates using every time period are in Table 2 of the article, Panels (A) and (B), column (2). The coefficient is -0.0401 for the full BDS, compared to a median coefficient of -0.413 in this exercise; for the QCEW, the estimate is -0.0074, compared to a median value of -0.011 in this exercise.

as negative as that we estimate using the full panel. Moreover, the point estimate for this randomly-drawn span of years would be negative 95% of the time.

## 5.6 Effects on Job Growth by Industry and Age Group

To this point in the article, we have presented results for virtually the entire workforce, including workers of all ages in all industries. In this final results section, we disaggregate the effect on job growth rates by industry and age group. The BDS does not report separate employment outcomes by state and industry, but the QCEW and QWI do. The QWI additionally reports outcomes by age group. In Table 4, we estimate the effects of the minimum wage on job growth in different industries (two-digit NAICS code).<sup>27</sup> Much of the literature focuses on one or several industries that are thought to be more responsive to changes in the minimum wage. We choose to show all industries as it is not necessarily clear which particular *industry codes* ought not to be sensitive to the minimum wage. For instance, some of the largest sub-industries within the manufacturing NAICS codes (31-33) are printing, apparel manufacturing, and retail bakeries. That said, industries that tend to have a higher concentration of low-wage jobs show more deleterious effects on job growth from higher minimum wages, and the results appear consistent between the QCEW and QWI. For example, the industries that have negative and statistically significant effects in both data sets are construction; retail trade; professional services; administrative support and waste management; arts, entertainment, and recreation; and accommodation and food service.<sup>28</sup> Of the 40 coefficients we report in the two data sets, only one (educational services, in the QCEW) is positive and statistically significant, and even then only at the 10% level.

The lower panel of Table 4 shows the effects of higher minimum wages by age bin reported in the QWI. As one might expect, the effects are by far the strongest on those aged 14 to 18, twice the size of the effect on those age 19 to 21 and over three times the size of the effect on those age 22 to 24. By age 35, the effects are very small and insignificant; they rise again but remain statistically insignificant for those over age 65. These results are in line with the expectation that the minimum wage will reduce job prospects for those with less skill and experience.

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<sup>27</sup>See <http://www.naics.com/search.htm> for a full list of the component industries of each category.

<sup>28</sup>It may seem anomalous that professional services would be negatively affected, but firms in this category span a broad array, from lawyers' offices to direct mail advertising. The large negative effect on the offices of holding companies ("management") is perhaps stranger; note, though, that the effect is only present in the QCEW and that this category has among the fewest firms of any industry.

## 6 Conclusion

We examine how a wage floor impacts employment by directly assessing employment dynamics. In a worker search and matching model (e.g. [Acemoglu, 2001](#); [Flinn, 2011](#)), a minimum wage has two opposing effects on employment: it reduces demand for new workers by raising the marginal cost of an employee, while inducing additional search effort from unemployed workers, potentially improving the employee-employer match quality. The theory shapes our understanding of how a minimum wage affects employment, but the equilibrium result is an empirical question.

We provide both theoretical and empirical reasons to believe that an effect of the minimum wage should be most pronounced on net job growth. In addition, we conduct a simulation showing that the common practice of including state-specific time trends will attenuate the measured effects of the minimum wage on employment if the true effect is in fact on the rate of job growth. We examine the effects in three separate data sets and find that the results are similar both qualitatively and quantitatively: the minimum wage reduces net job growth.

The results for job creation show that, in equilibrium, any supply-side effects on search (and the potential increase in the quality of employer-employee matches) do not overcome the negative demand-side effects of higher labor costs. The lack of strong effects on job destruction is in line with the literature on the fixed costs of labor and firing aversion. More importantly, we find that on net the minimum wage meaningfully affects employment via a reduction in the rate of long run job growth.

Our results have implications for the recent proposals to index the minimum wage to inflation. We show that the effects on employment are limited by the erosion due to inflation. Permanent real increases in the minimum wage are likely to have substantially greater impacts than the nominal changes we study.

Following the recent recession, unemployment remains disproportionately high for less educated and inexperienced workers ([United States Bureau of Labor Statistics, 2013](#)). In the long run, this group of workers faces substantially longer periods of unemployment or delays in hiring, thus bearing more of the cost from minimum wages. This phenomenon is particularly important given the evidence that minimum wage jobs often result in relatively rapid transitions to higher-paying jobs.

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## 7 Figures and Tables

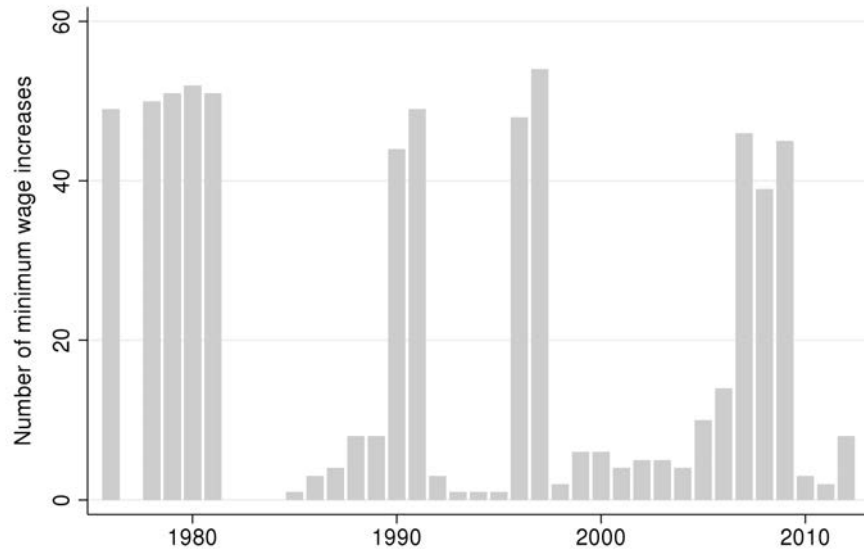


Figure 1: Frequency of increases to effective state real minimum wages (1976-2012)

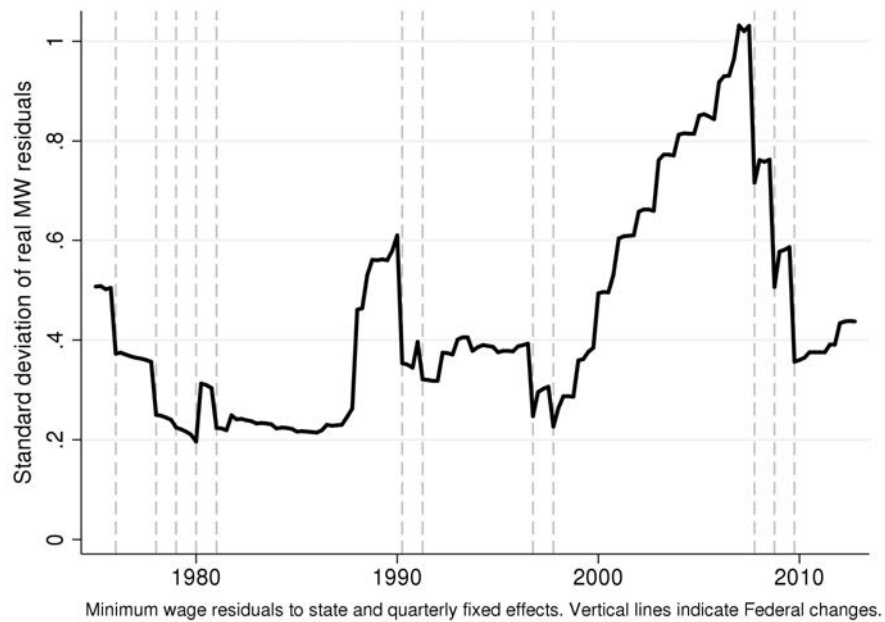


Figure 2: Standard deviation of residual state real minimum wages (1975-2012)

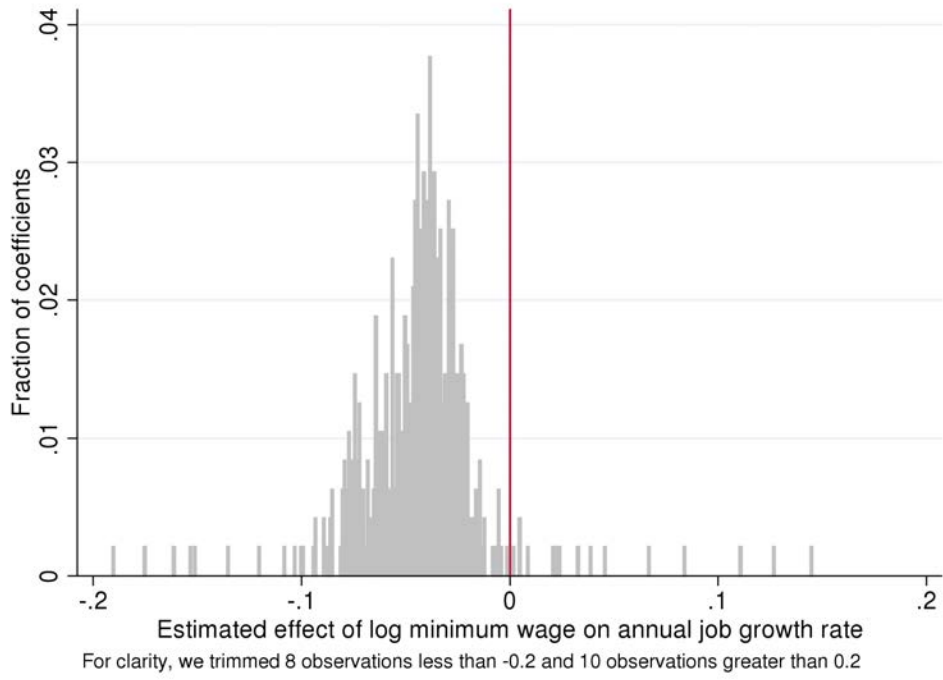


Figure 3: Distribution of point estimates from subsamples of time-spans in the BDS

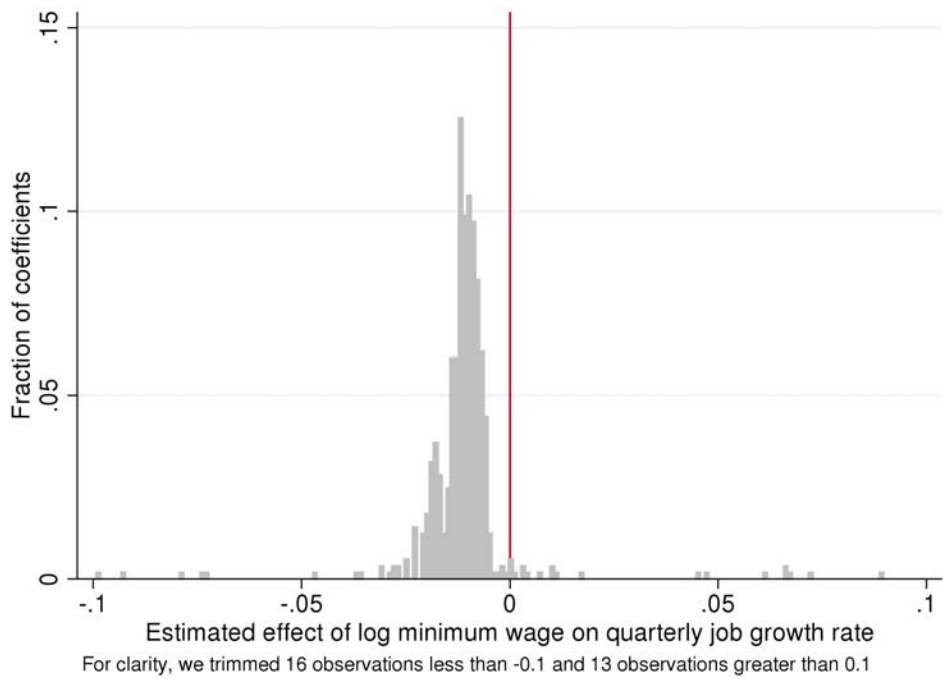


Figure 4: Distribution of point estimates from subsamples of time-spans in the QCEW

Table 1: Summary statistics for state characteristics and employment outcomes in three administrative data sets

	<b>BDS</b>			<b>QCEW</b>			<b>QWI</b>		
	Annual, 1977 - 2011			Quarterly, 1975 - 2012			Quarterly, varies - 2012*		
	Mean	Std. Dev.	Median	Mean	Std. Dev.	Median	Mean	Std. Dev.	Median
State minimum wage (\$)	4.40	1.360	4.25	4.53	1.535	4.25	5.86	1.094	5.15
State minimum wage (\$real)	7.09	0.916	6.89	7.28	0.975	7.05	6.89	0.729	6.85
Employment variables:									
Job growth rate	0.021	0.0346	0.022	0.0051	0.0256	0.0049	0.0019	0.0241	0.0061
Jobs (thousands)	1887.6	2103.0	1224.9	2167.9	2402.9	1441.7	5242.8	5588.4	3527.4
Job creation (thousands)	314.8	370.4	206.5				1147.8	1266.7	769.4
Job destruction (thousands)	282.0	337.3	180.1				1097.1	1226.6	734.7
State annual covariates:									
Population (thousands)	5160.6	5725.6	3513.4	5138.0	5704.7	3502.0	6136.5	6784.5	4343.4
Share aged 15-59	0.62	0.0196	0.62	0.62	0.0199	0.62	0.62	0.0145	0.62
GSP/capita (\$real)	41,592	16,310	38,447	41,302	16,334	38,148	45,345	8384	43,969
Observations	1785			7752			3029		

Notes: We define each state's minimum wage annually as of March 12 in the BDS, and as of the first date for each quarter in the QCEW and QWI. We use the maximum of the federal minimum wage and the state's minimum wage each period, drawn from state-level sources. Employment statistics are computed for the aggregate population of non-agricultural employees in each state for each of the three listed data sets. The job growth rate is defined as employment at time  $t$  minus employment at time  $t - 1$ , divided by the average of employment at times  $t$  and  $t - 1$ . We use the job growth rate and employment outcomes annually for the BDS, and quarterly for the QCEW and QWI. The QCEW does not report gross job creation or destruction. All real dollar amounts are indexed to \$2011 using the CPI-Urban. The QWI is a highly unbalanced panel, beginning with only four states in 1990 and gradually expanding until forty-nine states had joined by 2004. We include all available state-quarters of the QWI.

Table 2: Effect of the minimum wage on employment outcomes in three administrative data sets

	<b>Panel (A): BDS</b> Annual, 1977 - 2011				<b>Panel (B): QCEW</b> Quarterly, 1975 - 2012				<b>Panel (C): QWI</b> Quarterly, varies - 2012*			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
[1] Job growth rate	-0.0382*** (0.0093)	-0.0401*** (0.0134)	-0.0547*** (0.0165)	-0.0587*** (0.0187)	-0.0053** (0.0024)	-0.0074** (0.0031)	-0.0137*** (0.0038)	-0.0150*** (0.0040)	-0.0041 (0.0030)	-0.0081** (0.0039)	-0.0106** (0.0040)	-0.0106** (0.0042)
[2] Log of employment	-0.1859 (0.1178)	-0.2411** (0.0970)	-0.0335 (0.0313)	-0.0125 (0.0160)	-0.1105 (0.1174)	-0.1542 (0.1003)	-0.0104 (0.0388)	-0.0027 (0.0170)	-0.0383 (0.0438)	-0.0640 (0.0403)	0.0003 (0.0344)	-0.0086 (0.0149)
[3] Log of job creation	-0.3384*** (0.1064)	-0.3557*** (0.1051)	-0.2524*** (0.0686)	-0.2332*** (0.0641)					-0.1958** (0.0810)	-0.2530*** (0.0759)	-0.1246 (0.0814)	-0.1366* (0.0714)
[4] Log of job destruction	-0.0993 (0.1056)	-0.1004 (0.1415)	0.0721 (0.0674)	0.1136* (0.0617)					-0.1607* (0.0807)	-0.2070*** (0.0749)	-0.0800 (0.0827)	-0.0924 (0.0698)
State fixed effects	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y
Time period FE	Y				Y				Y			
Region * time FE		Y	Y	Y		Y	Y	Y		Y	Y	Y
State time trends			Y	Y			Y	Y			Y	Y
Other controls				Y				Y				Y
Observations	1785	1785	1785	1785	7675	7675	7675	7675	2980	2980	2980	2980

\*  $p < 0.1$  \*\*  $p < 0.05$  \*\*\*  $p < 0.01$  Notes: Each coefficient represents a separate regression of the dependent variable on the natural log of a state's real minimum wage. The column numbers correspond to Specifications (1) - (4) in the text. Robust standard errors are clustered by state and reported in parentheses. Regressions include observations for the aggregate population of employers at the state level as reported in each of the three data sets. The job growth rate is defined as employment at time  $t$  minus employment at time  $t - 1$ , divided by the average of employment at times  $t$  and  $t - 1$ . Time period fixed effects and time trends are annual for the BDS, and quarterly for the QCEW and QWI. "Other controls" consist of the natural log of annual state population, the share of state population aged 15-59, and the natural log of annual real gross state product per capita. All dollar amounts are indexed to \$2011 using the CPI-Urban. The QCEW does not report gross job creation or destruction. The QWI is a highly unbalanced panel, beginning with only four states in 1990 and gradually expanding until forty-nine states had joined by 2004. We include all available state-quarters of the QWI.

Table 3: Robustness checks for the effect of the minimum wage on job growth rates

	Baseline results (1)	Leading values tests			Division time FE (6)	Quadratic trends (7)	Inflation indexing (8)	Pre-2008 only (9)	
		Indicator (2)	$t + 2$ (3)	$t + 3$ (4)	$t + 4$ (5)				
<b>BDS: Job growth rate</b>									
Log of minimum wage	-0.0587*** (0.019)	-0.0636*** (0.020)	-0.0510** (0.020)	-0.0575** (0.023)	-0.0532** (0.023)	-0.0608*** (0.023)	-0.0423*** (0.016)	-0.0580*** (0.020)	-0.0660** (0.028)
I( $\Delta MW_{t+1}$ )		-0.0037* (0.002)							
Lead of Ln(MW)			-0.0159 (0.013)	-0.0076 (0.014)	0.0203 (0.013)				
Observations	1785	1785	1734	1683	1632	1785	1785	1740	1581
<b>QCEW: Job growth rate</b>									
Log of minimum wage	-0.0150*** (0.004)	-0.0137*** (0.004)	-0.0255*** (0.009)	-0.0173*** (0.005)	-0.0148*** (0.004)	-0.0173*** (0.006)	-0.0132*** (0.004)	-0.0129*** (0.004)	-0.0154*** (0.005)
I( $\Delta MW_{t+1}$ )		0.0024 (0.002)							
Lead of Ln(MW)			0.0131* (0.008)	0.0034 (0.004)	-0.0002 (0.004)				
Observations	7675	7622	7569	7516	7463	7675	7675	7455	6655
<b>QWI: Job growth rate</b>									
Log of minimum wage	-0.0106** (0.004)	-0.0062 (0.004)	-0.0134 (0.008)	-0.0120* (0.007)	-0.0124** (0.005)	-0.0159** (0.006)	-0.0143*** (0.005)	-0.0071* (0.004)	-0.0119** (0.006)
I( $\Delta MW_{t+1}$ )		0.0049* (0.003)							
Lead of Ln(MW)			0.0036 (0.008)	-0.0009 (0.006)	0.0014 (0.005)				
Observations	2980	2931	2882	2833	2784	2980	2980	2760	2100

\*  $p < 0.1$  \*\*  $p < 0.05$  \*\*\*  $p < 0.01$  Notes: Column (1) replicates specification (4) from Table 2. Separately: Column (2) adds an indicator equal to one if the nominal minimum wage increases in the following period. Columns (3) - (5) include, respectively, the leading value of the log minimum wage at time  $t+2$ ,  $t+3$ , or  $t+4$ . Column (6) uses Division-by-time fixed effects, rather than Region-by-time. Column (7) adds quadratic state time trends. Column(8) drops the observations with an inflation-indexed state minimum wage, and Column (9) uses only pre-2008 data.

Table 4: Effect of the minimum wage on job growth rates by industry and age

		QCEW		QWI	
		Quarterly, 1990 - 2012		Quarterly, varies - 2012	
		Coef.	Std. Err.	Coef.	Std. Err.
<b>Industry</b> <b>(NAICS)</b>	All	-0.0116***	(0.0037)	-0.0106**	(0.0042)
	11: Agriculture and wildlife	-0.0156	(0.0316)	0.0073	(0.0238)
	21: Mining	-0.0157	(0.0401)	-0.0022	(0.0211)
	22: Utilities	-0.0193	(0.0196)	0.0009	(0.0144)
	23: Construction	-0.0296***	(0.0110)	-0.0311**	(0.0143)
	31-33: Manufacturing	-0.0121	(0.0111)	0.0032	(0.0085)
	42: Wholesale trade	-0.0042	(0.0040)	-0.0083	(0.0070)
	44-45: Retail trade	-0.0092**	(0.0039)	-0.0131**	(0.0052)
	48-49: Transportation and warehouse	-0.0229**	(0.0109)	-0.0014	(0.0056)
	51: Information service	-0.0056	(0.0119)	0.0056	(0.0092)
	52: Finance and insurance	-0.0015	(0.0044)	-0.0119**	(0.0057)
	53: Real estate	-0.0095*	(0.0056)	-0.0075	(0.0068)
	54: Professional service	-0.0229***	(0.0062)	-0.0217***	(0.0068)
	55: Management	-0.0790***	(0.0280)	-0.0302	(0.0380)
	56: Administrative support	-0.0245***	(0.0087)	-0.0243**	(0.0102)
	61: Education related	0.0365*	(0.0203)	0.0152	(0.0148)
	62: Health care	-0.0009	(0.0025)	-0.0009	(0.0053)
	71: Arts and entertainment	-0.0521**	(0.0195)	-0.0364*	(0.0199)
	72: Accommodation and food	-0.0217**	(0.0086)	-0.0188**	(0.0077)
81: Other service	-0.0322*	(0.0181)	-0.0103	(0.0086)	
92: Public administration	-0.0085	(0.0117)	-0.0145	(0.0111)	
<b>Age group</b>	14-18			-0.0457***	(0.0160)
	19-21			-0.0241**	(0.0117)
	22-24			-0.0140**	(0.0068)
	25-34			-0.0105**	(0.0044)
	35-44			-0.0041	(0.0038)
	45-54			-0.0039	(0.0033)
	55-64			-0.0045	(0.0032)
	65-99			-0.0097	(0.0061)
Observations		4641		2980	

\*  $p < 0.1$     \*\*  $p < 0.05$     \*\*\*  $p < 0.01$     Notes: Each coefficient represents a separate regression of the job growth rate for that industry or age group on the natural log of a state's real minimum wage. Robust standard errors are clustered by state and reported in parentheses. All regressions include state fixed effects, Region-by-quarterly fixed effects, state linear time trends, and the additional controls. The BDS does not report state-level employment by industry or age, and the QCEW does not report by age. The QCEW provides consistent (NAICS) industry codes beginning in 1990. See the notes for Table 2 for additional information.

## A Illustration of Staggered Treatment

This appendix section briefly demonstrates how inference can go awry for situations in which there are staggered treatments. Consider two states, A and B. State A is treated at some time  $t$ , while State B is treated at some time  $s$ , with  $s > t$ . The magnitude of the treatment is the same for both states (the simplest case is for treatment to take on a value of either 0 or 1). Prior to any treatment, both states exhibit similar trends in employment. This is illustrated in Figure 5. In panel (a), there is no treatment effect. In panel (b), treatment has a *discrete* and symmetric negative effect on the employment level. And, in panel (c), the treatment has a symmetric negative effect on employment growth, but does not discretely alter the employment level. For simplicity of exposition, let the time spanned prior to treatment in State A equal the time spanned following treatment in State B.

Now, consider two outcomes in a difference-in-differences (DiD) framework: (1) the employment level, and (2) employment growth. In panel (a), the DiD estimate using either outcome will be zero. For panel (b), DiD estimates a negative treatment effect for the employment level, but estimates a null treatment effect on job growth. In panel (c), DiD estimates a null treatment effect for the employment level, but a negative treatment effect for employment growth. (These null results are partly due to the assumed equivalence of the time period prior to “A treated” with that following “B treated.” But, more generally, if the true treatment effect is on the employment level, then DiD estimates for the effect on job growth will be close to zero, and vice versa.)

The null effect identified for the employment level in Figure 5(c) at first seems counter-intuitive to the difference-in-differences paradigm: after all, the outcomes clearly move closer together following treatment in state A. The discrepancy results from the staggering of treatment across the two states. Recall the basic DiD equation:

$$Y_{it} = \delta_B + \tau_t + \beta * I(\text{Treatment}_{it} = 1) + u_{it}$$

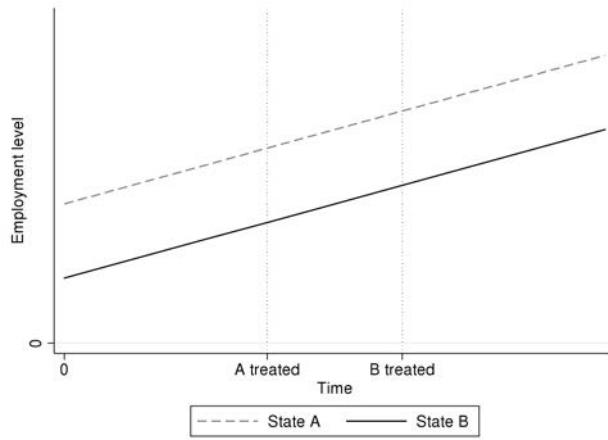
In this example, the only time period(s) in which  $\beta$  may be identified separately from  $\tau_t$  are those during which *only* State A is treated. For all other time periods,  $I(\text{Treatment}_{it} = 1)$  takes the same value for both states.

In cases for which the true treatment effect is null, as in panel (a), then the conclusions drawn from either outcome will be the same. In contrast, the inference from DiD results for panels (b) and (c) will be very different depending on which outcome is used. In either of these cases, estimations using the “wrong” outcome will be substantially attenuated (and

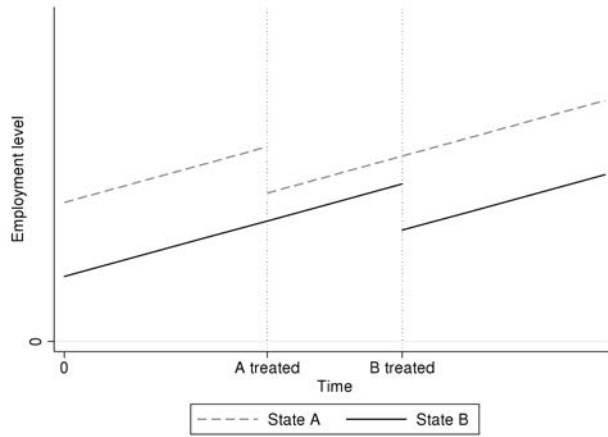
the same would hold if the treatment effect were positive instead of negative).

This type of staggering of state treatments is common in empirical settings for which difference-in-differences techniques are employed. As shown in Figures 1 and 2 (and 8 and 9), this is certainly the case for state minimum wages. This simple illustration has shown how examining only one outcome – either the employment level or employment growth – could lead to dramatically incorrect inference about the treatment effect of the minimum wage. And, if the treatment truly has no effect, then both outcomes will show no effect.

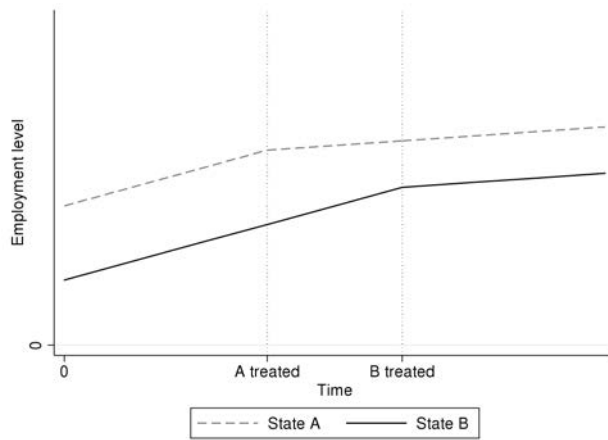




(a) No treatment effect



(b) Treatment effect discrete in levels



(c) Treatment effect discrete in growth

Figure 5: Illustration of three types of staggered treatment effects

## B Attenuated Estimates for Employment Effects

In this appendix section, we illustrate the potential attenuation problem induced by including state-specific time trends as a control. We begin with a very simple example to illustrate how – if a policy’s true effect is in a state’s growth rate – including state time trends as controls yields biased and misleading difference-in-differences results for the outcome in levels. Then, we conduct a more formal Monte Carlo simulation exercise that underscores this attenuation problem and shows how re-defining the outcome as a growth rate offers a compelling solution. Throughout this appendix, we selected numerical values that as closely as possible match the statistical properties of the actual U.S. employment data used in our study.

### B.1 Example of a Disemployment Effect in the Growth Rate

Here, we provide a simple illustration of how including state time trends as controls can sharply attenuate difference-in-differences results, misleading inference. We construct eleven periods (-5 through 5) for two hypothetical states, one of which is treated following period zero. For simplicity, we set the baseline log-employment to zero for the control state and 0.03 for the treated state. During the “pre” periods (through zero), neither state is treated and both states have an employment growth rate of 0.02 log-points per period. During the “post” periods (one through five) the control state maintains this same growth rate, but growth in the treated state drops to 0.015 log-points. Figure 6 illustrates these time paths for the treated state with a solid black line and the control state in grey. The time trend for the treated state is also plotted by a dashed black line. By construction, the control state’s time trend perfectly coincides with its log-employment level.

In Figure 7, we compare the difference-in-differences of outcomes between these hypothetical states, without versus with state linear time trends. Panel (a) graphically presents the canonical difference-in-differences model. As is immediately evident in Figure 6, the difference between the states throughout the pre-treatment period is constant at 0.03 log-points. Following treatment, employment between states steadily converges, and the average difference during the post-treatment period is 0.015 log-points. The simple difference-in-differences treatment effect for employment is thus the vertical difference between the two dashed lines: 0.015 minus 0.03, or 50% of the initial difference in state employment.

Panel (b) of Figure 7, rather than showing differences in levels, plots the differences in residuals about each state’s time trend. For the control state, which exhibits perfectly linear employment growth, the residuals about trend are always zero, so the heights of

the bars mirror residuals in the treated state. As in Panel (a), the dashed lines show the average difference for the pre- and post-treatment periods. In sharp contrast to the average differences in levels shown Panel (a), the average difference in residuals-to-trend is negligibly small in both periods. On average, the difference is 0.00057 pre-treatment and -0.00068 post-treatment. The difference-in-differences is -0.00125 log-points, roughly eight percent of the magnitude in Panel (a), or 4% of the initial difference in state employment.

The inclusion of state time trends in this toy example leads to strikingly misleading inference for the disemployment effect. Moreover, the pre-treatment time paths for the two states are identical in trend by construction, so including state time trends in this model cannot correct for any confounding selection into treatment. In real data, such similarity in pre-treatment trends is not always the case, as we discuss in Section 2, such that omitting state time trends is not a universally advisable solution. Instead, we argue for examining growth rates directly. This approach is supported by the Monte Carlo simulation below.

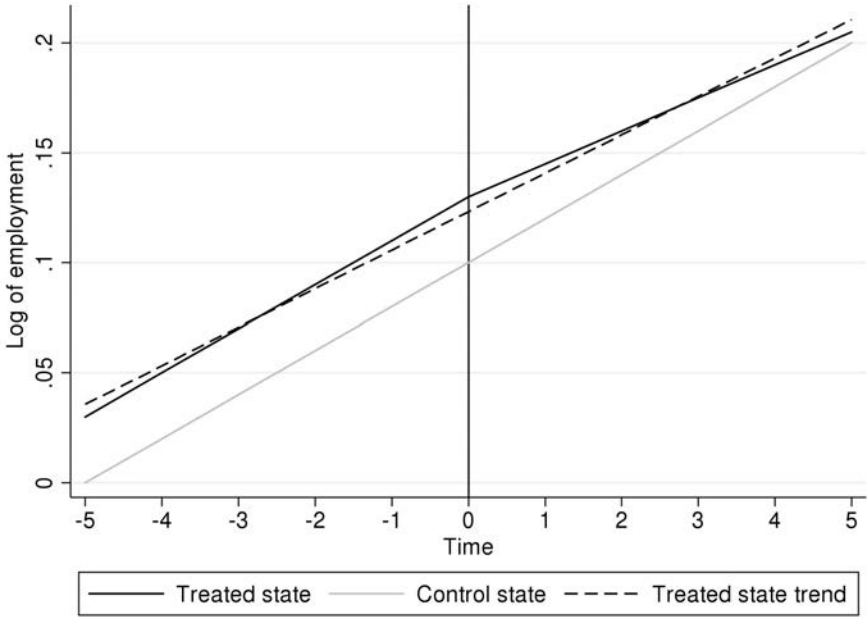


Figure 6: Simple example of disemployment effect in growth rate

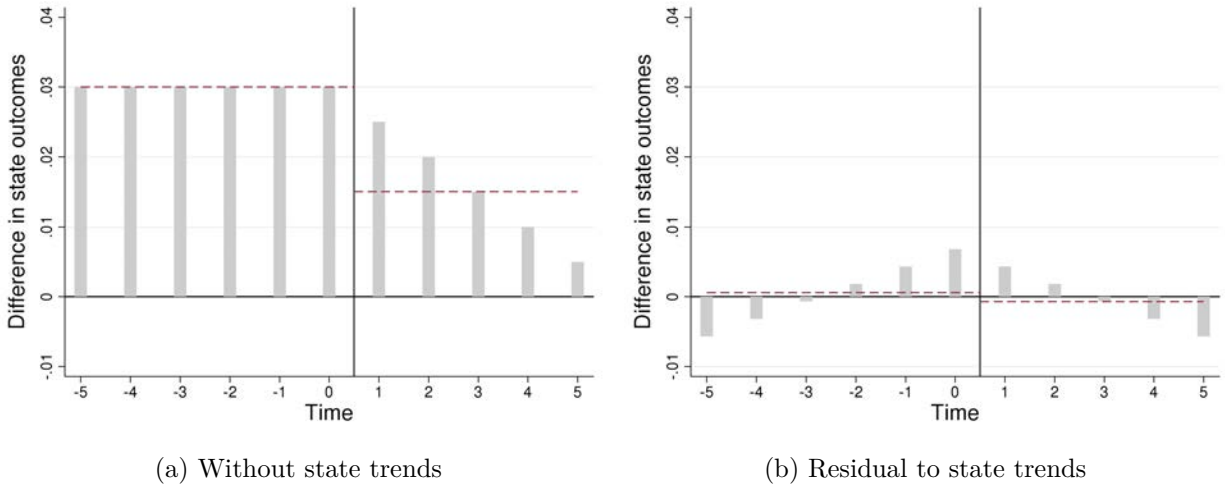


Figure 7: Difference-in-differences without versus with state linear trends

## B.2 Monte Carlo Simulation Exercise

This simulation further illustrates a particular type of attenuation problem for estimates of the employment effect of the minimum wage. The attenuation occurs when the following three conditions hold: (1) the true effect of the policy is on the growth rate of employment, as depicted in the previous example; (2) the employment (log) level is used as the regression outcome; and, (3) the econometric specification includes state time trends as controls. Support for the first condition is discussed at length in Section 2 of the article. The latter two conditions are fairly common in the literature. Section 2 of the article and the previous sub-section of this appendix additionally include discussion of *how* state time trends induce attenuation. Here, we focus on the mechanics of this simulation exercise.

### B.2.1 Data Generating Process

We use the actual distribution of changes in state minimum wages and employment as a starting point for our simulated data. Specifically, we compute the first difference in each year for each state's real log minimum wage and log employment. We include values regardless of whether there was a nominal increase in the state's minimum wage, so the sign of many changes in minimum wage is negative. This yields 34 periods of observations on 51 state entities, or 1734 total observed changes in log-employment and minimum wages. Next, we strip these data values of their state and year identifiers and unlink changes in minimum wages from their respective changes in employment. This leaves us with two independent

distributions of changes, one for real log minimum wages and one for log-employment; each has 1734 data values.

## B.2.2 Steps in Each Monte Carlo Repetition

Within each Monte Carlo repetition, we perform the following steps. First, we draw values without replacement from the distributions of first differences in employment and minimum wages. We assign these values randomly to 34 periods of 51 states, thereby forming a simulated panel of state-year data.  $\Delta\ln(\text{employment})_{st}$  and  $\Delta\ln(\text{MW})_{st}$  denote the first differences in (log) employment and real minimum wage, respectively, in state  $s$  in period  $t$ , relative to the previous period ( $t-1$ ).

Next, we impose a treatment effect relating the minimum wage to the growth rate of employment. To prevent the effect from being purely deterministic, we draw a parameter  $\phi_{st}$  from a  $Normal(-0.05, 0.02)$  distribution for each state-year observation. That is, each 10% increase in a state's real minimum wage causes, in expectation, a 0.5 percentage point reduction in employment growth.

Because the effect is on the employment growth rate, the treatment effect in a state in one year persists throughout all future years, although it may be eroded by treatment effects associated with future real decreases in that state's minimum wage. Letting  $\alpha_{st}$  denote the treatment effect in a given state-year, this is:

$$\alpha_{st} = \sum_{r=1}^t \phi_{sr} \cdot \Delta\ln(\text{MW})_{sr} \quad \phi_{sr} \sim N(-0.05, 0.02) \quad \forall t \in [1, 34]$$

For each state-year, we then add  $\alpha_{st}$  to the previously assigned first difference in employment, forming a new growth pattern which encompasses these treatment effects. Although we include state fixed effects in each specification below, we use the actual employment level for each state in 1977 to set  $\ln(\text{employment})_{s0}$  for each of the 51 simulated states. We set  $\ln(\text{MW})_{s0}$  for all states equal to zero. Formally:

$$\ln(\text{employment})_{st} = \ln(\text{employment})_{st-1} + \Delta\ln(\text{employment})_{st} + \alpha_{st}$$

$$\ln(\text{MW})_{st} = \ln(\text{MW})_{st-1} + \Delta\ln(\text{MW})_{st}$$

Using these equations, we simulate a panel of log-employment and minimum wages in 51 states in 34 periods, which encompasses the stipulated treatment effect on employment growth. We use this panel to estimate four specifications relating the minimum wage to

employment outcomes:

$$\text{employment growth}_{st} = \beta \cdot \ln(\text{MW})_{st} + \phi_s + \tau_t + \epsilon_{st} \quad (1)$$

$$\text{employment growth}_{st} = \beta \cdot \ln(\text{MW})_{st} + \phi_s + \tau_t + \psi_s \cdot t + \epsilon_{st} \quad (2)$$

$$\ln(\text{employment})_{st} = \beta \cdot \ln(\text{MW})_{st} + \phi_s + \tau_t + \epsilon_{st} \quad (3)$$

$$\ln(\text{employment})_{st} = \beta \cdot \ln(\text{MW})_{st} + \phi_s + \tau_t + \psi_s \cdot t + \epsilon_{st} \quad (4)$$

Equations (1) and (3) do not include state time trends, while equations (2) and (4) do. In each Monte Carlo repetition, we store the point estimate for each  $\beta$  in each of these specifications, ignoring the standard errors.

### B.2.3 Results of Simulation Exercise

We conducted 10,000 repetitions of the above steps, which yields 10,000 point estimates for each  $\beta$  value in equations (1) - (4). In Table 5, we present coefficients at the first percentile, the median, and the 99th percentile of these distributions.<sup>29</sup>

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<sup>29</sup>The full code used in this simulation, along with all other code and data included in this study, is available from the authors online and by request.

Table 5: Effects of the Minimum Wage on Simulated Employment Outcomes

Coefficients:	1 <sup>st</sup> Pctl.	Median	99 <sup>th</sup> Pctl.
Job growth			
[0] Simulated true effect	-.0505	-.05	-.0495
[1] Estimate without linear trends	-.0669	-.05	-.0332
[2] Estimate including linear trends	-.0752	-.0502	-.0244
Log of employment			
[3] Estimate without linear trends	-.659	-.446	-.227
[4] Estimate including linear trends	-.141	-.0248	.09035

Notes: Estimated coefficients result from regressing the outcomes (in rows) on the log of the minimum wage in the simulated data. Reported values are at the first percentile, the median, and the 99th percentile from a Monte Carlo simulation of 10,000 repetitions. The true elasticity for growth is  $Normal(-0.05, 0.02)$  by construction in the simulated data.

Row [0] of the coefficients in Table 5 presents the true effect of the minimum wage on job growth by construction in the simulated data. Note that because we draw the coefficient from a  $Normal(-0.05, 0.02)$  distribution, there is a standard error on the coefficient, but the resulting distribution of coefficients for the true effect is compact. Row [1] corresponds to Equation (1) above, relating the minimum wage to job growth in a specification without state time trends. The median coefficient is identical to the true simulated effect, and we can rule out an estimated null effect of the minimum wage on simulated job growth with high confidence. We add a state time trend to the specification in Row [2], but again the median coefficient is nearly identical to the true effect.

Results for the employment level are presented in Rows [3] and [4]. Because of the randomization that was used in generating the simulated data, there is no “true” coefficient for the effect of the minimum wage on employment in these simulated data; the exact extent to which the effect on job growth is reflected in the employment level depends partly on the (random) ordering of the simulated changes to state minimum wages. However, given that in this simulation the minimum wage has a large negative effect on job growth by construction, it is reasonable to expect a fairly large – and certainly non-zero – effect on the employment level. This is indeed the case in Row [3], which omits state time trends. But, including state time trends in Row [4] attenuates the estimate to a (small and statistically insignificant) zero. This attenuation occurs despite: (1) a large true effect on job growth by construction;

and (2) no systematic correlation of changes to state minimum wages with state effects, year effects, or state time trends.



## C Historical Minimum Wage Increases and Erosion

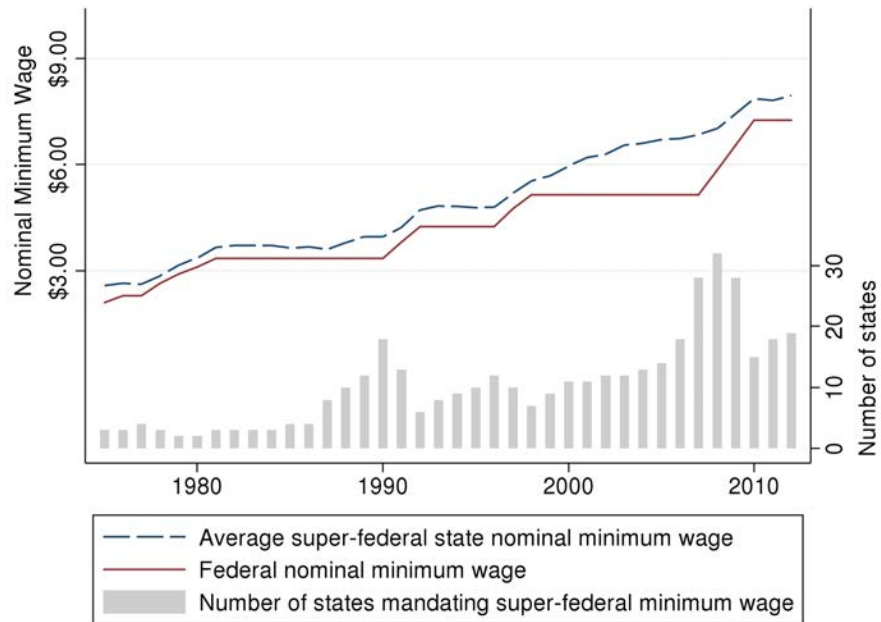


Figure 8: Comparison of federal to state nominal minimum wages (January, 1975-2012)

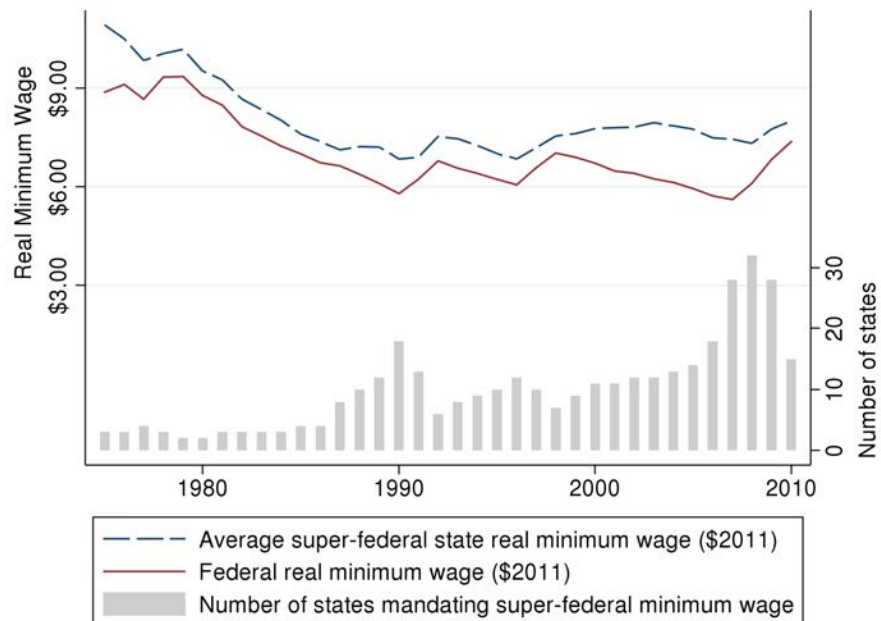


Figure 9: Comparison of federal to state real minimum wages (January, 1975-2012))

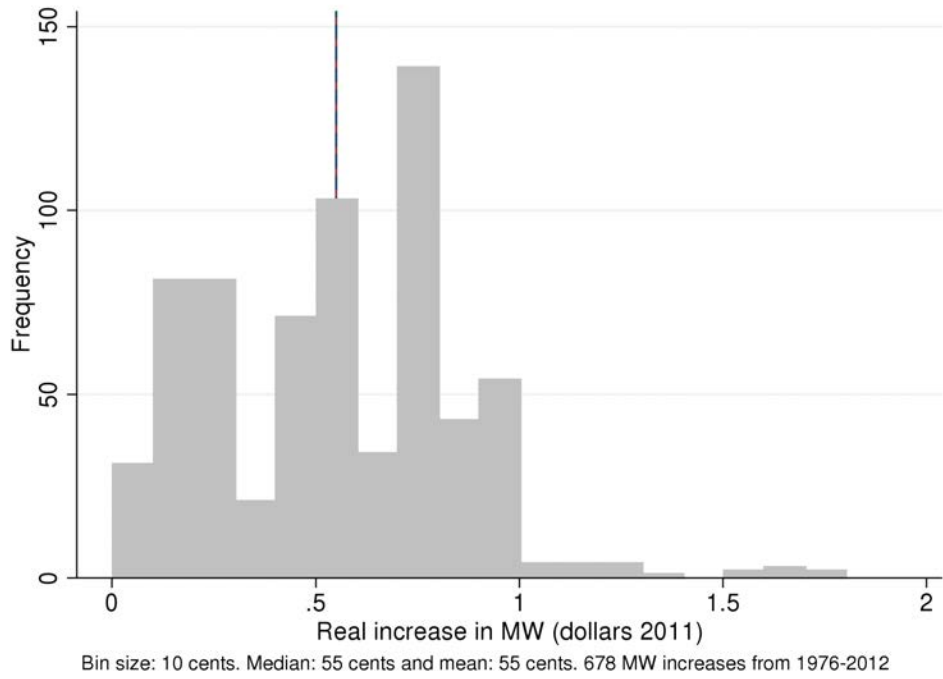


Figure 10: Distribution of real minimum wage increases

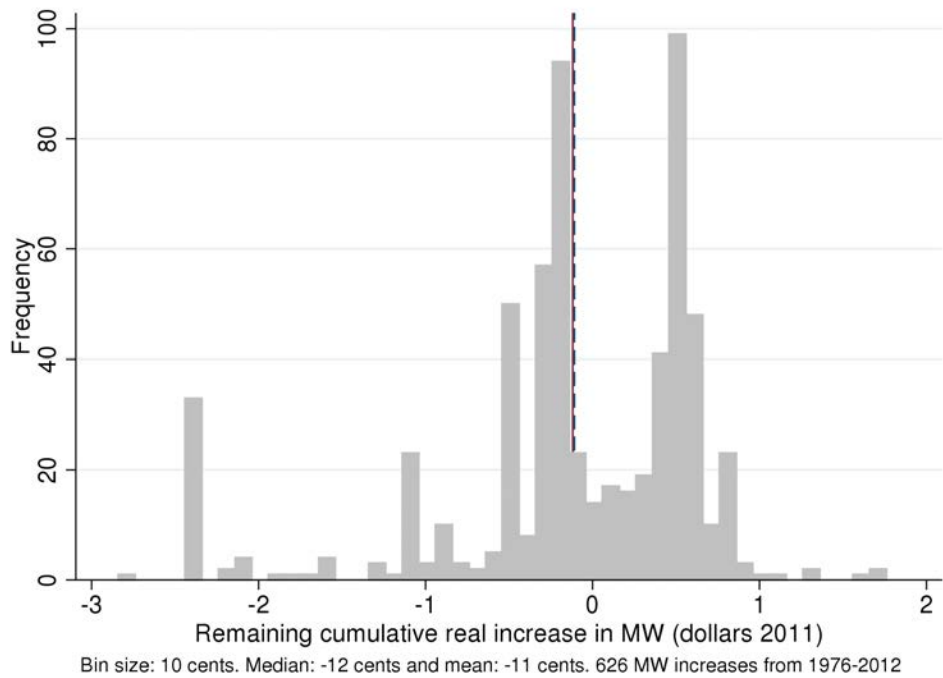


Figure 11: Cumulative difference in real minimum wage prior to a new increase

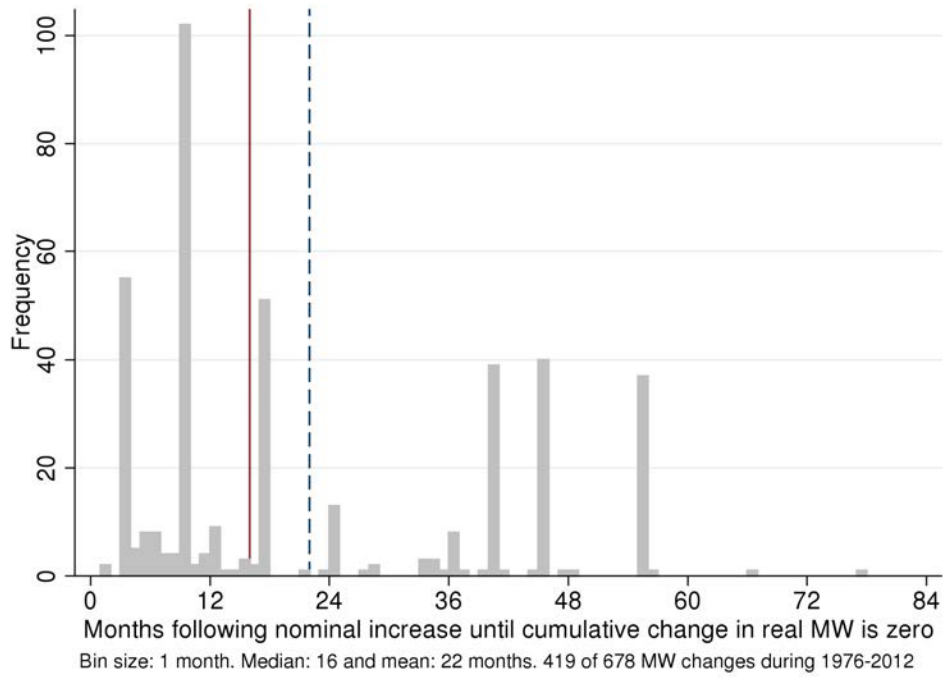


Figure 12: Erosion of real increases in minimum wage

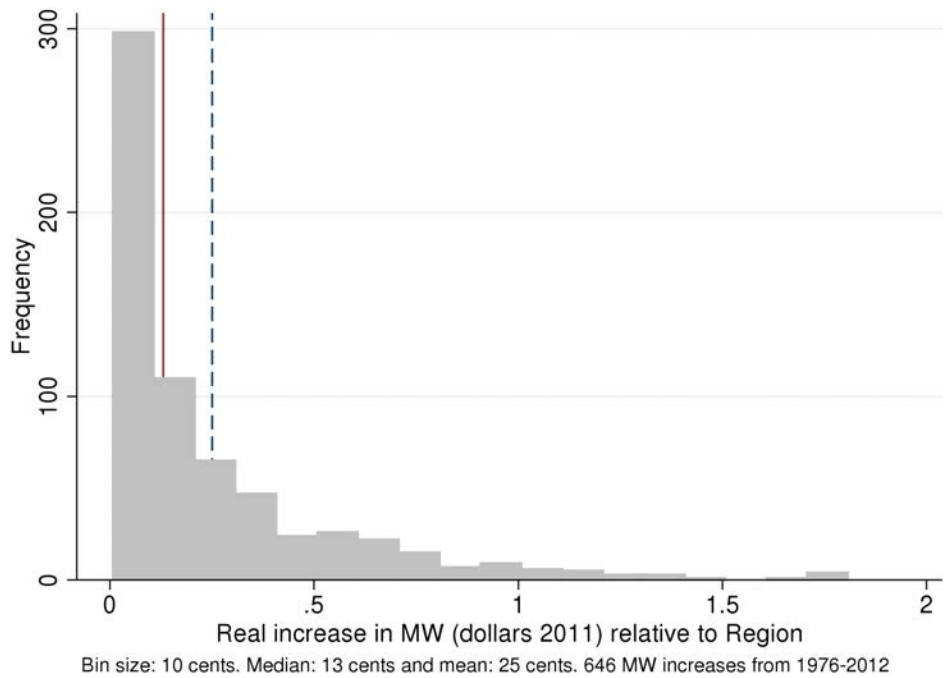


Figure 13: Distribution of relative minimum wage increases

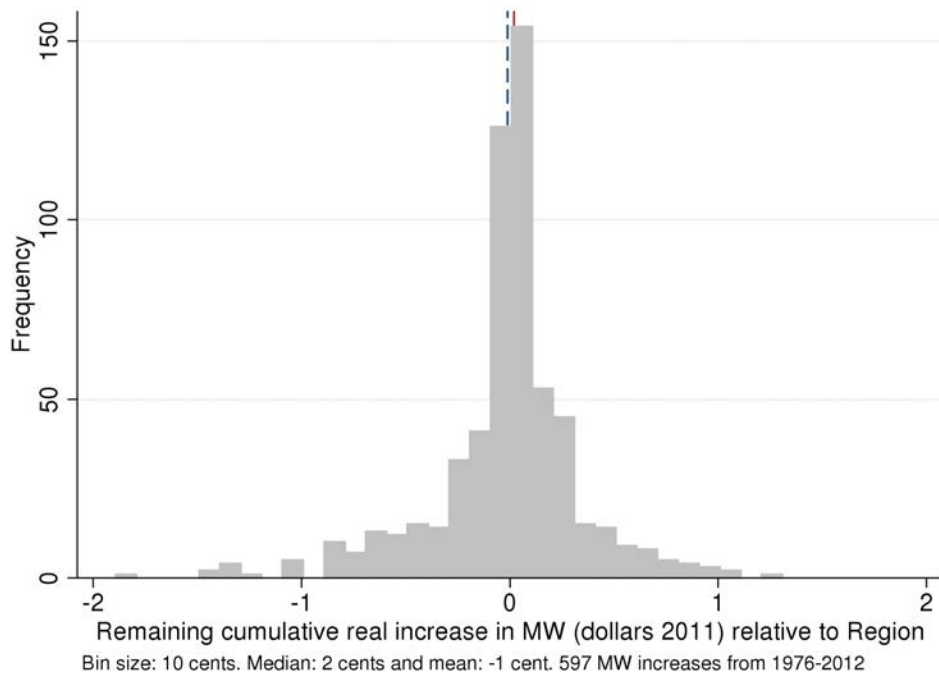


Figure 14: Cumulative difference in relative minimum wage prior to a new increase

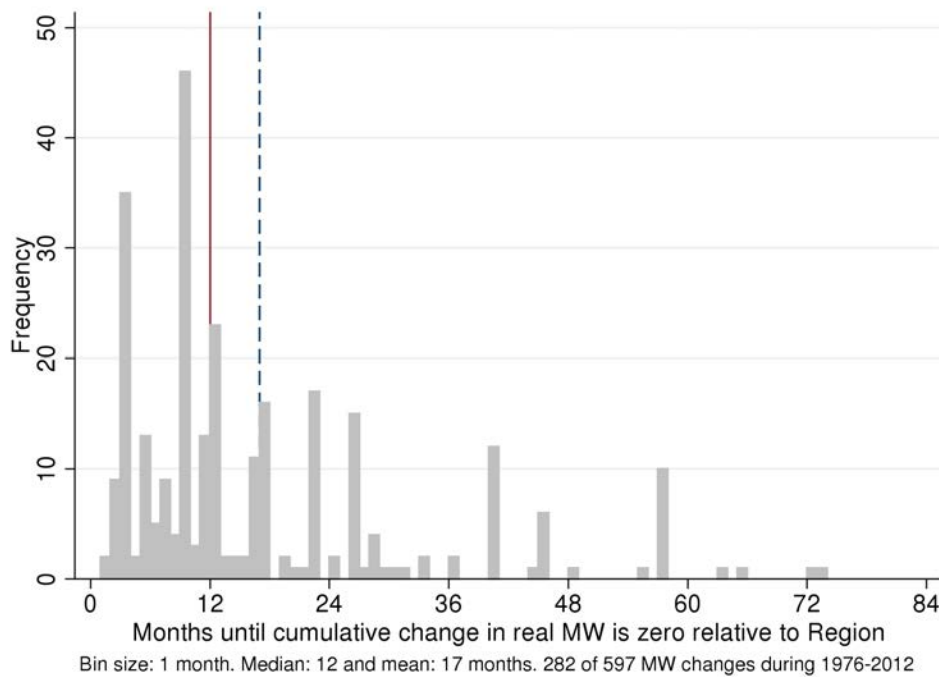


Figure 15: Erosion of relative increases in minimum wage