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THE MINIMUM WAGE AND THE GREAT RECESSION: EVIDENCE OF EFFECTS ON THE EMPLOYMENT AND INCOME TRAJECTORIES OF LOW-SKILLED WORKERS

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ABSTRACT

We estimate the minimum wage's effects on low-skilled workers' employment and income trajectories. Our approach exploits two dimensions of the data we analyze. First, we compare workers in states that were bound by recent increases in the federal minimum wage to workers in states that were not. Second, we use 12 months of baseline data to divide low-skilled workers into a "target" group, whose baseline wage rates were directly affected, and a "within-state control" group with slightly higher baseline wage rates. Over three subsequent years, we find that binding minimum wage increases had significant, negative effects on the employment and income growth of targeted workers. Lost income reflects contributions from employment declines, increased probabilities of working without pay (i.e., an "internship" effect), and lost wage growth associated with reductions in experience accumulation. Methodologically, we show that our approach identifies targeted workers more precisely than the demographic and industrial proxies used regularly in the literature. Additionally, because we identify targeted workers on a population-wide basis, our approach is relatively well suited for extrapolating to estimates of the minimum wage's effects on aggregate employment. Over the late 2000s, the average effective minimum wage rose by 30 percent across the United States. We estimate that these minimum wage increases reduced the national employment-to-population ratio by 0.7 percentage point.

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Between July 23, 2007, and July 24, 2009, the federal minimum wage rose from $5.15 to $7.25 per hour. Over a similar time period, the employment-to-population ratio declined by 4 percentage points among adults aged 25 to 54 and by 8 percentage points among those aged 15 to 24. Both ratios remain well below their pre-recession peaks. The empirical literature is quite far from consensus, however, regarding the minimum wage's potential contribution to these employment changes (Card and Krueger, 1995; Neumark and Wascher, 2008; Dube, Lester, and Reich, 2010; Neumark, Salas, and Wascher, 2013; Meer and West, 2013). In this paper, we analyze the minimum wage's effects on the employment and income trajectories of low-skilled workers during the Great Recession and subsequent recovery.

Our analysis harnesses the fact that the 2007 through 2009 increases in the federal minimum wage were differentially binding across states. Between December 2007 and July 2009, the effective minimum wage rose by $1.31 in the states we designate as "bound" and by $0.43 in the states we designate as "unbound." Of the $0.88 differential, $0.58 took effect on July 24, 2009. We analyze the effects of these differentially binding minimum wage increases using monthly, individual-level panel data from the 2008 panel of the Survey of Income and Program Participation (SIPP). The SIPP allows us to use 12 months of individual-level wage data, from August 2008 through July 2009, to further divide low-skilled individuals into those whose wages were directly targeted by the new federal minimum and those whose wages were moderately above. These rich baseline data yield a number of advantages over the literature's standard approaches.[1]

[1] Analyses of individual-level panel data are not as common in the minimum wage literature as one might expect. Examples include Currie and Fallick (1996), who analyze teenage employment in the 1979 National Longitudinal Survey of Youth, Neumark, Schweitzer, and Wascher (2004) and Neumark and Wascher (2002), who use the short panels made possible by the matched monthly outgoing rotation files of the Current Population Survey (CPS), and Linneman (1982), who analyzed the minimum wage using 1973-1975 data from the Panel Study of Income Dynamics. Burkhauser, Couch, and Wittenburg (2000) analyze the minimum wage using the 1990 SIPP, but adopt the conventional state-panel approach of analyzing its effects on the employment of low-wage demographic groups rather than isolating samples of targeted individuals on the basis of baseline wage data.

Our approach's first advantage is its capacity to describe the minimum wage's effects on a broad population of targeted workers. Past work focuses primarily on the minimum wage's effects on particular demographic groups, such as teenagers (Card, 1992a,b; Currie and Fallick, 1996), and/or specific industries, like food service and retail (Katz and Krueger, 1992; Card and Krueger, 1994; Kim and Taylor, 1995; Dube, Lester, and Reich, 2010; Addison, Blackburn, and Cotti, 2013; Giuliano, 2013). While minimum and sub-minimum wage workers are disproportionately represented among these groups, both are selected snapshots of the relevant population. Furthermore, it is primarily low-skilled adults, rather than teenage dependents, who are the intended beneficiaries of anti-poverty efforts (Burkhauser and Sabia, 2007; Sabia and Burkhauser, 2010). Assessing the minimum wage from an anti-poverty perspective thus requires characterizing its effects on the broader population of low-skilled workers, which we are able to do.[2]

Econometrically, our setting has several advantages. One benefit of our rich baseline data is that they allow us to limit the extent to which our "target" group contains unaffected individuals. Second, the data allow us to identify relatively low-skilled workers whose wage distributions were not directly bound by the new federal minimum. We use these workers' employment trajectories to construct a set of within-state counterfactuals. The experience of these workers allows us to control for the form of time varying, state-specific shocks that are a source of contention in the recent literature (Dube, Lester, and Reich, 2010; Meer and West, 2013; Allegretto, Dube, Reich, and Zipperer, 2013). Third,

[2]Linneman (1982) similarly discusses this benefit of analyzing individual-level panel data in the context of minimum wage increases during the 1970s. A drawback of the individual-panel approach is that the resulting samples of targeted workers exclude individuals who were not employed when baseline data were collected. It is best suited for estimating the effects of minimum wage increases on the employment trajectories of those directly targeted. An advantage of this study's analysis of monthly panel data is that our estimates capture the minimum wage's effects on both the regularly employed and on highly marginal labor force participants. Specifically, the only individuals we are unable classify on the basis of baseline wages are those who were unemployed for all 12 baseline months. Additionally, although we do not have a reported wage for such individuals, we can directly estimate the effect of binding minimum wage increases on this group's subsequent employment; the estimated effect is negative, economically small, and statistically indistinguishable from 0.

our research design allows for transparent, graphical presentations of the employment and income trajectories underlying our regression estimates.

We begin by assessing the extent to which minimum wage increases affected the wage distributions of low-skilled workers. Among workers with average baseline wages less than $7.50, the probability of reporting a wage between $5.15 and $7.25 declined substantially. We find that the wage distributions of low-skilled workers in bound and unbound states fully converge along this dimension. Further, we estimate that the minimum wage's bite on our target group's wage distribution is nearly twice its bite for a comparison sample of food service workers and teenagers.

We next estimate the minimum wage's effects on employment. We find that increases in the minimum wage significantly reduced the employment of low-skilled workers. By the second year following the $7.25 minimum's implementation, we estimate that targeted workers' employment rates had fallen by 6 percentage points (8 percent) more in bound states than in unbound states.

We further analyze a sample of teenagers and food service workers to compare our approach with approaches commonly used in the literature. For this sample, we estimate an employment decline of just under 4 percentage points. The estimated employment effects thus scale roughly in proportion to the minimum wage's bite on these groups' wage distributions. All else equal, Sabia, Burkhauser, and Hansen (2012) note that estimates of the minimum wage's effects on employment will scale with the extent to which an analysis sample contains unaffected workers. Their insight thus points to a partial line of reconciliation between the disemployment effects we observe and the statistical null results found in some of the recent literature. The magnitude of our estimated employment effects also likely reflects the setting we analyze, namely the Great Recession and subsequently sluggish recovery.

The primary threat to our estimation framework is the possibility that low-skilled workers in the bound and unbound states were differentially affected by the Great Recession. We show graphically that the housing crisis was more severe in unbound states, potentially biasing our estimates towards zero. In our baseline specification, we control directly for a proxy for the severity of the crisis. As noted above, our use of monthly, individual-level panel data enables us to more systematically construct within-state control groups with baseline wages only moderately higher than the new federal minimum. Our initial results are robust to netting out changes in the employment of these slightly higher skilled workers. We further find our estimates to be robust to a broad range of potentially relevant changes to our baseline specification.

In addition to reducing employment, we find that binding minimum wage increases increased the likelihood that targeted individuals work without pay (by 2 percentage points or 12 percent). This novel effect is concentrated among individuals with at least some college education. We take this as suggestive that such workers' entry level jobs are relatively readily posted as internships. For low-skilled, low-education workers, the entire change in the probability of having no earnings comes through unemployment.

We next estimate the effects of binding minimum wage increases on low-skilled workers' incomes and income *trajectories*. Our data provide a unique opportunity to investigate such effects, as the SIPP's monthly, individual-level panel extends for 3 years following the July 2009 increase in the federal minimum. To the best of our knowledge, this enables us to provide the first direct estimates of the minimum wage's effects on medium-run economic mobility. Given longstanding and widespread concern over developments in inequality (Katz and Murphy, 1992; Autor, Katz, and Kearney, 2008; Kopczuk, Saez, and Song, 2010), such effects may be of significant interest.

We find that this period's binding minimum wage increases reduced low-skilled individuals' average monthly incomes. Relative to low-skilled workers in unbound states, targeted workers' average incomes fell by $100 over the first year and by an additional

$50 over the following 2 years. While surprising at first glance, we show that the short-run estimate follows directly from our estimated effects on employment and the likelihood of working without pay. The medium-run estimate reflects additional contributions from lost wage growth associated with lost experience. Because most minimum wage workers are on the steep portion of the wage-experience profile (Murphy and Welch, 1990; Smith and Vavrichek, 1992), this effect can be substantial. We directly estimate, for example, that targeted workers experienced a 5 percentage point decline in their medium-run probability of reaching earnings greater than $1500 per month.[3] Like previous results, these estimates are robust to netting out the experience of workers with average baseline wages just above the new federal minimum. As in Kahn (2010) and Oreopoulos, von Wachter, and Heisz's (2012) analyses of the effects of graduating during recessions, we thus find that early-career opportunities have persistent effects.

After presenting these results, we compare our estimates of the minimum wage's effects with results from research on the effects of the Earned Income Tax Credit (EITC), an alternative policy for increasing the incomes of low-skilled workers. The EITC has been found to significantly increase the employment and incomes of low-skilled workers (Eissa and Liebman, 1996; Meyer and Rosenbaum, 2001; Eissa and Hoynes, 2006), reduce inequality (Liebman, 1998), and reduce tax-inclusive measures of poverty (Hoynes, Page, and Stevens, 2006). The results thus contrast rather sharply.

We conclude by assessing our estimates' implications for the effects of this period's minimum wage increases on aggregate employment. Over the late 2000s, the average effective minimum wage rose by nearly 30 percent across the United States. Our best estimate is that these minimum wage increases reduced the employment-to-population ratio of working-age adults by 0.7 percentage point. This accounts for 14 percent of the

[3] Earning $1500 would require working full time (40 hours per week for 4.33 weeks per month) at a wage of $8.80. We characterize $1500 as a "lower middle class" earnings threshold.

total decline over the relevant time period.

1 The Minimum Wage's Potential Effects

This section briefly motivates the outcomes on which we focus our empirical analysis.[4] The minimum wage can affect firm operations and workers' well-being through many channels. Following is a non-exhaustive list of potential margins of interest:

1. Wages of low-skilled workers.
2. Employment of low-skilled workers.
3. Income of low-skilled workers.
4. Income trajectories of low-skilled workers.
5. Firm offerings of benefits including health insurance.
6. Firm spending on the quality of workplace conditions.
7. Firm substitution between low-skilled labor, high-skilled labor, and capital.
8. Firm utilization of inputs with which low-skilled labor is complementary.
9. Incomes of firm owners.
10. Prices of goods produced by firms that employ minimum wage workers.

The list highlights the potential complementarity of firm-, industry-, and individual-level analyses of the minimum wage. Our analysis of individual-level panel data speaks to the minimum wage's most direct effects on its targeted beneficiaries. These include

[4]Because the relevant theoretical insights have been made in a literature extending from Stigler (1946) through Gramlich (1976) to Lee and Saez (2012), we do not intend to break new ground.

its intended and unintended effects on the wage distributions, employment, income, and income trajectories of low-skilled workers. Firm- and industry-level analyses may best speak to outcomes related to the quality of workplace conditions, the utilization of inputs that complement or substitute for low-skilled labor, the incomes of firm owners, and output prices.

2 Background on the Late 2000s Increases in the Federal Minimum Wage

We estimate the minimum wage's effects on employment and income trajectories using variation driven by federally mandated increases in the minimum wage rates applicable across the U.S. states. On May 25, 2007, Congress legislated a series of minimum wage increases through the "U.S. Troop Readiness, Veterans' Care, Katrina Recovery, and Iraq Accountability Appropriations Act." Increases went into effect on July 24th of 2007, 2008, and 2009. In July 2007, the federal minimum rose from \$5.15 to \$5.85; in July 2008 it rose to \$6.55, and in July 2009 it rose to \$7.25.

Figure 1 shows our division of states into those that were and were not bound by changes in the federal minimum wage. We base this designation on whether a state's January 2008 minimum was below \$6.55, rendering it partially bound by the July 2008 increase and fully bound by the July 2009 increase. Using Bureau of Labor Statistics (BLS) data on states' prevailing minimum wage rates, we designate 27 states as fitting this description.

Figure 2 shows the time paths of the average effective minimum wages in the states to which we do and do not apply our "bound" designation. Two characteristics of the paths of the minimum wage rates in unbound states are worth noting. First, their average minimum wage exceeded the minimum applicable in the bound states prior to the passage of the 2007 to 2009 federal increases. Second, these states voluntarily increased their minimums well ahead of the required schedule. On average, the effective minimum across these states had surpassed \$7.25 by January of 2008. This group's effective minimums rose, on average, by roughly 20 cents over the period we study, which extends for 4 years beginning in August 2008. By contrast, bound states saw their effective minimums rise by nearly the full, legislated \$0.70 on July 24, 2009. As borne in mind throughout, these states experienced differentially binding minimum wage increases in both July 2008 and July 2009. Our estimation framework, which we describe in the following section, may thus capture both the July 2009 increase's full effect and some dynamic effects of the increase from July 2008.

Allegretto, Dube, Reich, and Zipperer (2013) emphasize several distinguishing characteristics of states that have typically maintained minimum wage rates higher than the federal minimum and have thus been less likely to be bound by federal increases. They find these to be states with relatively liberal voting publics, relatively volatile business cycles, and relatively high degrees of job polarization. Figure 1 confirms the political divisions one might expect, as Republican-leaning states were much more likely to be bound by the recent increases in the federal minimum.

In past studies, business cycles have been particularly relevant due to the timing of state initiated increases in the minimum wage, which often take place near the business cycle's peak. This typically raises the possibility of an upward bias to estimated unemployment effects; states' economies tend to be in decline as their minimum wage increases go into effect. Because we estimate the effects of a binding federal minimum, the endogeneity of the timing with which any one state enacts an increase is of less concern. We must be wary of differences, however, in the Great Recession's severity in bound and unbound states. If unbound states experienced relatively severe housing bubbles, our estimates would potentially be biased towards 0.

Figure 3 presents data from the BLS, the Bureau of Economic Analysis (BEA), and the Federal Housing Finance Agency (FHFA) on the macroeconomic experiences of bound and unbound states during the Great Recession.[5] Throughout this time period, unbound states have higher per capita incomes, but lower employment-to-population ratios, than do bound states. While the economic indicators of both groups turned significantly for the worse over the recession's course, bound states were less severely impacted by the Great Recession than were unbound states. It is particularly apparent that unbound states had relatively severe housing bubbles (Panel C). These macroeconomic factors would, if controlled for insufficiently, tend to bias the magnitudes of our estimated employment impacts towards 0. The following section describes our empirical strategy for addressing this concern.

3 Data Sources and Estimation Framework

We estimate the effects of minimum wage increases using data from the 2008 panel of the Survey of Income and Program Participation (SIPP). We analyze a sample restricted to individuals aged 16 to 64 for whom the relevant employment and earnings data are available for at least 36 months between August 2008 and July 2012. For each individual, this yields up to 12 months of data preceding the July 2009 increase in the minimum wage. In the low-wage samples on which we focus, hourly wage rates are reported directly for 77 percent of the observations with positive earnings. For the remaining 23 percent, we impute hourly wages as earnings divided the individual's usual hours per week times their reported number of weeks worked. We use these 12 months of baseline wage, hours, and earnings data to divide low-skilled workers into 3 groups.

The first group we analyze includes those most directly impacted by the federal

[5]All series are weighted by state population so as to reflect the weighting implicit in our individual-level regression analysis.

minimum wage. Specifically, it includes those whose average wage, when employed during the baseline period, was less than \$7.50.[6] An essential early step of the analysis is to confirm that the increase in the federal minimum wage shifted this group's wage distribution as intended. The second group includes individuals whose average baseline wages were between \$7.50 and \$8.50. Because the employment situations of low-skilled workers are relatively volatile, this group's workers had non-trivial probabilities of working in minimum wage jobs in any given month. The extent of the minimum wage increase's effect on this group's wage distribution is an empirical question to which we allow the data to speak. The third group includes individuals whose average baseline wages were between \$8.50 and \$10.00. Guided by the baseline wage data, we characterize these workers as a comparison group of low-skilled workers for whom increases in the effective minimum wage had no direct effect.

Our initial estimates, conducted on a sample consisting of group 1 individuals, take the following, dynamic difference-in-differences form:

$$\begin{aligned} Y_{i,s,t} = & \sum_{p(t)\neq 0} \beta_{p(t)} \text{Bound}_s \times \text{Period}_{p(t)} \\ & + \alpha_{1_s} \text{State}_s + \alpha_{2_t} \text{Time}_t + \alpha_{3_i} \text{Individual}_i \\ & + \mathbf{X_{s,t}} \gamma + \mathbf{D_i} \times \text{Trend}_t \phi + \varepsilon_{i,s,t}. \end{aligned} \tag{1}$$

We control for the standard features of difference-in-differences estimation, namely sets of state, State_s, and time, Time_t, fixed effects. Our ability to control for individual fixed effects, Individual_i, renders controls for individual-level, time-invariant characteristics

[6]The average is calculated over months in which the individual was employed, excluding months when unemployed. The measure's intent is to capture the individual's average marginal product as remunerated by the firms for which he or she works. One consequence of this approach is that individuals who were unemployed throughout the baseline period are excluded from all samples. Because we estimate average wages using 12 months of baseline data, however, our samples include marginally attached individuals so long as they worked for at least one month between August 2008 and July 2009.

redundant. The vector $\mathbf{X_{s,t}}$ contains time varying controls for each state's macroeconomic conditions. In our baseline specification, $\mathbf{X_{s,t}}$ includes the FHFA housing price index, which proxies for the state-level severity of the housing crisis.[7]

Equation (1) allows for dynamics motivated by graphical evidence reported in Section 4. Specifically, we show in Section 4 that the prevalence of wages between the old and new federal minimum declined rapidly beginning in April 2009. We thus characterize May to July 2009 as a "Transition" period. Prior months correspond to the baseline, or period $p = 0$. We characterize August 2009 through July 2010 as period Post 1 and all subsequent months as period Post 2. The primary coefficients of interest are $\beta_{\text{Post 1}(t)}$ and $\beta_{\text{Post 2}(t)}$, which characterize the differential evolution of the dependent variable in states that were bound by the new federal minimum relative to states that were not bound. We calculate the standard errors on these coefficients allowing for the errors, $\varepsilon_{i,s,t}$, to be correlated at the state level. Because our treatment group contains 27 states, we do not face common inference concerns associated with imbalance between the number of treatment and control states (Bertrand, Duflo, and Mullainathan, 2004).[8]

We initially use equation (1) to confirm that binding minimum wage increases shift the distribution of wages as intended. For this analysis, we construct a set of outcome variables of the following form:

$$Y^{j}_{i,s,t} = 1\{W^{j-1} < \text{Hourly Wage}_{i,s,t} < W^{j}\}. \tag{2}$$

[7]It is not uncommon for minimum wage studies to control directly for a region's overall employment or unemployment rate. Conceptually, we find it preferable to exclude such variables because they may be affected by the policy change of interest. The housing price index is a conceptually cleaner, though still imperfect, proxy for time varying economic conditions that were not directly affected by minimum wage changes. Our results are essentially unaffected by the inclusion of additional state macroeconomic aggregates in $\mathbf{X_{s,t}}$. An analysis of our baseline result's robustness along this margin can be found in appendix table A6.

[8]We have confirmed that our standard errors change little when estimated using a block-bootstrap procedure with samples drawn at the state level. We conducted this exercise on a sample restricted to the 94 percent of the group 1 individuals that live in the same state throughout the sample.

These $Y^j_{i,s,t}$ are indicators that are set equal to 1 if an individual's hourly wage is between W^{j-1} and W^j. In practice each band is a 50 cent interval. The $\beta_{p(t)}$ from these regressions thus trace out the short and medium run shifts in the wage distribution's probability mass function that were associated with binding minimum wage increases.

We then move to our primary outcome of interest, namely the likelihood that an individual is employed. There are standard threats to interpreting the resulting $\beta_{p(t)}$ as unbiased, causal estimates of the effect of binding minimum wage increases. Most importantly, our estimates could be biased by differences in the Great Recession's severity in bound states relative to unbound states.

Within the difference-in-differences specification, we directly control for proxies for the macroeconomic experiences of each state. Recent debate within the minimum wage literature suggests that such controls may be insufficient.[9] Although we find our estimates of equation (1) to be robust to a range of approaches to controlling for heterogeneity in macroeconomic conditions, we additionally implement a triple-difference model. In this framework, displayed below, we use workers whose average baseline wages were between \$8.50 and \$10.00 to construct a set of within-state control groups:

$$\begin{aligned} Y_{i,s,t} = &\sum_{p(t)\neq 0} \beta_{p(t)} \text{Period}_{p(t)} \times \text{Bound}_s \times \text{Target}_{g(i)} \\ &+ \alpha_{1_{s,p(t)}} \text{State}_s \times \text{Period}_{p(t)} + \alpha_{2_{s,g(i)}} \text{State}_s \times \text{Target}_{g(i)} + \alpha_{3_{t,g(i)}} \text{Time}_t \times \text{Target}_{g(i)} \\ &+ \alpha_{4_s} \text{State}_s + \alpha_{5_t} \text{Time}_t + \alpha_{6_i} \text{Individual}_i + \mathbf{X_{s,t,g(i)}}\gamma + \mathbf{D_i} \times \text{Trend}_t \phi + \varepsilon_{i,s,t}. \end{aligned} \quad (3)$$

Equation (3) augments equation (1) with the standard components of triple-difference estimation. These include group-by-time-period effects, group-by-state effects, and state-

[9]Specifically, in criticizing work by Neumark and Wascher (2008) and Meer and West (2013), Allegretto, Dube, Reich, and Zipperer (2013) argue that their estimates of the minimum wage's effects are biased due to time varying spatial heterogeneity in economic conditions.

by-time-period effects. These controls account for differential changes in the employment of the target and within-state control groups over time, cross-state differences in the relative employment of these groups at baseline, and time varying spatial heterogeneity in economic conditions.

A shortcoming of the triple-difference approach involves the possibility of employer substitution of "within state control" workers for "target group" workers. Substitution of this form would lead the triple-difference estimates to overstate minimum wage increases' total employment impacts. In our context, we find that the estimated effects of minimum wage increases are relatively insensitive to shifting from the difference-in-differences framework to this triple-difference framework.

Table 1 presents summary statistics characterizing the samples on which we estimate equations (1) and (3). Focusing on columns 1 and 2, which characterize group 1 individuals, several differences between the samples from bound and unbound states are apparent. Individuals in bound states are moderately more likely to be employed and less likely to work without pay than are individuals in unbound states. They also tend to be slightly younger and less likely to obtain at least some college eduction.

By construction, our bound and unbound states differ in terms of their baseline minimum wage rates. Their policy environments converge upon the enactment of the new federal minimum. Baseline employment differences should thus not be surprising.[10] Demographic differences create the risk, however, that one might expect the employment trajectories of individuals in bound and unbound states to differ. Consequently, we test our specifications' robustness to the inclusion of $\mathbf{D_i} \times \text{Trend}_t$, an extensive set of demographic dummy variables interacted with linear time trends. We similarly confirm that our estimates are robust to controlling for a set of linear trends interacted with

[10]In a standard experimental setting, treatment and control groups are in similar environments at baseline, after which the treatment group is exposed to the treatment. In our setting, effective minimum wage rates differ at baseline and converge upon the implementation of the higher new minimum.

dummy variables associated with each individual's modal industry of employment over the baseline period. We further check the robustness of our estimates to a variety of additional specification modifications. Before presenting our estimates of equations (1) and (3), we use the following section to graphically present the raw data underlying our results.

4 Graphical View of the Wages, Employment, and Incomes of Low-Skilled Workers

Figure 4 presents time series tabulations of the raw data underlying our estimates of equations (1) and (3). In the panels of column 1, the sample consists of individuals whose average baseline wages were less than $7.50 per hour. In the panels of column 2, the sample consists of individuals whose average baseline wages were between $7.50 and $10.00 per hour.

The panels in row 1 plot the fraction of individuals that, in any given month, had an hourly wage between $5.15 and $7.25. Prior to the implementation of the $7.25 federal minimum, individuals in states that were bound by the federal minimum were much more likely to have wages in this range than individuals in unbound states. Those in bound states spent roughly 40 percent of their months in jobs with wages between $5.15 and $7.25, 22 percent of their months unemployed, 17 percent of their months in unpaid work, and their remaining months in sub-minimum wage jobs (e.g., tipped work) or in jobs paying more than $7.25. By contrast, individuals in unbound states spent 22 percent of their baseline months in jobs with hourly wages between $5.15 and $7.25. These fractions began converging in April 2009, three months before the new federal minimum took effect.[11] The observed transition period motivates our accounting for

[11]The transition window likely reflects a combination of real economic factors and measurement arti-

dynamics when estimating equations (1) and (3). By November 2009, individuals in the bound and unbound states have equal likelihoods of being in jobs with wages between \$5.15 and \$7.25.

Panel B shows that the wages of individuals with average baseline wages between \$7.50 and \$10.00 per hour were largely unaffected by the increase in the federal minimum. Their probability of having a wage between \$5.15 and \$7.25 in any given month was around 5 percent. Prior to the increase in the federal minimum, individuals in bound states had marginally higher probabilities of having such wages.

The panels in row 2 plot our initial outcome of interest, namely the fraction of individuals who are employed. Low-skilled workers in states with low minimum wages initially had moderately higher employment rates, by about 3 percentage points, than those in states with higher minimums. As wages adjusted to the new federal minimum, this baseline difference narrows. Over subsequent years, the employment of those in bound states is, on average, roughly 1 percentage point less than that of low-skilled individuals in unbound states. Relative to the baseline period, the differential employment change observable in the raw data is 3 percentage points in the first year and 4 percentage points in subsequent years. The data exhibit the seasonality one would expect in the employment patterns of the relevant populations. The employment series' convergence appears linked, at least initially and in part, to a relatively weak summer hiring season in states bound by the July 2009 increase in the federal minimum.

If these employment changes were driven primarily by cross-state differences in the severity of the Great Recession, similar (perhaps slightly smaller) changes would be

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facts. Employers hiring workers in May and June 2009 may simply have found it sensible to post positions at the wage which would apply by mid-summer rather than at the contemporaneous minimum. The measurement issue involves the SIPP's 4 month recall windows. Individuals interviewed about their May and June wages in August 2009 may have mistakenly reported their August wage as their wage throughout the recall window. Our response to both potential explanations is to allow for flexible dynamics when estimating the minimum wage's effects on employment.

expected among workers with modestly greater skills. Panel D shows that such changes did not occur. Comparing the bound and unbound states, the employment of workers with average baseline wages between $7.50 and $10.00 moved in parallel over this period. These data reveal that estimates of equations (1) and (3) will yield similar results.

The panels in row 3 show similar patterns for trends in average monthly income. During the baseline period, the average incomes of low-skilled individuals in bound and unbound states evolve similarly. Several months following the increase in the federal minimum wage, the income growth of low-skilled individuals in the bound states begins to lag the income growth of low-skilled individuals in unbound states. No such divergence is apparent among individuals with baseline wages between $7.50 and $10.00. The data reported in Panel E suggest that the wage gains and employment declines of targeted workers initially offset one another. Subsequently, declines in employment and experience accumulation appear to have led the income growth of low-skilled individuals in bound states to lag that of low-skilled workers in unbound states. In Section 5.5 we present a detailed analysis of the factors contributing to these differential income trajectories.

Figures 5 and 6 present these data in regression-adjusted form. Each marker in the figures is an estimate of a coefficient of the form $\beta_{p(t)}$ from equations (1) and (3), where each $p(t)$ corresponds with an individual month; period $p = 0$ is April 2009, the month immediately preceding the transition period. The regression-adjusted changes in employment and income are largely as one would expect based on the raw data presented in Figure 4. Adjusting for the housing bubble's greater severity in unbound states relative to bound states moderately increases the estimated magnitudes. The following section presents these and other results in a more summary, tabular fashion.

5 Regression Analysis of the Minimum Wage's Effects

This section presents our estimates of equations (1) and (3). We begin by verifying that the enacted minimum wage increases shifted the wage distributions of workers with average baseline wages below $7.50 as intended. We then estimate the minimum wage's effect on employment, after which we explore several additional outcomes relevant to the welfare of affected individuals and their families.

5.1 Effects on Low-Skilled Workers' Wage Distributions

This section first presents data on the baseline wage distributions of low-skilled workers. It then presents estimates of the extent to which these distributions shift following binding minimum wage increases. Figure 7 characterizes the wage distributions of workers with average baseline wages below $7.50 (Panel A), average baseline wages between $7.50 and $8.50 (Panel B), and average baseline wages between $8.50 and $10.00 (Panel C). The histogram in each panel presents the distribution of each group's wages during the baseline period. This distribution, and in particular the frequency of wage rates in the affected region, is the basis upon which we select our "target" and "within-state control" groups.[12] Note that the histograms exclude the large mass of observations with no earnings, which includes months spent either unemployed or working without pay.

The histogram for workers with average baseline wages below $7.50 has substantial mass associated with monthly wage rates between $6.50 and $7.50, as shown in Panel A. Panel B shows that workers with average baseline wages between $7.50 and $8.50 have far less mass in the affected region. Nonetheless, this groups' employment and earnings

[12] Specifically, we choose our "target" group to be a group with significant baseline mass in the affected region and our "within-state control" group to be the lowest-skilled group that spends essentially no baseline months with wage rates in the affected region. The estimated effects of binding minimum wage increases on these distributions confirms that the former's distribution shifted significantly while the latter's did not.

are sufficiently volatile that they appear to spend non-trivial numbers of months in minimum or near-minimum wage jobs. Panel C reveals workers with average baseline wages between \$8.50 and \$10.00 to be low-skilled workers who spent essentially none of their baseline months at affected wage rates. Because these individuals were not directly affected by the increased federal minimum, we view them as an ideal sample for constructing within-state employment counterfactuals for estimating equation (3).

Table 2 and the panels of Figure 7 present estimates of the wage distribution shifts that were associated with binding minimum wage increases. Each of the relevant markers represents a point estimate from a separate estimate of equation (1). The marker just to the right of the dashed line at \$6.55, for example, represents the change in the probability of having a wage between \$6.50 and \$7.00. As in all of this paper's specifications, the sample includes observations whether or not an individual was unemployed.[13]

Panel A shows that, for individuals with average baseline wages below \$7.50, the wage distribution shifted significantly out of precisely the targeted region. As summarized in Table 2's column 1, the probability of having a wage between \$5.15 and \$7.25 declined by just over 16 percentage points. This mass does not shift exclusively to the new federal minimum; a portion collects between \$7.50 and \$8.00.[14]

Appendix Figure A1 and Table A1 characterize the bite of binding minimum wage increases on the wage distributions of groups of workers that are commonly analyzed in the literature. Figure A1's Panels A and B display the wage distributions of teenagers and food service workers. As summarized in Table A1, the minimum wage's bite on these groups' wage distributions is just over half the size of its bite on the distribution for workers with average baseline wages below \$7.50. Relative to our analysis of workers

[13]The panels of this figure do not report a marker associated with having a wage of 0, which would correspond to the "No Earnings" outcome analyzed in Section 5.5.

[14]We take this evidence as being consistent with that found in Katz and Krueger's (1992) longitudinal survey of Texas food service establishments.

with average baseline wages in the affected range, analyses of these groups will thus have an attenuated ability to detect any effects of minimum wage increases on employment.

In Figure 7's Panels B and C we characterize the extent to which the wage distributions of workers higher up the skill distribution were bound by the increase in the minimum wage. For Panel B we repeat the exercise conducted for Panel A, but on the sample of workers with average baseline wages between \$7.50 and \$8.50. Because this group's members spent non-trivial numbers of months in jobs with directly affected wages, its wage distribution shifts non-trivially out of the affected region. As reported in Table 2's column 2, the direct effect of the increased minimum wage was to reduce this group's probability of having a wage between \$5.15 and \$7.25 by 4 percentage points. The shifted mass collects entirely in the lower half of this group's average baseline wage range, i.e., between \$7.50 and \$8.00.

Finally, we characterize the minimum wage's bite on the distribution of wages for those with average baseline wages between \$8.50 and \$10.00. Figure 7's Panel C reveals no evidence of systematic movements in this group's wage distribution. Table 2's column 3 confirms that the minimum wage increase had an economically negligible effect on this group's wages; the upper bound of the 95 percent confidence interval suggests that the reduction in the probability of earning a wage between \$5.15 and \$7.25 was less than 2 percentage points.

The histograms in Figures 7 and A1 display our approach's suitability for identifying both targeted workers and workers who were low-skilled but unaffected, making them attractive as within-state controls. As desired, the baseline wage distribution for workers with average baseline wages less than \$7.50 has significant mass between \$6.50 and \$7.50. Our within-state control group has a baseline wage distribution tightly clustered between \$8.00 and \$10.00. As illustrated in figure A1's panel C, comparison samples drawn based on industries will tend to contain many much higher skilled, and thus less directly comparable, individuals. Figure A1's panels A and B show that analysis samples of teenagers and food service workers similarly have baseline wage distributions more diffuse than that of our target sample.

5.2 Baseline Results on Employment

Table 2's columns 4 through 6 present estimates of equation (1) in which the outcome is an indicator for being employed. Column 4 reports the result for individuals with average baseline wages less than \$7.50. The coefficient in row 1 implies that binding increases in the federal minimum wage resulted in a 4.3 percentage point decline between the baseline period and the following year. The decline relative to baseline averaged 6.3 percentage points over the two subsequent years (the "medium run").

Column 5 shows the result for the group with average baseline wages between \$7.50 and \$8.50. The estimated effect of the minimum wage on this group's employment is statistically indistinguishable from 0, with a medium-run point estimate of negative 1.1 percentage points. Finally, column 6 shows the result for the group with average baseline wages between \$8.50 and \$10.00. The estimated effect on this group's employment is again quite close to zero. Like the raw data from Figure 4, these results reveal that estimates of equations (1) and (3) will yield similar results.

Appendix Table A5 further fleshes out our estimates of the effect of binding minimum wage increases on employment among the adult population. To the results reported in Table 2, it adds estimates associated with adults who were either unemployed throughout the baseline period (column 1) or whose average baseline wages were equal to or greater than \$10.00 (column 5). The estimates for both groups are economically and statistically indistinguishable from 0. The point estimate for the effect of binding minimum wage increases on the subsequent employment of those who were unemployed at baseline is modestly negative.

5.3 Contrasting Approaches To Evaluating the Minimum Wage

In further analysis, we estimate the minimum wage's effects on the employment of populations studied frequently in the literature, namely teenagers and food service workers. More specifically, we estimate equation (1) on a sample selected to include individuals who were teenagers or for whom food service was the modal industry of employment during the baseline period. Column 5 of Appendix Table A1 reports our estimate that binding minimum wage increases reduced this sample's medium-run employment by 3.9 percentage points. Column 6 reports an estimate near 0 for the minimum wage increase's effect on the employment of manufacturing workers, whose wage distribution was unaffected. Our specification thus passes the primary falsification test emphasized in a recent exchange involving Dube, Lester, and Reich (2010), Meer and West (2013), and Dube (2013).

We draw two lessons from comparing the estimates associated with our baseline sample and the sample of teenagers and food service workers. First, the estimates associated with teenagers and food service workers reinforce the conclusion that this period's minimum wage increases reduced the employment of low-skilled workers. Second, they point to a potential line of reconciliation between some of the literature's null results and our finding of significant disemployment effects.

As emphasized by Sabia, Burkhauser, and Hansen (2012), cross-study comparisons require scaling estimates by the extent to which alternative analysis samples are actually affected by the minimum wage. Comparisons involving estimates from industry-level studies are particularly difficult because such studies typically lack the individual-level data required to directly estimate the minimum wage's bite on the underlying workers' wage distribution.[15] We estimate that the wage distribution of our target sample was

[15]The extent of the minimum wage's bite on populations under study is often inferred from CPS data. A variety of measurement issues make it rare, however, to have directly comparable estimates of the minimum wage's effects on the wage distributions of alternative study populations. Relevant

nearly twice as affected as the wage distribution of teenagers and food service workers. Our estimates of the minimum wage increase's effects on these groups' employment were similarly proportioned. It is thus important to note that, all else equal, estimates of a minimum wage increase's effects on relatively untargeted groups will be attenuated and, as a result, more prone to type II error.

Appendix Table A4 provides a further line of comparison between our results and the findings of industry-specific analyses of the minimum wage. In our baseline analysis and our analysis of teenagers and food service workers, we estimate the minimum wage's effects on the employment of low-skilled *individuals*. By contrast, analyses of industry-level data estimate the minimum wage's effects on total employment in low-skill-intensive *industries*. In Table A4 we present estimates of the minimum wage's effect on the probability that any given individual is employed in the food service sector. For the full sample of individuals aged 16 to 64, the estimated effect on food service employment is economically negligible and statistically indistinguishable from 0. As revealed in column 2, this masks a 3 percentage point decline in food service employment among individuals with average baseline wages below $7.50. Column 3 reports an offsetting increase in the food service employment of workers with higher baseline wage rates.[16]

We draw two additional lessons from this analysis. First, we note that the minimum wage's effects may vary significantly across industries, making it difficult to extrapolate from industry-specific estimates to aggregate employment. In a standard model, the

measurement issues include survey reporting error and variation in the minimum wage's applicability due to exceptions such as those made for tipped workers. Sabia, Burkhauser, and Hansen (2012) and the present study's appendix materials are the only recent examples of such analyses of which we are aware. An alternative approach to inferring the minimum wage increase's direct effect involves using industry- or firm-level data to estimate its effect on average earnings per worker, as in Dube, Lester, and Reich (2010). In such data, however, increases in earnings per worker may reflect either increases in the earnings of the low-skilled or substitution of high-skilled workers for low-skilled workers. Absent additional information, such data will not enable researchers to distinguish between these outcomes.

[16]Because the sample in column 3 is roughly 10 times the size of the sample in column 1, the -0.03 employment effect from column 2 is essentially fully offset by the estimate of 0.003 from column 3.

determinants of an industry's adaptation to a minimum wage change include its ability to substitute between low-skilled workers, high-skilled workers, and capital, as well as the elasticity of demand for its output. The results in Table A4 provide evidence that, during the period we study, food-service employers had significant scope for substituting between low- and high-skilled workers.

Second, the results highlight that substitution between low- and high-skilled workers can complicate efforts to evaluate the minimum wage's effects using data on industry-level wage bills and employment. In such data, the results in Table A4 would be indistinguishable from an outcome in which an increase in the minimum wage non-trivially increased per-worker earnings and had minimal effects on employment. In the setting we analyze, this mistaken interpretation would leave the impression that the minimum wage had achieved its objective of increasing low-skilled workers' incomes at little cost.

5.4 Robustness of the Estimated Employment Effects

Table 2's primary result of interest is column 4's estimate of the minimum wage increase's effect on the employment of targeted workers. Tables 3 and A6 present an analysis of this result's robustness. In Table 3, estimates in Panel A are of equation (1)'s difference-in-differences model. Estimates in Panel B are of equation (3)'s triple difference model, in which we use workers with average baseline wages between $8.50 and $10.00 as a within state control group.

The result in column 1 of Panel A replicates the finding from Table 2's column 4. The result in column 1 of Panel B shows this result to be robust to estimating the minimum wage's effect using the triple-difference framework. The medium run estimate implies that binding minimum wage increases reduced the target group's employment rate by 5.9 percentage points.

Column 2 presents results in which we exclude our controls for states' macroeconomic conditions. Not controlling for variation in the housing bubble's severity across states reduces the estimated coefficients by 2 percentage points. The estimated employment decline is 4.2 percentage points in the difference-in-differences model and 3.8 percentage points in the triple-difference model. This reflects the fact that, as shown in Figure 3, the housing bubble was more severe in unbound states than in bound states. We further explore the relevance of macroeconomic controls in Table A5, which we discussion momentarily.

Column 3 shows that our results are robust to controlling for state-specific linear time trends.[17] In the difference-in-differences specification, including these controls increases the estimated medium-run coefficient from 6.3 to 7.2 percentage points. Column 4 shows that our results are relatively insensitive to controlling for exhaustive sets of age, education, and family-size indicators interacted with linear time trends.[18] Differential trajectories linked to moderate differences in the demographic characteristics of the group 1 samples in bound and unbound states thus appear unlikely to underlie our estimates. We find the same to be true of differences associated with bound and unbound states' industrial compositions.

The remaining columns involve changes in our sample inclusion criteria. Column 5 shows that our results are robust to requiring that, for inclusion in the final sample,

[17]We share Meer and West's (2013) concern that, because of the dynamics with which minimum-wage induced employment losses may unfold, direct inclusion of state-specific trends is not a particularly attractive method for controlling for the possibility of differential changes in the economic conditions of each state over time. The dynamics allowed for by our Transition, Post 1, and Post 2 periods turn out, in this context, to be sufficient to render state-specific trends largely irrelevant. This is less true in later analysis of the minimum wage's effects on income. The minimum wage may affect income through direct disemployment effects, subsequent effects on experience accumulation, and related effects on training opportunities. The latter effects will be realized as effects on income growth, making Meer and West's (2013) critique particularly pertinent.

[18]We similarly find our results to be robust to controlling for time trends interacted with dummy variables for 20 cent bins in our measure of average baseline wages (result not shown). This check is addressed at the concern that, because minimum wage workers in unbound states had relatively high wages at baseline, their employment and earnings trajectories might differ for reasons related to mean reversion.

individuals appear in the sample for at least 42 months rather than our baseline requirement of 36 months. The specifications in columns 6 and 7 involve modifications to our criteria for categorizing the bound and unbound states. Column 6 drops unbound states in which the January 2008 minimum wage was less than $7.00, as such states were moderately bound by subsequent increases in the federal minimum. Removing these 4 states (Arizona, Florida, Missouri, and West Virginia) from the control group modestly increases the medium-run point estimate to 6.7 percentage points in the difference-in-differences model and leaves the triple-difference estimate unchanged at 5.9 percentage points. Finally, column 7 removes from the sample any bound state with a January 2009 minimum wage above $6.55. Our baseline designation uses states' January 2008 minimum wage rates to ensure that it is based on decisions made before our sample begins. We observe that 4 states (Montana, Nevada, New Hampshire, and New Mexico) with January 2008 minimum wage rates below $6.55 voluntarily increased their minimums before they were required to do so. Dropping these states from the sample modestly decreases the medium-run estimates in both the difference-in-differences and triple difference specifications (by several tenths of a percentage point in each case).

Appendix Table A6 provides additional evidence regarding the relevance of controls for differences in the severity of the Great Recession in bound and unbound states. Columns 1 and 2 replicate columns 1 and 2 from panel A of Table 3. As an alternative to controlling for the housing price index, column 3 adds controls for state level income and employment per capita. Column 4 adds controls for stimulus spending per capita and two additional variables. The first, "Predicted State Income," is a projection of state-specific changes in aggregate output that are predictable on the basis of each state's historical relationship with the national business cycle. The second, "Predicted State Employment," is a projected change in employment based on each states' baseline industrial composition and subsequent industry-specific employment growth at the national level (Bartik, 1991; Blanchard and Katz, 1992). The inclusion of these alternative macroeconomic control variables increases the estimated effect of binding minimum wage increases relative to specifications that include no such controls. When these variables are included alongside the housing price index, the estimates are essentially unchanged from the baseline. The housing price index consistently emerges as a stronger predictor of employment among low-skilled individuals than the alternative macroeconomic control variables. The specifications in columns 6 and 7 incorporate state-specific trends, the full sets of trends in various demographic characteristics, and trends specific to each individual's modal industry of employment at baseline. In both of these specifications, we estimate that binding minimum wage increases resulted in 7 percentage point declines in the employment of low-skilled workers.

5.5 Further Employment Outcomes, Average Income, and Poverty

Tables 4 and 5 report the results of a more in depth analysis of the minimum wage's effects on employment and income related outcomes. In Table 4's second column we present evidence of a novel channel through which job markets may respond to minimum wage increases. Specifically, we show that binding minimum wage increases resulted in an increase in the probability that targeted individuals work without pay, perhaps in internships, by 2 percentage points. Between disemployment and work without pay, column 3 reports a combined 8 percentage point reduction in paid employment. Appendix Tables A7 and A8 show that our estimates of both the "internship" effect and the total effect on paid employment are robust to the same set of specification changes as our estimate of the traditional disemployment effect. Estimates of the medium-run effect on the probability of working without pay range from 1.3 to 2.5 percentage points. Estimates of the total effect on the probability of paid employment range from 6.0 to 9.4 percentage points.

Table 5 presents estimates of these effects separately for individuals with and without at least some college education. Here it is important to keep in mind that all of the individuals in the sample have average baseline wages less than $7.50. Those with at least some college education are relatively likely to be current students or very early in their careers.

We find that low-skilled workers with at least some college education underlie the entirety of the minimum wage's effect on the likelihood of working without pay. The increase in the federal minimum wage made low-skilled workers with at least some college education 4 percentage points (roughly 20 percent) more likely to work without pay. For individuals with less education, the entirety of the minimum wage's effect on paid employment comes through unemployment. These findings are suggestive of a difference in the entry-level positions of high- and low-education workers. Entry positions sought by high-education workers appear relatively interchangeable with unpaid internships. Appendix Tables A11 and A12 report further analysis of demographic heterogeneity in our estimates of the minimum wage's effects.

Returning to outcomes more central to the minimum wage's redistributive properties, Table 4's columns 4 and 5 report the effect of binding minimum wage increases on average monthly incomes. Column 4 reports the effect on individual-level income while column 5 reports the effect on family-level income. We censor these outcomes at $7,500 and $22,500 per month respectively; this affects fewer than 1 percent of observations, which are associated with incomes far beyond those attainable through minimum wage employment. In our difference-in-differences specification, we estimate that binding minimum wage increases reduced the average monthly income of low-skilled workers by $97 in the short-run and $153 in the medium-run. Results are slightly larger, though estimated with significantly less precision, in our triple-difference specification. Robustness across these specifications is particularly relevant for outcomes involving income.

Specifically, it reassures us that the results are not spuriously driven by growth in the control-group workers' incomes towards the relatively high per capita incomes associated with unbound states (recall Panel D of Figure 3).

Figure 6 more fully highlights the dynamics underlying these results. In the figure it is apparent that employment losses and wage gains offset one another over the transition months. Accumulating employment losses and lost wage gains associated with lost experience begin outstripping the legislated wage gains in subsequent periods. Appendix Table A9 reports the robustness of the estimated effects on average income to the same set of specification checks as the outcomes previously analyzed.

To better understand these estimates, note that targeted individuals in bound states had positive earnings in 61 percent of baseline months. In 22 percent they were unemployed and in 17 percent they worked with 0 earnings. Average income for the target sample was $770 across all baseline months, and thus roughly $1,260 in months with positive earnings. For the short run (i.e., year 1), we estimated a 6 percentage point decline in the probability of having positive earnings. This effect is thus directly associated with an average decline of roughly $75, or $\$1,260 \times 0.06$. The decline in months with positive earnings rises to 8 percentage points over the following two years, implying a direct earnings decline of $100. Gains for workers successfully shifted from the old minimum to the new minimum offset very little of this decline.[19]

[19]Recall that we estimated a 16 percentage point decline in the probability of having a wage between $5.15 and $7.25. Nearly half of this turns out to involve shifts into unemployment or unpaid work. The wage increase for the remaining 8 percentage points was roughly 10 percent (from the $6.55 minimum for 2008 to the $7.25 minimum for 2009). A 10 percent increase on the $1,260 base, realized by 8 percent of workers, averages to a gain of $10. Measurement error in self-reported wage rates likely leads this approach to understate the true gain; it likely attenuates our estimates of the minimum wage's bite on the wage distributions of low-skilled workers. An alternative approach, likely generating an upper bound, is to infer the minimum wage's bite from the data displayed in Figure 4. Figure 4's panel A showed that low-skilled workers in bound states saw their probability of reporting a wage between $5.15 and $7.25 decline by roughly 35 percentage points from a base of just over 40 percentage points. Even the 35 percentage points of bite one could maximally infer from figure 4 implies quite modest offsets of the income losses associated with disemployment, work without pay, and lost experience accumulation.

The effects of lost employment rise over time due to lost experience. Minimum wage workers tend to be on the steep portion of the wage-experience profile (Murphy and Welch, 1990). Using mid-1980s SIPP data, Smith and Vavrichek (1992) found that 40 percent of minimum wage workers experienced wage gains within 4 months and that nearly two-thirds did so within 12 months. The median gain among the one-year gainers was a substantial 20 percent. Among those unemployed or working without pay, foregone wage growth of these magnitudes brings the implied medium-run earnings decline to $130.[20] Targeted workers who maintain employment may also experience slow earnings growth if employers reduce opportunities for on the job training.

Our estimates of the minimum wage increase's effect on income are initially somewhat surprising. As illustrated above, however, they follow from the magnitude of our estimated employment effects coupled with three more conceptually novel factors. These factors include our finding of an "internship" effect, effects on income growth through reduced experience accumulation, and the fact that direct effects on wages were smaller than typically assumed.

We emphasize that our analysis involves increases in the minimum wage that took effect during a period of significant labor shedding. The employment, internship, and experience-accumulation effects are thus likely to have been particularly potent during this historical episode. Finally we note that our income point estimates come with considerable uncertainty. The associated standard errors are sufficiently large that we cannot rule out relatively modest declines in average income.

Returning to Table 4, we estimate the minimum wage's effects on family-level outcomes. On average in our sample, each targeted worker is in a family with 1.3 targeted workers. This is roughly the average of the ratio of our estimates of the minimum wage

[20]Two years of early-career earnings growth at 15 percent per year would bring earnings from a baseline of $1,260 to $1,670. An 8 percentage point decline in months at such earnings implies an average reduction of $133.

increase's effect on family-level income to its effect on individual-level income. In the triple-difference specifications, for example, the short-run effect on individual-level income is $112 per month while the estimated effect on family-level income is $120 (the medium-run estimates are $185 and $300).

Finally, column 6 shows that the effect of binding minimum wage increases on the incidence of poverty was statistically indistinguishable from 0. Unsurprisingly, given our finding on family-level earnings, the point estimate for the medium-run effect on the likelihood of being in poverty is positive. The absence of a decline in poverty echoes findings by Burkhauser and Sabia (2007), Sabia and Burkhauser (2010), Neumark and Wascher (2002), and Neumark, Schweitzer, and Wascher (2005), as well as a summary of earlier evidence by Brown (1999).

5.6 Transitions from Low-Wage Work into Middle Class Earnings

We next analyze income growth through the lens of economic mobility, a topic of significant recent interest (Kopczuk, Saez, and Song, 2010; Chetty, Hendren, Kline, and Saez, 2014; Chetty, Hendren, Kline, Saez, and Turner, 2014). Concern regarding the minimum wage's effects on upward mobility has a long history (Feldstein, 1973). A potential mechanism for such effects, namely the availability of on-the-job training, has received some attention in the literature (Hashimoto, 1982; Arulampalam, Booth, and Bryan, 2004). We are not aware, however, of direct evidence of the minimum wage's effects on individuals' transitions into employment at higher wages and earnings levels.

Because we observe individuals for four years, we are able to track transitions of low-wage workers into middle and lower middle class earnings. The data reveal that initially low-wage workers spend non-trivial numbers of months with earnings exceeding those of a full time, minimum wage worker. Consider earnings above $1500, which could be generated by full time work at $8.80 per hour. During the first year of our sample, workers with average baseline wages less than $7.50 earn more than $1500 in 8 percent of months. By the sample's last two years this rises, adjusting for inflation, to 18 percent. We investigate the minimum wage's effects on the likelihood of reaching such earnings.

Table 6 reports the results. We find significant declines in economic mobility, in particular for transitions into lower middle class earnings. For the full sample with average baseline wages less than $7.50, the difference-in-differences estimate implies that binding minimum wage increases reduced the probability of reaching earnings above $1500 by 4.9 percentage points. This represents a 24 percent reduction relative to the control group's medium-run probability of attaining such earnings. As with previous results, this finding cannot readily be explained by cross-state differences in economic conditions. Netting out the experience of individuals with baseline wages between $8.50 and $10.00 moderately increases the point estimate to 5.4 percentage points (26 percent).

The estimated reductions in the probability of reaching lower middle class earnings levels are particularly meaningful for low-skilled workers with no college education. In the difference-in-differences specification, the estimated decline in this group's probability of earning more than $1500 per month is 4.9 percentage points (see column 2). In the triple-difference specification the estimate is 8.2 percentage points. Declines of these magnitudes represent 32 and 54 percent declines relative to the control group's probability of reaching such earnings. For those with at least some college education, the estimated declines average a more moderate 4 percentage points, equivalent to 17 percent of the control group's probability of reaching such earnings. Figure 8 presents the raw data underlying these results, and Appendix Table A10 reports the robustness of the estimated effects to the same set of specifications checks as the outcomes previously analyzed.

We next examine the probability of reaching the middle-income threshold of $3000 per month. For the full sample, we estimate that binding minimum wage increases reduced this probability by 1.8 percentage points. In the difference-in-differences specification, this estimate is statistically distinguishable from 0 at the 10 percent level; in the triple-difference specification this is not the case, although the point estimate is essentially unchanged. Though our sub-sample analysis has little precision, the average medium-run effect appears to be driven primarily by those with at least some college education. The full sample decline of 1.7 percentage points is a non-trivial 26 percent of the control group's medium-run probability of reaching such earnings.

We interpret the evidence as implying that binding minimum wage increases reduced the medium-run class mobility of low-skilled workers. Such workers became significantly less likely to rise to the lower middle class earnings threshold of $1500 per month. The reduction was particularly large for low-skilled workers with relatively little education.

The dynamics of our estimated employment and class mobility results are suggestive of the underlying mechanisms. Our employment results emerge largely during the first year following the increase in the federal minimum wage. By construction, our mobility outcomes are not outcomes that can be affected by the loss of a full time minimum wage job. Effects on mobility into lower middle class earnings only emerge over subsequent years. It appears that binding minimum wage increases blunted these workers' prospects for medium-run economic mobility by reducing their short-run access to opportunities for accumulating experience and developing skills. This period's minimum wage increases may thus have made the first rung on the earnings ladder more difficult for low-skilled workers to reach.

5.7 Contrasting the Minimum Wage and the Earned Income Tax Credit

The minimum wage is one of many policies implemented with an objective of increasing the effective wage rates or earnings of low-skilled workers. In the U.S. context, the policy most obviously interchangeable with the minimum wage is the Earned Income Tax Credit (EITC).[21] Analyses of the relative effectiveness of redistributing via wage regulation versus the tax code date at least as far back as Stigler (1946), whose discussion ranged from potential employment effects to target-efficiency and administrative complexity. Our estimates speak to the minimum wage's effectiveness in achieving its direct objective of increasing the incomes of targeted workers. In the paragraph below, we contrast our estimates with the relevant results from the literature on the EITC.

Our estimates provide evidence that binding minimum wage increases reduced the employment, average income, and income growth of low-skilled workers over short- and medium-run time horizons. By contrast, analyses of the EITC have found it to increase both the employment of low-skilled adults and the incomes available to their families (Eissa and Liebman, 1996; Meyer and Rosenbaum, 2001; Eissa and Hoynes, 2006). The EITC has also been found to significantly reduce both inequality (Liebman, 1998) and tax-inclusive poverty metrics, in particular for children (Hoynes, Page, and Stevens, 2006). Evidence on outcomes with long-run implications further suggest that the EITC has tended to have its intended effects. Dahl and Lochner (2012), for example, find that influxes of EITC dollars improve the academic performance of recipient households' children. This too contrasts with our evidence on the minimum wage's effects on medium-run economic mobility.

[21]The potential substitutability and complementarity of regulatory and tax-financed redistributive measures is an underlying theme in Clemens's (2014) analysis of community rating regulations in insurance markets.

6 Implications for Changes in Aggregate Employment Over the Great Recession

Between December 2006 and December 2012, the average effective minimum wage rose from $5.88 to $7.56 across the United States. Over this same time period, the employment-to-population ratio for adults aged 16 to 64 declined by nearly 5 percentage points. Clemens and Wither (2014) more fully characterize this period's employment declines, including its demographic and cross-country dimensions. Sustained U.S. employment declines were particularly dramatic for young adults aged 15 to 24. As of September 2014, this group's employment remained down by 7.5 percentage points from its pre-recession peak. Additionally, U.S. employment declines generally exceeded those that occurred in other advanced economies.[22] These dimensions of the data suggest that U.S.-specific developments in low-skilled labor markets underlie much of the slump in U.S. employment. We thus conclude by considering our results' implications for the relationship between minimum wage increases and these sustained employment declines.

Panel B of Appendix Table A5 presents the full set of results required to infer aggregate employments effects from our estimates. The table presents estimates of equation (1) on sub-samples that fully partition the set of adults aged 16 to 64. Column 2 replicates our baseline estimate that binding minimum wage increases reduced the target population's employment rate by 6 percentage points. The remaining estimates provide evidence that these minimum wage increases had little if any effect on the employment of other groups, which is the assumption maintained for this section's calculation.[23]

[22]See Hoffmann and Lemieux (2014) for related characterizations of cross-country developments in unemployment rates.

[23]Note that for this exercise the estimated effects on other groups' employment must be directly interpreted as estimates of the minimum wage increase's effects rather than as evidence supporting the validity

Several factors make extrapolating our results into changes in national employment-to-population ratios somewhat speculative. First, as emphasized earlier, we have likely estimated both the full effect of the July 2009 increase and some dynamic effects of the July 2008 increase in the federal minimum. In July 2009, the differential increase between bound and unbound states was $0.58. We may more plausibly be detecting the effect of an increase equivalent to $0.80 or $0.90 (roughly 13 percent of initial levels). Second, the employment effects of minimum wage increases may be non-linear, with larger effects for final increments than for initial increases. Third, the nominal value of the minimum wage has more purchasing power, and hence more bite, in the typical bound state than in the typical unbound state. Figure 3 showed bound states to have relatively low incomes and housing costs, making it unsurprising that they had not voluntarily brought their minimums to the levels seen in the unbound states. These issues inform the lower and upper bounds we place on our estimates of the minimum wage's aggregate effects.

The calculation proceeds as follows. Our baseline estimate is that binding minimum wage increases reduced the employment of workers with average baseline wages below $7.50 by 6 percentage points (8 percent).[24] For this section's purposes, the estimates reported in Appendix Table A5 make us reasonably confident in treating the decline in the target group's employment as the sole effect of this period's binding minimum wage increases on aggregate employment. Applying the relevant weights, the target sample represents 7.4 percent of the U.S. population aged 16 to 64. A 6 percentage point decline

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of our estimates of the minimum wage's effect on the target group.

[24]The implied elasticity of the sample-group's employment with respect to the minimum wage is thus around two-thirds (8 percent divided by the 13 percent from the previous paragraph). Comparing elasticities across studies is more difficult than one might expect due to the need to scale appropriately by the extent to which each study's "target" group was actually affected by the relevant minimum wage increase (Sabia, Burkhauser, and Hansen, 2012). This is made difficult by the fact that most minimum wage studies do not directly estimate the extent to which legislated minimum wage changes shift the wage distributions of the populations whose employment they study. As shown in Appendix Figure 1, standard approaches identify targeted individuals with significantly less precision than the approach employed in the current study. While our full-group elasticity is thus larger than those estimated in most past research, it appears more moderately sized when compared on an appropriately scaled basis.

in this group's employment thus implies a $7.4 \times 0.06 = 0.45$ percentage point decline in the employment-to-population ratio. We take this to be a conservative lower bound estimate for the total effect of the $1.68 average rise in the effective minimum wage across the country. We consider twice this amount, or 0.90 percentage points, to be a reasonable upper bound. Our best estimate is that this period's minimum wage increases resulted in a 0.7 percentage point decline in the national employment-to-population ratio for adults aged 16 to 64. This accounts for 14 percent of the total decline in the employment-to-population ratio over this time period.

7 Conclusion

We investigate the effects of recent federal minimum wage increases on the employment and income trajectories of low-skilled workers. While the wage distribution of low-skilled workers shifts as intended, the estimated effects on employment, income, and income growth are negative. We infer from our employment estimates that minimum wage increases reduced the national employment-to-population ratio by 0.7 percentage point between December 2006 and December 2012. As noted above, this accounts for 14 percent of the national decline in the employment-to-population ratio over this period.

We also present evidence of the minimum wage's effects on low-skilled workers' economic mobility. We find that binding minimum wage increases significantly reduced the likelihood that low-skilled workers rose to what we characterize as lower middle class earnings. This curtailment of transitions into lower middle class earnings began to emerge roughly one year following initial declines in low wage employment. Reductions in upward mobility thus appear to follow reductions in access to opportunities for accumulating work experience.

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Tables and Figures

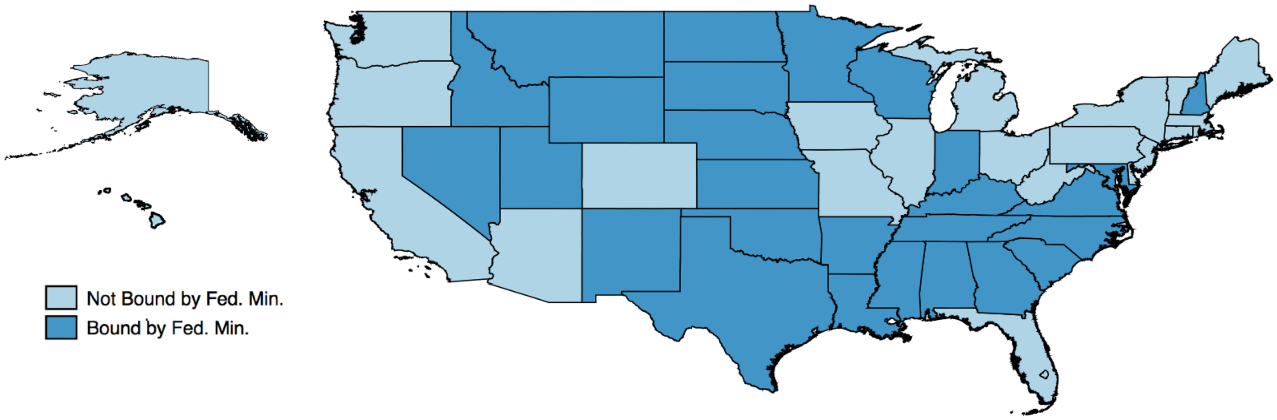

Figure 1: **States Bound by the 2008 and 2009 Federal Minimum Wage Increase:**
The map labels states on the basis of whether we characterize them as bound by the July 2008 and July 2009 increases in the federal minimum wage. We define bound states as states reported by the Bureau of Labor Statistics (BLS) to have had a minimum wage less than $6.55 in January 2008. Such states were at least partially bound by the July 2008 increase in the federal minimum and fully bound by the July 2009 increase from $6.55 to $7.25.

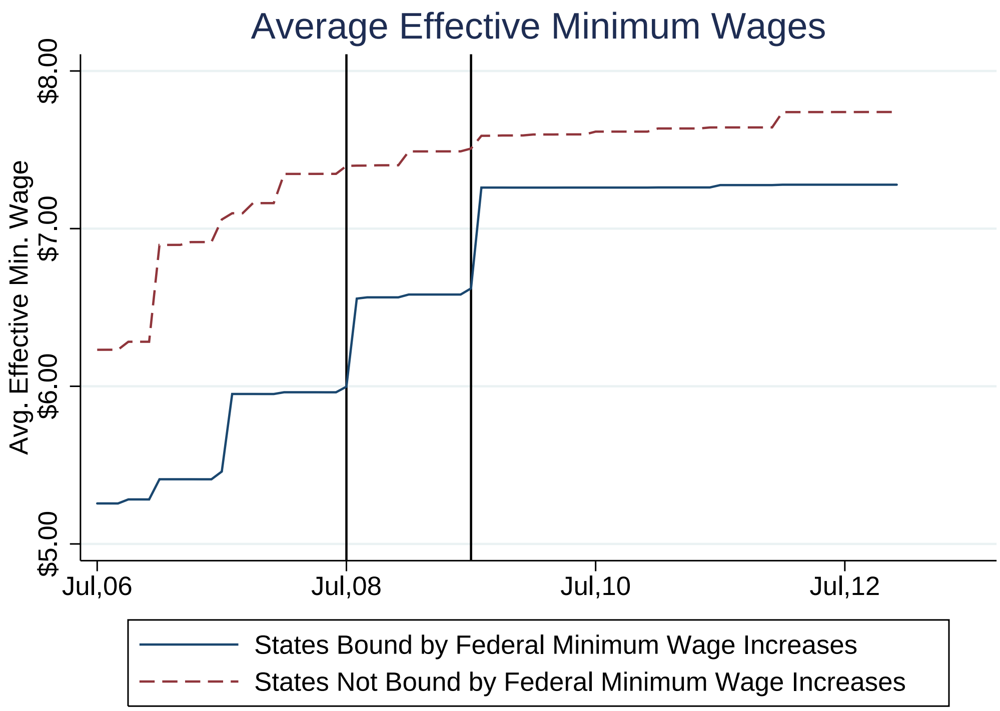

Figure 2: **Evolution of the Average Minimum Wage in Bound and Unbound States:** As in the previous figure, we define bound states as states reported by the Bureau of Labor Statistics (BLS) to have had a minimum wage less than $6.55 in January 2008. Such states were at least partially bound by the July 2008 increase in the federal minimum and fully bound by the July 2009 increase from $6.55 to $7.25. Effective monthly minimum wage data were taken from the detailed replication materials associated with Meer and West (2014). Within each group, the average effective minimum wage is weighted by state population. The first solid vertical line indicates the timing of the July 2009 increase in the federal minimum wage as well as the first month of data available in our samples from the 2008 panel of the Survey of Income and Program Participation. The second solid vertical line indicates the timing of the July 2009 increase in the federal minimum wage.

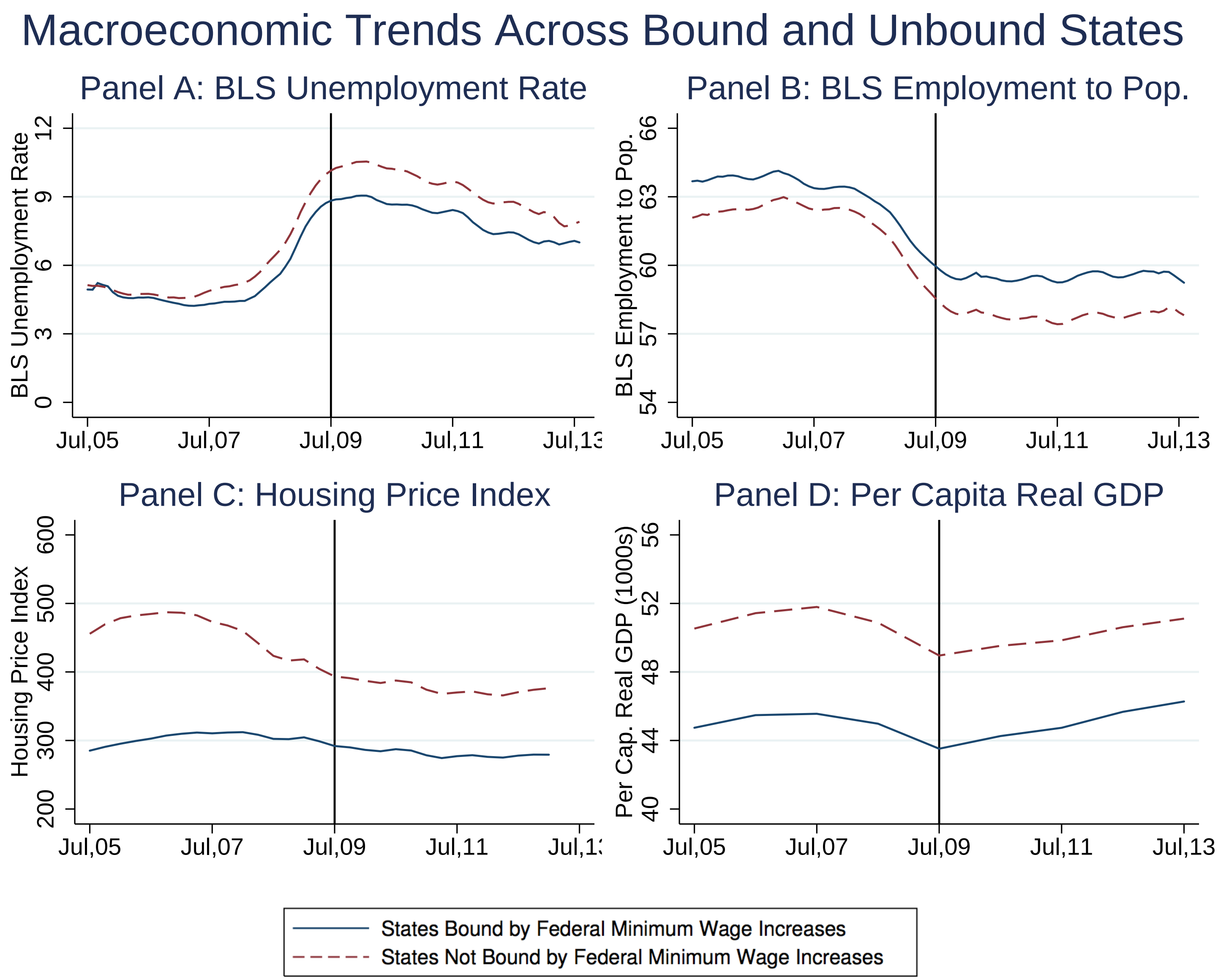

Figure 3: **Macroeconomic Trends in Bound and Unbound States:**
Bound and unbound states are defined as in previous figures. This figure's panels plot the evolution of macroeconomic indicators over the course of the housing bubble and Great Recession. All series are weighted by state population so as to reflect the weighting implicit in our individual-level regression analysis. Panel A plots the average monthly unemployment rate, as reported by the BLS. Panel B plots the average monthly employment to population ratio, also as reported by the BLS. Panel C plots the average of the quarterly Federal Housing Finance Agency's housing price index. Panel D plots the average of annual real per capita GDP, as reported by the Bureau of Economic Analysis (BEA). In each panel, the solid vertical line indicates the timing of the July 2009 increase in the federal minimum wage.

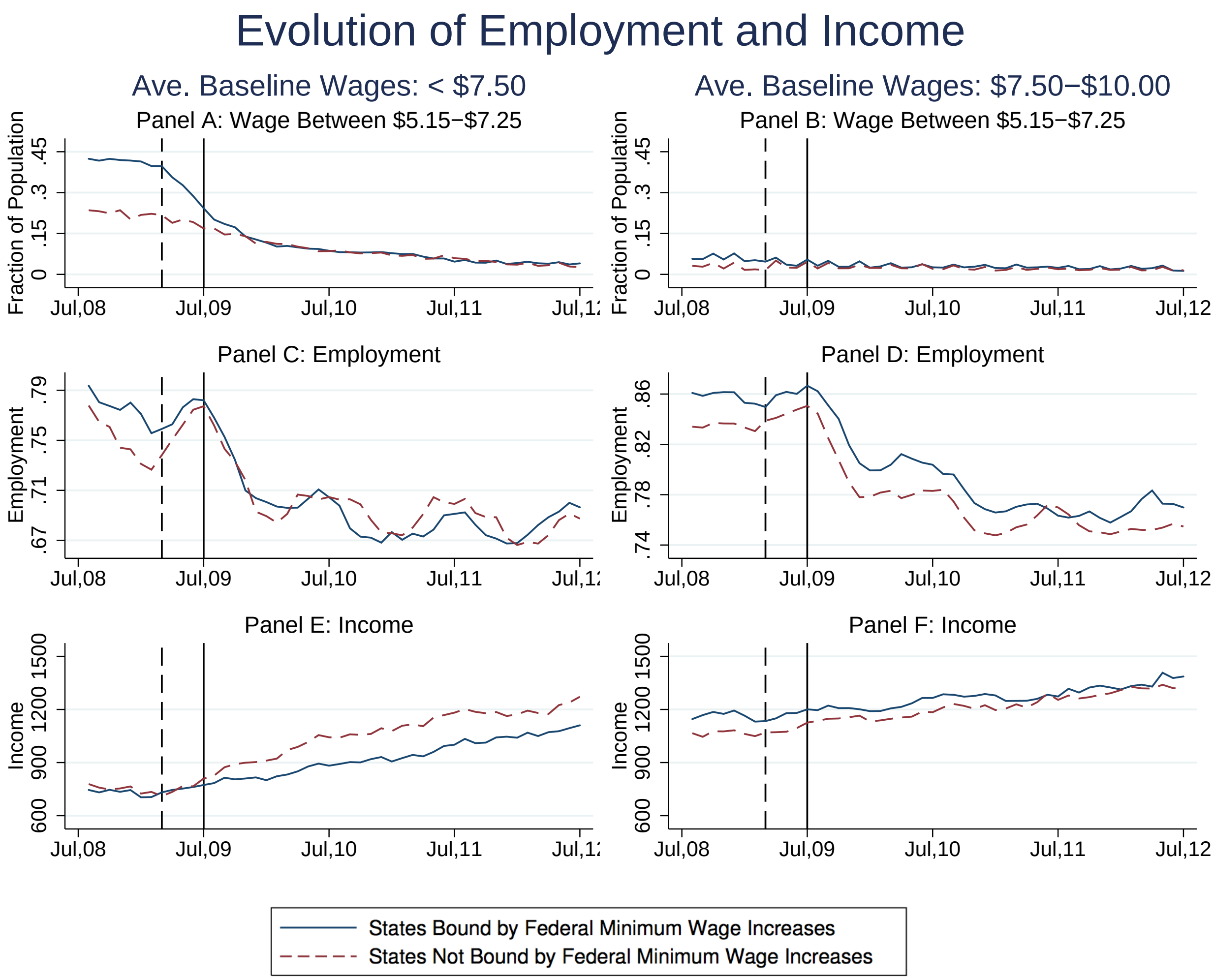

Figure 4: **Employment and Income Trends in Bound and Unbound States:** Bound and unbound states are defined as in previous figures. The figure plots the evolution of three wage, employment, and earnings related outcomes for groups of low-skilled workers. In all cases the series are constructed by the authors using data from the 2008 panel of the Survey of Income and Program Participation (SIPP). In column 1, the samples in each panel consist of individuals whose average baseline wages (meaning wages when employed between August 2008 and July 2009) are less than $7.50. In column 2, the samples in each panel consist of individuals whose average baseline wages are between $7.50 and $10.00. In row 1, the reported outcome is the fraction of observations for which an individual's wage falls between $5.15 and $7.25. In row 2, the reported outcome is the fraction of observations for which an individual is employed. In row 3, the reported outcome is the average earnings of all in-sample individuals. In each panel, the solid vertical line indicates the timing of the July 2009 increase in the federal minimum wage. The dashed vertical line indicates the April 2009 beginning of the transition of wages out of the range between the old and new federal minimum; the date for the latter designation is driven by the data displayed in this figure's Panel A.

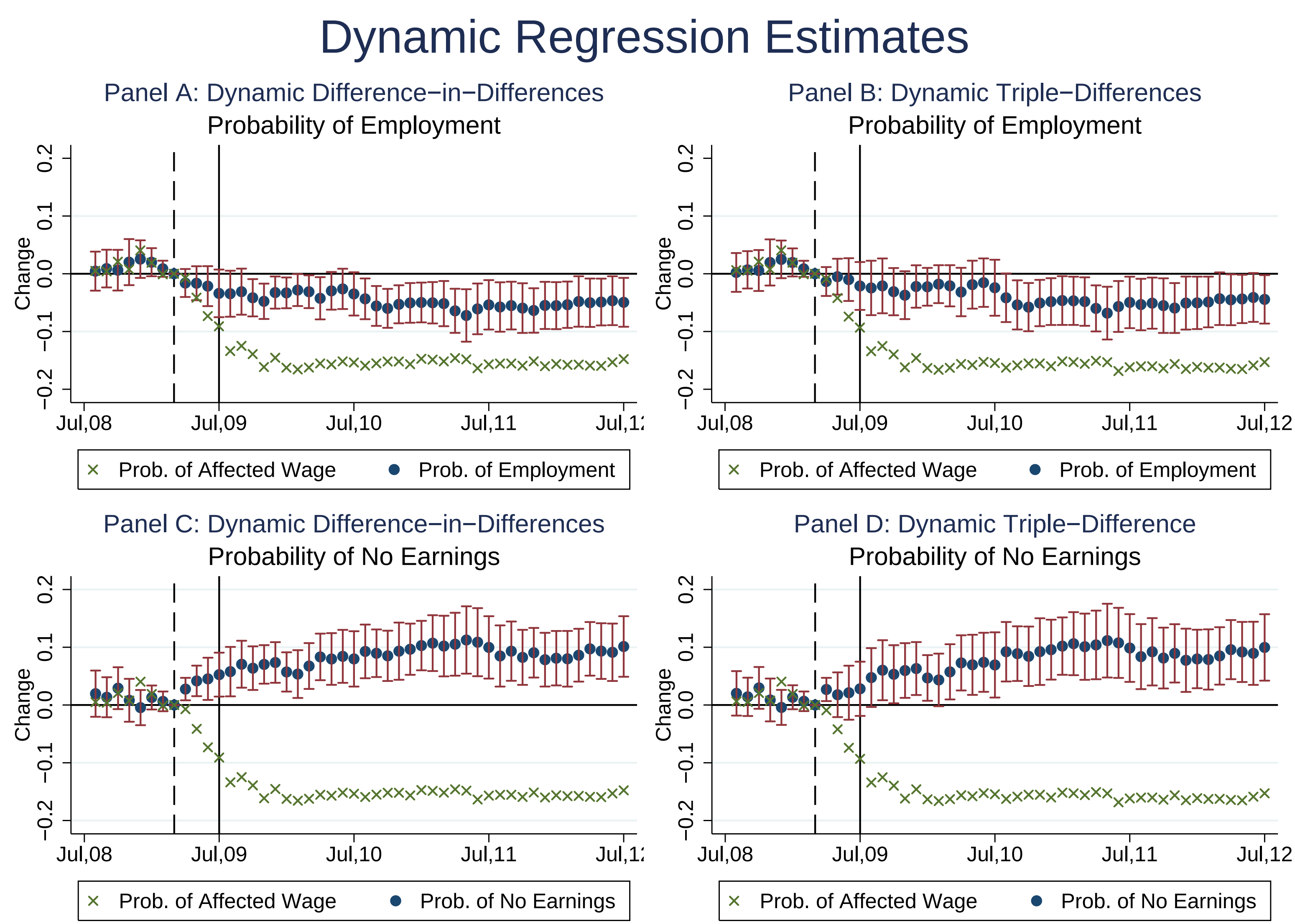

Figure 5: **Dynamic Regression Estimates:** The figure reports fully dynamic estimates of the minimum wage's effects on the wages and employment of low-skilled workers. Each marker is an estimate of a coefficient of the form $\beta_{p(t)}$ from equations (1) and (3), where the relevant $p(t)$ correspond with individual months and period $p = 0$ is April 2009, the month immediately preceding the transition period. Panels A and C present estimates of the difference-in-differences model of equation (1), while panels B and D present estimates of the triple-difference model of equation (3). In each panel, the green X's are estimates of the effect of binding minimum wage changes on the probability of having a wage between $5.15 and $7.25. In panels A and B, the blue dots are estimates of the effect of binding minimum wage increases on the probability of employment (with accompanying 95 percent confidence intervals). In panels C and D, the blue dots are estimates of the effect of binding minimum wage increases on the probability of having no earnings (with accompanying 95 percent confidence intervals).

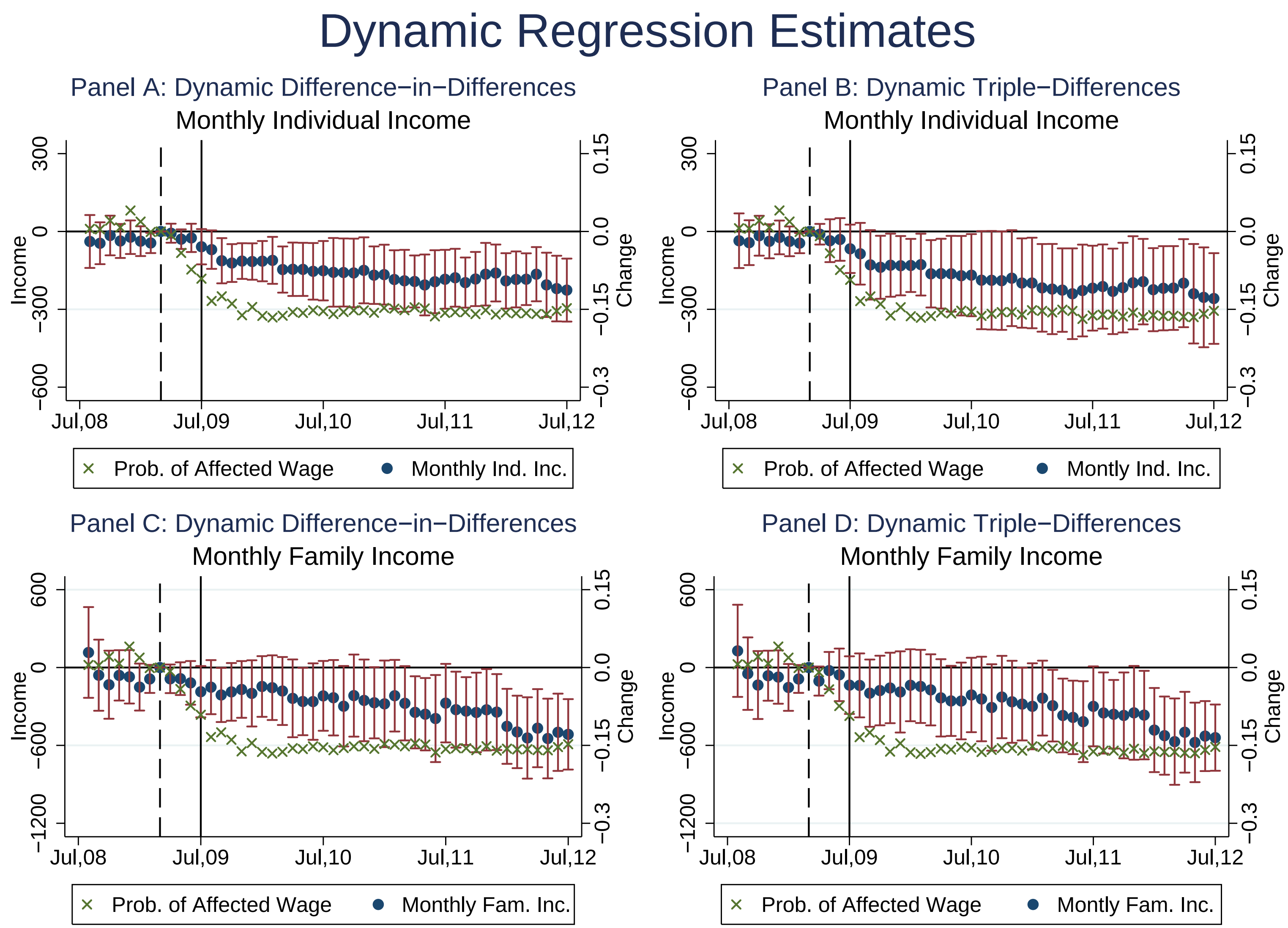

Figure 6: **Dynamic Regression Estimates:** The figure reports fully dynamic estimates of the minimum wage's effects on the wages and income of low-skilled workers. Each marker is an estimate of a coefficient of the form $\beta_{p(t)}$ from equations (1) and (3), where the relevant $p(t)$ correspond with individual months and period $p = 0$ is April 2009, the month immediately preceding the transition period. Panels A and C present estimates of the difference-in-differences model of equation (1), while panels B and D present estimates of the triple-difference model of equation (3). In each panel, the green X's are estimates of the effect of binding minimum wage changes on the probability of having a wage between $5.15 and $7.25. In panels A and B, the blue dots are estimates of the effect of binding minimum wage increases on individual-level monthly income (with accompanying 95 percent confidence intervals). In panels C and D, the blue dots are estimates of the effect of binding minimum wage increases on family-level monthly income (with accompanying 95 percent confidence intervals).

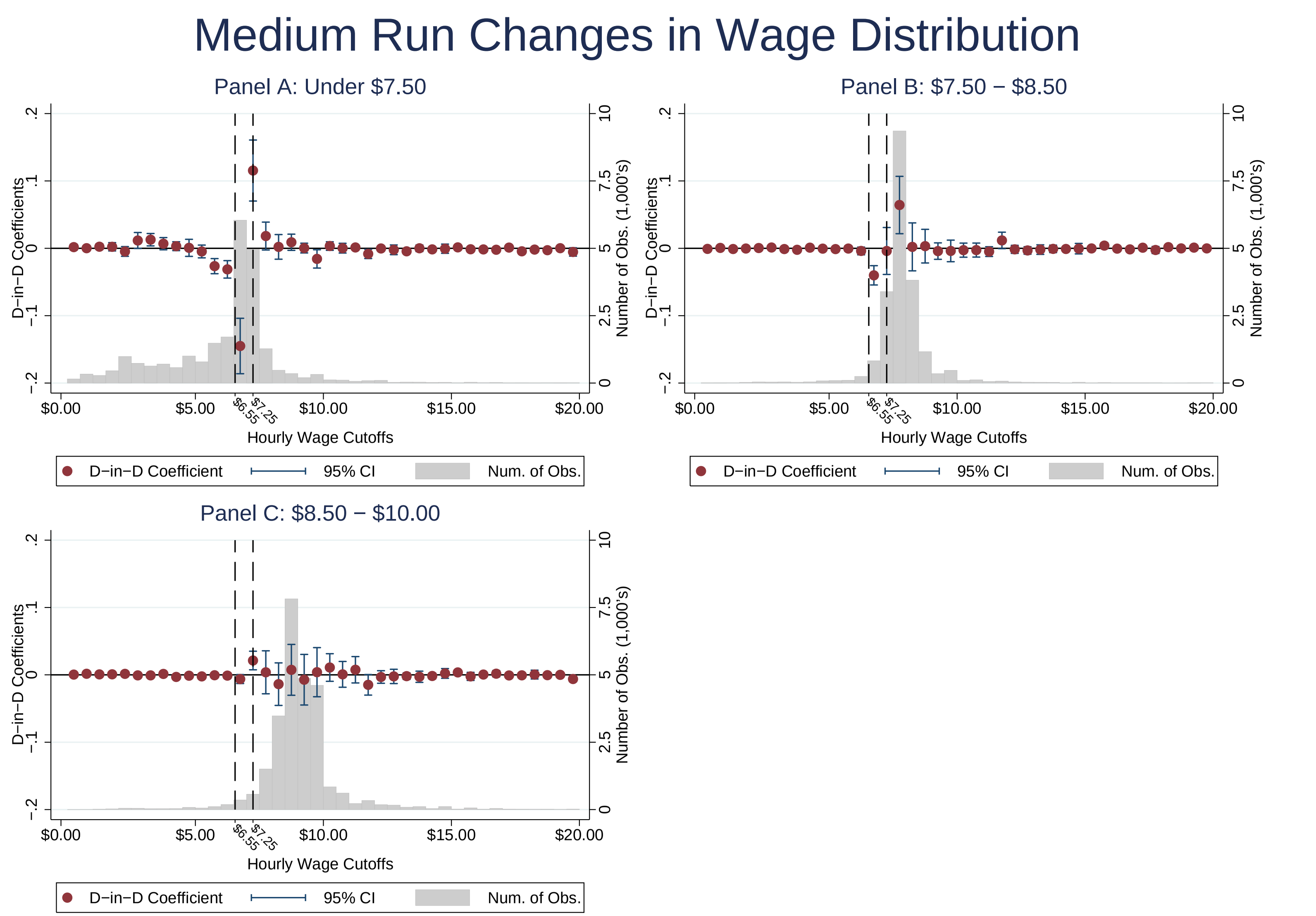

Figure 7: **Estimated Effects of the Minimum Wage on Hourly Wage Distributions:** The figure reports estimates of binding minimum wage increase's medium run effects on the wage distributions of three groups of low-skilled earners. More specifically, each dot is an estimate of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. The dependent variables in each specification take the form $Y^j_{i,s,t} = 1\{W^{j-1} < \text{Hourly Wage}_{i,s,t} < W^j\}$. These $Y_{i,s,t}$ are indicators equal to 1 if an individual's hourly wage is in the band between W^{j-1} and W^j, where each band is a 50 cent interval. The results can thus be described as estimates of the minimum wage's effect on the wage distribution's probability mass function. In Panel A, the sample consists of individuals whose average baseline wages (meaning wages when employed between August 2008 and July 2009) are less than $7.50. In Panel B, the sample consists of individuals whose average baseline wages are between $7.50 and $8.50. In Panel C, the sample consists of individuals whose average baseline wages are between $8.50 and $10.00. In the background of each panel is a histogram characterizing the frequency distribution of hourly wages during the sample's baseline period.

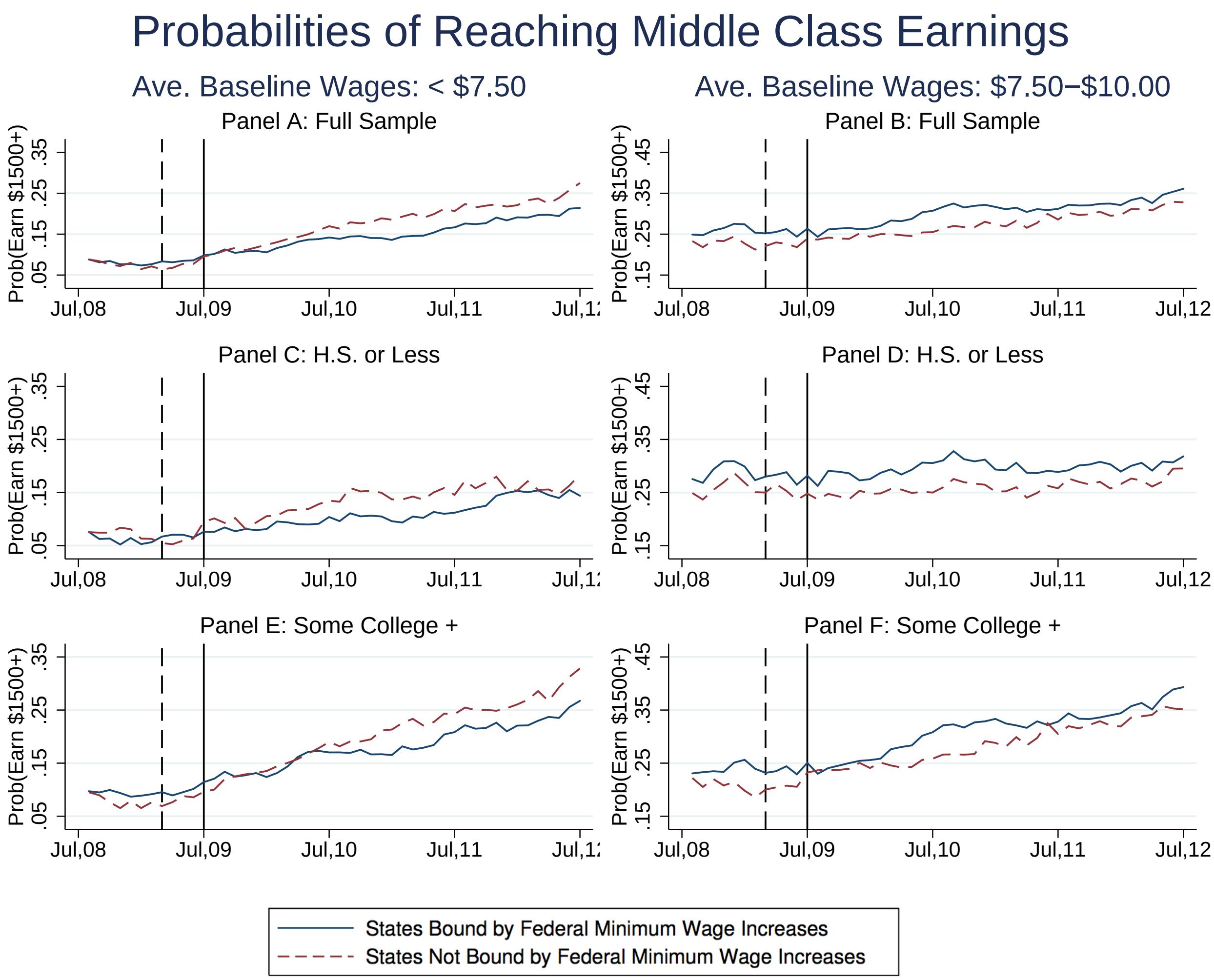

Figure 8: **Probabilities of Reaching Middle Class Earnings:**
Bound and unbound states are defined as in previous figures. In all panels, the figure plots the evolution the fraction of all in-sample individuals with earnings greater than $1500, which is equivalent to full time work at a wage of $8.82. The series are constructed by the authors using data from the 2008 panel of the Survey of Income and Program Participation (SIPP). In column 1, the samples in each panel consist of individuals whose average baseline wages (meaning wages when employed between August 2008 and July 2009) are less than $7.50. In column 2, the samples in each panel consist of individuals whose average baseline wages are between $7.50 and $10.00. Row 1 presents tabulations of the outcome of interest for the full sample of individuals as defined above. In row 2 the sample is limited to individuals with no college education, while in row 3 the sample is limited to individuals with at least some college education. In each panel, the solid vertical line indicates the timing of the July 2009 increase in the federal minimum wage. The dashed vertical line indicates the April 2009 beginning of the transition of wages out of the range between the old and new federal minimum; the date for the latter designation is driven by the data displayed in figure 4's Panel A.

Table 1: **Baseline Summary Statistics by Treatment Status and Average Baseline Wages**

	(1)	(2)	(3)	(4)	(5)	(6)
Ave. Baseline Wage	Wage < \$7.50		\$7.50-\$8.49		\$8.50-\$9.99	
Treatment Status	Bound	Not Bound	Bound	Not Bound	Bound	Not Bound
Earn \$5.15-\$7.25	0.374	0.218	0.0806	0.0405	0.0330	0.0240
	(0.484)	(0.413)	(0.272)	(0.197)	(0.179)	(0.153)
Employed	0.777	0.751	0.825	0.810	0.889	0.865
	(0.416)	(0.433)	(0.380)	(0.393)	(0.314)	(0.341)
Unpaid Work	0.168	0.211	0.101	0.116	0.0842	0.0922
	(0.374)	(0.408)	(0.302)	(0.320)	(0.278)	(0.289)
No Earnings	0.391	0.461	0.276	0.306	0.195	0.227
	(0.488)	(0.498)	(0.447)	(0.461)	(0.396)	(0.419)
Num hours worked/week	24.36	23.57	27.13	24.21	31.83	29.60
	(18.29)	(19.13)	(17.43)	(17.07)	(15.98)	(16.52)
Income	770.1	748.2	993.5	886.7	1319.3	1269.9
	(1409.0)	(1115.4)	(949.8)	(1037.0)	(1012.4)	(1179.9)
Below FPL	0.474	0.439	0.393	0.411	0.307	0.344
	(0.499)	(0.496)	(0.489)	(0.492)	(0.461)	(0.475)
Age	35.79	37.05	36.92	34.89	40.58	38.02
	(13.92)	(14.51)	(13.56)	(13.65)	(13.10)	(13.29)
Num. of Children	1.345	1.265	1.263	1.333	1.110	1.085
	(1.408)	(1.423)	(1.340)	(1.350)	(1.352)	(1.311)
More than H.S. Deg.	0.580	0.635	0.584	0.591	0.585	0.595
	(0.494)	(0.481)	(0.493)	(0.492)	(0.493)	(0.491)
Num. of Individuals	1876	1546	1038	1335	1249	1589
Observations	22493	18566	12561	16061	15211	19273

Sources: Baseline summary statistics were calculated by the authors using data from the 2008 panel of the Survey of Income and Program Participation. The baseline corresponds with the period extending from August 2008 through July 2009. Columns 1, 3, and 5 report summary statistics for individuals in states we designate as bound by increases in the federal minimum, as described in the note to Figure 1. Column 2, 4, and 6 report summary statistics for individuals in the remaining states, which we designate as unbound. In Columns 1 and 2, the sample consists of individuals whose average baseline wages (meaning wages when employed between August 2008 and July 2009) are less than \$7.50. In Columns 3 and 4, the sample consists of individuals whose average baseline wages are between \$7.50 and \$8.50. In Columns 5 and 6, the sample consists of individuals whose average baseline wages are between \$8.50 and \$10.00.

Table 2: **Effects on Wage Distributions and Employment by Average Baseline Wages**

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable	Wage between $5.15 and $7.25				Employed	
Bound x Post 1	-0.161***	-0.036***	-0.008*	-0.043*	0.020	-0.011
	(0.022)	(0.010)	(0.004)	(0.018)	(0.022)	(0.011)
Bound x Post 2	-0.164***	-0.042***	-0.007	-0.063**	-0.011	-0.008
	(0.024)	(0.008)	(0.005)	(0.018)	(0.023)	(0.013)
Housing Price Index	-0.623	-0.098	0.109	1.022***	0.416	-0.035
	(0.424)	(0.167)	(0.088)	(0.234)	(0.504)	(0.390)
N	146,933	102,154	122,594	146,933	102,154	122,594
Mean of Dep. Var.	0.304	0.057	0.027	0.766	0.813	0.877
Estimation Framework	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D
Sample	Under $7.50	$7.50-$8.49	$8.50-$9.99	Under $7.50	$7.50-$8.49	$8.50-$9.99

Note: +, *, **, and *** indicate statistical significance at the 0.10, 0.05, 0.01, and 0.001 levels respectively. The table reports estimates of the minimum wage's short and medium run effects on the relevant dependent variables, which are named in the heading of each column. More specifically, the estimates in row 1 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning in August 2009 and extending through July 2010. The estimates in row 2 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. In columns 1-3, the dependent variable is an indicator for whether an individual's hourly wage is $5.15 and $7.25. In columns 4-6, the dependent variable is an indicator for whether an individual is employed. In Columns 1 and 4, the sample consists of individuals whose average baseline wages (meaning wages when employed between August 2008 and July 2009) are less than $7.50. In Columns 2 and 5, the sample consists of individuals whose average baseline wages are between $7.50 and $8.50. In Columns 3 and 6, the sample consists of individuals whose average baseline wages are between $8.50 and $10.00. Standard errors are clustered at the state level.

Table 3: **Robustness of Estimated Effects on Employment**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent Variable				Employed			
Panel A:				Difference-in-Differences Specifications			
Bound x Post 1	-0.043*	-0.029	-0.051*	-0.041*	-0.042*	-0.049*	-0.042*
	(0.018)	(0.021)	(0.020)	(0.018)	(0.020)	(0.020)	(0.019)
Bound x Post 2	-0.063**	-0.042*	-0.072**	-0.061***	-0.061**	-0.067**	-0.059**
	(0.018)	(0.019)	(0.025)	(0.016)	(0.017)	(0.021)	(0.018)
N	146,933	146,933	146,933	146,933	121,365	128,728	143,973
Mean of Dep. Var.	0.766	0.766	0.766	0.766	0.769	0.765	0.764
Estimation Framework	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D
Panel B:				Triple Difference Specifications			
Bound x Post 1 x Target	-0.032	-0.018	-0.032	-0.033	-0.027	-0.040	-0.033
	(0.022)	(0.025)	(0.022)	(0.023)	(0.024)	(0.024)	(0.023)
Bound x Post 2 x Target	-0.059**	-0.038+	-0.058**	-0.064**	-0.051*	-0.059**	-0.053**
	(0.019)	(0.021)	(0.019)	(0.020)	(0.023)	(0.021)	(0.019)
N	269,527	269,527	269,527	269,527	223,148	238,727	263,782
Mean of Dep. Var.	0.817	0.817	0.817	0.817	0.820	0.817	0.816
Estimation Framework	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D
Macro Covariates	Yes	No	Yes	Yes	Yes	Yes	Yes
State Trends	No	No	Yes	No	No	No	No
Trends In Demographics	No	No	No	Yes	No	No	No
Minimum Sample Inclusion	3 yrs	3 yrs	3 yrs	3 yrs	3.5 yrs	3 yrs	3 yrs
Excluded States	None	None	None	None	None	N.B.<$7.00	B.>$6.55

Note: +, *, **, and *** indicate statistical significance at the 0.10, 0.05, 0.01, and 0.001 levels respectively. Panel A reports estimates of the minimum wage's short and medium run effects on an indicator for whether or not an individual is employed. More specifically, the estimates in row 1 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning in August 2009 and extending through July 2010. The estimates in row 2 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. Panel B reports analogous estimates of $\beta_{p(t)}$ from equation (3), namely our triple-difference specification. In Panel A the sample consists exclusively of individuals with average baseline wages less than $7.50. In Panel B the sample is augmented to include individuals whose average baseline wages are between $8.50 and $10.00 as a within-state control group. The columns explore our baseline results' (column 1) robustness to a variety of specification changes, which are further described in the main text and within the table itself. Standard errors are clustered at the state level.

Table 4: **Effects on Employment Status, Income, and Poverty Status**

Dependent Variable	(1) Employed	(2) Unpaid Work	(3) No Earnings	(4) Ind. Income	(5) Fam. Income	(6) Below FPL
Panel A:	Difference-in-Differences Specifications					
Bound x Post 1	-0.043*	0.015	0.057**	-96.996*	-131.955	0.008
	(0.018)	(0.010)	(0.019)	(37.581)	(85.091)	(0.015)
Bound x Post 2	-0.063**	0.019*	0.081***	-152.976**	-285.200*	0.020
	(0.018)	(0.009)	(0.022)	(44.511)	(119.009)	(0.022)
N	146,933	146,933	146,933	146,933	146,933	146,933
Mean of Dep. Var.	0.766	0.187	0.421	747.633	4,194.831	0.458
Panel B:	Triple Difference Specifications					
Bound x Post 1 x Target	-0.032	0.015	0.047*	-112.728+	-120.412	-0.000
	(0.022)	(0.014)	(0.023)	(56.538)	(105.210)	(0.023)
Bound x Post 2 x Target	-0.059**	0.021+	0.080**	-184.768*	-303.234*	0.027
	(0.019)	(0.013)	(0.023)	(74.359)	(127.642)	(0.035)
N	269,527	269,527	269,527	269,527	269,527	269,527
Mean of Dep. Var.	0.817	0.141	0.324	995.022	4,252.687	0.399

Note: +, *, **, and *** indicate statistical significance at the 0.10, 0.05, 0.01, and 0.001 levels respectively. Panel A reports estimates of the minimum wage's short and medium run effects on the relevant dependent variables, which are named in the heading of each column. More specifically, the estimates in row 1 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning in August 2009 and extending through July 2010. The estimates in row 2 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. Panel B reports analogous estimates of $\beta_{p(t)}$ from equation (3), namely our triple-difference specification. In Panel A the sample consists exclusively of individuals with average baseline wages less than $7.50. In Panel B the sample is augmented to include individuals whose average baseline wages are between $8.50 and $10.00 as a within-state control group. Standard errors are clustered at the state level.

Table 5: **Heterogeneity of Effects on Employment and Unpaid Work by Education Group**

Dependent Variable	(1) Employed	(2) Employed	(3) Unpaid Work	(4) Unpaid Work	(5) No Earnings	(6) No Earnings
Panel A:	Difference-in-Differences Specifications					
Bound x Post 1	-0.046	-0.041*	0.001	0.022+	0.048+	0.063**
	(0.030)	(0.020)	(0.015)	(0.013)	(0.024)	(0.021)
Bound x Post 2	-0.069**	-0.057*	-0.005	0.033**	0.065*	0.090**
	(0.025)	(0.024)	(0.015)	(0.012)	(0.027)	(0.028)
N	60,282	86,651	60,282	86,651	60,282	86,651
Mean of Dep. Var.	0.772	0.762	0.180	0.191	0.408	0.429
Panel B:	Triple Difference Specifications					
Bound x Post 1 x Target	-0.060+	-0.020	-0.007	0.026	0.053+	0.045
	(0.031)	(0.028)	(0.019)	(0.016)	(0.027)	(0.030)
Bound x Post 2 x Target	-0.102***	-0.037	-0.010	0.040**	0.092**	0.077*
	(0.028)	(0.025)	(0.019)	(0.015)	(0.029)	(0.030)
N	111,292	158,235	111,292	158,235	111,292	158,235
Mean of Dep. Var.	0.824	0.811	0.132	0.147	0.308	0.336
Sample	H.S. or Less	Some Coll.+	H.S. or Less	Some Coll.+	H.S. or Less	Some Coll.+

Note: +, *, **, and *** indicate statistical significance at the 0.10, 0.05, 0.01, and 0.001 levels respectively. Panel A reports estimates of the minimum wage's short and medium run effects on the relevant dependent variables, which are named in the heading of each column. More specifically, the estimates in row 1 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning in August 2009 and extending through July 2010. The estimates in row 2 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. Panel B reports analogous estimates of $\beta_{p(t)}$ from equation (3), namely our triple-difference specification. In Panel A the sample consists exclusively of individuals with average baseline wages less than $7.50. In Panel B the sample is augmented to include individuals whose average baseline wages are between $8.50 and $10.00 as a within-state control group. In columns 1, 3, and 5 the samples consist of individuals with a high school degree or less, while in columns 2, 4, and 6 the samples consist of those with at least some college education. Additional details are provided in the main text and within the table itself. Standard errors are clustered at the state level.

Table 6: **Effects on the Probability of Reaching Lower Middle or Middle Class Earnings Levels**

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable		Earn $1,500+			Earn $3,000+	
Panel A:	Difference-in-Differences Specifications					
Bound x Post 1	-0.016	-0.022	-0.012	-0.004	-0.006	-0.002
	(0.011)	(0.015)	(0.013)	(0.007)	(0.008)	(0.009)
Bound x Post 2	-0.049***	-0.049**	-0.044*	-0.018+	-0.010	-0.021
	(0.013)	(0.015)	(0.017)	(0.009)	(0.009)	(0.013)
N	146,933	60,282	86,651	146,933	60,282	86,651
Mean of Dep. Var.	0.206	0.153	0.237	0.068	0.032	0.090
Panel B:	Triple Difference Specifications					
Bound x Post 1 x Target	-0.002	-0.035	0.019	-0.013	-0.021+	-0.008
	(0.018)	(0.027)	(0.024)	(0.011)	(0.012)	(0.014)
Bound x Post 2 x Target	-0.054*	-0.082**	-0.035	-0.017	-0.010	-0.022
	(0.024)	(0.030)	(0.030)	(0.015)	(0.015)	(0.019)
N	269,527	111,292	158,235	269,527	111,292	158,235
Mean of Dep. Var.	0.206	0.153	0.237	0.068	0.032	0.090
Sample	Full	H.S. or Less	Some Coll.+	Full	H.S. or Less	Some Coll.+

Note: +, *, **, and *** indicate statistical significance at the 0.10, 0.05, 0.01, and 0.001 levels respectively. Panel A reports estimates of the minimum wage's short and medium run effects on the relevant dependent variables, which are named in the heading of each column. More specifically, the estimates in row 1 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning in August 2009 and extending through July 2010. The estimates in row 2 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. Panel B reports analogous estimates of $\beta_{p(t)}$ from equation (3), namely our triple-difference specification. In Panel A the sample consists exclusively of individuals with average baseline wages less than $7.50. In Panel B the sample is augmented to include individuals whose average baseline wages are between $8.50 and $10.00 as a within-state control group. Standard errors are clustered at the state level.

Appendix (Intended For Online Publication Only)

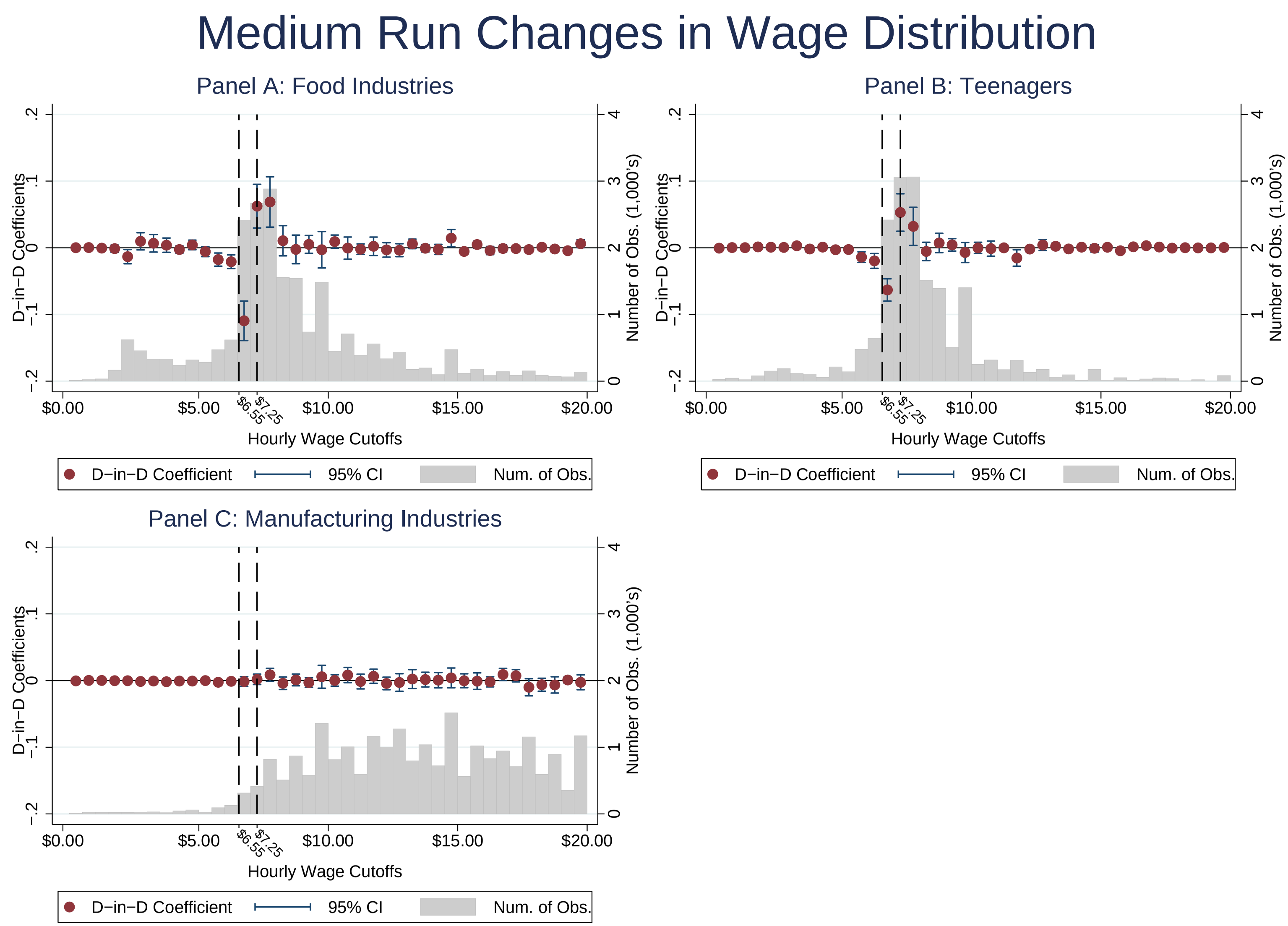

Figure A.1: **Estimated Effects of the Minimum Wage on Hourly Wage Distributions:** The figure reports estimates of binding minimum wage increase's medium run effects on the wage distributions of three groups of low-skilled earners. More specifically, each dot is an estimate of the coefficient $\beta_p(t)$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. The dependent variables in each specification take the form $Y^j_{i,s,t} = 1\{W^{j-1} < \text{Hourly Wage}_{i,s,t} < W^j\}$. These $Y_{i,s,t}$ are indicators equal to 1 if an individual's hourly wage is in the band between W^{j-1} and W^j, where each band is a 50 cent interval. The results can thus be described as estimates of the minimum wage's effect on the wage distribution's probability mass function. In Panel A, the sample consists of individuals whose modal industry of employment over the baseline was food service. In Panel B, the sample consists of individuals who were teenagers at the beginning of the sample period. In Panel C, the sample consists of individuals whose modal industry of employment over the baseline was manufacturing. In the background of each panel is a histogram characterizing the frequency distribution of hourly wages during the sample's baseline period.

Table A.1: **Cross-Sample Comparison of Effects on Wages and Employment**

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable	Wage between \$5.15 and \$7.25			Employed		
Bound x Post 1	-0.161***	-0.082***	-0.005	-0.043*	-0.020	0.003
	(0.022)	(0.014)	(0.004)	(0.018)	(0.013)	(0.010)
Bound x Post 2	-0.164***	-0.090***	-0.006	-0.063**	-0.039**	0.008
	(0.024)	(0.016)	(0.004)	(0.018)	(0.013)	(0.013)
N	146,933	275,042	166,555	146,933	275,042	166,555
Mean of Dep. Var.	0.304	0.104	0.015	0.766	0.557	0.934
Estimation Framework	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D
Sample	Under \$7.50	Food Service and Teens	Manufacturing	Under \$7.50	Food Service and Teens	Manufacturing

Note: +, *, **, and *** indicate statistical significance at the 0.10, 0.05, 0.01, and 0.001 levels respectively. The table reports estimates of the minimum wage's short and medium run effects on the relevant dependent variables, which are named in the heading of each column. More specifically, the estimates in row 1 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning in August 2009 and extending through July 2010. The estimates in row 2 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. In columns 1-3, the dependent variable is an indicator for whether an individual's hourly wage is \$5.15 and \$7.25. In columns 4-6, the dependent variable is an indicator for whether an individual employed. In Columns 1 and 4, the sample consists of individuals whose average baseline wages (meaning wages when employed between August 2008 and July 2009) are less than \$7.50. In Columns 2 and 5, the sample consists of individuals who were teenagers or whose modal industry was food service when employed at baseline. In Columns 3 and 6, the sample consists of individuals whose modal industry was manufacturing when employed at baseline. Standard errors are clustered at the state level.

Table A.2: **Cross-Sample Comparison of Effects on Wages and the Probability of Working Without Pay**

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable	Wage between $5.15 and $7.25				Unpaid Work	
Bound x Post 1	-0.161***	-0.082***	-0.005	0.015	0.005	0.001
	(0.022)	(0.014)	(0.004)	(0.010)	(0.005)	(0.004)
Bound x Post 2	-0.164***	-0.090***	-0.006	0.019*	0.007	-0.003
	(0.024)	(0.016)	(0.004)	(0.009)	(0.006)	(0.005)
N	146,933	275,042	166,555	146,933	275,042	166,555
Mean of Dep. Var.	0.304	0.104	0.015	0.187	0.082	0.042
Estimation Framework	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D
Sample	Under $7.50	Food Service and Teens	Manufacturing	Under $7.50	Food Service and Teens	Manufacturing

Note: +, *, **, and *** indicate statistical significance at the 0.10, 0.05, 0.01, and 0.001 levels respectively. The table reports estimates of the minimum wage's short and medium run effects on the relevant dependent variables, which are named in the heading of each column. More specifically, the estimates in row 1 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning in August 2009 and extending through July 2010. The estimates in row 2 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. In columns 1-3, the dependent variable is an indicator for whether an individual's hourly wage is $5.15 and $7.25. In columns 4-6, the dependent variable is an indicator for whether an individual worked with no pay. In Columns 1 and 4, the sample consists of individuals whose average baseline wages (meaning wages when employed between August 2008 and July 2009) are less than $7.50. In Columns 2 and 5, the sample consists of individuals who were teenagers or whose modal industry was food service when employed at baseline. In Columns 3 and 6, the sample consists of individuals whose modal industry was manufacturing when employed at baseline. Standard errors are clustered at the state level.

Table A.3: **Cross-Sample Comparison of Effects on Wages and the Probability of Having Positive Earnings**

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable	Wage between \$5.15 and \$7.25			No Earnings		
Bound x Post 1	-0.161***	-0.082***	-0.005	0.057**	0.025*	-0.002
	(0.022)	(0.014)	(0.004)	(0.019)	(0.012)	(0.010)
Bound x Post 2	-0.164***	-0.090***	-0.006	0.081***	0.046***	-0.010
	(0.024)	(0.016)	(0.004)	(0.022)	(0.012)	(0.012)
N	146,933	275,042	166,555	146,933	275,042	166,555
Mean of Dep. Var.	0.304	0.104	0.015	0.421	0.525	0.108
Estimation Framework	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D
Sample	Under \$7.50	Food Service and Teens	Manufacturing	Under \$7.50	Food Service and Teens	Manufacturing

Note: +, *, **, and *** indicate statistical significance at the 0.10, 0.05, 0.01, and 0.001 levels respectively. The table reports estimates of the minimum wage's short and medium run effects on the relevant dependent variables, which are named in the heading of each column. More specifically, the estimates in row 1 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning in August 2009 and extending through July 2010. The estimates in row 2 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. In columns 1-3, the dependent variable is an indicator for whether an individual's hourly wage is \$5.15 and \$7.25. In columns 4-6, the dependent variable is an indicator for whether an individual has no earnings, characterized as either being unemployed or employed without pay. In Columns 1 and 4, the sample consists of individuals whose average baseline wages (meaning wages when employed between August 2008 and July 2009) are less than \$7.50. In Columns 2 and 5, the sample consists of individuals who were teenagers or whose modal industry was food service when employed at baseline. In Columns 3 and 6, the sample consists of individuals whose modal industry was manufacturing when employed at baseline. Standard errors are clustered at the state level.

Table A.4: **Effects on Food Service Employment**

	(1)	(2)	(3)
		Food Service	
Bound x Post 1	-0.000	-0.023*	0.002+
	(0.001)	(0.011)	(0.001)
Bound x Post 2	-0.001	-0.033*	0.003*
	(0.002)	(0.013)	(0.001)
N	1,966,935	146,933	1,820,002
Mean of Dep. Var.	0.047	0.217	0.033
Estimation Framework	D-in-D	D-in-D	D-in-D
Weighted	No	No	No
Individual Fixed Effects	Yes	Yes	Yes
Sample	Full Sample	Under $7.50	All Other

Note: +, *, **, and *** indicate statistical significance at the 0.10, 0.05, 0.01, and 0.001 levels respectively. The table reports estimates of the minimum wage's short and medium run effects on the relevant dependent variables, which are named in the heading of each column. More specifically, the estimates in row 1 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning in August 2009 and extending through July 2010. The estimates in row 2 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. In column 1, the sample contains all individuals aged 16 to 64 for whom the relevant earnings and employment data were available for at least 36 months between August 2008 and July 2012. In column 2, the sample consists of individuals from the sample in column 1 whose average baseline wages (meaning wages when employed between August 2008 and July 2009) were less than $7.50. The sample in column 3 is the complement of the sample in column 2. Standard errors are clustered at the state level.

Table A.5: **First Stage and Employment Effects for All Adults**

	(1)	(2)	(3)	(4)	(5)
Panel A:	Dependent Variable: Affected Wage				
Bound x Post 1	0.001	-0.161***	-0.036***	-0.008*	-0.000
	(0.001)	(0.022)	(0.010)	(0.004)	(0.001)
Bound x Post 2	0.001	-0.164***	-0.042***	-0.007	-0.000
	(0.001)	(0.024)	(0.008)	(0.005)	(0.001)
N	520,074	146,933	102,154	122,594	1,075,180
Mean of Dep. Var.	0.000	0.304	0.057	0.027	0.004
Estimation Framework	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D
Panel B:	Dependent Variable: Employment				
Bound x Post 1	-0.004	-0.043*	0.020	-0.011	0.005
	(0.004)	(0.018)	(0.022)	(0.011)	(0.005)
Bound x Post 2	-0.013	-0.063**	-0.011	-0.008	0.000
	(0.009)	(0.018)	(0.023)	(0.013)	(0.007)
N	520,074	146,933	102,154	122,594	1,075,180
Mean of Dep. Var.	0.000	0.766	0.813	0.877	0.946
Estimation Framework	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D
Sample	Unemp. at Base	Under $7.50	$7.50-$8.49	$8.50-$9.99	Over $10.00

Note: +, *, **, and *** indicate statistical significance at the 0.10, 0.05, 0.01, and 0.001 levels respectively. In panel A the dependent variable in this panel is an indicator equal to 1 if an individual reports a wage between $5.15 and $7.25 in the relevant month. In panel B the dependent variable is an indicator equal to 1 if an individual is employed. In both panels, the estimates in row 1 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning in July 2009 and extending through July 2010. The samples used across columns 1 through 5 fully partition the set of all individuals aged 16 to 64 for whom the relevant earnings and employment data were available for at least 36 months between August 2008 and July 2012. The estimates in row 2 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. Standard errors are clustered at the state level.

Table A.6: **Further Robustness of the Estimate Employment Effects**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent Variable				Employed			
Bound x Post 1	-0.043*	-0.029	-0.031	-0.032	-0.042*	-0.050*	-0.049*
	(0.018)	(0.021)	(0.020)	(0.023)	(0.019)	(0.020)	(0.022)
Bound x Post 2	-0.063**	-0.042*	-0.048*	-0.049*	-0.061**	-0.071**	-0.070*
	(0.018)	(0.019)	(0.018)	(0.022)	(0.019)	(0.025)	(0.026)
Housing Price Index	1.022***				0.965**	1.787*	1.827**
	(0.234)				(0.298)	(0.712)	(0.610)
State Employment Rate			1.242+	1.207	0.794		1.107
			(0.740)	(0.745)	(0.734)		(0.790)
State Inc. Per Cap. (1000s)			0.005	0.005	0.002		0.004
			(0.004)	(0.004)	(0.003)		(0.004)
Stimulus Per Cap. (1000s)				0.015	0.016		0.051
				(0.048)	(0.049)		(0.046)
Predicted State Income				0.088	-0.311		-1.052
				(0.474)	(0.478)		(1.250)
Predicted State Employment				-0.201	-0.442		-0.850
				(0.696)	(0.683)		(0.826)
N	146,933	146,933	146,933	146,933	146,933	145,749	145,749
Mean of Dep. Var.	0.766	0.766	0.766	0.766	0.766	0.769	0.769
Estimation Framework	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D
State Trends	No	No	No	No	No	Yes	Yes
Trends In Demographics	No	No	No	No	No	Yes	Yes
Trends In Baseline Ind.	No	No	No	No	No	Yes	Yes

Note: +, *, **, and *** indicate statistical significance at the 0.10, 0.05, 0.01, and 0.001 levels respectively. Each column reports estimates of the minimum wage's short and medium run effects on employment. More specifically, the estimates in row 1 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning in August 2009 and extending through July 2010. The estimates in row 2 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. The stimulus spending variable was taken from Chodorow-Reich, Feiveson, Liscow, and Woolston (2012). "Predict State Income" is a projection of state-specific changes in aggregate output that are predictable on the basis of each state's historical relationship with the national business cycle. "Predicted State Employment" is a projected change in employment based on each states' baseline industrial composition and subsequent industry-specific employment growth at the national level (Bartik, 1991; Blanchard and Katz, 1992). Standard errors are clustered at the state level.

Table A.7: **Robustness of Estimated Effects on Working Without Pay**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent Variable				Unpaid Work			
Panel A:				Difference-in-Differences Specifications			
Bound x Post 1	0.015	0.019+	0.018	0.013	0.017+	0.017	0.014
	(0.010)	(0.010)	(0.012)	(0.010)	(0.009)	(0.012)	(0.010)
Bound x Post 2	0.019*	0.025*	0.023	0.013	0.020*	0.019+	0.019+
	(0.009)	(0.010)	(0.015)	(0.010)	(0.008)	(0.010)	(0.009)
N	146,933	146,933	146,933	146,933	121,365	128,728	143,973
Mean of Dep. Var.	0.187	0.187	0.187	0.187	0.187	0.186	0.187
Estimation Framework	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D
Panel B:				Triple Difference Specifications			
Bound x Post 1 x Target	0.015	0.015	0.015	0.013	0.017	0.015	0.014
	(0.014)	(0.014)	(0.014)	(0.014)	(0.014)	(0.016)	(0.014)
Bound x Post 2 x Target	0.021+	0.022+	0.021+	0.016	0.020	0.020	0.023+
	(0.013)	(0.013)	(0.013)	(0.013)	(0.014)	(0.014)	(0.013)
N	269,527	269,527	269,527	269,527	223,148	238,727	263,782
Mean of Dep. Var.	0.141	0.141	0.141	0.141	0.140	0.140	0.141
Estimation Framework	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D
Macro Covariates	Yes	No	Yes	Yes	Yes	Yes	Yes
State Trends	No	No	Yes	No	No	No	No
Trends In Demographics	No	No	No	Yes	No	No	No
Minimum Sample Inclusion	3 yrs	3 yrs	3 yrs	3 yrs	3.5 yrs	3 yrs	3 yrs
Excluded States	None	None	None	None	None	N.B.<$7.00	B.>$6.55

Note: +, *, **, and *** indicate statistical significance at the 0.10, 0.05, 0.01, and 0.001 levels respectively. Panel A reports estimates of the minimum wage's short and medium run effects on the probability that an individual works without pay. More specifically, the estimates in row 1 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning in August 2009 and extending through July 2010. The estimates in row 2 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. Panel B reports analogous estimates of $\beta_{p(t)}$ from equation (3), namely our triple-difference specification. In Panel A the sample consists exclusively of individuals with average baseline wages less than $7.50. In Panel B the sample is augmented to include individuals whose average baseline wages are between $8.50 and $10.00 as a within-state control group. Standard errors are clustered at the state level.

Table A.8: **Robustness of Estimated Effects on Having No Earnings**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent Variable				No Earnings			
Panel A:				Difference-in-Differences Specifications			
Bound x Post 1	0.057**	0.047*	0.069***	0.054**	0.059**	0.066**	0.055**
	(0.019)	(0.021)	(0.019)	(0.019)	(0.020)	(0.019)	(0.019)
Bound x Post 2	0.081***	0.067**	0.094***	0.074**	0.081***	0.086**	0.077***
	(0.022)	(0.021)	(0.025)	(0.021)	(0.018)	(0.026)	(0.022)
N	146,933	146,933	146,933	146,933	121,365	128,728	143,973
Mean of Dep. Var.	0.421	0.421	0.421	0.421	0.418	0.421	0.422
Estimation Framework	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D
Panel B:				Triple Difference Specifications			
Bound x Post 1 x Target	0.047*	0.033	0.046+	0.046+	0.044+	0.055*	0.048*
	(0.023)	(0.026)	(0.023)	(0.024)	(0.024)	(0.024)	(0.023)
Bound x Post 2 x Target	0.080**	0.060*	0.080**	0.080**	0.072**	0.080**	0.076**
	(0.023)	(0.025)	(0.023)	(0.025)	(0.023)	(0.025)	(0.023)
N	269,527	269,527	269,527	269,527	223,148	238,727	263,782
Mean of Dep. Var.	0.324	0.324	0.324	0.324	0.320	0.323	0.326
Estimation Framework	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D
Macro Covariates	Yes	No	Yes	Yes	Yes	Yes	Yes
State Trends	No	No	Yes	No	No	No	No
Trends In Demographics	No	No	No	Yes	No	No	No
Minimum Sample Inclusion	3 yrs	3 yrs	3 yrs	3 yrs	3.5 yrs	3 yrs	3 yrs
Excluded States	None	None	None	None	None	N.B.<$7.00	B.>$6.55

Note: +, *, **, and *** indicate statistical significance at the 0.10, 0.05, 0.01, and 0.001 levels respectively. Panel A reports estimates of the minimum wage's short and medium run effects on the probability that an individual has no earnings. More specifically, the estimates in row 1 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning in August 2009 and extending through July 2010. The estimates in row 2 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. Panel B reports analogous estimates of $\beta_{p(t)}$ from equation (3), namely our triple-difference specification. In Panel A the sample consists exclusively of individuals with average baseline wages less than $7.50. In Panel B the sample is augmented to include individuals whose average baseline wages are between $8.50 and $10.00 as a within-state control group. Standard errors are clustered at the state level.

Table A.9: **Robustness of Estimated Effects on Average Income**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent Variable	Average Individual Income						
Panel A:	Difference-in-Differences Specifications						
Bound x Post 1	-96.996*	-87.599*	-84.541	-87.660*	-96.980*	-124.336**	-93.847*
	(37.581)	(36.696)	(53.457)	(38.776)	(38.372)	(43.543)	(38.195)
Bound x Post 2	-152.976**	-139.535**	-106.873	-123.230**	-158.234***	-185.557***	-144.592**
	(44.511)	(42.081)	(83.908)	(45.438)	(36.943)	(51.477)	(45.290)
N	146,933	146,933	146,933	146,933	121,365	128,728	143,973
Mean of Dep. Var.	747.633	747.633	747.633	747.633	753.252	753.153	743.266
Estimation Framework	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D
Panel B:	Triple Difference Specifications						
Bound x Post 1 x Target	-112.728+	-92.376	-113.532+	-115.456+	-119.105*	-114.755+	-122.473*
	(56.538)	(56.956)	(56.811)	(57.693)	(57.780)	(65.054)	(57.033)
Bound x Post 2 x Target	-184.768*	-155.007*	-185.196*	-186.519*	-200.613*	-198.313*	-186.548*
	(74.359)	(70.847)	(74.983)	(74.020)	(75.638)	(86.382)	(76.135)
N	269,527	269,527	269,527	269,527	223,148	238,727	263,782
Mean of Dep. Var.	995.022	995.022	995.022	995.022	1,001.254	1,000.081	991.508
Estimation Framework	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D
Macro Covariates	Yes	No	Yes	Yes	Yes	Yes	Yes
State Trends	No	No	Yes	No	No	No	No
Trends In Demographics	No	No	No	Yes	No	No	No
Minimum Sample Inclusion	3 yrs	3 yrs	3 yrs	3 yrs	3.5 yrs	3 yrs	3 yrs
Excluded States	None	None	None	None	None	N.B.<$7.00	B.>$6.55

Note: +, *, **, and *** indicate statistical significance at the 0.10, 0.05, 0.01, and 0.001 levels respectively. Panel A reports estimates of the minimum wage's short and medium run effects on monthly income. More specifically, the estimates in row 1 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning in August 2009 and extending through July 2010. The estimates in row 2 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. Panel B reports analogous estimates of $\beta_{p(t)}$ from equation (3), namely our triple-difference specification. In Panel A the sample consists exclusively of individuals with average baseline wages less than $7.50. In Panel B the sample is augmented to include individuals whose average baseline wages are between $8.50 and $10.00 as a within-state control group. Standard errors are clustered at the state level.

Table A.10: **Robustness of Estimated Effects on Transitions into Lower Middle Class Earnings**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent Variable				Probability of Earning $1500+			
Panel A:				Difference-in-Differences Specifications			
Bound x Post 1	-0.016	-0.014	-0.007	-0.015	-0.011	-0.019	-0.017
	(0.011)	(0.011)	(0.018)	(0.011)	(0.013)	(0.013)	(0.011)
Bound x Post 2	-0.049***	-0.046**	-0.021	-0.043**	-0.053***	-0.056***	-0.048***
	(0.013)	(0.013)	(0.025)	(0.013)	(0.014)	(0.014)	(0.013)
N	146,933	146,933	146,933	146,933	121,365	128,728	143,973
Mean of Dep. Var.	0.206	0.206	0.206	0.206	0.203	0.220	0.206
Estimation Framework	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D
Panel B:				Triple Difference Specifications			
Bound x Post 1 x Target	-0.002	0.001	-0.002	-0.005	-0.017	0.002	-0.008
	(0.018)	(0.016)	(0.018)	(0.018)	(0.022)	(0.020)	(0.018)
Bound x Post 2 x Target	-0.054*	-0.049*	-0.053*	-0.059*	-0.075**	-0.050+	-0.058*
	(0.024)	(0.022)	(0.024)	(0.023)	(0.027)	(0.028)	(0.024)
N	269,527	269,527	269,527	269,527	223,148	238,727	263,782
Mean of Dep. Var.	0.283	0.283	0.283	0.283	0.284	0.293	0.283
Estimation Framework	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D	D-in-D-in-D
Macro Covariates	Yes	No	Yes	Yes	Yes	Yes	Yes
State Trends	No	No	Yes	No	No	No	No
Trends In Demographics	No	No	No	Yes	No	No	No
Minimum Sample Inclusion	3 yrs	3 yrs	3 yrs	3 yrs	3.5 yrs	3 yrs	3 yrs
Excluded States	None	None	None	None	None	N.B.<$7.00	B.>$6.55

Note: +, *, **, and *** indicate statistical significance at the 0.10, 0.05, 0.01, and 0.001 levels respectively. Panel A reports estimates of the minimum wage's short and medium run effects on the probability that an individual has earnings greater than $1500 in a month. More specifically, the estimates in row 1 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning in August 2009 and extending through July 2010. The estimates in row 2 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. Panel B reports analogous estimates of $\beta_{p(t)}$ from equation (3), namely our triple-difference specification. In Panel A the sample consists exclusively of individuals with average baseline wages less than $7.50. In Panel B the sample is augmented to include individuals whose average baseline wages are between $8.50 and $10.00 as a within-state control group. Standard errors are clustered at the state level.

Table A.11: **Heterogeneity of Employment Effects by Age, Family Structure, and Gender**

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable	Employed					
Panel A:	Difference-in-Differences Specifications					
Bound x Post 1	-0.041	-0.040+	-0.034	-0.056*	-0.036	-0.057*
	(0.026)	(0.020)	(0.023)	(0.028)	(0.021)	(0.025)
Bound x Post 2	-0.068**	-0.055**	-0.064**	-0.061*	-0.052*	-0.082**
	(0.024)	(0.020)	(0.023)	(0.028)	(0.022)	(0.029)
N	82,224	64,709	89,963	56,970	91,257	55,676
Mean of Dep. Var.	0.738	0.800	0.739	0.808	0.779	0.745
Panel B:	Triple Difference Specifications					
Bound x Post 1 x Target	-0.025	-0.027	-0.041	-0.031	-0.038	-0.022
	(0.034)	(0.025)	(0.029)	(0.028)	(0.026)	(0.037)
Bound x Post 2 x Target	-0.057+	-0.051+	-0.096**	-0.020	-0.059*	-0.063+
	(0.030)	(0.027)	(0.028)	(0.031)	(0.023)	(0.037)
N	141,035	128,492	154,353	115,174	164,419	105,108
Mean of Dep. Var.	0.785	0.851	0.792	0.850	0.827	0.800
Sample	Under 35	Age 35+	Kids in HH	Childless	Female	Male

Note: +, *, **, and *** indicate statistical significance at the 0.10, 0.05, 0.01, and 0.001 levels respectively. Panel A reports estimates of the minimum wage's short and medium run effects on the probability that an individual is employed. More specifically, the estimates in row 1 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning in August 2009 and extending through July 2010. The estimates in row 2 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. Panel B reports analogous estimates of $\beta_{p(t)}$ from equation (3), namely our triple-difference specification. In Panel A the sample consists exclusively of individuals with average baseline wages less than \$7.50. In Panel B the sample is augmented to include individuals whose average baseline wages are between \$8.50 and \$10.00 as a within-state control group. Columns 1 and 2 split the baseline sample on the basis of age, columns 3 and 4 on the basis of household structure, and columns 5 and 6 on the basis of gender. Additional details are provided in the main text and within the table itself. Standard errors are clustered at the state level.

Table A.12: **Heterogeneity of Paid Employment Effects by Age, Family Structure, and Gender**

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable	No Earnings					
Panel A:	Difference-in-Differences Specifications					
Bound x Post 1	0.079**	0.027	0.057*	0.058*	0.046+	0.076**
	(0.026)	(0.019)	(0.024)	(0.025)	(0.023)	(0.025)
Bound x Post 2	0.109***	0.046+	0.091**	0.066*	0.071**	0.098**
	(0.026)	(0.024)	(0.026)	(0.026)	(0.024)	(0.032)
N	82,224	64,709	89,963	56,970	91,257	55,676
Mean of Dep. Var.	0.427	0.413	0.443	0.386	0.409	0.440
Panel B:	Triple Difference Specifications					
Bound x Post 1 x Target	0.063+	0.017	0.064*	0.031	0.051+	0.037
	(0.034)	(0.026)	(0.030)	(0.028)	(0.029)	(0.034)
Bound x Post 2 x Target	0.096**	0.054+	0.121***	0.034	0.085**	0.071
	(0.035)	(0.030)	(0.030)	(0.026)	(0.024)	(0.043)
N	141,035	128,492	154,353	115,174	164,419	105,108
Mean of Dep. Var.	0.354	0.293	0.353	0.286	0.312	0.344
Sample	Under 35	Age 35+	Kids in HH	Childless	Female	Male

Note: +, *, **, and *** indicate statistical significance at the 0.10, 0.05, 0.01, and 0.001 levels respectively. Panel A reports estimates of the minimum wage's short and medium run effects on on the probability that an individual has no earnings. More specifically, the estimates in row 1 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning in August 2009 and extending through July 2010. The estimates in row 2 are of the coefficient $\beta_{p(t)}$ from equation (1), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. Panel B reports analogous estimates of $\beta_{p(t)}$ from equation (3), namely our triple-difference specification. In Panel A the sample consists exclusively of individuals with average baseline wages less than \$7.50. In Panel B the sample is augmented to include individuals whose average baseline wages are between \$8.50 and \$10.00 as a within-state control group. Columns 1 and 2 split the baseline sample on the basis of age, columns 3 and 4 on the basis of household structure, and columns 5 and 6 on the basis of gender. Additional details are provided in the main text and within the table itself. Standard errors are clustered at the state level.