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Minimum Wage Increases in a Recessionary Environment[†]

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Abstract

Do seemingly large minimum-wage increases in an environment of deep recession produce clearer evidence of disemployment than is often observed in the modern minimum wage literature? This paper uses three data sets to examine the employment effects of the most recent increases in the U.S. minimum wage. We focus on two high-risk groups – restaurant-and-bar employees and teenagers – for the years 2005-2010. Although the evidence for a general disemployment effect is not uniform, estimates do suggest the presence of a negative minimum wage effect in states hardest hit by the recession.

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“Based on 20 years of research, I doubt there is ever a good time to raise the minimum wage. However, with the aggregate unemployment rate at 9.4 percent, the teen unemployment rate exceeding 22 percent, and the unemployment rate for black teens nearing 40 percent, next month’s increase seems like the worst timing possible”
(David Neumark, 2009).

The earliest of the modern literature tackling the impact of minimum wages on labor-market outcomes tended to point fairly uniformly to the existence of small, negative effects of higher minimum wages on employment (see the review in Brown, 1999). This result was supported in both time-series studies and state-level panel data research focusing on teenagers. By contrast, a series of studies that focused on employment outcomes within the fast-food sector (most notably, Card and Krueger, 1995) failed to reproduce this finding of a negative employment effect. Even though the robustness of the fast-food estimates to data-measurement issues and statistical approach was soon to be called into question,¹ some later studies using panel data from the restaurant industry (Addison et al., 2012; Dube et al., 2010) have tended to corroborate Card and Krueger’s lack of evidence of a negative employment effect. In addition, some contemporary studies considering teenagers have found less evidence of a general negative effect for teenagers.²

What accounts for the difficulty in uncovering adverse minimum wage effects in recent studies? One possibility is that the underlying effect of minimum wages on employment has diminished over the years. The minimum wage has fallen, both in real terms and relative to the average wage in the economy, so that increases in the minimum wage directly affect a smaller number of low-wage workers.³ It is also the case that the recent literature has added data from years in which the general health of the U.S. labor market has been strong. Unlike the high unemployment years of the later 1970s and

¹Kennan (1995) and Neumark and Wascher (2000) questioned the data used by Card and Krueger, with Card and Krueger (2000) responding that such issues did not perturb their primary finding. In addition, Donald and Lang (2007) later criticized the necessary statistical assumptions that underlie the inherent small-scale experiment in the Card and Krueger approach.

² See, in particular, Allegretto et al. (2011). Sabia (2009) does still find evidence of a negative employment effect for teenagers, a result Allegretto et al. argue is not robust to specification choice. As discussed below, Neumark et al. (2012) have countered that the Allegretto et al. finding is also not robust.

³ Analogously, the opposite circumstance was used as justification by Castillo-Freeman and Freeman (1992) for their finding of acutely adverse minimum wage effects in Puerto Rico.

1980s, the 15-year period after the 1991 federal minimum wage increases was marked by relatively tight labor markets.

The performance of the U.S. economy in the last few years has been poor, indexed by a doubling of the unemployment rate to over 10 percent at the peak of the recent recession (October 2009). At the same time, the Fair Minimum Wage Act of May 2007 raised the federal minimum wage by over 40 percent from July 2007 to July 2009. Although many states had minimum wages above the federal minimum prior to these increases, the hike had some effect in the large majority of states, and was fully effective in over 70 percent of them by August 2009. As our opening quote illustrates, a substantial increase in the federal minimum wage during a significant recession drew frequent criticism from economists and other observers. If the 1991-2007 period was characterized by decreasing relative minimum wages and low unemployment, a reversal of these trends might be expected to produce a larger disemployment effect than typically estimated in the recent literature.

In what follows, we use data from 2005-2010 to focus on minimum-wage employment effects during the most recent recession. As we discuss below, this allows us to handle concerns about geographically-disparate trends in employment while at the same time addressing recent criticisms raised by Neumark et al. (2012) concerning the assumption of constant trend patterns over a long time period. We examine both subsets of the labor market that have received primary attention in previous studies – restaurant employees and teenagers – using three different data sets. To preview our findings, we report limited overall evidence of reduced employment within restaurants and among teenagers, but stronger evidence of a disemployment effect for both groups in states with particularly high unemployment rates.

I. Recent Minimum Wage Increases and the Recessionary Impact

Concern over the impact of the recent increases in the federal minimum wage has most probably centered on the conjunction of two fears: first, the perception that the minimum wage was being increased very substantially; and, second, its particular timing when the economy might not be able to absorb such increases easily. Infrequent adjustments to the minimum wage over the 1980-2005 interval had indeed caused it to

fall relative to the average wage. But in a historical context, just how sizeable was the 40 percent increase in the mandate? The top line in Figure 1 plots the real value of the federal minimum wage relative to the average real value of the federal minimum over the 1939-2009 period. Values greater than one signify that the minimum wage exceeds the period average. The final increase in July 2009 did lead to a federal minimum that was higher than the historical average, but just barely. Indeed, the real value of the federal minimum was actually higher throughout the period from the late 1950s to the middle 1980s. On this metric, then, the recent episode is not particularly notable.

The bottom line in the figure charts the federal minimum wage relative to the average manufacturing wage. From this perspective, we can see that the federal minimum relative to the average manufacturing wage is only slightly higher in 2009 than it was after the last federal increase in 1997. On the other hand, since the increase in the late 2000s was from a base that had not been that low since 1950, in an absolute sense the increases of 2007-2009 were non-trivial.⁴

Of course, the period since 1980 has witnessed a plethora of state legislation adopting state minima that exceed federal levels. Figure 2 repeats the plot of the federal minimum relative to the average manufacturing wage for the last three decades, but now also traces that same average when the state minima are also taken into account.⁵ Although state mandates had also caused the overall state average to diverge from the federal minimum in the late 1980s and mid 1990s (both periods in which the federal minimum was not increased over several years), the importance of state minimum wage legislation is noticeably different over the last 10 years. That is to say, once we take state minima into account, the average prevailing minimum wage in 2009 was at its highest relative level since 1985.

Should we expect more severe minimum-wage impacts during a recession? Minimum-wage positions do tend to be unskilled jobs characterized by high turnover that may be most prone to disappear during a downturn. The theoretical framework contained

⁴ The relative value of the minimum wage had fallen so much that by 2006 only 2.2 percent of hourly workers were paid at or below the federal minimum wage. By 2010, this share had increased to 6.0 percent.

⁵ In calculating the average minimum wage to be used in forming the relative state minimum wage, we weight the minimum in each state by that state's population. The relative state minimum wage calculation includes all states, including those paying the federal minimum.

in Addison et al. (2012) derives a coefficient on the minimum wage (in a log-linear employment model) equal to the product of the labor-demand elasticity and the proportion of workers covered by the minimum wage. To the extent that this latter proportion increases during recessions, we could expect a bigger impact of minimum wages even if the labor-demand elasticity is unchanged.

However compelling the argument behind this possibility might be, it has not been the focus of much previous research. One exception is the analysis of Neumark and Wascher (2002), who present a disequilibrium treatment and explicitly model whether labor market outcomes lie at the equilibrium of labor supply and demand or on the labor demand curve alone; only in the latter case is the minimum wage effective. The authors implement an endogenous-switching regression model with employment data from the CPS for 16 to 24 year olds from the years 1973-1989. The model provides estimates of minimum wage effects when the minimum is binding, as well as the probabilities that observations are on the labor demand curve only. For our purposes, their crucial result is that the probability that the minimum wage will be binding increases when the prime-age male unemployment rate is higher.

Holmes et al. (2009) provide a different framework to study the question of recessionary impacts by estimating separate minimum wage equations for NBER-defined contractions and expansions of the economy, using aggregate Bureau of Labor Statistics (BLS) time-series data for 1954-2008. They find evidence of a negative employment effect from minimum wages only during economic contractions. Specifically, their results suggest that a 10 percent increase in the minimum wage would decrease white (black) teenage employment by 3.1 (5.0) percent. On the other hand, a dissenting result is provided by Allegretto et al. (2011) using CPS data on teens for 1990-2009. The authors test for the role of the cycle using both cross section and time series variation in the unemployment rate. Neither interaction terms of minimum wages with unemployment rates nor the joint effect of the minimum wage and the unemployment rate proves statistically significant.⁶

⁶ In an international context, Dolton and Bondibene (2012) use cross-country evidence from 1976-2008 to study minimum wage impacts for several demographic groups, finding a more detrimental impact in the presence of an economic downturn for youths (aged 15-24 years) but not in the cases of teenagers (15-19 years) or adults (25-64 years).

Finally, in our search for potentially stronger evidence of minimum wage effects, we are not alone in concentrating on disemployment effects during the recent recession. Evans and Macpherson (2010) use CPS data from the 2005-2009 period to estimate standard minimum-wage employment equations, and report negative minimum wage effects on employment for teenagers during this interval. Gittings and Schmutte (2010), using both CPS and Quarterly Workforce Indicators data, find a negative effect on teenage employment for the longer sample period 2000-2009. Hoffman and Ke (2011) focus on the final installment of the latest federal minimum wage increase in 2009. They estimate cross-state differenced models of the minimum-wage impact, finding no significant disemployment effects among teenagers but a significantly negative effect for workers aged 20 to 24. On the other hand, Hirsch et al. (2011), who investigate establishment-level data for quick-service restaurants in Georgia and Alabama, 2007-2009, fail to uncover any evidence of a negative employment effect. The authors conjecture from their surveys of managers that minimum wage increases may elevate technical efficiency along a number of dimensions of business operation (including higher performance standards).

II. Econometric Models and Data

A. Models

Our empirical modeling is based on measures of employment across different counties (or states) in the U.S, using the panel nature of the data to account for both cross-sectional and time-series effects in employment determination. Our “basic model” is specified as

$$y_{it} = \beta_1 \log(MW_{it}) + \beta_2' x_{it} + \gamma_i + \lambda_t + \varepsilon_{it}, \quad (1)$$

where y is the log of employment (or earnings), MW is the state-level minimum wage, x is a vector of other covariates potentially affecting y , and i refers to county or state depending on the data set. Fixed-effects estimation directly controls for geographic (γ) and time (λ) effects. Trends in the dependent variable that vary across counties are handled by using “clustered” standard errors robust to the presence of heteroskedasticity and to any correlation in errors between observations in the same state.

The basic model requires that such trends are uncorrelated with the state minimum (and x) in order to ensure consistent estimates of the minimum-wage effect.

Estimates of minimum wage elasticities can be sensitive to the particular time period examined, a point highlighted in the criticism of Neumark et al. (2012). In our primary extension to the above basic model, we directly incorporate geographic-specific linear trends, so that our estimating equation becomes

$$y_{it} = \beta_1 \log(MW_{it}) + \beta_2' x_{it} + \gamma_i + \lambda_t + \delta_i t + \varepsilon_{it}. \quad (2)$$

Both Addison et al. (2012) and Allegretto et al. (2011) show that incorporating such trends can materially affect minimum wage estimates (for the 1990-early 2000s period). In a reanalysis of the latter paper, Neumark, et al. (2012) argue that linear trends are not a sufficient geographic-specific control for the time period examined, and show that results are sensitive to allowing for nonlinearities in the geographic-specific patterns. They report evidence of significant minimum-wage elasticities for teenagers once nonlinearities are allowed, arguing that the combination of recessionary and boom periods in the Allegretto et al. study led to linear trends being inadequate controls.⁷ By focusing on a shorter time period, our analysis is not affected by these concerns. We still allow for geographic-specific differences in the underlying employment trends, reflecting the potential for heterogeneity in the severity of the recession across the U.S. not otherwise captured by our other controls.

We have also considered an alternative approach to handling geographically-disparate trends that focuses on county-border pairs (as in Dube et al., 2010). In this estimation, every pair of counties bordering each other but in different states is designated as a separate county pair (denoted with the subscript p), and the model is specified as

$$y_{ipt} = \beta_1 \log(MW_{ipt}) + \beta_2' x_{ipt} + \gamma_i + \eta_{pt} + \varepsilon_{ipt}. \quad (3)$$

Estimation of this model by fixed effects essentially becomes the estimation of a model that differences the dependent and independent variables (including the county fixed effect) across the two counties in any given pairing. This approach suffers from an inconsistent modeling of any given county's outcome when the county has more than one

⁷ Their results with nonlinear geographic-specific trends then end up being similar to results without any trends, but there is no reason to expect this to always be the case.

border pair.⁸ Neumark et al. (2012) have also argued that border counties are not in fact particularly good matches for this kind of analysis, as other economic characteristics do not seem to match more closely for border counties than for any other randomly chosen counties. We present estimates using this approach primarily for comparison purposes.

Our models are estimated for employment in segments in the labor market regarded as low-wage. Following both Addison et al. (2012) and Dube et al. (2010), our industry-level analysis will focus on *Food Services and Drinking Places*. As is common in the prior literature, our analysis of demographic groups treats teenagers (aged 16-19 years) as the group most likely to be affected by minimum wages.

B. Data Sets

We use three data sets to estimate our models. Two of these have been prominently used in prior minimum-wage research – the Quarterly Census of Employment and Wages (QCEW), and the Current Population Survey (CPS). The QCEW data are collected by the BLS from paperwork filed by employers with the unemployment insurance program, and have geographic information to the county level. Industry codes using the North American Industrial Coding System (NAICS) allow us to construct measures of total employment and average weekly earnings for specific sectors of the economy. Our analysis focuses on the *Food Services and Drinking Places* sector. Some sector/county combinations are censored by the BLS due to a small number of establishments, so we construct a panel of counties available in every quarter of our sample period. The only earnings variable is a weekly measure.⁹ The QCEW data allow us to disaggregate employment within this general sector, and so we also estimate separate models for the *Full-Service Restaurants* and the *Limited-Service Eating Places* subsectors.¹⁰ A more detailed description of the QCEW data is provided by Addison et al. (2012).

⁸ In the Dube et al. approach, the same county is included in the data more than once if it borders more than one county in the other state, implying the underlying nonparametric trend for any border county can vary each time the county is included in the sample. To avoid overemphasizing counties in multiple pairs, we weight each observation by the inverse of the number of times that county appears in the data set paired with a border county.

⁹ This measure includes most wage-like compensation, including tips, bonuses, stock options, and employer contributions to retirement plans.

¹⁰ A share of the workers in the *full-service* sector includes tipped employees, for whom minimum wage legislation still generally applies. Specifically, the federal requirement is that all employees receive a “minimum cash wage.” This minimum cash wage plus tips must at least equal the minimum wage;

The independent variables in our models include relatively standard measures of demand/supply factors that could reasonably affect employment within low-wage sectors in a county. Most importantly, our minimum wage variable is the effective minimum wage measured at the state level (discussed above). For quarters in which the state raised the minimum wage in the middle of the quarter, we calculate the minimum wage for that quarter as the weighted average of the older and newer minimum (with weights depending on the percentage of the time in that quarter in which the particular minimum is in effect). Population is included to capture effects of the size of the county on employment, while total employment in the county (from the QCEW) reflects effects from the size of the labor market. Average weekly earnings across all sectors is also included as a control.¹¹ The unemployment rate of the county (taken from the Local Area Unemployment Statistics) is included as a measure of the severity of the recession, while the school enrollment rate (for individuals aged 16-24 years and measured at the state level using the Current Population Survey) is included as a potential demand/supply measure.

The monthly data from the Current Population Survey (CPS) has been the most widely-used source for data in minimum wage studies – at least since the work of Card (1992) and Neumark and Wascher (1992). The bulk of this research has considered minimum-wage effects on particular demographic groups – most commonly teenagers aged 16 to 19 years – but has also ranged further afield to include young adults aged 20 to

otherwise the firm must increase the wage payment so that total hourly earnings equal the federal minimum. This implies that increases in the minimum wage (but not the minimum cash wage) could still lead to an increased wage payment by the firm if the old cash wage plus tips paid to a worker was less than the new higher minimum wage (assuming tips are unchanged). State laws lead to variation in the minimum cash wage and how much “tip credit” is allowed, although in all states the hourly wage including tips for tipped employees must still equal (or in some cases exceed) the highest minimum wage enforced in that state. (In seven states, no tip credit is allowed to employers, so the state minimum wage applies to all employees regardless of their tip status.) Therefore, for employees in the *full-service* sector most tipped employees will be affected by changes in minimum wages laws as the mandate will, at the very least, impact the tip-plus-wage-rate they must legally be paid. And, of course, there are several low-wage occupations in full-service restaurants where employees do not receive tips. There may still be concerns that minimum-wage changes are less important when the base cash wage is not changed, so we check below the robustness of our results for the *full-service* sector by alternatively treating the minimum wage as fixed for states where the minimum cash wage for tipped employees has not varied in recent years.

¹¹ A restricted version of our model – in which the minimum wage variable is entered as the log of the minimum/average wage ratio (similar to the Kaitz index) – would follow if the coefficient on the minimum and average log wages were equal in magnitude but of opposite signs. However, we do not impose this restriction as there may be other avenues for the average wage to affect employment; it is also not supported in most of our estimated models.

24 years and/or high-school dropouts in some studies (see the surveys by Neumark and Wascher, 2007, 2008). Less research has focused on industrial sectors using the CPS (although the study by Lang and Kahn, 1998, is a notable exception). Since January 2003, the CPS has used the NAICS coding system, allowing us to focus on the same general *Food Services and Drinking Places* sector examined with the QCEW (though disaggregation of the sector is not possible with the CPS). At the time of writing, data through 2010 were available from the CPS, so we extend our sample period by one year relative to the 2005-2009 period considered with the QCEW.

When using the CPS, the complete set of respondents to the basic monthly survey is used in measuring employment for a particular sector or demographic group. However, as earnings information is asked of the approximately one quarter of the sample that is part of the Outgoing Rotation Group (ORG), we have considerably fewer observations when measuring average earnings.¹² There are minor changes in the definition for some of the independent variables in our CPS analysis. Unemployment rates are now measured at the state-level, leading us to select the more traditional prime-age male unemployment rate as a control for recessionary conditions. The average wage variable is also calculated using prime-age men only.

Given our focus on minimum-wage effects during a period beginning in 2005, we are able to make the first use of the American Community Survey (ACS) in minimum-wage research. The U.S. Census Bureau instituted the ACS as an ongoing survey that obtains data similar to the former long form of the decennial census (which it replaced). The ACS surveys individuals every month, with a sample size that is roughly five times larger than that available from the CPS. Given that earnings and labor-supply questions are asked in the ACS, it potentially serves as a useful alternative to CPS data in labor economics research. The sampling frame for the first years of the ACS was relatively small but, beginning with the 2005 survey, this increased to a sample of roughly one percent of the population. We begin our analysis with the 2006 data, however, as only in that year did the ACS start to include individuals living in group quarters (such as college

¹²We find it odd that much of the recent research appears to use only the ORG samples to measure both employment and earnings, when more accurate measures of employment are available using the full monthly CPS. In calculating average wages, we drop all imputed observations as the CPS does not use state in imputing earnings (see Hirsch and Schumacher, 2004, for a discussion of the problems associated with CPS imputation).

dorms). Our analysis ends with the 2009 data, the latest year in which an ACS public use sample was available at the time of writing. The analysis with the ACS is parallel to that of the CPS, only with variables measured more accurately due to the larger underlying sample. The ACS does have one weakness relative to the CPS: it does not identify the month in which the interview takes place, only the year. We therefore use annual measures of employment (and other variables) and assign a weighted average minimum wage (similar to the quarter-based minimum-wage averages used with the QCEW). As with the CPS, the ACS allows us to analyze employment outcomes for either demographic groups or industrial sectors, with only the general food services and drinking places sector identified.¹³

Descriptive statistics are reported in Table 1, where some of the differences in the data sets are highlighted. Employment and population numbers in the QCEW are considerably smaller than in the CPS and the ACS, given the county-level focus of the QCEW. The unemployment rate is also higher in the QCEW, as it is not restricted to prime-age men. The average of hourly earnings in the CPS, among restaurant-and-bar workers and among teenagers, is about \$2 higher than the average minimum wage over the period studied. Average weekly earnings from the QCEW is slightly higher among *full-service* than *limited-service* restaurants, but the difference is quite small suggesting both of these sectors could be affected by minimum-wage changes.

III. Empirical Findings

A. Results from the QCEW

Our initial set of estimates focuses on minimum wage effects on both employment and earnings using data from the QCEW for the food services and drinking places sector. Our expectation is that an increase in the minimum wage should lead to an increase in average earnings among those still employed after the minimum increase. This effect would partly stem from workers whose wages are directly increased by the increased minimum, but it could also reflect spillover effects for higher-wage workers (as well as

¹³ Earnings are only available through a retrospective question on annual earnings, and so assignment of reported earnings to a particular year is problematic. As a result, we use the prime-age male hourly earnings for that state and year from the CPS, simply averaging across the 12 months for any state to obtain our annual measure of this variable.

compositional changes that shift demand towards those workers).¹⁴ If an empirical approach has difficulty in supporting an increasing minimum-wage impact on earnings, it may cast doubt on its ability to provide reliable information about employment effects.

All three methods for handling geographic trends are used in obtaining the estimates from the QCEW in Table 2. Our “basic” specification includes time dummies for each quarter in our sampled period. It thus controls nonparametrically for a general national trend in employment, but allows for no direct controls for county-specific variation in that trend (see equation 1). The “county-trend” specification (detailed in equation 2) allows for a separate linear trend in each county (in addition to the general nonparametric national trend). Finally, the “border-county” specification allows for a shared nonparametric trend between any two border counties (see equation 3). The first row of the table provides minimum-wage coefficients using data from all establishments in the food services and drinking places sector.¹⁵ Reported in the first three columns, the estimated minimum wage coefficients in the earnings equations are as expected, with a 10 percent increase in the minimum wage predicted to increase weekly earnings by about 1.5 percent. This result is largely insensitive to the method for handling trends, and is of a similar magnitude to the results in Addison et al. (2012) and Dube et al. (2010).¹⁶ These earlier papers failed to find statistically significant effects on employment in the broad restaurant-and-bar sector, and the same is again true here for all three specifications. The final three columns report estimated employment effects, and the results for the general sector do not provide statistically significant evidence of a minimum-wage effect. For reasons discussed in section II, our preference is to directly model county-specific trends, and with this approach the estimates are quite precise. For example, a 95 percent confidence interval for the minimum-wage effect on employment ranges from -0.043 to

¹⁴ In the case of the QCEW, the one mitigating factor is that we study weekly rather than hourly earnings, and weekly earnings could fall if employers lowered hours sufficiently for their workers. An hours elasticity greater than one would be necessary for this to occur, however, and this seems very unlikely given prior work on this issue (for example, see Zavodny, 2000).

¹⁵ Only coefficients on the log of the minimum wage are presented in Table 2. The full set of results is available in Appendix Table A

¹⁶ In Addison et al.’s preferred specification with county-specific trends, the estimated earnings elasticity is 0.171, while in Dube et al.’s preferred specification using border-county pairs the estimated elasticity is 0.188. Both estimates are statistically significant at conventional levels.

0.021 – clearly excluding the elasticities close to -0.2 often alluded to in the literature.¹⁷ The second row of Table 2 allows for potential lagged effects of the minimum wage by using a five-quarter moving average of the minimum wage, with little change in the results.

Separate estimates for the *full-service* and *limited-service* restaurant sectors are reported in the third and fourth rows of Table 2. Our expectation is that as the use of minimum wage labor is presumably more prevalent in the *limited-service* sector, the minimum wage impact on earnings should also be higher. This expectation is consistent with our results for the country-trend specification (but not for the border-county results). Perhaps contrary to expectations, however, the only (marginally) statistically significant minimum-wage employment effect is found for the *full-service* sector (and only after including county trends). This estimate is nonetheless quite small – an elasticity of around -0.04 – and the 95 percent confidence interval of [-0.079, -0.003] again excludes effects close to -0.2. This small estimate is significant due to the very small standard error provided by the county-trend analysis. Indeed, most of the standard errors are smaller in the current set of results than with the parallel findings in Addison et al. (2012) and Dube et al. (2010), despite the shorter time period. This outcome likely stems from the greater variation in the minimum wage variable in the later period compared to the earlier data.¹⁸

B. Results from the CPS and ACS

For analyses at the industry level, the QCEW would seem preferable to the CPS given its larger and more complete sample of employment, with the possibility of controlling for county-level factors (rather than the state-level factors with the CPS). Perhaps the only advantage of the CPS is that employment can be measured at the monthly level, so that the correlation with minimum-wage changes is more accurate. For analyses at the demographic-level, however, alternative data sets are instrumental as

¹⁷ Neumark and Wascher (2007) identify the “consensus range” of the earliest minimum-wage literature as suggesting elasticities between -0.1 and -0.3, and argue that the preponderance of studies in the new minimum-wage research support a finding of negative employment effects.

¹⁸ To explore the sensitivity of the results in the *full-service* sector to the treatment of minimum wages for tipped employees (see footnote 10), we constructed a new minimum wage variable that holds the nominal minimum wage constant in states where changes in the regular minimum wage occurred but where the cash minimum wage was not changed. Although the employment coefficient estimate for the baseline method is now larger in magnitude and statistically significant when we use this alternative measure of the state minimum wage (an estimate of -0.09, with a p-value of 0.001), conclusions from the county-trend and border-county results are not affected.

information on the demographics of employees is not available with the QCEW. We report estimates from the CPS for the food services and drinking places sector, allowing comparison with results from the QCEW. But our primary use of the CPS is to estimate models for teenagers, the main focus of earlier research with the CPS.¹⁹

The estimated employment elasticities for the CPS are reported in the first two columns of Table 3, with the first row showing the food services and drinking places results. The CPS analysis of this sector provides considerably less precise estimates than the county-level analysis with the QCEW. Both specifications provide negative estimated effects, but it is not possible to reject the null that the coefficients equal zero. However, the standard errors are almost four times as large with the CPS, and an elasticity of -0.2 would be in the 95 percent confidence interval even with the state-trends specification. The results for teenagers (in the second row) are also imprecise, but in this case we do have marginal significance of the negative estimated elasticity when state trends are incorporated.²⁰ The estimate of -0.18 is also in the range typically suggested by Neumark and Wascher as to be expected for teenagers. The bottom two rows of Table 3 provide estimates using a 13-month moving average of the minimum wage to allow for lagged effects. These estimates are naturally less precise (given the lower variation in the moving average variable). The teenager elasticity estimate is now insignificant, although the actual coefficient estimate is not that much smaller than when the contemporaneous minimum-wage variable was used.²¹

We also estimated similar equations as for the CPS but using the ACS as the source of data. These results, presented in the last two columns of Table 3, do not provide statistically significant estimates for either food services and drinking places or

¹⁹ Only minimum-wage estimates for the employment equations are reported. The full specifications for the employment equations are reported in Appendix Table B. We have also estimated earnings equations with the CPS, generally finding increasing impacts on earnings from higher minimum wages. Earnings estimates are available from the authors upon request.

²⁰ The result from our basic specification without state-level trends may seem at odds with that of Even and Macpherson (2010), who report a statistically significant negative employment elasticity for teenagers using the CPS over a similar time period when using a somewhat similar specification. If we replicate the Evans and Macpherson specification, we too find a negative and marginally significant coefficient estimate. However, this result is no longer significant when we either (a) change the national-trend controls from annual dummies (which they use) to monthly dummies (which we use), or (b) weight by state population. (Results are available from the authors upon request.)

²¹ We also estimated employment equations for young adults (aged 20-24 years), and for junior-high-school dropouts (25 and older). In neither case were the coefficient estimates significant (other than a positive estimate for dropouts when state trends are not included).

teenagers. Indeed, the estimated coefficients for food services and drinking places are positive. On the other hand, the estimated coefficient for teenagers is significant, and a 95 percent confidence interval would include elasticities as large in magnitude as -0.2. It is noteworthy that the ACS estimates for teenagers have lower standard errors than those from the CPS – even though the ACS only provides annual counts. The considerably larger sample size with the ACS allows for calculation of employment numbers with less sampling error, which would seem to be the primary reason for the apparent increase in precision.

In sum, the CPS and ACS do not provide evidence of disemployment effects in the most recent period within food services and drinking places. However, the QCEW appears to be substantially better for studying this sector, as the estimated effects found to be significant in the QCEW would not be found so with the CPS or ACS. On the other hand, the CPS (and ACS) is necessary if we are to consider effects for teenagers, and in this case there is at least limited evidence of a negative minimum-wage effect on employment.

C. Did the Recession Enhance the Minimum-Wage Impact?

Our results estimate how minimum wages affected the employment of low-wage groups during a period in which the labor market is sharply contracting. By considering such a period, our analysis is less likely to be affected by the concerns raised by Neumark et al. (2012) to the effect that underlying trends deviate substantially from linear assumptions. While we have found some limited evidence of disemployment effects with this more concentrated focus, we are not seeing the large increases in estimated elasticities that might be expected if the recession substantially enhanced the relevance of minimum wages. As part of the period studied here includes a distinct pre-recession interval, we next directly assess whether loose labor markets enhance the minimum-wage effect.

Our first approach to determining whether minimum wage effects were larger during the recession was to include an interaction of the minimum-wage variable with a dummy for the observation being from the post-2007 sub-period. As the recession is dated as starting in December 2007, the interaction allows us to test whether the minimum-wage impact varied between the pre- and post-recession intervals of the sample

period. Although not reported here, these estimates did not provide evidence of more substantial minimum-wage effects during the recession.²² Of course, the brunt of the recession was not felt equally throughout the U.S., so that a more focused examination of potential recessionary impacts would be to consider whether the minimum-wage impact was larger in high unemployment-rate areas. Although Neumark and Wascher (2002) reported results consistent with higher unemployment making it more likely that the minimum wage was binding, both Allegretto et al (2011) and Addison et al. (2012) failed to find evidence of a minimum-wage/unemployment interactive effect for teenagers or for restaurant-and-bar workers (using data primarily from the 1990s and early-to-mid 2000s).

To examine whether the recent experiences with considerably higher unemployment rates led to larger minimum-wage impacts, we estimated the following extension of equation (2)

$$y_{it} = \beta_1 \log(MW_{it}) + \beta_2 \log(MW_{it})(UR_{it} - \overline{UR}_{05-07}) + \beta_3' x_{it} + \gamma_i + \lambda_t + \delta_i t + \varepsilon_{it} , \quad (4)$$

where UR is the unemployment rate (also included in x) and $\overline{UR}_{05-07}$ is the average unemployment rate in the 2005-2007 period. This specification allows us to interpret β_1 as the impact of the minimum wage when the unemployment rate is at its average level during the pre-recession years in the data, while β_2 reflects the tendency for the minimum-wage impact to vary with the unemployment rate. A “basic” version of equation (4) without geographic-specific trends was also estimated for comparison purposes.

Estimates for food services and drinking places employment using the QCEW are reported in the top panel of Table 4. Results for the broad sector do suggest that the minimum-wage impact primarily occurs when the unemployment rate is above the pre-recession average, as the minimum-wage coefficient is insignificant but the interaction is significant. For the QCEW county-level data, the average unemployment rate increased by 2.3 percentage points from 2005-2007 to 2008-2009; given the estimates in the second

²² This result occurs as long as geographic-specific trends are included in the specification, and is reported in an earlier working-paper version of this paper available from the authors upon request.

column, this would suggest an estimated minimum-wage elasticity of about -0.05 at the average unemployment rate during the recession years. While evidence of a minimum-wage impact is clear, it would appear that the magnitude for restaurants and bars is still quite small even during recessions. It is interesting that the increased effect during high-unemployment periods seems to be largely driven by responses among *limited-service* restaurants (reported in the final two columns). There is still evidence of a small negative minimum-wage elasticity among *full-service* restaurants (in the middle two columns), but no evidence that this becomes more important with increases in unemployment. The greater recession sensitivity of the *limited-service* sector may reflect a reduced ability to pass on minimum-wage-related cost increases through product-price increases in the midst of a recession, if recessions make product demand among fast-food restaurants particularly price sensitive.²³ As a result, somewhat more of the response shows up in employment responses than in normal times.

Results using the CPS are reported in the bottom panel of Table 4.²⁴ The estimated interaction coefficient for food services and drinking places is not statistically significant (though an effect of the magnitude estimated with the QCEW would not be significant given the standard error of the CPS estimate). On the other hand, there is strong evidence of a larger minimum-wage impact among teenagers when the unemployment rate is high. In this case, the observed increase in the unemployment rate of 2.8 percentage points would suggest a minimum-wage elasticity of -0.34 at the average unemployment level during 2008-2009.²⁵ In this case, it does appear that the recession did make teenagers particularly susceptible to increases in the minimum wage. Part of the response may have been a substitution away from teenagers and towards older, more-skilled workers made available by general job losses during the recession. As a result, sector-specific results fail to show a particularly large increase in employment losses,

²³ Aaronson (2001) presents evidence that minimum wage increases do appear to be at least partially passed on through prices in the restaurant sector.

²⁴ The ACS is potentially less useful for identifying unemployment-rate interactions, as its annual basis may not capture the worsening state of the labor market as accurately as the monthly measures available from the CPS. This limitation might explain why the estimated models with unemployment interactions using the ACS provided statistically insignificant results (results provided upon request).

²⁵ The increase in the unemployment rate is different between the two data sets as the QCEW analysis is based on county-level unemployment rates for all workers, while the CPS uses state-level unemployment rates just for prime-age men.

even though changes may have been occurring in the demographic-group composition of those sectors.²⁶

V. Conclusions

Much of the recent research on the employment effects of minimum wages has failed to provide convincing evidence that increases in the minimum wage are associated with material reductions in employment. One possible explanation for negative findings of this type might be that the typical increases in wage minima over 1980-2005 have simply been too small, taking wages to levels that are still quite low in both real and relative terms. Combined with potentially low own-wage elasticities for low-skilled labor in many service-type sectors, such modest increases may have not provided sufficient identifying information or traction for the researcher to detect statistically any negative effects of minimum wages on low-wage employment.

In the present paper, we have focused on contemporary economic developments in the U.S., motivated by the perception that recent minimum wage increases have been altogether more substantial than in the past. This perception, although not strongly borne out in practice, does receive some support in the data. More importantly, our time frame encompasses a period of major recession that might also be expected to provide a more severe response of employers to minimum-wage hikes. By focusing on this recent period, we are able to evaluate the possibility that more recent effects of minimum wages have been larger than the typically small elasticities reported in the literature, especially as the unemployment rate increases. To this end, we have used three different data sets (including one that has never been used before for this purpose) to estimate the employment effects of minimum wages for low wage groups during the 2005-2010 period, applying three principal estimation strategies employed in the literature.

Our findings produce limited evidence of a negative minimum wage impact on employment. The results support a small negative minimum-wage elasticity among *full-service* restaurants, and a larger estimated elasticity among teenagers. When focusing on

²⁶ We have attempted to test this possibility directly by estimating models in which employment of teenagers in the food services and drinking places sector is the dependent variable. Unfortunately, the small numbers problem with the CPS when focusing on such specific groups renders the estimates very imprecise. The estimated minimum-wage elasticity for employment was negative (and larger in magnitude than the overall food services and drinking places elasticity), but it was not statistically significant.

elasticities in states with particularly high unemployment rates, we report a negative estimated minimum-wage elasticity within the general food services and drinking places sector. More importantly, we find evidence of a substantially higher minimum-wage elasticity for teenagers in states with the highest unemployment rates during our sample period. The suggestion is that the large increase in teenage unemployment during the recession is at least partly explained by the accompanying increases in the minimum wage.

The bottom line of this investigation is that even when implemented during a significant economic downturn, minimum-wage increases do not appear to have particularly strong effects in reducing employment within the sector of the economy most likely to be affected by the minimum wage. However, minimum-wage elasticities for teenagers do seem to have increased substantially in states where the unemployment rate was particularly high. If the minimum-wage impact was largely concentrated among teenagers, this outcome may be less of a concern than if it had been felt more generally (on the reasoning that comparatively few minimum-wage teenagers are to be found in poor families). But in any event the warning that a severe recession is the worst time for a minimum-wage increase does appear warranted.

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Figure 1: Relative Federal Minimum Wage, 1939-2009

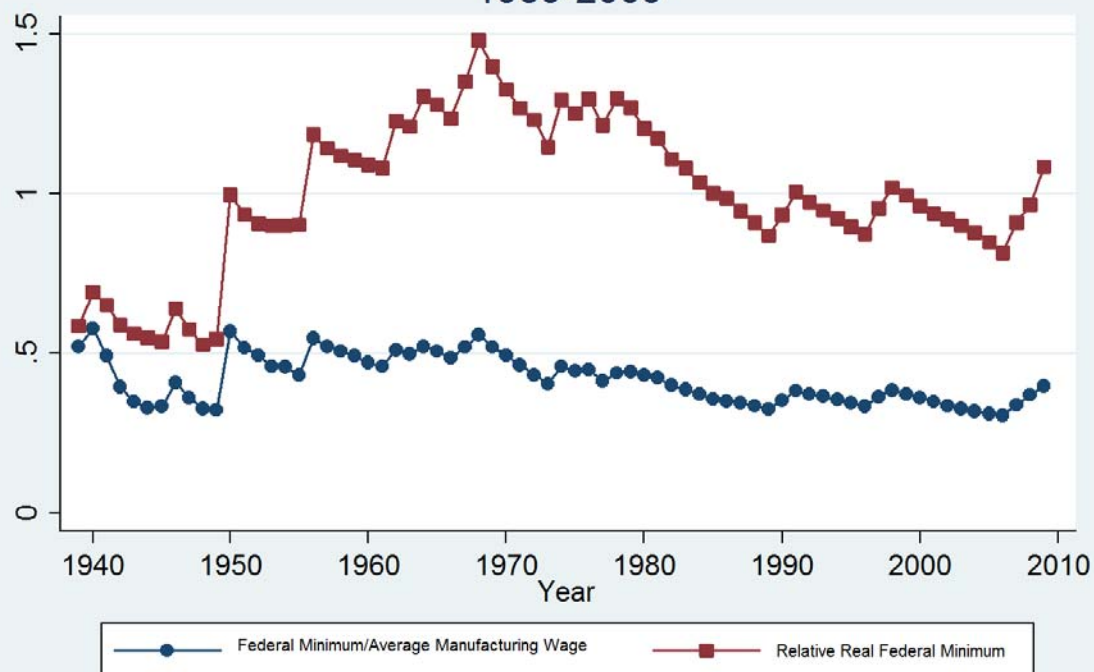


Figure 2: Minimum Wages Relative to the Average Manufacturing Wage, 1980-2009

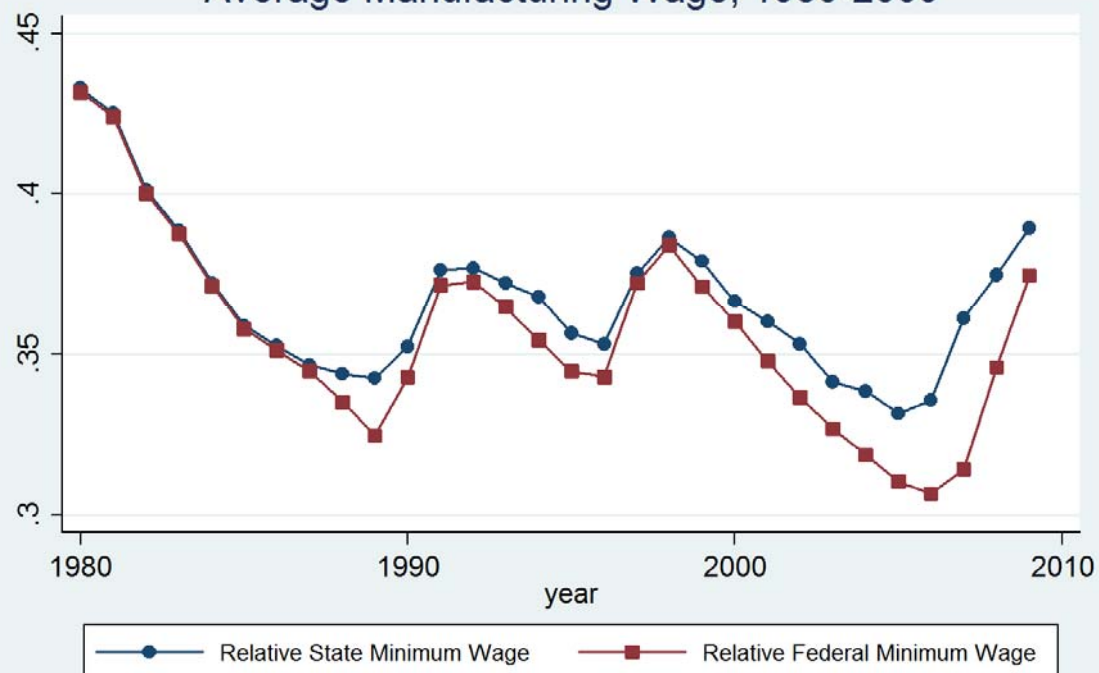


Table 1	Means and Standard Deviations		
	QCEW: 2005-2009	CPS: 2005 - 2010	ACS: 2006 - 2009
<i>Food Services and Drinking Places</i>			
Employment	2,302 (3,199)	164,030 (181,132)	159,615 (178,256)
Earnings (QCEW -- weekly; CPS -- hourly)	224 (46)	8.96 (2.36)	
<i>Limited-Service Restaurants</i>			
Employment	1,023 (1,360)	...	
Weekly Earnings	211 (35)	...	
<i>Full-Service Restaurants</i>			
Employment	1,157 (1,660)	...	
Weekly Earnings	228 (59)	...	
<i>Teenagers</i>			
Employment		110,643 (111,440)	114,667 (113,871)
Hourly Earnings		8.29 (1.77)	
<i>Independent Variables</i>			
Minimum Wage	6.45 (0.81)	6.51 (0.97)	6.71 (0.83)
Population	76,067 (96,906)	5,760,757 (6,452,120)	5,861,520 (6,560,818)
Total Employment (QCEW - Private; CPS/ACS - Adult)	24,368 (33,365)	2,811,635 (3,061,593)	2,799,421 (3,064,235)
Private Weekly Earnings	629 (129)		
Unemployment Rate (QCEW-All Participants; CPS/ACS- Prime-Age Male)	6.0 (2.6)	4.9 (2.7)	4.9 (1.9)
Enrollment Rate, 16-24 year olds	51.0 (8.0)	50.6 (11.9)	59.2 (4.2)
Teen Population		330,501 (374,835)	344,334 (392,554)
Prime-Age Male Wage		22.42 (3.76)	22.44 (3.00)
<i>Notes:</i> 1) QCEW values are per-county averages (and standard deviations) of the variables over the 2005-2009 sample period. The averages for the independent variables are for the sample of counties with complete information for food services and drinking places employment. 2) CPS measures are per-state averages (and standard deviations) of the variables over the 2005-2010 sample period. The averages for the independent variables are for the sample of state/month observations included in the food services and drinking places regressions. 3) ACS values are per-state averages (and std. devs.) of the variables over the 2006-2010 sample period. 4) All dollar measures are expressed in May 2010 dollars, adjusting for inflation using the CPI.			

Table 2	Minimum Wage Estimates of Employment and Earnings Equations, QCEW, 2005-2009
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		<i>Earnings</i>				<i>Employment</i>		
		Basic	County Trends	Border County		Basic	County Trends	Border County
(1)	Food Services and Drinking Places	0.134*** (0.017)	0.131*** (0.014)	0.163*** (0.031)		-0.015 (0.019)	-0.011 (0.016)	-0.076 (0.051)
(2)	Food Service and Drinking Places: MW Moving Average	0.129*** (0.025)	0.153*** (0.018)	0.153*** (0.051)		-0.037 (0.028)	-0.023 (0.027)	-0.090 (0.072)
(3)	Full-Service Restaurants	0.134*** (0.021)	0.118*** (0.017)	0.181*** (0.050)		-0.031 (0.028)	-0.038* (0.021)	0.035 (0.041)
(4)	Limited-Service Restaurants	0.152*** (0.016)	0.149*** (0.019)	0.147*** (0.042)		0.002 (0.021)	-0.003 (0.022)	-0.017 (0.065)

Notes: Only minimum-wage coefficient estimates are reported, although each regression contains all covariates discussed in equation (1). Clustered standard errors are in parentheses. See text for details on the estimation procedure for coefficients and clustered standard errors for each specification. All regressions include fixed-effects for county and quarter-year, and are weighted by the average county population over the 2005-2009 period. Estimates in row 2 use an equally weighted 5-quarter moving average of the current minimum wage with 4 lags of the minimum wage. ***, **, * denote statistical significance at the 0.01, 0.05, and 0.10 levels, respectively

Table 3	Minimum Wage Estimates for Employment, CPS (2005-2010) and ACS (2006-2009)
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		CPS: 2005-2010			ACS: 2006 - 2009	
		Basic	State Trends		Basic	State Trends
(1)	Food Services and Drinking Places	-0.090 (0.081)	-0.044 (0.082)		0.041 (0.058)	0.124 (0.114)
(2)	Teenagers	0.002 (0.101)	-0.178* (0.099)		0.035 (0.076)	-0.059 (0.085)
(3)	Food Service and Drinking Places: MW Moving Average	-0.014 (0.112)	0.065 (0.132)		...	...
(4)	Teenagers: MW Moving Average	0.013 (0.139)	-0.158 (0.117)		...	...

Notes: Only minimum-wage coefficient estimates are reported, although each regression contains all covariates discussed in equation (1). Clustered standard errors are in parentheses. See text for details on the estimation procedure for coefficients and clustered standard errors for each specification. All regressions include fixed-effects for state and quarter-year, and are weighted by the average state population over the corresponding time period. Estimates in rows 3 and 4 use an equally weighted 13-month moving average of the current minimum wage with 12 lags of the minimum wage. ***, **, * denote statistical significance at the 0.01, 0.05, and 0.10 levels, respectively

Appendix Table B		Estimates of Employment Equations for Low-Wage Sectors, CPS, 2005-2010			
<i>Independent Variables</i>	<i>Food Service and Drinking Places</i>			<i>Teenagers</i>	
	Basic	State Trends		Basic	State Trends
Minimum Wage	-0.090 (0.081)	-0.044 (0.082)		0.002 (0.101)	-0.178* (0.099)
Population	0.650 (0.408)	0.735 (0.542)		0.899*** (0.064)	0.952*** (0.059)
Total Employment	-0.174 (0.165)	-0.213 (0.171)		-0.342 (0.571)	1.128*** (0.422)
Average Prime-Age Male Wage	-0.058* (0.030)	-0.052* (0.029)		-0.124 (0.196)	-0.046 (0.202)
Unemployment Rate	0.028 (0.352)	0.089 (0.343)		0.020 (0.028)	0.029 (0.023)
Teen Population				-0.019*** (0.005)	-0.005 (0.004)
Enrollment Rate/100	0.081 (0.082)	0.191 (0.083)			
N	3,665	3,665		3,634	3,634
Notes: All dependent and independent variables are in logarithmic form, with the exception of the unemployment rate and the enrollment rate. Standard errors are in parentheses. See text for details on the estimation procedure for coefficients and clustered standard errors for each specification. All regressions include fixed-effects for state and month-year, and are weighted by the average state population over the 2006-2009 period. The specification in the lower panel includes the same additional covariates and fixed effects as in the top panel, although only the minimum-wage coefficient estimate is reported. ***, **, * denote statistical significance at the 0.01, 0.05, and 0.10 levels, respectively					

Appendix Table C		Estimates of Employment Equations for Low-Wage Sectors, ACS, 2006-2009			
<i>Independent Variables</i>	<i>Food Service and Drinking Places</i>			<i>Teenagers</i>	
	Basic	State Trends		Basic	State Trends
Minimum Wage	0.041 (0.058)	0.124 (0.114)		0.035 (0.076)	-0.059 (0.085)
Population	0.923 (0.602)	0.538 (2.670)		-0.370 (1.263)	3.404* (2.009)
Total Employment	0.332 (0.567)	0.843 (1.133)		0.061 (1.020)	0.302 (0.842)
Average Prime-Age Male Wage	-0.180* (0.104)	-0.152 (0.157)		0.028 (0.141)	-0.007 (0.176)
Unemployment Rate	0.012 (0.008)	0.007 (0.014)		-0.029* (0.016)	0.015 (0.012)
Teen Population				0.989*** (.243)	1.186*** (0.224)
Enrollment Rate/100	-0.215 (0.387)	-0.193 (0.705)			
N	204	204		204	204
<i>Notes:</i> All dependent and independent variables are in logarithmic form, with the exception of the unemployment rate and the enrollment rate. Standard errors are in parentheses. See text for details on the estimation procedure for coefficients and clustered standard errors for each specification. All regressions include fixed-effects for state and month-year, and are weighted by the average state population over the 2006-2009 period. The specification in the lower panel includes the same additional covariates and fixed effects as in the top panel, although only the minimum-wage coefficient estimate is reported. ***, **, * denote statistical significance at the 0.01, 0.05, and 0.10 levels, respectively					

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