MINIMUM WAGES AND THE BUSINESS CYCLE:

Does a Wage Hike Hurt More in a Weak Economy?

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Employment Policies
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Executive Summary

In his 2013 State of the Union address, President Obama called for an increase in the federal minimum wage to $9 an hour. The following month, Senator Tom Harkin (D-IA) and Representative George Miller (D-CA) introduced a bill to raise the wage even higher to $10.10 an hour.

One year later, President Obama as well as Congressmen Harkin and Miller have agreed on the $10.10 figure. Less clear, however, is whether the empirical evidence shows that a higher minimum wage is good public policy.

Some policymakers have expressed concern about forcing higher labor costs on low-margin businesses in a still-recovering economy. Economists have questioned the impact on the entry-level labor market: A summary of the last two decades of research on the subject, authored by economists at University of California-Irvine and the Federal Reserve Board, finds that raising the minimum wage reduces employment for the least-skilled and least-experienced employees.

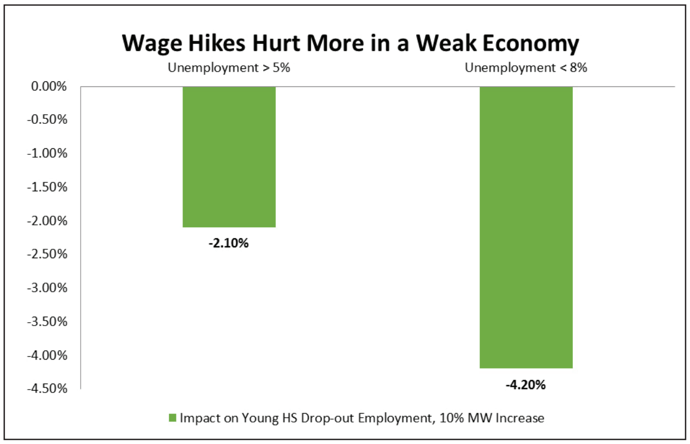
In this new study, Dr. Joseph Sabia of San Diego State University analyzes Census Bureau data to measure the impact of a higher minimum wage in periods of strong and weak economic growth. His results suggest that raising the minimum wage is rarely a good idea, and that it's particularly risky in times of weak economic growth.

Sabia's research draws on data from 1990 through 2010, which covers the last three increases in the federal minimum wage and numerous increases in state minimum wages. Sabia makes use of this state-level variation to determine the impact of a higher minimum wage on the employment of less-skilled workers, carefully controlling for other factors that could affect their employment.

Sabia finds that, over the past two decades, each ten percent increase in the minimum wage has reduced employment for less-educated young adults by as much as 2.3 percent. However, this top-line result masks important variation in the effects of the minimum wage that depends on the economy.

For instance, in tight labor markets—when prime-age (age 25-54) male unemployment is below five percent

—the minimum wage reduces employment for young drop-outs by roughly two percent. However, in weak labor markets—when the prime-age male unemployment jumps above eight percent—the impact of a higher minimum wage more than doubles. Specifically, each 10 percent minimum wage increase reduces employment for young drop-outs by over four percent.



Particularly at a time when unemployment for young adults has been above 20 percent for more than five years, Sabia's results further support the idea that raising the minimum wage would hurt those employees it's intended to help. And his finding that the effects of a higher minimum wage are exacerbated in a weak economy suggests that the drawbacks of setting the minimum wage to rise automatically—regardless of the state of the economy—should be considered carefully by policymakers on the local, state, and federal level.

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Introduction

A recent review of the minimum wage literature by Neumark and Wascher (2007, 2008) concluded that most credible recent studies of the low-skilled employment effects of minimum wage increases point to a return to the "old consensus" of disemployment elasticities in a range from -0.1 to -0.3. However, since the publication of this review, a flurry of new studies by Dube, Lester, and Reich (2011); Addison, Blackburn and Cotti (2009, 2011); and Allegretto, Dube, and Reich (2011) find that minimum wage increases have no effect on low-skilled employment. These authors suggest that failing to account for heterogeneous macroeconomic trends across labor markets may result in an overstatement of the low-skilled employment effects of minimum wage increases. While separating the low-skilled employment effects of minimum wage increases from trends due to macroeconomic shocks is not a new concern in this literature (see, for example, Deere et al., 1995; Burkhauser, Couch, and Wittenberg 2000), the question of how best to address heterogeneous economic trends across U.S. states has reached a new salience with the current set of studies (Neumark Salas, and Wascher 2013).

Among the recent approaches to address heterogeneous economic trends across regions (spatial heterogeneity) include using more highly-skilled control groups along with a synthetic cohort design (Sabia, Burkhauser, and Hansen 2012), comparing contiguous counties across state borders with different minimum wages (Dube, Lester, and Reich 2011), and including region- or state-specific time trends as additional controls (Allegretto, Dube, and Reich 2011). Using the second and third approaches, Dube, Lester, and Reich (hereafter DLR) and Allegretto, Dube, and Reich (hereafter ADR) conclude that prior estimates of the low-skilled employment effects of minimum wage increases obtained using a standard difference-in-difference approach are biased upward in absolute magnitude.

One interpretation of the findings of DLR and ADR is that minimum wage increases are more likely to be enacted by states during periods of negative shocks to low-skilled labor markets and that the use of "better" control groups can ameliorate this bias due to policy endogeneity. However, this interpretation is at odds with substantial evidence that minimum wage increases are enacted pro-cyclically (Reich 2009, Baskaya and Rubinstein 2011) and new evidence that the control groups chosen by DLR and ADR may actually be worse than those used in the traditional minimum wage literature (Neumark, Salas, and Wascher 2013). An alternate interpretation of DLR's and ADR's findings is that the authors "throw the baby out with the bath water" by eliminating valid sources of identifying variation in state minimum wages (Neumark, Salas, and Wascher 2013).

Using data drawn from the Current Population Survey (CPS) from 1990 to 2010, the current study re-examines the low-skilled employment effects of minimum wage increases, with careful attention to both spatial heterogeneity and how minimum wage effects differ across the state business cycle. I find that adding observable controls for state-specific economic shock produces a pattern of results that is inconsistent with the hypothesis that minimum wage increases are enacted anti-cyclically. Controlling for spatial heterogeneity via state-specific productivity shocks to low-skilled sectors and state-specific non-linear time trends, I find that minimum wage increases between 1990 and 2010 reduced employment of teenagers and younger high school dropouts. While the average estimated employment elasticities over this two decade period are in the consensus range of -0.1 to -0.3, these average elasticities mask potentially important differences across peaks and troughs in the state business cycle.

To measure trends in the state business cycle that are unrelated to the minimum wage, I use two measures: the prime-age (ages 25-to-54) male state unemployment rate and year-to-year growth in state gross domestic product (GDP) generated by the finance sector. The results show relatively larger negative low-skilled employment elasticities during periods of high unemployment and sluggish GDP growth than during peaks in the state business cycle. Low-skilled employment elasticities during troughs are as large as -0.45.

Background

Because minimum wages are usually endogenously determined by policymakers, separating the labor market effects of minimum wage increases from broader economic trends has been a concern in the minimum wage literature for some time (Card and Krueger, 1995; Deere et al. 1995; Burkhauser et al. 2000).

Disentangling minimum wage effects from the effects of macroeconomic trends was at the heart of an important study by Burkhauser et al (2000). These authors attempted to resolve the findings of two Current Population Survey (CPS)-based studies—Card and Krueger (1995) and Deere et al. (1995). Using within-state variation in state and federal minimum wages, Card and Krueger (1995) found no evidence that minimum wage increases between 1979 and 1992 reduced employment of teenagers; however, using the same data source, Deere et al. (1995) found that the 1990-91 federal minimum wage increase reduced employment of teenagers.

Burkhauser et al. (2000) concluded that difference in their findings could be explained by the respective authors' treatment of year dummies in their specifications. While Card and Krueger (1995) included year dummies to, in part, account for national macroeconomic trends that could be correlated with minimum wages and low-skilled employment, Deere et al. (1995) interpreted year effects during the 1990-91 period as the impact of federal minimum wage changes. Burkhauser et al. (2000) showed that in the sample period examined by Card and Krueger (1995), year effects explained over 90 percent of the variation in the minimum wage. Thus, the estimation of employment effects of the minimum wage in regressions that include year effects were hampered by limited identifying variation and hence, low precision. Burkhauser et al. (2000) further showed that in their preferred specification—in which year effects were omitted, but controls for national macroeconomic recessions were included—minimum wage increases reduced employment of teenagers.

The tradeoff between bias and efficiency in estimates of the employment effects of minimum wages was also raised by Sabia (2009). He showed that with increased variation in state minimum wages in the late 1990s and early 2000s, the inclusion of year effects did not diminish identifying variation sufficiently so as to preclude finding significant negative employment effects for teenagers. Sabia (2009) found teenage employment elasticities with respect to the minimum wage of around -0.2 in the specification preferred by Card and Krueger (1995).

With a general consensus that controlling for national macroeconomic trends via year effects is appropriate in the estimation of minimum wage effects in state panels (Neumark and Wascher 2005), recent attention has now turned to whether local geographic trends may be correlated with minimum wages and low-skilled employment. To address concerns about heterogeneous geographic trends, DLR (2011) restrict their samples to pairs of counties across state borders in the hopes of comparing comparable labor markets with different minimum wages. With this approach, the authors find that minimum wage increases are associated with no change in employment in low-skilled sectors, including the restaurant industry. The authors conclude that failing to select "treatment" and "control" regions with similar underlying macroeconomic trends can lead to misattribution of negative employment trends to the minimum wage.

Several critiques have arisen in response to the work of DLR (2011). First, as Neumark and Wascher (2008) note, examining industry-wide employment effects can mask important labor-labor substitution across heterogeneous skill levels within industries. Second, Addison, Blackburn, and Cotti (2011) point out that the authors use some border counties multiple times, with each being assigned its own county fixed effect.[1] Finally, Neumark, Salas, and Wascher (2013) note several important shortcomings of the approach taken by DLR: (i) the data fail to justify DLR's exclusion of alternate non-border counties or regions as controls, as many non-border counties look more similar on observables to treatment counties, (ii) DLR's findings are quite sensitive to the number of leads and lags of the minimum wage included in their model, and (iii) DLR's selection of matching counties often produces matched pairs that are quite dissimilar across an important set of observables; when the matched pairs of nearby counties and states are restricted to better controls, negative employment effects reemerge.

Another approach to address spatial heterogeneity in employment trends has been to include geographic-specific time trends. ADR (2011) draw data on teenagers from the CPS and find that in the standard difference-in-difference model estimated by Sabia (2009), minimum wages are associated with modest declines in teen employment. However, after the addition of controls for census division-specific time shocks or state-specific linear time trends, estimated employment elasticities are small and statistically indistinguishable from zero. An interpretation of these findings offered by ADR is that state minimum wages are more likely to be enacted during times of negative economic shocks to low-skilled labor markets. But this interpretation is at odds with much of the legislative history and with prior evidence that minimum wages are pro-cyclical (Neumark, Salas, and Wascher 2013). For instance, Reich (2009) notes:

> "Minimum-wage increases are voted almost without exception *and are mostly implemented* in times of growing employment. This pattern holds for both federal and state increases." (p. 366)

Baskaya and Rubinstein (2011) also find evidence that state minimum wage increases are pro-cyclical. These findings imply that the adverse low-skilled employment effects of state minimum wage increases should be *stronger* (that is, larger in absolute magnitude) rather than weaker after controlling for special heterogeneity, as failure to control for such heterogeneity should bias estimates *toward* zero.

Neumark, Salas, and Wascher (2013) instead suggest a potentially more plausible interpretation of ADR's findings: the null effects can be explained by their eliminating potentially valid sources of identifying variation with the inclusion of controls for state-specific linear time trends. They show that controlling for higher-order polynomials in state time trends rather than linear time trends results in negative teen employment elasticities in line with consensus estimates. In addition, Neumark, Salas and Wascher (2013) critique ADR's assertion that within-census division state comparisons are more credible that cross-census division state comparisons by showing that (i) states outside a census division are often a better match for "treatment" states when one examines low-skilled economic conditions prior to a minimum wage hike, and (ii) even within census divisions, minimum wage increases still adversely affect teen employment within four of nine divisions; in the remaining five

[1]For instance, the authors note that it is unclear how to conceptually interpret differences in estimates of low-skilled employment levels for the same county depending on the county-pair to which it is matched.

divisions, produce employment elasticities that are too imprecisely estimated to reject a finding of negative employment effects.

Taken together, the recent work of ADR and the critique by Neumark, Salas, and Wascher (2013) highlight the importance of empirically addressing the role of spatial heterogeneity, while not eliminating potentially valid sources of identifying variation in the minimum wage.

While much attention has been paid to separating the effects of minimum wage increases from other underlying state economic trends, relatively little work has been done on whether the employment effects of minimum wage increases differ over the state business cycle. If, for instance, when labor markets are more slack, employers are likely to first lay off less low-skilled workers, the effects of minimum wage increases during troughs of the business cycle may be larger (in absolute magnitude) than during peaks. Along the same lines, during tight labor markets with a growing economy, increases in aggregate demand could blunt the adverse employment effects of minimum wage hikes.

ADR explore heterogeneity in the effects of the minimum wage across the business cycle using the aggregate state unemployment rate as a measure of the health of the macroeconomy. With their preferred controls for special heterogeneity (state-specific linear time trends and census division-specific year effects) they find no effects of the minimum wage during peaks or toughs in the state business cycle. One concern about their estimates is, as noted above, the restrictive form of their geographic-specific time shocks to address spatial heterogeneity. A second concern is that their aggregate unemployment rate includes teen and other low-skilled employment, which could be affected by increases in the minimum wage.

The current study builds on the work of ADR and the critique by Neumark, Salas, and Wascher (2013) by carefully addressing the role of spatial heterogeneity and (i) extending the examination of low-skilled populations affected by the minimum wage to include younger high school dropouts and (ii) examining whether there is heterogeneity in the employment effects of the minimum wage across the state business cycle using measures of the business cycle that are more plausibly exogenous to minimum wage increases.

Data and Methods

This analysis draws data from the CPS Merged Outgoing Rotation Group MORG from 1990 to 2010 to estimate the effects of minimum wage increases on low-skilled employment. Following Card and Krueger (1995) and Sabia (2009), state-by-year aggregate variables are generated from the repeated cross sections. First, as a benchmark, a standard difference-in-difference model is estimated:

$$E_{st} = \alpha + \delta_1 MW_{st} + \delta_2' X_{st} + \alpha_s + \tau_t + \varepsilon_{st}, \qquad (1)$$

where E_{st} is the natural log of the employment to population ratio of teenagers ages 16 to 19 (or individuals ages 16 to 24 without a high school diploma) in state *s* at time *t*; X_{st} is a vector of state-specific socio-demographic controls for the (natural log of the) percent of the population ages 16-to-19, the percent of the population ages 55 to 65, and the percent of the population ages 16 to 64 who are married; MW_{st} is the natural log of the higher of the state or federal minimum wage; α_s is a vector of state fixed effects; and τ_t is a set of year fixed effects. In equation (1), identification of δ_1, which can be interpreted as the low-skilled employment elasticity with respect to the minimum wage, comes from within-state variation in the minimum wage over time, generally via state changes in the minimum wage but also from differential state-specific increases due to federal minimum wage changes. During the 1990-2010 period, there were three federal minimum wage changes (1990-91, 1996-97, and 2007-09) and over 30 states changed their minimum wages (see Department of Labor 2012 and Neumark, Salas, and Wascher 2013 for a discussion of these state minimum wage changes).

Next, to explore the role of observable spatial heterogeneity related to state economic conditions, I add a set of state-specific time-varying economic controls:

$$E_{st} = \alpha + \delta_1 MW_{st} + \delta_2\text{'}\boldsymbol{X}_{st} + \delta_3\text{'}\boldsymbol{Z}_{st} + \alpha_s + \tau_t + \varepsilon_{st}, \qquad (2)$$

where $\boldsymbol{Z}_{st}$ is a vector of economic controls commonly employed in the literature (Card and Krueger 1995; Burkhauser et al. 2000; Sabia 2009; ADR 2011), including the high school completion rate, the prime-age (ages 25-54) male unemployment rate, and the prime age wage rate. In addition, I also control for state-specific economic shocks that may affect low-skilled workers by controlling for state gross domestic product generated by the retail and manufacturing sectors.[2] These controls are designed to capture spatial heterogeneity specifically related to low-skilled productivity shocks.

A comparison of the estimates of δ_1 from equations (1) and (2) should shed light on whether minimum wages are likely to be implemented during times of negative economic shocks as ADR imply, or pro-cyclically as others suggest (Reich 2009, Baskaya and Rubinstein 2011). That is, if the absolute magnitude of the estimated employment elasticity is smaller in equation (2) than in equation (1), this would tend to support ADR's hypothesis.

To examine whether the employment effects of minimum wage hikes differ across the state business cycle, I use two measures of the state business cycle that are plausibly exogenous to the minimum wage: the prime-age male state unemployment rate and state-specific year-to-year growth in gross domestic product (GDP) generated by the finance sector. A main effect for each of these measures is included on the right hand-side of equation (2) along with an interaction of each measure and the minimum wage. I explore whether minimum wage increases have different effects across three threshold unemployment rates to capture the ranges explored by ADR (2011): less than or equal to 5.0 percent, between 5.0 and 8.0 percent, and greater than 8.0 percent. For year-to-year state GDP growth generated by the financial sector I use three thresholds for nominal growth: greater than 6.5 percent annual growth, between 2.0 and 6.5 percent annual growth, and less than 2.0 percent annual growth.

Finally, to address the role of spatial heterogeneity, I take three approaches. First, as noted above, I control for state-specific productivity in low-skilled sectors to capture geographic specific economic shocks. Second, following Neumark, Salas, and Wascher (2013) and ADR (2011), I examine states within the same census division over time to account for the fact that time shocks may differ across these regions. And finally, I allow for state-specific non-linear time trends to control for state-specific shocks to low-skilled employment unrelated to the minimum wage.

Results

Table 1 presents weighted descriptive statistics for the sample period. During the period 1990-2010, the employment-to-population ratio for teenagers was 0.394 while for 16-to-24 year-olds without a high school diploma was 0.364. The average nominal wage rate for each group was approximately \$6.50. The average nominal minimum wage during the sample period was \$5.38

[2]Note, however, that the inclusion of these GDP controls may understate employment effects of the minimum wage if minimum wage hikes reduce low-skilled GDP through adverse employment effects (Sabia 2011).

and the average prime-age male unemployment rate was 4.8 percent.

The regression results presented in the tables below focus on estimates of δ_1, which can be interpreted as elasticities given the log-log nature of the specification. Appendix Table 1 shows coefficient estimates on the controls. All models are weighted by the relevant state population and standard errors are corrected for clustering at the state-level (Bertrand et al., 2004).

Wage Effects

Before turning to employment effects, I first examine whether minimum wage increases were binding by examining the effects of such hikes on the natural log of hourly earnings of teenagers ages 16-to-19 and 16-to-24 year-old high school dropouts. Wages are measured using the directly reported wage for those who are paid hourly; for non-hourly workers, the wage rate is calculated by dividing the usual weekly earnings by usual hours worked per week. The key findings appear in Table 2. Column (1) shows difference-in-difference estimates using the basic demographic controls (as in equation 1); column (2) adds economic controls, including state-specific low-skilled GDP (as in equation 2); column (3) adds state-specific linear time trends; and column (4) adds census division-specific year effects. Thus, columns (3) and (4) focus on the spatial heterogeneity controls preferred by ADR.

The findings show that across specifications (columns 1-4) and across the samples of low-skilled workers (Panels I and II), minimum wage increases were associated with increases in hourly earnings. For teenagers ages 16-to-19, the estimated wage elasticities range from 0.071 to 0.127 and for 16-to-24 year-old high school dropouts, the wage elasticities range from 0.067 to 0.123. Thus, the evidence does suggest that minimum wage hikes were binding over the sample period for low-skilled workers. Next, I turn to employment effects.

Benchmark Employment Estimates

Table 3A presents the main findings for low-skilled employment effects. Panel I shows the results for teenagers. In column (1), I show the results of equation (1), where a 10 percent increase in the minimum wage is associated with a statistically insignificant 1 percent reduction in the teen employment to population ratio. The addition of observable state-specific economic controls (column 2) actually *increases* the absolute magnitude of the low-skilled employment elasticity by 54 percent to -0.154 elasticity, which is statistically distinguishable from zero at the 10 percent level. This finding is inconsistent with ADR's interpretation of minimum wage increases being enacted anti-cyclically. While the inclusion of a state-specific linear time trend as in ADR (column 3) diminishes the magnitude of the elasticity, I find that, like Neumark, Salas, and Wascher (2013), allowing a more flexible state-specific time trend more consistent with aggregate teen employment trend data (which includes a third-order polynomial) actually *increases* the absolute magnitude of the teen employment elasticity to -0.233 and produces an estimate statistically distinguishable from zero at the 10 percent level.

The findings for 16-to-24 year-old high school dropouts in Panel II of Table 3A show a similar pattern of results. A comparison of columns (1) and (2) shows that the low-skilled employment elasticity with respect to the minimum wage remains statistically equivalent (range of -0.230 to -0.251) with the inclusion of state-specific economic controls, providing little evidence for the hypothesis that minimum wages are more likely to be enacted during negative shocks to low-skilled employment. While the inclusion of state-specific linear time trends slightly reduces the magnitude of the estimated elasticity (column 3), a more flexible functional form to the state time trend increases the estimated employment elasticity to -0.309. These results are, therefore, consistent with the recent findings of Neumark, Salas, and Wascher (2013).

In Table 3B, I present separate estimates by census division as in Neumark, Salas, and Wascher (2013). The odd-numbered columns present estimates of equation (2) by census division and the even-numbered columns add state-specific non-linear time trends. The findings are also consistent with Neumark, Salas and Wascher (2013). Across several census divisions (New England, Middle Atlantic, and East North Central), I find evidence of significant negative employment effects for both teenagers and 16-to-24 year-olds without a high school diploma, with elasticities ranging from -0.3 to -0.8. Across other census divisions, there are no significant adverse employment effects, but the estimates are generally too imprecise to reject employment elasticities in the consensus range.

Taken together, the results in Tables 3A and 3B point to adverse low-skilled employment effects of the minimum wage that are generally robust to spatial heterogeneity as measured by (i) state-specific shocks to low-skilled productivity, (ii) state-specific non-linear time shocks, and (iii) census division-specific time effects. These findings suggest that in contrast to the conclusion offered by ADR, it is far too soon to conclude that minimum wages have no low-skilled employment effects after controlling for spatial heterogeneity.

Heterogeneity in Low-Skilled Employment Effects Over State Business Cycle

To examine whether minimum wages interact with tight or slack labor markets to produce differential low-skilled employment effects, ADR interact the overall state unemployment rate with the minimum wage. In contrast, I use the prime-age (ages 25-to-54) male unemployment rate, which is plausibly unaffected by minimum wages (Burkhauser et al. 2000). Specifically, I measure three phases of the business cycle: prime-age male unemployment rate under 5.0 percent (tight labor market), between 5.0 and 8.0 percent, and higher than 8.0 percent (slack labor market).[3] These generally mirror cutoffs tested by ADR. The model estimated then includes the prime-age male unemployment rate as a main effect and focuses on the interaction between the minimum wage and the set of indicators for different phases of the business cycle

Table 4A shows results of the effect of the interaction of the minimum wage across the business cycle. Column (1) of Panel I shows the main difference-in-difference results for teenagers ages 16 to 19. The results suggest that in tight labor markets (unemployment rate < 5.0 percent), a 10 percent increase in the minimum wage is associated with a statistically insignificant 1.24 percent decrease in teen employment. However, relative to tight labor markets, in troughs of the state business cycle, when the prime-age male unemployment rate is 8.0 percent or higher, the estimated elasticity is nearly three times larger in absolute magnitude at (-0.124 + -0.220) -0.344.

In column (2), state-specific non-linear time trends are included and the result is largely unchanged. In tight labor markets, the estimated teen employment elasticity is -0.197 and rises in absolute magnitude as the state unemployment rate rises to -0.410 in slack labor markets. Finally, the comparison of only states within census divisions (columns 3 and 4)—a specification preferred by ADR, but with potentially important limitations (Neumark, Salas, and Wascher 2013)—reduces the magnitude of all of the estimated employment effects, but still produces a pattern consistent with the hypothesis of larger adverse teen employment effects in troughs as compared to peaks (33 percent to 100 percent larger) in the state business cycle.

[3]When I estimate an ordered probit model of the effect of minimum wage increases on the prime-age male unemployment rate (including state and year effects), I find no evidence of a statistically significant relationship, with a point estimate of -0.949 and a standard error of 2.73. This finding holds with the inclusion of state-specific non-linear time trends (third-order polynomial time effects).

Panel II of Table 4A shows comparable results for those ages 16-to-24 without a high school diploma. Column (1) shows that in tight labor markets, the low-skilled employment elasticity is -0.214, but rises to -0.414 during times of high state unemployment. The addition of state-specific non-linear time effects (column 2) produces a similar pattern of results, with an estimated employment elasticity with respect to the minimum wage of -0.447 during troughs of the state business cycle. And as with teenagers, comparing only those states within census divisions (columns 3 and 4) reduces the magnitude and precision of the estimated employment elasticities, but continues to point to a pattern of findings consistent with larger employment effects in troughs as compared to peaks of the state business cycle.

Table 4B uses an alternate measure of the state business cycle that is plausibly exogenous to state minimum wages: year-to-year nominal growth in state GDP generated by the finance sector. As noted above, the indicators for growth generated were growth of 6.5 percent or greater (peaks of the state business cycle), growth between 2.0 and 6.5 percent, and growth of less than 2.0 percent (troughs of the state business cycle). Again, main effects for growth are included on the right hand-side of equation (2) and the results in Table 4B focus on the interactions of indicators of growth with the minimum wage. [4] Note from the means of this growth measure as compared to the unemployment measure (see Table 1), troughs of the state business cycle are measured somewhat more liberally with the growth measure than with the unemployment measure.

The pattern of findings in Table 4B is generally similar to that seen in Table 4A using the prime-age male unemployment rate. For teenagers, I find that in peaks of the state business cycle, the teen employment elasticity ranged from approximately -0.089 to -0.196, only statistically distinguishable from zero at the 15 percent level. During troughs in the state business cycle, the teen employment elasticity is 50 percent to 100 percent larger than during peaks in the basic difference-in-difference model (column 1) and in the specification comparing states within census divisions (columns 3 and 4). But caution should be taken given that the model including state-specific time trends, however, produces no significant differences across the business cycle (column 2).

For younger individuals without a high school diploma, there is more consistent evidence that low-skilled employment effects of minimum wage hikes are greater during periods of weaker GDP growth than stronger GDP growth. During peaks of the business cycle, estimated low-skilled employment elasticities range from -0.098 to -0.228. However, during periods of weak growth (less than 2.0 percent in nominal growth), estimated elasticities range from -0.195 to -0.355. Moreover, even in the specification preferred by ADR with both state time trends and census division year effects (column 4), there is evidence that weaker growth (between 2 percent and 6.5 percent) is associated with relatively larger (in absolute magnitude) adverse low-skilled employment effects.

In summary, the findings in Tables 4A and 4B suggest that even after controlling for spatial heterogeneity in a variety of ways, minimum wages may have larger adverse employment effects during slack as compared to tight labor markets.

[4] When I estimate an ordered probit model of the effect of minimum wage increases on the GDP growth rate (including state and year effects), I obtain a point estimate of -2.30 and a standard error of 1.42. The result holds with the inclusion of state-specific non-linear time trends (third-order polynomial time effects).

Conclusions

This study re-examines the low-skilled employment effects of minimum wage increases across the business cycle, with more careful attention to the role of spatial heterogeneity. The results suggest that across the business cycle, minimum wage increases reduce employment of teenagers and younger individuals without a high school diploma. The estimated employment elasticities are in the consensus range of -0.1 to -0.3, and are robust to controls for spatial heterogeneity—state-specific productivity shocks to low-skilled sectors and unmeasured state-specific non-linear time trends—that do not eliminate potentially valid sources of identifying variation, consistent with Neumark, Salas, and Wascher (2013).

I then examine whether minimum wage hikes have differential effects on low-skilled employment at different phases of the state business cycle using two arguably exogenous measures of the business cycle: the prime-age male unemployment rate and annual growth in state GDP generated by the finance sector. The results suggest that the adverse low-skilled employment effects of minimum wage increases may be larger during troughs as opposed to peaks in the business cycle. This finding is consistent with the hypothesis that employers are most likely to shed the least experienced, least skilled workers during slack labor markets when minimum wages rise.

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Table 1. Descriptive Statistics, 1990-2010	
	Mean (StDev)
Dependent Variables	
Wage Rate for Teenage Workers Ages 16 to 19 (dollars $)	6.57(1.29)
Employment to Population Ratio for Teenagers Ages 16 to 19	0.394 (0.091)
Wage Rate for High School Dropout Workers Ages 16 to 24 (dollars $)	6.50 (1.20)
Employment to Population Ratio for High School Dropouts Ages 16 to 24	0.364 (0.085)
Independent Variables	
Minimum Wage	5.38 (1.10)
Unemployment Rate for Prime-Age Males Ages 25 to 54	0.048 (0.022)
Average Wage Rate for Adults Ages 25 to 54 (dollars $)	14.14 (3.07)
Percent of Population Ages 16 to 19	0.088 (0.007)
Percent of Population Ages 55 to 64	0.144 (0.023)
Percent of Population Married	0.551 (0.032)
High School Completion Rate	0.795 (0.042)
Per Capita Retail GDP in millions $	0.001 (0.0002)
Per Capita Manufacturing GDP in millions $	0.003 (0.0009)
Prime-Age Male UR < 5%	0.645 (0.479)
Prime-Age Male UR of 5% to 7.9%	0.264 (0.441)
Prime-Age Male UR ≥ 8 %	0.091 (0.287)
Finance GDP Growth of > 6.5%	0.338 (0.473)
GDP Growth of 2% to 6.5%	0.389 (0.488)
GDP Growth of < 2 %	0.273 (0.446)
State Dummies	—
Year Dummies	—
N	1,071

Notes: Weighted OLS estimates obtained using CPS Merged Outgoing Rotation Group files from 1990 to 2010.

Table 2. Estimates of the Relationship Between Minimum Wage Increases and Low-Skilled Workers' Hourly Earnings, 1990-2010				
	(1)	(2)	(3)	(4)
Panel I: Teenagers Ages 16 to 19				
Log (Min Wage)	0.085** (0.033)	0.071*** (0.023)	0.074** (0.034)	0.127** (0.062)
Panel II: Individuals Ages 16 to 24 without HS Diploma				
Log (Min Wage)	0.067+ (0.042)	0.078** (0.033)	0.084** (0.036)	0.123* (0.069)
State Effects?	Y	Y	Y	Y
Year Effects?	Y	Y	Y	Y
Demographic Controls?	Y	Y	Y	Y
Economic Controls?	N	Y	Y	Y
State Linear Trends?	N	N	Y	N
Census Division Time Shocks?	N	N	N	Y
N	1,071	1,071	1,071	1,071

***Statistically significant at 1% level **at 5% level *at 10% level +at 15% level
Notes: Weighted OLS estimates obtained using CPS Merged Outgoing Rotation Group files from 1990 to 2010. All models include controls for the prime-age male unemployment rate, the average adult wage rate, percent of population ages 16 to 19, youth high school graduation rate, state fixed effects, and year fixed effects. Standard errors corrected for clustering on the state are in parentheses.

Table 3A. Estimates of the Relationship Between Minimum Wage Increases and Low-Skilled Employment, 1990-2010				
	(1)	(2)	(3)	(4)
Panel I: Teenagers Ages 16 to 19				
Log (Min Wage)	-0.100 (0.118)	-0.154* (0.091)	-0.109 (0.118)	-0.233* (0.137)
Panel II: Individuals Ages 16 to 24 without HS Diploma				
Log (Min Wage)	-0.251*** (0.116)	-0.230* (0.134)	-0.194 (0.159)	-0.309+ (0.204)
State Effects?	Y	Y	Y	Y
Year Effects?	Y	Y	Y	Y
Demographic Controls?	Y	Y	Y	Y
Economic Controls?	N	Y	Y	Y
State Linear Trends?	N	N	Y	N
State 3rd Order Polynomial Trends?	N	N	N	Y
N	1,071	1,071	1,071	1,071

***Statistically significant at 1% level **at 5% level *at 10% level +at 15% level
Notes: Weighted OLS estimates obtained using CPS Merged Outgoing Rotation Group files from 1990 to 2010. All models include controls for the prime-age male unemployment rate, the average adult wage rate, percent of population ages 16 to 19, youth high school graduation rate, state fixed effects, and year fixed effects. Standard errors corrected for clustering on the state are in parentheses.

Table 3B. Census Division-Specific Employment Estimates				
	Teenagers Ages 16-to-19		Ages 16-to-29 without HS Diploma	
	(1)	(2)	(3)	(4)
New England *[N = 126]*	-0.122 (0.122)	-0.200 (0.208)	-0.471* (0.234)	-0.790** (0.313)
Middle Atlantic *[N = 63]*	-0.089 (0.099)	-0.332** (0.134)	-0.345** (0.143)	-0.402** (0.171)
East North Central *[N = 105]*	-0.357** (0.157)	-0.341** (0.104)	-0.357** (0.157)	-0.341** (0.102)
West North Central *[N = 147]*	-0.030 (0.183)	-0.019 (0.164)	0.039 (0.178)	0.015 (0.256)
South Atlantic *[N = 189]*	0.125 (0.315)	0.366 (0.255)	0.167 (0.192)	0.533* (0.262)
East South Central *[N = 84]*	0.419 (0.278)	0.316 (0.520)	0.076 (0.597)	0.194 (0.999)
West South Central *[N = 84]*	-0.004 (0.265)	-0.064 (0.548)	0.033 (0.394)	0.317 (0.401)
Mountain *[N = 168]*	0.111 (0.219)	0.243* (0.125)	-0.245 (0.197)	-0.056 (0.191)
Pacific *[N = 105]*	-0.001 (0.280)	0.142 (0.242)	-0.021 (0.354)	-0.138 (0.234)
State Effects?	Y	Y	Y	Y
Year Effects?	Y	Y	Y	Y
Demographic Controls?	Y	Y	Y	Y
Economic Controls?	Y	Y	Y	Y
State 3rd Order Polynomial Trends?	N	Y	N	Y
N	1,071	1,071	1,071	1,071

***Statistically significant at 1% level **at 5% level *at 10% level +at 15% level

Notes: Weighted OLS estimates obtained using CPS Merged Outgoing Rotation Group files from 1990 to 2010. All models include controls for the prime-age male unemployment rate, the average adult wage rate, percent of population ages 16 to 19, youth high school graduation rate, state fixed effects, and year fixed effects. Standard errors corrected for clustering on the state are in parentheses.

Table 4A. Heterogeneity in the Low-Skilled Employment Effects of the Minimum Wage Across the State Business Cycle, Using Prime-Age Male Unemployment Rate				
	(1)	(2)	(3)	(4)
Panel I: Teenagers Ages 16 to 19				
Log (Min Wage)	-0.124 (0.090)	-0.197 (0.123)	-0.116 (0.098)	-0.104 (0.102)
UR of 5% to 7.9%*Log (Min Wage)	-0.035 (0.017)	-0.057 (0.045)	-0.0001 (0.048)	-0.016 (0.057)
UR ≥ 8 %*Log (Min Wage)	-0.220** (0.0287)	-0.213** (0.096)	-0.115 (0.088)	-0.116 (0.111)
Panel II: High School Dropouts Ages 16 to 24				
Log (Min Wage)	-0.214* (0.110)	-0.266 (0.198)	-0.197* (0.110)	-0.173 (0.155)
UR of 5% to 7.9%*Log (Min Wage)	-0.021 (0.047)	-0.066 (0.041)	-0.009 (0.063)	-0.048 (0.067)
UR ≥ 8 %*Log (Min Wage)	-0.204** (0.093)	-0.181** (0.076)	-0.065 (0.131)	-0.125 (0.136)
State Effects?	Y	Y	Y	Y
Year Effects?	Y	Y	Y	Y
Demographic Controls?	Y	Y	Y	Y
Economic Controls?	Y	Y	Y	Y
State 3rd order polynomial trends?	N	Y	N	Y
Census Division Time Shocks?	N	N	Y	Y
N	1,071	1,071	1,071	1,071

***Statistically significant at 1% level **at 5% level *at 10% level +at 15% level

Notes: Weighted OLS estimates obtained using CPS Merged Outgoing Rotation Group files from 1990 to 2010. All models include controls for the prime-age male unemployment rate, the average adult wage rate, percent of population ages 16 to 19, youth high school graduation rate, state fixed effects, and year fixed effects. Standard errors corrected for clustering on the state are in parentheses.

Table 4B. Heterogeneity in the Low-Skilled Employment Effects of the Minimum Wage Across the State Business Cycle, Using State GDP Growth in Finance Sector				
	(1)	(2)	(3)	(4)
Panel I: Teenagers Ages 16 to 19				
Log (Min Wage)	-0.140+ (0.090)	-0.196+ (0.125)	-0.089 (0.098)	-0.040 (0.101)
GDP Growth of 2% to 6.5%*Log (Min Wage)	0.016 (0.035)	0.024 (0.036)	-0.069** (0.034)	-0.051 (0.037)
GDP Growth of < 2 %*Log (Min Wage)	-0.071** (0.036)	0.016 (0.053)	-0.093+ (0.088)	-0.043 (0.060)
Panel II: High School Dropouts Ages 16 to 24				
Log (Min Wage)	-0.228** (0.113)	-0.228 (0.194)	-0.174+ (0.198)	-0.098 (0.132)
GDP Growth of 2% to 6.5%*Log (Min Wage)	0.002 (0.046)	-0.034 (0.043)	-0.077* (0.041)	-0.102** (0.046)
GDP Growth of < 2 %*Log (Min Wage)	-0.082* (0.046)	-0.090 (0.077)	-0.181** (0.076)	-0.087 (0.071)
State Effects?	Y	Y	Y	Y
Year Effects?	Y	Y	Y	Y
Demographic Controls?	Y	Y	Y	Y
Economic Controls?	Y	Y	Y	Y
State 3rd order polynomial trends?	N	Y	N	Y
Census Division Time Shocks?	N	N	Y	Y
N	1,017	1,017	1,017	1,017

***Statistically significant at 1% level **at 5% level *at 10% level +at 15% level

Notes: Weighted OLS estimates obtained using CPS Merged Outgoing Rotation Group files from 1990 to 2010. All models include controls for the prime-age male unemployment rate, the average adult wage rate, percent of population ages 16 to 19, youth high school graduation rate, state fixed effects, and year fixed effects. Standard errors corrected for clustering on the state are in parentheses.

Appendix Table 1. Estimated Coefficients on Control Variables for Teenage Wage and Employment Regressions				
	Teenagers Ages 16-to-19		Ages 16-to-29 without HS Diploma	
	Wages	*Employ*	*Wages*	*Employ*
Log (Min Wage)	0.071*** (0.023)	-0.154* (0.091)	0.078** (0.033)	-0.237** (0.115)
Log (Prime-Age Male UR)	-0.019* (0.010)	-0.091*** (0.020)	-0.031*** (0.011)	-0.146*** (0.024)
Log (Prime-Age Wage)	0.310*** (0.059)	0.099 (0.173)	0.312*** (0.078)	0.187 (0.205)
Log (% Population Ages 16-19)	0.003 (0.047)	-0.051 (0.059)	0.005 (0.054)	-0.266*** (0.071)
Log (% Population Ages 55-64)	0.004 (0.036)	-0.087 (0.078)	0.009 (0.033)	-0.110 (0.090)
Log (% Population Ages 16-64 Married)	-0.171 (0.168)	-0.576*** (0.148)	-0.383** (0.177)	-0.824*** (0.200)
Log (High School Completion Rate)	0.043 (0.088)	0.037 (0.155)	-0.211** (0.082)	-0.624*** (0.194)
Log (Per Capita Retail GDP)	0.104 (0.074)	0.432*** (0.088)	0.130* (0.078)	0.376*** (0.095)
Log (Per Capita Manufacturing GDP)	0.031* (0.019)	-0.018 (0.030)	0.053** (0.022)	-0.040 (0.033)
N	1,071	1,071	1,071	1,071
State Effects?	Y	Y	Y	Y
Year Effects?	Y	Y	Y	Y

***Statistically significant at 1% level **at 5% level *at 10% level +at 15% level

Notes: Weighted OLS estimates obtained using CPS Merged Outgoing Rotation Group files from 1990 to 2010. All models include controls for the prime-age male unemployment rate, the average adult wage rate, percent of population ages 16 to 19, youth high school graduation rate, state fixed effects, and year fixed effects. Standard errors corrected for clustering on the state are in parentheses.

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