The Minimum Wage and the Great Recession: Evidence of Effects on the Employment and Income Trajectories of Low-Skilled Workers

Jeffrey Clemens and Michael Wither*

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Abstract:

We estimate the minimum wage's effects on low-skilled individuals' employment and income trajectories following the Great Recession. Our approach exploits two dimensions of the data we analyze. First, we compare individuals in states that were fully bound by the 2007 to 2009 increases in the federal minimum wage to individuals in states that were not. Second, we use variation in the minimum wage's bite across skill groups to separate our samples into "target" and "within-state control" groups. Using the 2008 panel of the Survey of Income and Program Participation and the Current Population Survey, we find that binding minimum wage increases had significant, negative effects on the employment and income growth of targeted workers. Estimates using both data sets are robust to adopting a range of alternative strategies, including matching on the size of states' housing declines, to account for variation in the Great Recession's severity across states. In aggregate, we estimate that this period's minimum wage increases reduced the national employment-to-population ratio by around 0.6 percentage point. We provide evidence that the $7.25 federal minimum wage had deeper and more sustained bite on low-skilled groups' wage distributions than the federal minimum wage increases of the 1990s. We interpret this key contextual fact, which stems in part from the effects of trade, technology, and the housing market on demand for low-skilled labor, through a framework that emphasizes how both bargaining frictions and labor demand's determinants shape the minimum wage's effects.

*Clemens: University of California at San Diego, Jeffrey Clemens, Economics Department, 9500 Gilman Drive #0508, La Jolla, CA 92093-0508, USA. Telephone: 1-509-570-2690. E-mail: *clemens.jeffrey@gmail.com*. Wither: University of California at San Diego. Email: *mwither@ucsd.edu*. We thank Jean Roth for greatly easing the navigation and analysis of SIPP data, as made accessible through NBER. We thank Prashant Bharadwaj, Marika Cabral, Julie Cullen, Gordon Dahl, Roger Gordon, Jim Hamilton, Neale Mahoney, Day Manoli, Jonathan Meer, Karthik Muralidharan, Michael Strain, Stan Veuger, Johannes Wieland, and seminar participants at Brown, the Census Bureau, CRA, Cornell-PAM, Ernst & Young, IMPAQ International, Texas A&M, NC State, UC Berkeley, Stanford, Columbia, the Federal Reserve Board, the Harris School, Kellogg, and the Chicago and San Francisco Federal Reserve Banks for helpful comments and suggestions. This revision subsumes portions of analyses initially distributed in Clemens (2016) and Clemens (2015).

Between July 23, 2007, and July 24, 2009, the U.S. federal minimum wage rose from \$5.15 to \$7.25 per hour. During the concurrent recession, the employment-to-population ratio declined by 4 percentage points among prime aged adults and by 10 percentage points among those aged 16 to 21. Both ratios recovered slowly following the recession's conclusion, and young adult employment remains well below its pre-recession peak. 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.

We begin in section 1 by presenting a framework for describing and interpreting the minimum wage's effects. The minimum wage's effect on employment depends crucially on the distribution of productivity across potential workers. This basic insight connects the minimum wage's effects to the wage distribution's demand-side determinants. If business cycle forces, trade patterns (Autor, Dorn, and Hanson, 2013), or technological change (Katz and Murphy, 1992; Autor, Levy, and Murnane, 2003) reduce the value of low-skilled workers' output, the employment loss linked to a given minimum wage will rise.[1] The framework further highlights that the effects of increasing the minimum wage will be large when the productivity of many workers lies just above its initial value. The employment declines we estimate suggest that many workers were so situated during the period we analyze. The evolution of low-skilled groups' wage distributions, as presented in section 1.2, provides evidence consistent with this interpretation.

[1]Notably, new-Keynesian accounts of the business cycle connect employment losses to sources of downward wage rigidity, of which the minimum wage is one. Neoclassical and new-Keynesian intuitions thus both suggest that the minimum wage is an institution that will mediate the transmission of demand and productivity shocks into changes in employment.

Our empirical analysis harnesses the fact that the 2007 through 2009 increases in the federal minimum wage were differentially binding across states. In the states we describe as “partially bound” or “unbound,” the effective minimum wage rose by $1.42 between 2006 and 2012. In the states we describe as “fully bound” or “bound,” the effective minimum wage rose by $2.04. Of the long-run differential, $0.58 took effect on July 24, 2009. Our primary analysis of these differentially binding minimum wage increases uses 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 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 give our estimates several novel features relative to the literature’s standard approaches.[2]

Our approach’s first feature of interest 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

[2]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.

beneficiaries of anti-poverty efforts (Burkhauser and Sabia, 2007; Sabia and Burkhauser, 2010). Assessing the minimum wage from an anti-poverty perspective thus requires estimating its effects on the broader population of low-skilled workers, which we are able to do.[3]

Econometrically, our setting has several advantages. One benefit of the SIPP's 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 thus provides a method for controlling 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; Neumark, Salas, and Wascher, 2013). Third, 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 min-

[3]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 to 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.

imum wage's bite on our target group's wage distribution is nearly twice its bite on comparison samples selected using approaches common in the literature.

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 wage's implementation, we estimate that targeted workers' employment rates had fallen by 6.6 percentage points (9 percent) more in bound states than in unbound states.

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 relatively severe in unbound states. In our baseline specification, we control directly for a proxy for the severity of the crisis. We also apply a matching framework in which we conduct analysis on samples restricted to states with similarly severe housing declines. Our baseline estimates are robust to these and several additional potentially relevant specification modifications. By contrast, we find that taking no steps to control for variations in states' housing declines biases estimates towards zero.

We conduct further analysis of employment using the Current Population Survey (CPS). While the CPS lacks the SIPP's longitudinal advantages, it allows us to initiate samples in the years preceding the May 2007 legislation behind this period's minimum wage increases. This makes it possible to investigate the full transitional dynamics associated with the law's implementation. The aggregate employment implications of our SIPP and CPS analyses are quite similar.

In both the SIPP and CPS analyses, we show that the differential employment changes in bound states relative unbound states occur among the skill groups the minimum wage targets. In the SIPP analysis, our estimate of the employment decline among targeted workers is essentially unchanged by netting out employment changes among groups with moderately higher baseline wage rates. In the CPS analysis, we similarly find that differential employment declines were concentrated among the youngest and lowest education skill groups.

We next estimate the effects of binding minimum wage increases on low-skilled workers' incomes and income *trajectories*. The 2008 SIPP panel provides a unique opportunity to investigate such effects, as its 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 monthly incomes fell by $90 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 (a possible "internship" effect). 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.[4] Like previous results, these estimates are robust to netting out the experience of workers with average baseline wages just above the new federal min-

[4]Earning $1500 would require working full time (40 hours per week for 4.33 weeks per month) at a wage of $8.66. We characterize $1500 as a "lower middle class" earnings threshold.

imum. 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.

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 0.6 percentage point. This accounts for 12 percent of the total decline from 2006 to 2012.

1 A Wage and Employment Determination Framework

This section develops a framework for understanding the minimum wage's effects on employment rates and wage distributions. Observed wage rates result from transactions in which a firm's wage offer exceeds an individual's reservation wage. Let individual i have reservation wage $v_{i,t}$ at time t.[5] Individual i's productivity, the product of the quantity and market price of his or her output, is $a_{i,t}$ per hour.[6]

In the spirit of Bound and Johnson (1992), we describe firms' wage offers as resulting from a combination of competitive market forces and bargaining power. Firms maximize profits by employing all individuals they can hire at wage rates less than or equal to the value of their output. Absent binding minimum wage regulation, firms offer individual i a wage of $\theta_{i,t}a_{i,t}$ with $\theta_{i,t} \leq 1$. If workers are paid precisely their marginal product, $\theta_{i,t} = 1$. The bargaining parameter $\theta_{i,t}$ can be modeled as an outcome of a variety of

[5]A variety of factors, including the generosity of social insurance programs, may determine $v_{i,t}$.

[6]We treat $a_{i,t}$ as an individual-level constant. Introducing a firm-worker match component would not alter the primary considerations we emphasize. The key labor demand-side components of our framework can be microfounded as arising from a market in which identical firms produce homogenous output with a technology that is additively separable in the labor employed.

labor market frictions (Manning, 2011). It is central to the minimum wage's potential to increase individuals' earnings.

The final determinant of wage offers is the statutory minimum wage, w_t^{min}. So long as $a_{i,t} \geq w_t^{min}$, so that the value of the individual's expected output exceeds the statutory minimum wage, firms will offer employment at w_t^{min} when $\theta_{i,t}a_{i,t} < w_t^{min}$. When $a_{i,t} < w_t^{min}$, on the other hand, firms will not offer the individual employment. The framework thus incorporates channels through which the minimum wage's primary intended and unintended effects may be realized.

Observed wage rates, $w_{i,t}$, can be summarized as follows:

$$w_{i,t} = \begin{cases} \theta_{i,t}a_{i,t} \text{ if } \theta_{i,t}a_{i,t} > w_t^{min} \text{ and } \theta_{i,t}a_{i,t} \geq v_{i,t} \\ w_t^{min} \text{ if } w_t^{min} > v_{i,t} \text{ and } \theta_{i,t}a_{i,t} < w_t^{min} \text{ and } a_{i,t} \geq w_t^{min} \\ 0 \text{ if } a_{i,t} < w_t^{min} \\ 0 \text{ if } w_t^{min} < v_{i,t} \text{ and } w_t^{min} > \theta_{i,t}a_{i,t} \\ 0 \text{ if } \theta_{i,t}a_{i,t} < v_{i,t} \text{ and } \theta_{i,t}a_{i,t} \geq w_t^{min}. \end{cases} \tag{1}$$

The first two rows of equation (1) describe wage rates among the employed while rows three through five describe those not employed. Row one describes individuals whose wage offers exceed their reservation values and are unbound by the legal minimum. Row two describes individuals paid the minimum who would otherwise receive less as a result of their bargaining position. Row three describes the involuntary unemployment that occurs when the legal minimum exceeds the value of an individual's expected output. Rows four and five describe individuals whose reservation wage rates exceed the wage rates firms offer.

In the subsection below we implicitly assume that the parameters of the above framework are determined independently of another. We note that the literature has long considered the minimum wage's potential effects on prices (Aaronson, 2001; Aaronson, French, and MacDonald, 2008) as well as on the bargaining positions of workers for which $\theta_{i,t}a_{i,t}$ exceeds w_t^{min} (Lee, 1999; Autor, Manning, and Smith, 2016). Such relationships can be important for understanding the entirety of the minimum wage's effects on the real wage distribution. Further, changes in individuals' productivities and/or the minimum wage may alter their reservation values. While these linkages can either dampen or augment the minimum wage's effects on employment, they do not qualitatively alter the considerations we emphasize.

1.1 Implications For Evaluation of Minimum Wage Policy

The above framework generates several insights relevant to minimum wage analyses. First, it succinctly captures when the minimum wage's primary intended and unintended consequences will be relatively large. The minimum wage's intended effect of increasing low-skilled individuals' wages can be large when their bargaining positions ($\theta_{i,t}$) are relatively weak. Its unintended effect of reducing low-skilled individuals' employment can be large when the full market value of their output ($a_{i,t}$) is relatively low. The minimum wage's effectiveness is thus intimately linked to the mix of market and institutional forces underlying observed wages among the low-skilled.

The minimum wage's effect on employment depends on where it falls in the productivity distribution. At time t, let a_i be distributed according to the probability density function $f_t(\cdot)$. The gross employment loss linked to a minimum wage of w_t^{min} is then

$$\int_0^{w_t^{min}} f_t(a_i) \times 1\{\theta_i a_i \geq v_i\} d(a_i). \tag{2}$$

Equation (2) describes the fraction of the population that would desire to work at firms' unconstrained wage offers ($\theta_i a_i \geq v_i$), but whose productivity falls below the statutory minimum wage.

The minimum wage may also increase employment among individuals who would be unwilling to work at firms' unconstrained wage offers. Letting $g_t(\cdot)$ be the density of $\theta_i a_i$, the extent of this entry effect is

$$\int_0^{w_t^{min}} g_t(\theta_i a_i) \times 1\{w_t^{min} \leq a_i\} \times 1\{w_t^{min} \geq v_i\} \times 1\{\theta_i a_i < v_i\} d(\theta_i a_i). \tag{3}$$

Equation (3) describes individuals who firms are willing to pay the minimum ($w_t^{min} \leq a_i$) and who are willing to work for the minimum ($w_t^{min} \geq v_i$), but who would be unwilling to work for firms' unconstrained wage offers ($\theta_i a_i < v_i$).[7]

Equations (2) and (3) make clear that the net employment effect of a given minimum wage will evolve as the productivity distribution evolves. Suppose, for example, that new labor-substituting technologies become available between periods 1 and 2. The resulting downward shift in the value of low-skilled workers' output implies an increase in the minimum wage's effect on employment among those who are willing to work at firms' unconstrained wage offers. This gross employment loss changes by

$$\int_0^{w^{min}} f_2(a_i) \times 1\{\theta_i a_i \geq v_i\} d(a_i) - \int_0^{w^{min}} f_1(a_i) \times 1\{\theta_i a_i \geq v_i\} d(a_i). \tag{4}$$

Equation (4) could similarly describe how the gross employment loss evolves following a demand-side shock that reduces output prices.

Equation (4) highlights that the minimum wage can mediate the extent to which a negative demand or productivity shock manifests itself through employment declines as opposed to declines in transacted wages. Empirically, such effects would be difficult to disentangle from voluntary labor force exits among individuals for whom $\theta_i a_i$ falls

[7]The pervasiveness of this entry effect may depend crucially on the institutional underpinnings of the bargaining parameter θ_i (Manning, 2011). A positive entry effect requires breakdowns of bilateral efficiency. That is, it requires that firms sometimes offer wage rates at which an individual is unwilling to work ($\theta_i a_i < v_i$) even when there are potential gains from trade ($a_i \geq v_i$). Such breakdowns of bilateral efficiency would not occur in standard "ex-post bargaining" models, in which employment materializes in all firm-worker matches for which the worker's productivity exceeds his or her reservation wage.

below v_i. Difficulty arises because accounting for voluntary labor supply effects requires knowledge of affected individuals' reservation values.[8] Employment changes described by equation (4) will thus not readily be isolated through standard program evaluation analyses. Further, a decline in a_i may push some individuals out of employment through *both* the involuntary ($w^{min} > a_i$) and voluntary ($\theta_i a_i < v_i$) margins.[9]

We next describe the effects of a change to the minimum wage. All else equal, the gross employment loss (among those who are willing to work at firms' unconstrained wage offers) that results from increasing the minimum wage from w_1^{min} to w_2^{min} is

$$\int_{w_1^{min}}^{w_2^{min}} f_t(a_i) \times 1\{\theta_i a_i \geq v_i\} d(a_i). \tag{5}$$

Equation (5) shows that the gross employment loss due to a change in the minimum wage depends on the density of the productivity distribution between the minimum's old and new levels.[10] In general, it should not be surprising for the elasticities implied by this employment effect to vary across settings. Minimum wage changes that move through thin portions of the productivity distribution will have small effects while changes that move through thick portions will have large effects. Further, such effects will be more sustained when productivity growth is slow than when it is rapid.[11] The fol-

[8] Chetty, Guren, Manoli, and Weber (2012) observe that microeconometric estimates of labor supply elasticities can be useful for attempting to calibrate this voluntary exit margin. As they write, "the marginal density of the reservation wage distribution that determines the impacts of macroeconomic variation on employment also determines the impacts of quasi-experiments such as tax policy changes on employment rates." Regarding micro and macro labor supply elasticities, they observe that calibrations based on elasticities estimated using microeconomic quasi-experiments cannot rationalize the magnitude of employment fluctuations over the business cycle. Reconciling macro and micro labor supply elasticity estimates thus requires factors like wage rigidities, of which the minimum wage is one, that move individuals off of their short-run labor supply curves.

[9] One implication of this point is that efforts to decompose the causes of an employment decline can be sensitive to the order of decomposition.

[10] The "net employment change" would net out entry of the sort described by equation (3).

[11] Recent minimum wage analyses by Sorkin (2015) and Aaronson, French, and Sorkin (2013) further highlight that the minimum wage's employment effects will be relatively large on time horizons that

lowing subsection illustrates how the minimum wage's movement through low-skilled workers' wage distributions has varied across recent historical episodes.

1.2 Minimum Wage Changes and Observed Wage Distributions

The panels of figure 1 present wage data from the Merged Outgoing Rotation Groups of the Current Population Survey. The presented series describe the wage distributions of young adults aged 16 to 21 (panels A and B) and of individuals aged 30 and under with less than a completed high school education (panels C and D). The samples are restricted to individuals in states that have historically maintained minimum wage rates in line with the federal minimum.

The series in the panels map directly into the framework summarized by equation (1). Markers correspond with positive wage quantiles, sorted in descending order.[12] For a given year, a group's employment rate can be inferred from the highest x-axis value associated with a positive wage. The note to the figure provides further details on the construction of the presented series.

The figure displays the evolution of low-skilled groups' employment rates and wage distributions over the two most recent historical episodes surrounding increases in the federal minimum wage. The distributions in panels A and C are from 2006 and 2014 while the distributions in panels B and D are from 1994 and 2002. We emphasize that these episodes differ starkly with respect to productivity growth. During the earlier period, productivity growth averaged 2.7 percent per year and was widely spread across skill groups (Autor, Katz, and Kearney, 2008). During the latter period, productivity growth averaged 1.3 percent per year and, due in part to expanding trade with China,

incorporate firms' capital investment and technology adoption decisions. We note that the depth of a minimum wage increase's bite and the amount of time over which it is sustained are relevant for whether firms choose to incur the costs associated with adjusting along these margins.

[12]To compress the y-axis range we suppress the top 2 percentiles of each distribution.

was less favorable to the low-skilled (Autor, Dorn, and Hanson, 2013). Further, work by Bosler, Daly, Fernald, and Hobijn (2016) shows that within-skill-group productivity growth was much lower following the recession than the aggregate productivity data suggest. From 2008 to 2010, for example, they find that overall productivity growth was buoyed by a 0.9 percentage point contribution from changes in "labor quality." That is, it resulted from declines in employment among low-skilled groups.[13]

Baseline productivity and subsequent productivity growth determine the extent to which a minimum wage increase binds the productivity distribution. In panels B an D, the 2002 wage distributions reflect substantial upward shifts relative to 1994. These upward shifts are in line with what one would obtain by straightforwardly adjusting the 1994 wage distribution to account for economy-wide inflation and productivity growth. Rapid productivity growth rendered the minimum wage's rise from $4.25 to $5.15 far less binding than one would infer from the 1994 wage distribution alone. By contrast, the upper quantiles of the 2014 wage distributions in panels A an C are essentially unshifted relative to the corresponding quantiles of the 2006 wage distributions. Low inflation and low productivity growth, in particular at the skill distribution's lower end, led this latter period's minimum wage increases to bind deeply and for a sustained period of time.

Interpreted through the lens of this section's framework, the data in figure 1 suggest that the minimum wage contributed meaningfully to the decline in low-skilled individuals' employment since 2006. In the remainder of this paper, we attempt to isolate and quantify the causal effects of this period's minimum wage increases using variation in the magnitude of these minimum wage changes across states. We bear in mind that such estimates will not fully capture the minimum wage's role as a mediator of negative

[13]As in our CPS analysis, Bosler, Daly, Fernald, and Hobijn (2016) define skill groups on the basis of observable characteristics, in particular age and education. Their estimates will thus tend to understate the full contribution of changes in labor quality because they will not capture shifts in the composition of unobservable skill levels within age-by-education cells.

demand and productivity shocks' employment consequences. With reference to section 1.1, our estimates of these minimum wage increases' effects can detect the employment changes described by equation (5), but not those described by equation (4).

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. Making good on commitments from the 2006 election campaign, the 110th Congress legislated a series of minimum wage increases through the "U.S. Troop Readiness, Veterans' Care, Katrina Recovery, and Iraq Accountability Appropriations Act" on May 25, 2007. 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 2 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.[14]

[14]Designating states on the basis of their early 2008 minimum wage rates is most suitable for our analysis using the 2008 SIPP panel. For consistency, we apply this designation in our analyses using both the SIPP and the CPS. An alternative designation based on states' January 2007 minimum wage rates is arguably more suitable for analysis using the CPS. The states' whose status differs across these criteria are Wisconsin, which had a minimum wage of \$6.50 in January 2008, North Carolina, which had a minimum wage of \$6.15 in January 2008, Montana, which had a minimum wage of \$6.25 in January 2008, Minnesota, which had a minimum wage of \$6.15 in January 2008, Maryland, which had a minimum wage of \$6.15 in January 2008, Arkansas, which had a minimum wage of \$6.25 in January 2008, and Iowa, which had

Figure 3 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 by roughly $1 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 between August 2008 and August 2012. By contrast, bound states saw their effective minimums rise by nearly the full, legislated $0.70 on July 24, 2009. From 2006 to 2012, the effective minimum wage rose by $1.42 in the "unbound" or "partially bound" states and by $2.04 in the "bound" or "fully bound" states.

The primary threat to our estimation framework is the possibility that bound and unbound states experienced housing crises of different average severity. Figure 4 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.[15] 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

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a minimum wage of $7.25 in January 2008. Iowa was in line with the $5.15 federal minimum wage as of January 2007, but enacted an accelerated increase to $7.25 that was effective as of January 2008. The remaining states with conflicted designations were states that exceeded the federal minimum wage in January 2007 and which waited for the July 2008 federal increase to bring their minimum wage rates above $6.55. Dropping states with conflicted designations from the samples has essentially no effect on the estimates. Coding states according to their January 2007 minimum wage rates moderately reduces the estimated employment effects on the CPS samples of teenagers and young high school dropouts and moderately increases the estimated employment effects on the CPS samples of individuals aged 16 to 21.

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

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) and the Current Population Survey (CPS). Subsections 3.1 and 3.2 describe the estimation frameworks we implement in the SIPP and CPS analyses respectively. Subsection 3.3 describes a complementary matching framework that we implement in our analyses of both the SIPP and CPS data.

3.1 SIPP Analysis

In the 2008 SIPP panel, 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 by 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 the working age population into several groups.

The first group we analyze includes those most directly impacted by the federal minimum wage. Specifically, it includes those whose average wage, when employed during the baseline period, was less than \$7.50.[16] 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. The remainder of the population consists of those unemployed throughout the baseline period and those employed at average baseline wage rates greater than \$10.00.

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

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

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

[16] 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.

effects, Individual_i, renders controls for individual-level, time-invariant characteristics 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.[17]

Equation (6) 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.[18]

We initially use equation (6) 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:

$$Y^j_{i,s,t} = 1\{W^{j-1} < \text{Hourly Wage}_{i,s,t} < W^j\}. \tag{7}$$

These $Y^j_{i,s,t}$ are indicators that are set equal to 1 if an individual's hourly wage is between

[17]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 A.7.

[18]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.

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.[19] Although we find our estimates of equation (6) to be robust to a range of approaches to controlling for heterogeneity in macroeconomic conditions, we additionally implement a triple-difference model and a matching framework. In the triple-difference model, 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:

$$\begin{aligned}
Y_{i,s,t} = &\sum_{p(t)\neq 0} \beta_{p(t)} \text{Period}_{p(t)} \times \text{Bound}_s \times \text{Target}_{g(i)} \\
&+ \alpha_{1_{s,p(t)}} \text{State}_s \times \text{Period}_{p(t)} + \alpha_{2_{s,g(i)}} \text{State}_s \times \text{Target}_{g(i)} + \alpha_{3_{t,g(i)}} \text{Time}_t \times \text{Target}_{g(i)} \\
&+ \alpha_{4_s} \text{State}_s + \alpha_{5_t} \text{Time}_t + \alpha_{6_i} \text{Individual}_i + \mathbf{X_{s,t,g(i)}}\gamma + \mathbf{D_i} \times \text{Trend}_t \phi + \varepsilon_{i,s,t}. \qquad (8)
\end{aligned}$$

Equation (8) augments equation (6) with the standard components of triple-difference estimation. These include group-by-time-period effects, group-by-state effects, and state-

[19]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 (6) and (8). Columns 1 and 2 characterize the 3,200 group 1 individuals in our bound and unbound states' samples. 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.

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.[20] 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 esti-

[20]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.

mates are robust to controlling for a set of linear trends interacted with dummy variables associated with each individual's modal industry of employment over the baseline period. We further address this potential concern through analysis using the CPS. Because the CPS is a repeated cross-section, "target" samples must be selected on the basis of age or age and education. They are thus demographically similar by construction.

3.2 CPS Analysis

We conduct further employment analysis using data from the Current Population Survey (CPS). The specification estimated in our CPS analysis can be found below:

$$\begin{aligned} Y_{i,s,t} = &\sum_{p(t)\neq 0} \beta_{p(t)} \text{Bound}_s \times \text{Period}_{p(t)} \\ &+ \alpha_{1_s} \text{State}_s + \alpha_{2_t} \text{Time}_t + \mathbf{X_{s,t}}\gamma + \varepsilon_{i,s,t}. \end{aligned} \tag{9}$$

Because the CPS lacks the SIPP's longitudinal structure, equation (9) differs from equation (6) in that it lacks individual level fixed effects. Because the CPS allows us to initiate samples in the years preceding the May 2007 legislation behind this period's minimum wage increases, our coding of time periods tracks the the full transitional dynamics associated with the law's implementation. We code May 2007 through July 2009 as the law's implementation period (period $p =$ Transition) and earlier months and years as the base period ($p = 0$). Periods Post 1(t) and Post 2(t) remain coded as before.

3.3 Analysis on Samples Matched on the Housing Decline's Severity

Within the difference-in-differences frameworks of equations (6) and (9), we examine our results' robustness to applying sample restrictions generated by a matching procedure. We match states on the size of their median house price declines between 2006 and 2012 (with values averaged across all months in these years). The matches are thus based on the extent of the housing decline from the first to the last year of the CPS analysis sample. To be more precise regarding the procedure, we apply nearest neighbor matching without replacement. We then restrict the sample on the basis of the quality of the resulting matches. For example, the baseline matching sample requires that the difference in matched states' housing declines be no greater than $20,000.[21] The presented results are robust to applying alternative thresholds of $5,000, $10,000, and 5 log points.

Appendix table A.9 presents summary statistics for the housing declines in fully and partially bound states for the analysis samples we utilize. Row 1 shows the disparity in the severity of the housing decline between the full sets of fully and partially bound states. Between 2006 and 2012, the FHFA's all-transactions median house price index declined, on average, by $72,000 in partially bound states and by $24,700 in fully bound states. On a population-weighted basis the difference between the bound and unbound state samples was an even more substantial $90,000. Row 2 shows the comparable means for the sample restricted to pairs with declines no more than $5,000 apart from one another. For this sample, which retains 20 states, the mean decline in partially bound states was $29,700 and the mean decline in fully bound states was $27,400. The $10,000 matching criterion retains 24 states and yields means of $32,000 and $29,700 respectively. The $20,000 matching criterion retains 32 states and yields means of $38,500 and $27,200 respectively. Finally, the requirement that the declines be within 5 log points of one another retains 30 states and yields means of $36,900 and $24,100. The difference in the means in the latter sample reflects the fact that the baseline *level* of median house prices is, on average, higher in partially bound states than in fully bound states.

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[21]This approach to sample selection is sometimes called the "caliper" method (Cochran and Rubin, 1973; Crump, Hotz, Imbens, and Mitnik, 2006).

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

Before presenting our estimates of equations (6) and (8), we now graphically present the raw data underlying our results. Figure 5 presents time series tabulations of the raw data underlying our estimates of equations (6) and (8). 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 37 percent of their months in jobs with wages between $5.15 and $7.25, 28 percent of their months unemployed, 11 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.[22] The observed transition period motivates our accounting for dynamics when estimating equations (6) and (8). By November 2009, individuals in the bound and unbound states have equal likelihoods of being in jobs with wages between

[22]The transition window likely reflects a combination of real economic factors and measurement artifacts. 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.

$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 nearly 4 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 4 percentage points in the first year and 5 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 expected among workers with modestly greater skills. Panel D shows that the employment of workers with average baseline wages between $7.50 and $10.00 changed similarly in bound and unbound states between the initial and later years of this period. These data reveal that the short- and medium-run estimates associated with equations (6) and (8) will yield similar results.[23]

[23]It appears that the employment of slightly higher skilled workers declined in bound states relative

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 more detailed analysis of the factors contributing to these differential income trajectories.

Figures 6 and 7 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 (6) and (8), 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 5. Adjusting for the housing bubble's greater severity in unbound states relative to bound states moderately increases the estimated magnitudes.

Figure A.1 similarly presents dynamic difference-in-differences estimates using the CPS. The figure shows that, as in the SIPP, differential employment changes were concentrated among the young and low education individuals most intensively targeted by this period's minimum wage increases. The following section presents these and other

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to unbound states during the April to July transition period, after which it quickly recovered. Our triple-difference estimates of changes in employment during the transition period will thus be smaller than our difference-in-differences estimates. The short- and medium-run effects, however, will be robust across these specifications.

results in a more summary, tabular fashion.

5 Regression Analysis of the Minimum Wage's Effects

This section presents estimates of equations (6), (8), and (9). 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 8 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.[24] 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.

[24]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.

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 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 a reasonable sample for constructing within-state employment counterfactuals for estimating equation (8).

Table 2 and the panels of Figure 8 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 (6). 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.[25]

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.[26]

In Figure 8'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

[25]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.

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

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 8'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.

5.2 Baseline Results on Employment

Table 2's columns 4 through 6 present estimates of equation (6) 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 4.4 percentage point decline between the baseline period and the following year. The decline relative to baseline averaged 6.6 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 2.6 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 negative 0.2 percentage points. Like the raw data from Figure 5, these results reveal that estimates of equation (8) will yield similar results.

Appendix Table A.1 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 individuals with relatively high baseline wages is economically and statistically indistinguishable from 0. The estimate for those unemployed at baseline is modestly negative, suggesting an increase in the difficulty of labor force entry. Further sub-sample analysis reveals this estimate to be driven primarily by the baseline unemployed between ages 16 and 21.

5.3 Robustness of Employment Estimates within the SIPP

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 A.7 present an analysis of this result's robustness. In Table 3, estimates in Panel A are of equation (6)'s difference-in-differences model. Estimates in Panel B are of equation (8)'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 6.8 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 difference-in-differences estimate to 5.1 percentage points and the triple-difference estimate to 4.8 percentage points. This reflects the fact that, as shown in Figure 4, the housing bubble was more severe in unbound states than in bound states. We further explore the relevance of macroeconomic controls in Table A.7.

Column 3 shows that our results are robust to controlling for state-specific linear time trends.[27] In the difference-in-differences specification, including these controls increases the estimated medium-run coefficient from 6.6 to 8.5 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.[28] 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, 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

[27]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.

[28]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.

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 7.7 percentage points in the difference-in-differences model and to 7.4 percentage points in the triple-difference model. 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.

For further analysis of the minimum wage's effects on employment in the SIPP, we refer readers to Appendix 2. Appendix 2's first sub-section reports further robustness analysis, with emphasis on the potential relevance of alternative strategies for controlling for heterogeneity in macroeconomic conditions, including the potential effects of stimulus spending through the American Recovery and Reinvestment Act (ARRA). Tables A.10 and A.11 show that the results in table 2 are robust to restricting the samples to individuals in states that could be closely matched on the basis of the housing decline's severity. Appendix 2's second sub-section presents analysis relevant to readers interested in comparisons between our methodology and approaches used regularly in the literature.

5.4 Employment Estimates in the CPS

Tables 4 and 5 present estimates of equation (9) on data from the Current Population Survey (CPS). The samples in both tables fully partition the CPS's universe of individuals aged 16 to 64 from 2006 to 2012. In table 4 we divide the samples on the basis of age alone while in table 5 we divide the samples on the basis of age and education.[29] Table 4's column 1 analyzes individuals aged 16 to 21, while column 2 analyzes those aged 22 to 45 and column 3 and analyzes those aged 46 to 64. In table 5, the sample in column 1 consists of individuals ages 30 and under with less than a completed high school education. Column 2 analyzes older dropouts and individuals ages 30 and under with exactly a completed high school education, while column 3 analyzes individuals with higher levels of experience and education. In both tables, columns 4 through 6 present analyses of samples that are further restricted to individuals in states that could be closely matched on the basis of the housing decline's severity.

The results in tables 4 and 5 are broadly consistent with our findings in the SIPP. Conditional on the size of the housing decline, employment among low-skilled individuals declined more in states that were fully bound by this period's minimum wage increases than in states that were only partially bound. From the period preceding the May 2007 minimum wage legislation to the periods following its implementation, the differential employment decline among young adults aged 16 to 21 was between 2 and 3 percentage points. The differential employment decline among individuals ages 30 and under with less than a completed high school education was between 3 and 4 percentage points. There is no evidence of differential employment changes among higher skill groups. The declines estimated in columns 1 and 4 are thus not driven by factors that simulta-

[29]Dividing the sample on the basis of both age and education improves our ability to isolate the very least skilled individuals. Because educational attainment decisions evolved over the Great Recession's course, however, it introduces a potential sample selection bias. It is not clear what direction this bias might run. Nonetheless, we report estimates for both the age-based sample and the age-by-education sample to confirm that the implied effects on aggregate employment are similar.

neously reduced employment among individuals with higher levels of experience and education.

The point estimates in columns 1 and 4 of tables 4 and 5 are several percentage points smaller than those from table 2. In section 7 we show that, because these samples represent larger fractions of the working age population, the CPS and SIPP estimates' aggregate employment implications differ modestly. This stems from the fact that, because the CPS is a repeated cross section, CPS samples must be selected on the basis of demographics rather than baseline wage histories. Consequently, the CPS samples contain larger fractions of un-targeted individuals. Put differently, the CPS samples were less intensively treated than were the SIPP samples.

5.5 Further Employment Outcomes, Average Income, and Poverty

Table 6 reports the results of a more in depth analysis of the minimum wage's effects on employment and income related outcomes. In Table 6's second column we present evidence of a novel channel through which job markets may respond to minimum wage increases. Specifically, we estimate that binding minimum wage increases modestly increased the probability that targeted individuals work without pay, perhaps in internships, by 1.7 percentage points. Between disemployment and work without pay, column 3 reports a combined 8 percentage point reduction in paid employment. Appendix Tables A.2 and A.3 explore the robustness of our estimates of the "internship" effect and the total effect on paid employment. The estimates of the medium-run effect on the probability of working without pay range from 1.0 to 2.0 percentage points, and are statistically indistinguishable from 0 in many of the specifications. Estimates of the total effect on the probability of paid employment range from 6.2 to 9.8 percentage points.

Table 6'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 $92 in the short-run and $144 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 4).

Figure 7 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 A.4 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 28 percent of months they were unemployed and in 11 percent they worked with 0 earnings. Average income for the target sample was $750 across all baseline months, and thus roughly $1,230 in months with positive earnings. For the short run (i.e., year 1), we estimated a 5.9 percentage point decline in the probability of having positive earnings. This effect is thus directly associated with an average decline of roughly $73, or $\$1,230 \times 0.059$. The decline in months with positive earnings rises to 8.2 percentage points over the following two years, implying a direct earnings decline of $101. Gains for workers successfully shifted from the old minimum to the new minimum offset very little of this decline.[30]

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.[31] 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 a modest "internship" effect, effects on income growth through reduced experience accumulation, and the fact that direct effects on wages were smaller than typically assumed.

We next estimate the minimum wage's effects on family-level outcomes. On average

[30]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 $1,213 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 5. Figure 5'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 5 implies quite modest offsets of the income losses associated with disemployment, work without pay, and lost experience accumulation.

[31]Two years of early-career earnings growth at 15 percent per year would bring earnings from a baseline of $1,230 to $1,627. An 8.2 percentage point decline in months at such earnings implies an average reduction of $133.

in our sample, each targeted worker is in a family with 1.2 targeted workers. This is roughly the average of the ratio of our estimates of the minimum wage increase's effect on family-level income to its effect on individual-level income. In the difference-in-differences specifications, for example, the short-run effect on individual-level income is $92 per month while the estimated effect on family-level income is $118 (the medium-run estimates are $144 and $273).

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).

We close this section by interpreting our estimates through the lens of the framework developed in section 1. The data from figure 1 showed that, due in part to shifts in demand for low-skilled labor, this period's federal minimum wage increases bound low-skilled groups' wage distributions more deeply and for a more sustained period of time than did the federal minimum wage increases of the 1990s. This is important because, as described by equation (5), the effects of a minimum wage change are tightly linked to the local density of the productivity distribution. Minimum wage changes that move through less dense portions of the productivity distribution should be expected to have smaller effects on employment and more favorable effects on the incomes of targeted workers.

6 Transitions out of Low-Wage Work

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 $8.66 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 7 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 4.7 percentage points. 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.

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 4.7 percentage points (see column 2). In the triple-difference specification the estimate is 8.2 percentage points. Declines of these magnitudes represent declines of one third to one half 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 9 presents the raw data underlying these results, and Appendix Table A.5 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 1.7 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.

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 reduced upward mobility by making the first rung on the earnings ladder more difficult for low-skilled workers to reach.

7 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. Through late 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.[32] These dimensions of the data suggest that U.S.-specific developments in low-skilled labor markets underlie a non-trivial portion of the slump in U.S. employment. We thus conclude by considering our results' implications for the relationship between the minimum wage and this period's employment

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

declines.

The framework developed in section 1 highlights why extrapolating our results into changes in national employment-to-population ratios is necessarily somewhat speculative. We have estimated local average treatment effects for the differentially binding portion of this period's minimum wage increases. The key question, as shown in equation (5), is whether these local average treatment effects are estimated on relatively thick or thin segments of low-skilled groups' productivity distributions. We emphasize two reasons why the differentially binding portion of this period's minimum wage increases likely moved through relatively thick segments of low-skilled groups' productivity distributions.

First, the final increment of this period's minimum wage increases involved deeper movement into low-skilled groups' productivity distributions than did the first and second increments. Second, the "treatment" states were states with relatively low costs of living. A given nominal value of the minimum wage will thus tend to bite more deeply into the productivity distributions of workers in these states than in high cost of living states. So long as the density of the productivity distribution ($f(a_i)$) is increasing near the distribution's lower tail, equation (5) implies that our local average treatment effects will exceed the effects associated with the remainder of this period's minimum wage increases.

On the other hand, section 1 provides a reason why our estimates will not capture the entirety of the minimum wage's effects over a period like the one we analyze. Because the minimum wage is a source of downward wage rigidity, it mediates the extent to which negative shocks to labor demand manifest themselves through employment declines rather than wage declines.[33] Equation (4) shows that our program evaluation

[33] As noted previously, both neoclassical and new-Keynesian intuitions suggest that the minimum wage would heighten the employment consequences of a given negative shock to either demand or productivity.

estimates will not detect this aspect of the minimum wage's effects. These issues inform our construction of estimates of the minimum wage's effects on aggregate employment during this time period.

Panel B of Appendix Table A.1 presents the full set of results required to infer aggregate employments effects from our SIPP estimates. The table presents estimates of equation (6) 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 just over 6 percentage points. The remaining estimates provide evidence that these minimum wage increases had little effect on the employment of other groups, which is the assumption maintained for this section's calculation.

The calculation proceeds as follows. Our in-sample estimate is that binding minimum wage increases reduced the employment of workers with average baseline wages below \$7.50 by 6.6 percentage points. The 95 percent confidence interval on this estimate extends from 2.6 to 10.6 percentage points. Applying the relevant weights, this group accounts for 7.4 percent of the U.S. population aged 16 to 64. A 6.6 percentage point decline in this group's employment thus implies a $7.4 \times 0.066 = 0.49$ percentage point decline in the employment-to-population ratio in fully bound states. Because fully bound states account for 41 percent of the full U.S. population, the purely in-sample decline in employment implies a $0.49 \times 0.41 \approx 0.20$ percentage point decline in the national, working-age employment to population ratio. The confidence interval extends from 0.08 to 0.32 percentage point.

This 0.20 percentage point decline in the national employment to population ratio is a purely in-sample estimate for the differential change in the minimum wage in fully bound states relative to partially bound states. That is, it is an estimate of the effect of a \$0.62 differential increase in the minimum wage in states that account for 41 percent of the working age population. On average across all states, the effective minimum wage rose by $1.72 over this time period. A purely linear extrapolation, both within and across states, thus implies an aggregate effect of roughly $0.20 \times \frac{1}{.41} \times \frac{1.72}{.62} \approx 1.3$ percentage point.

Our baseline extrapolation effectively splits the difference between a purely linear extrapolation and failing to extrapolate at all. On the basis of our SIPP analysis, we estimate that this period's minimum wage increases account for a 0.75 percentage point decline in the employment to population ratio among *all* individuals ages 16 to 64. A similar extrapolation based on our CPS estimates implies an aggregate employment decline of just under 0.5 percentage points.[34] A simple pooling of the SIPP and CPS estimates yields an estimated aggregate employment decline of 0.6 percentage point with a 95 percent confidence interval extending from 0.35 to 0.85 percentage points.[35] This accounts for 12 percent of the total decline in the national employment-to-population ratio from 2006 to 2012.

[34] In the CPS analysis, the baseline estimates are of the effect of a $0.62 differential increase in the minimum wage in states that, as in the SIPP, account for 41 percent of the working age population. In the analysis of individuals ages 30 and under with less than a completed high school education, which account for 8.4 percent of the working age population, the purely in sample decline implied by the point estimate of 3.7 percentage points is $3.7 \times 0.084 \times 0.41 \approx 0.13$. The purely linear extrapolation yields $0.13 \times \frac{1}{.41} \times \frac{1.72}{.62} \approx 0.86$ percentage point, and the mid-point between the in-sample estimate and the linear extrapolation is 0.49. In the analysis of all young adults aged 16 to 21, the sample accounts for 12.4 percent of the working age population and the baseline point estimate is 2.3 percentage points. The purely in sample decline implied by this estimate is $2.3 \times 0.124 \times 0.41 \approx 0.12$. The purely linear extrapolation yields $0.12 \times \frac{1}{.41} \times \frac{1.72}{.62} \approx 0.79$ percentage point, and the mid-point between the in-sample estimate and the linear extrapolation is 0.45.

[35] We note that this confidence interval does not account for the difficult to quantify uncertainty associated with extrapolation. The pooled estimate of the purely in sample decline in aggregate employment is $\frac{\beta_{CPS}+\beta_{SIPP}}{2} = \frac{0.13+0.2}{2} = 0.165$ with a standard error of $\frac{\sqrt{SE^2_{CPS}+SE^2_{SIPP}}}{2} = \frac{\sqrt{0.038^2+0.061^2}}{2} = 0.036$. This pooling of estimates is similar to that conducted by Aaronson, Agarwal, and French (2012), who obtain an even greater precision gain by accounting for each estimate's precision when weighting estimates across data sources. We arrive at the point estimate and confidence interval in the main text by scaling the purely in-sample point estimate and standard error by 3.5. As noted above, this gives us the mid-point between the purely in-sample estimate and the estimate obtained by extrapolating linearly both across and within states.

8 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 shifted as intended, the estimated effects on employment, income, and income growth are negative. Our best estimate is that this period's minimum wage increases reduced the national employment-to-population ratio by 0.6 percentage point between 2006 and 2012.

To interpret the minimum wage's effects, we develop a framework that distills the neoclassical and bargaining-centric forces that underlie its intended and unintended consequences. The framework highlights that the minimum wage's effects depend crucially on the economic factors underlying low-skilled individuals' wages. Its intended effects can be large when low wage rates reflect weaknesses in low-skilled individuals' bargaining positions. Its employment effects can be large when low wage rates reflect low demand for low-skilled individuals' output.

Paired with data on low-skilled groups' wages and productivity growth, our framework sheds light on how the minimum wage's effects can vary across settings. Wage and productivity data reveal that the minimum wage increases we analyze had much deeper and more sustained bite on low-skilled groups' wage distributions than did prior minimum wage increases. This stems from both the magnitude of the minimum wage increases and the evolving effects of trade, technology, and the housing market on demand for low-skilled labor. We conclude that future minimum wage increases should likewise be analyzed with reference to low-skilled labor demand's evolution.

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Tables and Figures

Evolution of Low-Skilled Individuals' Wage Distributions

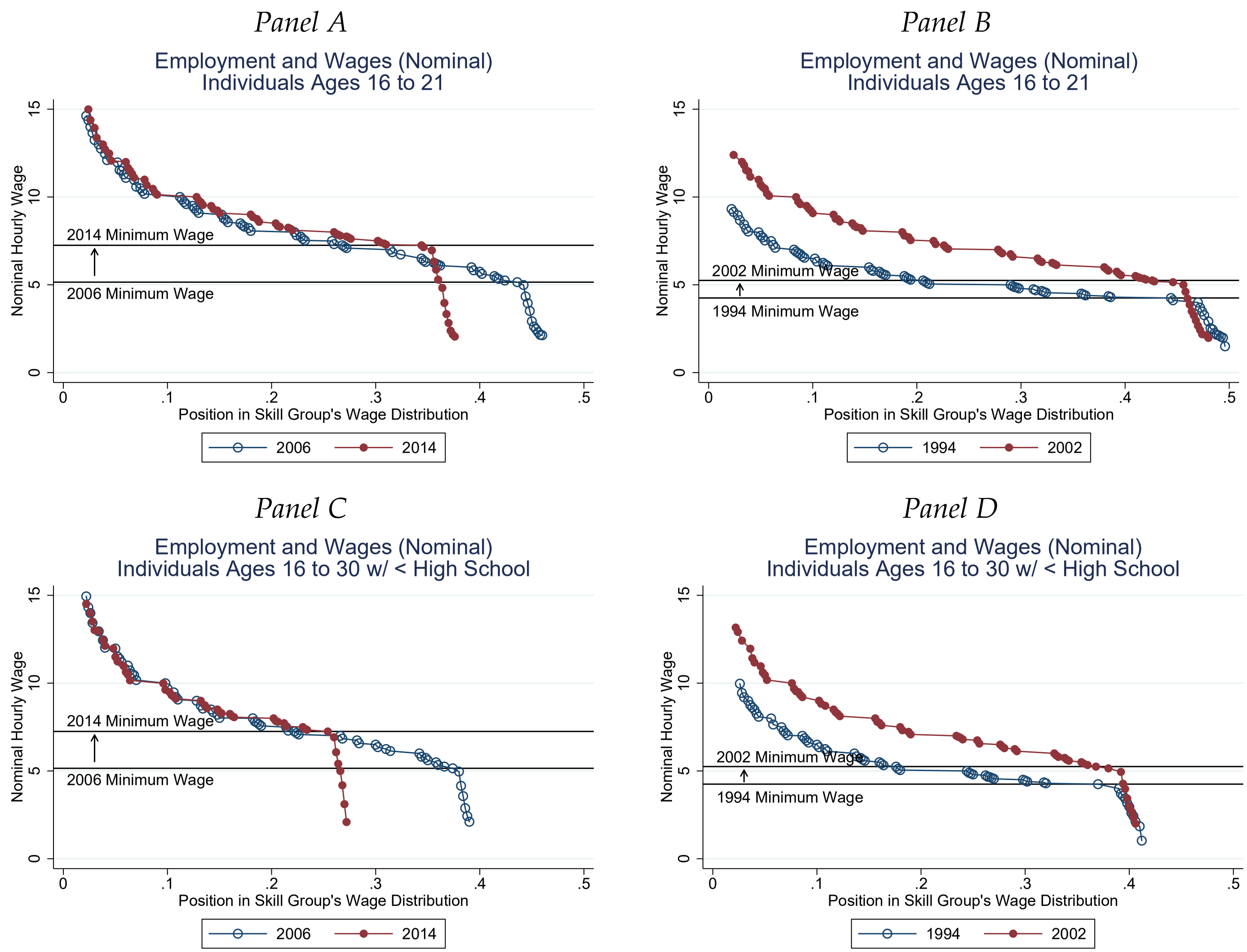

Figure 1: **Low-Skilled Individuals' Wage Distributions Surrounding Federal Minimum Wage Increases**
Note: The panels of the figure present wage distributions constructed using data from the NBER's CPS-MORG files for 1994, 2002, 2006, and 2014. The presented series describe the wage distributions of young adults ages 16 to 21 (panels A and B) and of individuals ages 30 and under with less than a completed high school education (panels C and D). The samples consist of individuals residing in states we categorize as "bound" or "fully bound" by the most recent federal minimum wage increases, as these states have historically maintained minimum wage rates quite close to the federal minimum. When available, individual-level wages are taken to be the reported values of the variable "earnhre" divided by 100. When "earnhre" is missing, individual-level wages are estimated as "earnwke/hours." Workers were sorted according to their wage rates, with unemployed individuals assigned wage rates of 0. The wage rates for each year were then divided into 500 quantiles with application of the CPS's population weights. The markers indicate all positive wage quantiles outside of the top 2 percentiles of each distribution.

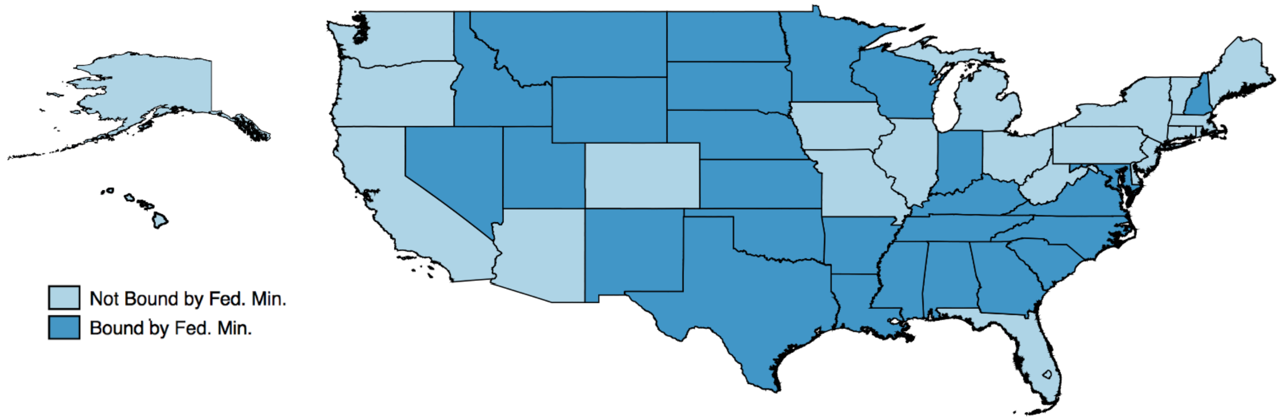

Figure 2: **States Bound by the 2008 and 2009 Federal Minimum Wage Increase:**
The map labels states on the basis of whether we characterize them as bound by the July 2008 and July 2009 increases in the federal minimum wage. We define bound states as states reported by the Bureau of Labor Statistics (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.

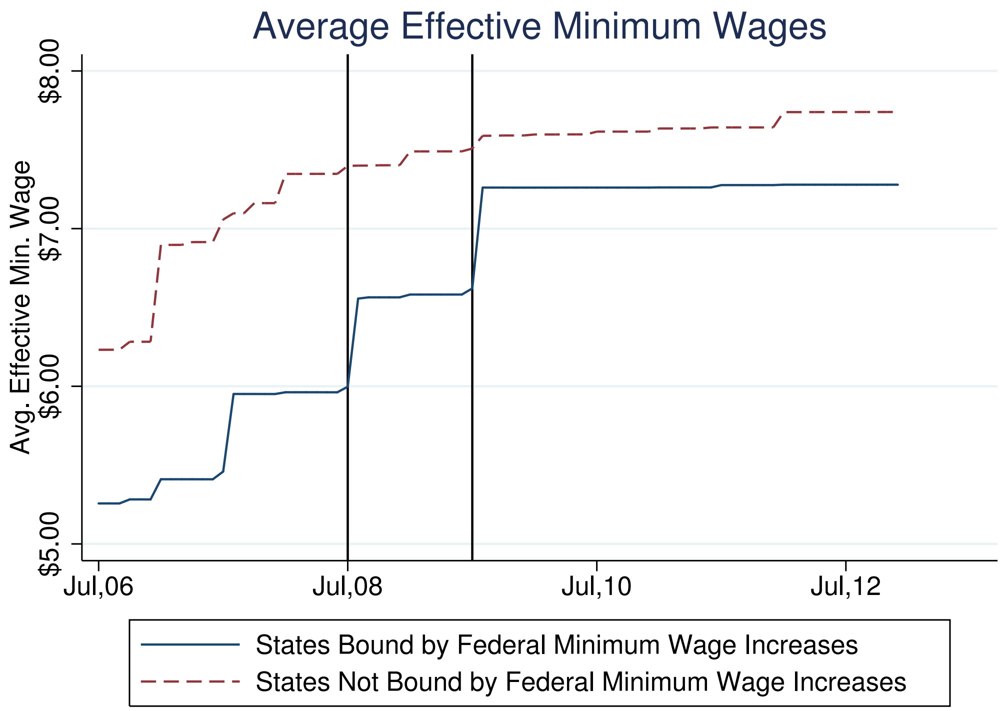

Figure 3: **Evolution of the Average Minimum Wage in Bound and Unbound States:** As in the previous figure, we define bound states as states reported by the Bureau of Labor Statistics (BLS) to have had a minimum wage less than $6.55 in January 2008. Such states were at least partially bound by the July 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 2008 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.

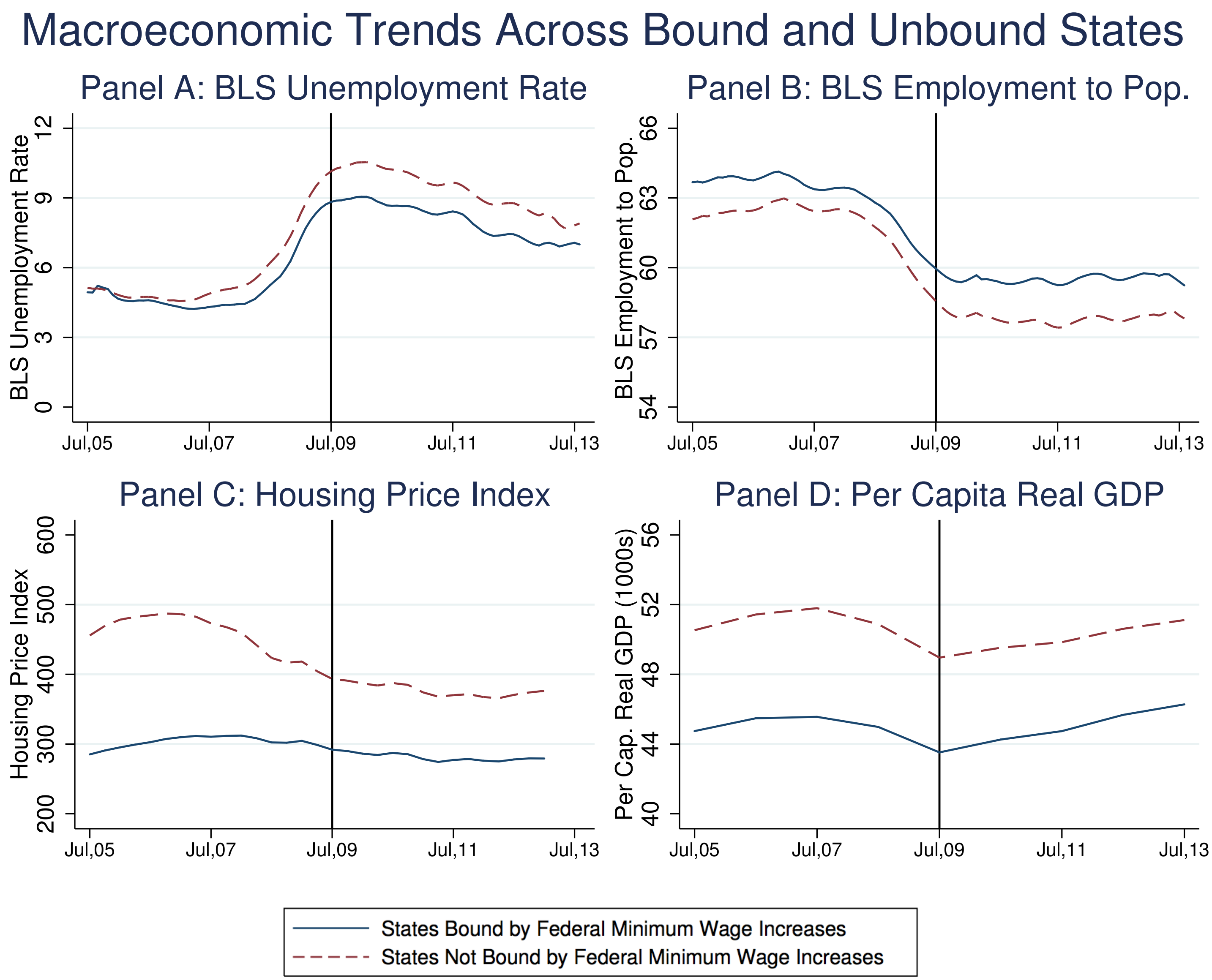

Figure 4: **Macroeconomic Trends in Bound and Unbound States:**
Bound and unbound states are defined as in previous figures. This figure's panels plot the evolution of macroeconomic indicators over the course of the housing bubble and Great Recession. All series are weighted by state population so as 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 monthtly 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.

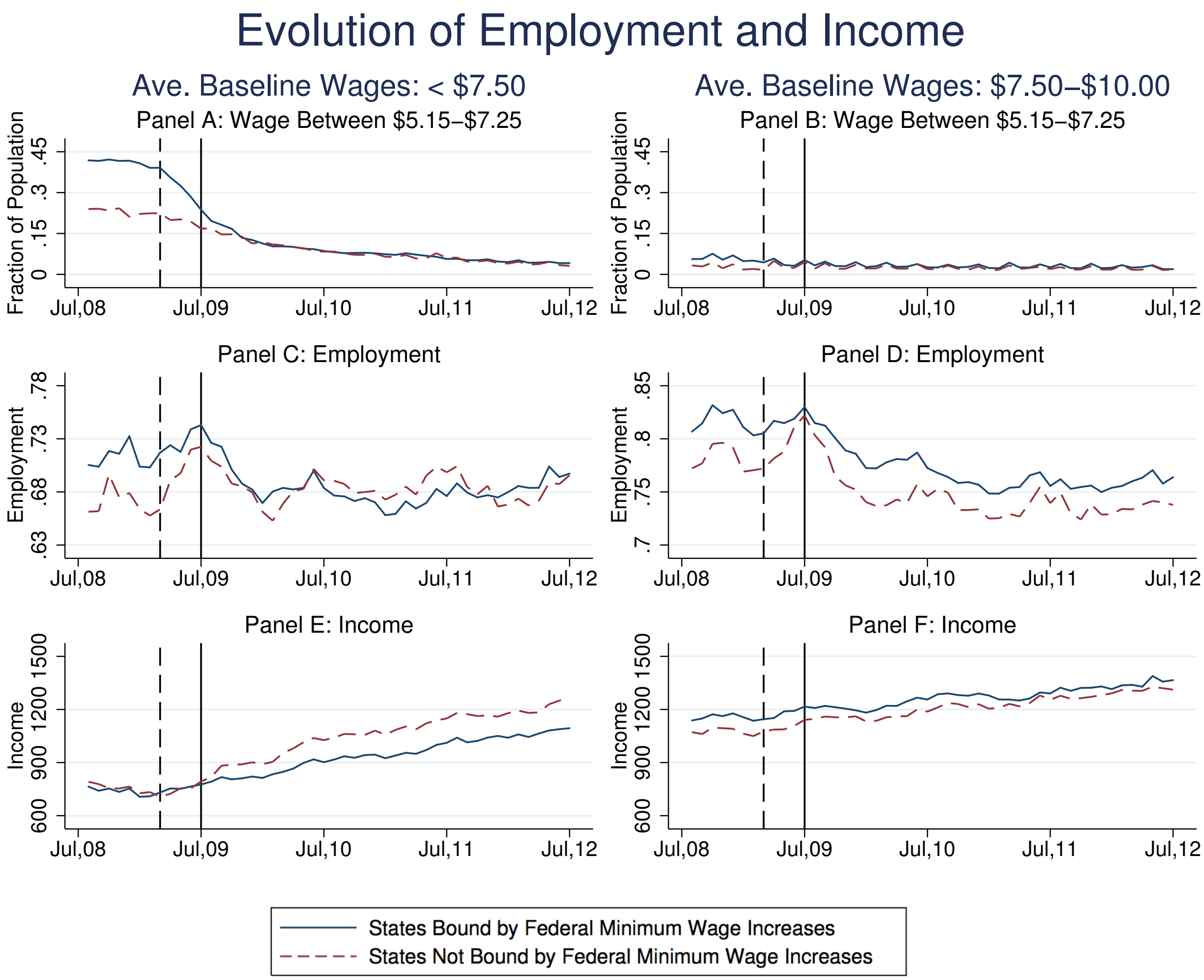

Figure 5: **Employment and Income Trends in Bound and Unbound States:** Bound and unbound states are defined as in previous figures. The figure plots the evolution of three wage, employment, and earnings related outcomes for groups of low-skilled workers. In all cases the series are constructed by the 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.

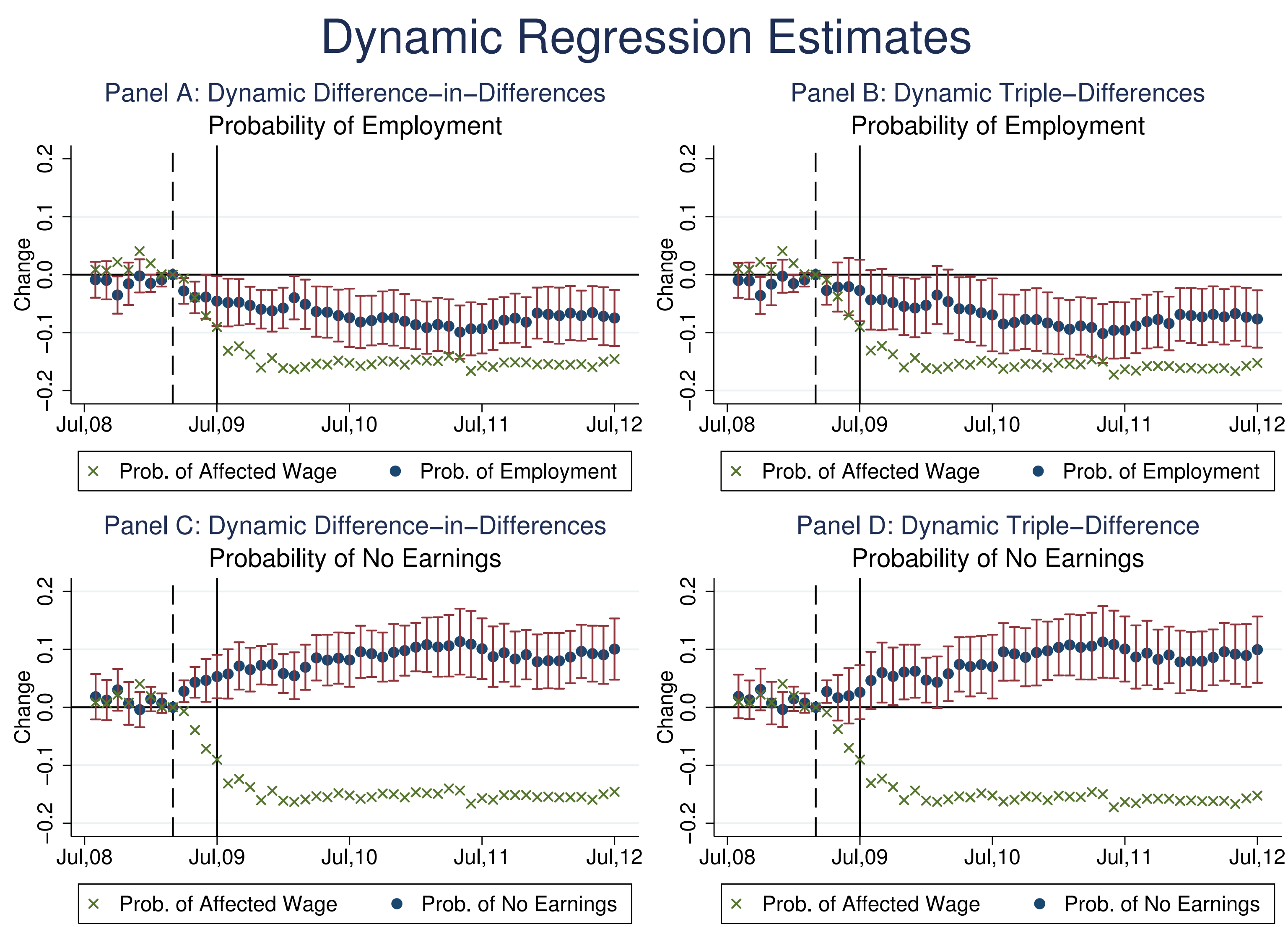

Figure 6: **Dynamic Regression Estimates:** The figure reports fully dynamic estimates of the minimum wage's effects on the wages and employment of low-skilled workers. Each marker is an estimate of a coefficient of the form $\beta_{p(t)}$ from equations (6) and (8), where the relevant $p(t)$ correspond 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 (6), while panels B and D present estimates of the triple-difference model of equation (8). 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).

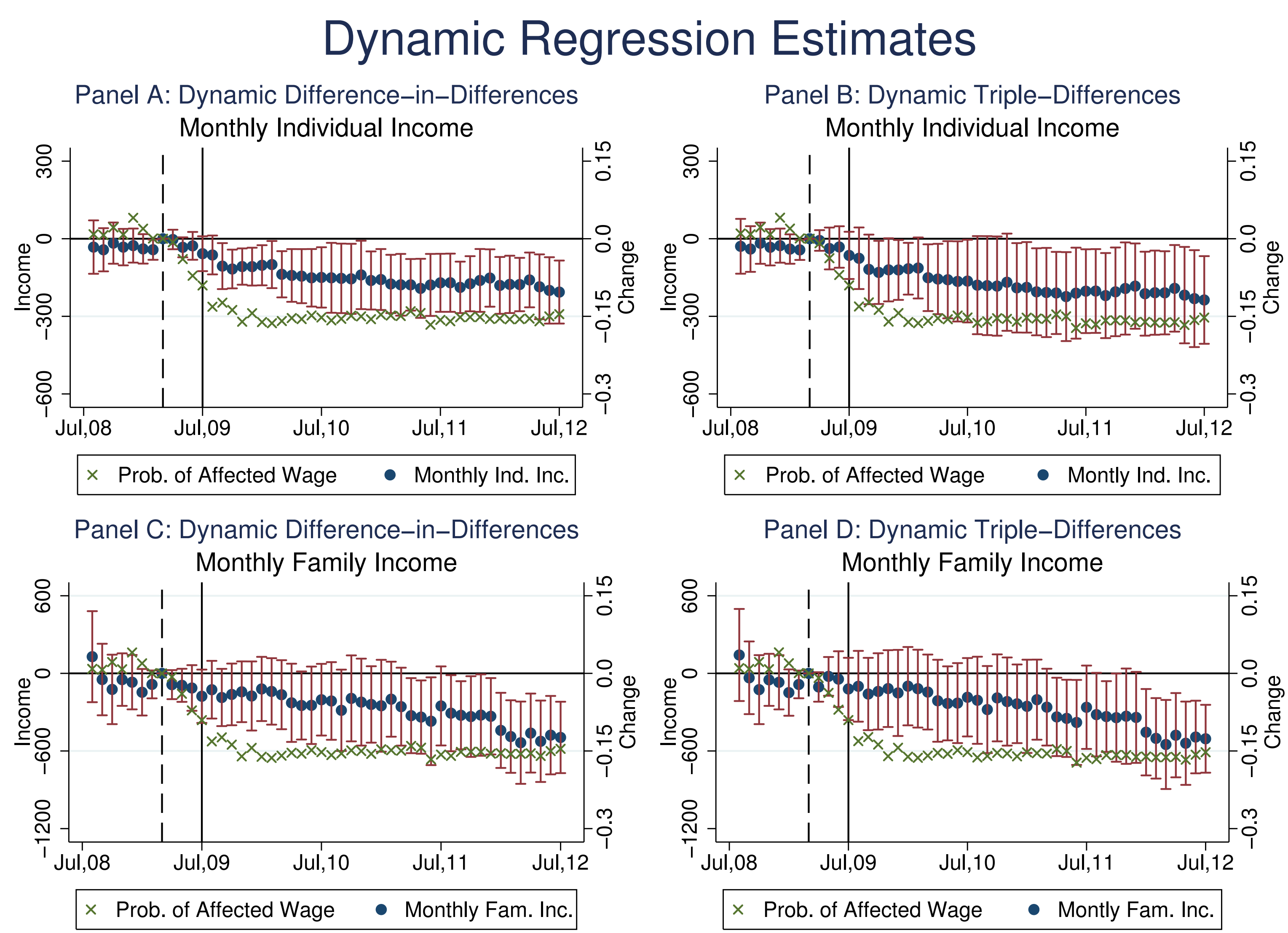

Figure 7: **Dynamic Regression Estimates:** The figure reports fully dynamic estimates of the minimum wage's effects on the wages and income of low-skilled workers. Each marker is an estimate of a coefficient of the form $\beta_{p(t)}$ from equations (6) and (8), where the relevant $p(t)$ correspond 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 (6), while panels B and D present estimates of the triple-difference model of equation (8). 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).

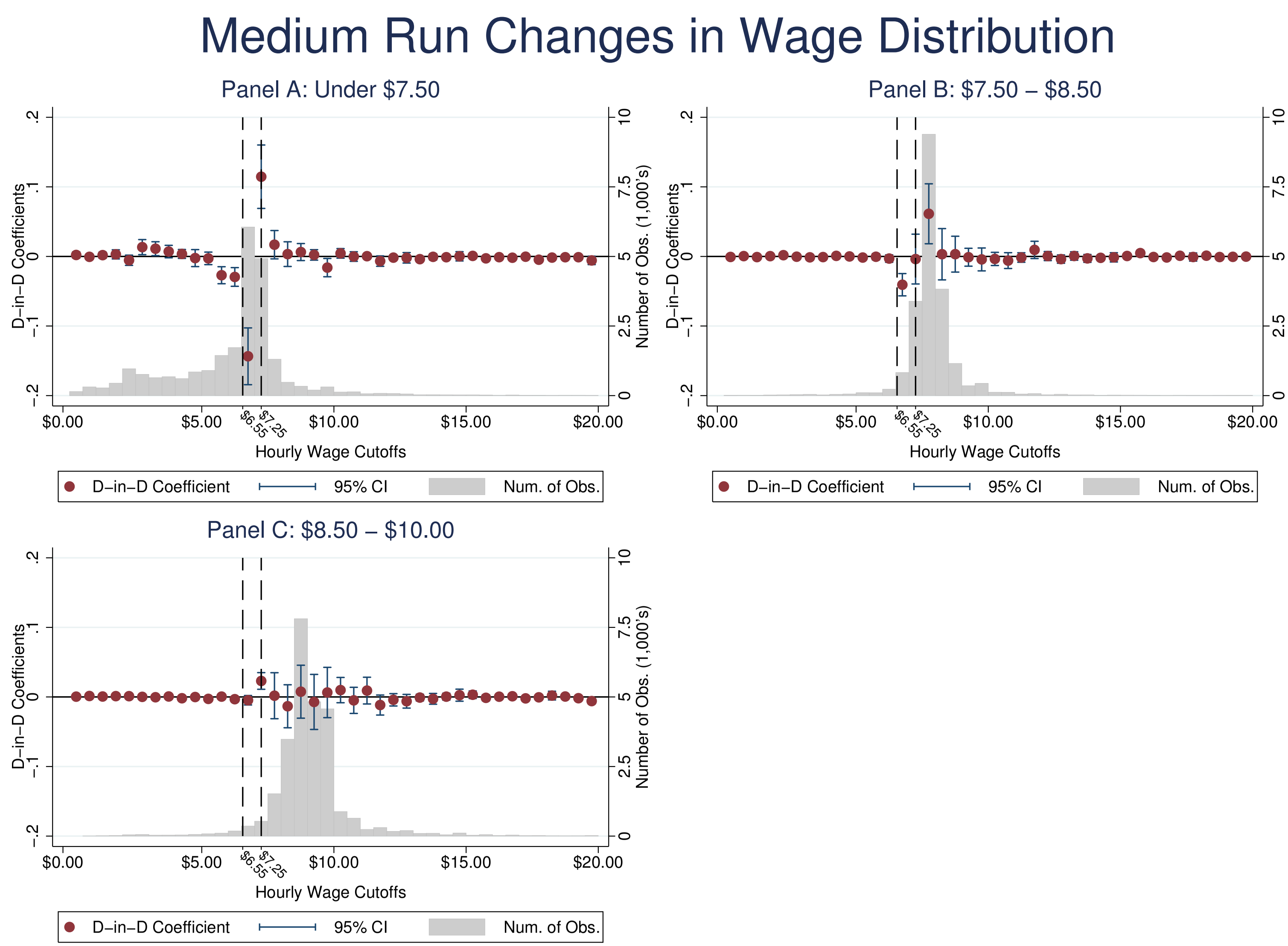

Figure 8: **Estimated Effects of the Minimum Wage on Hourly Wage Distributions:** The figure reports estimates of binding minimum wage increase's medium run effects on the wage distributions of three groups of low-skilled earners. More specifically, each dot is an estimate of the coefficient $\beta_{p(t)}$ from equation (6), 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.

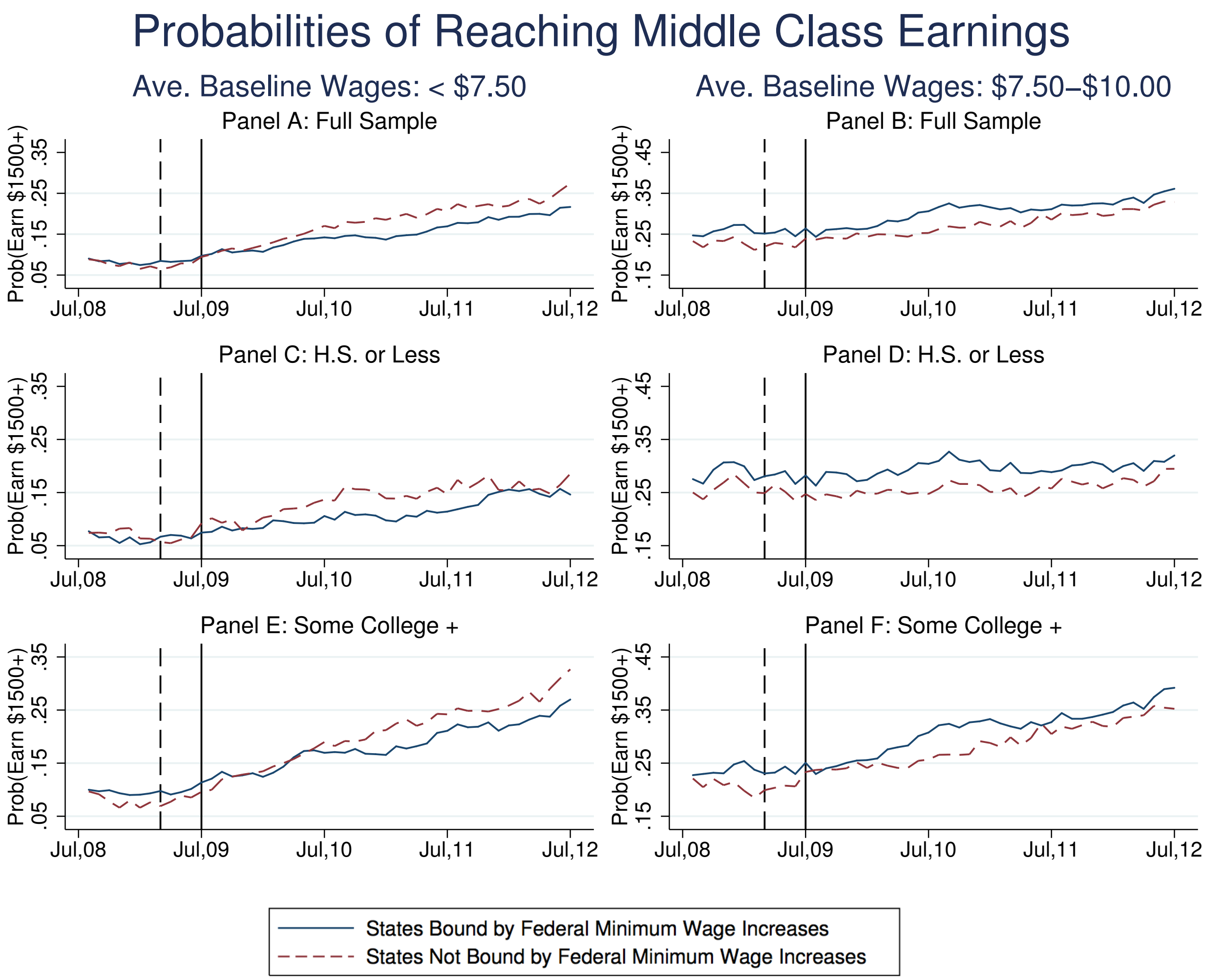

Figure 9: **Probabilities of Reaching Middle Class Earnings:**
Bound and unbound states are defined as in previous figures. In all panels, the figure plots the evolution the fraction of all in-sample individuals with earnings greater than $1500, which is equivalent to full time work at a wage of $8.66. 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 5'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
Wage \$5.15-\$7.25	0.373	0.217	0.0775	0.0402	0.0320	0.0220
	(0.484)	(0.412)	(0.267)	(0.196)	(0.176)	(0.147)
Employed	0.718	0.684	0.775	0.743	0.851	0.824
	(0.450)	(0.465)	(0.418)	(0.437)	(0.356)	(0.381)
Unpaid Work	0.110	0.142	0.0536	0.0527	0.0448	0.0492
	(0.313)	(0.349)	(0.225)	(0.223)	(0.207)	(0.216)
No Earnings	0.392	0.459	0.279	0.310	0.193	0.225
	(0.488)	(0.498)	(0.448)	(0.462)	(0.395)	(0.418)
Num hours worked/week	24.44	23.76	27.00	23.89	31.57	29.66
	(18.50)	(19.23)	(17.61)	(16.93)	(15.89)	(16.63)
Income	743.7	754.2	980.5	866.5	1317.9	1267.4
	(962.0)	(1008.1)	(911.1)	(911.6)	(968.9)	(1030.7)
Below FPL	0.294	0.256	0.217	0.237	0.177	0.170
	(0.456)	(0.436)	(0.412)	(0.425)	(0.381)	(0.376)
Age	31.58	33.02	32.51	30.30	36.24	33.65
	(13.96)	(14.56)	(13.54)	(13.47)	(13.09)	(13.31)
Num. of Children	1.091	1.015	1.053	1.055	0.921	0.920
	(1.302)	(1.281)	(1.279)	(1.247)	(1.275)	(1.187)
More than H.S. Deg.	0.564	0.628	0.572	0.569	0.584	0.589
	(0.496)	(0.483)	(0.495)	(0.495)	(0.493)	(0.492)
Same Job 6+ Months	0.489	0.486	0.545	0.517	0.614	0.568
	(0.500)	(0.500)	(0.498)	(0.500)	(0.487)	(0.495)
Emp. Entire Baseline	0.478	0.425	0.544	0.495	0.671	0.620
	(0.500)	(0.494)	(0.498)	(0.500)	(0.470)	(0.486)
Emp. Preceding Hike	0.703	0.671	0.758	0.713	0.834	0.809
	(0.457)	(0.470)	(0.428)	(0.452)	(0.373)	(0.393)
Num. of Individuals	1783	1477	1000	1262	1185	1526
Observations	20241	16857	11394	14406	13649	17526

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 2. 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	-0.160***	-0.034**	-0.008*	-0.044*	0.004	-0.008
	(0.021)	(0.010)	(0.004)	(0.019)	(0.021)	(0.012)
Bound x Post 2	-0.163***	-0.042***	-0.005	-0.066**	-0.026	-0.002
	(0.024)	(0.009)	(0.005)	(0.020)	(0.021)	(0.013)
Housing Price Index	-0.616	-0.072	0.100	0.755*	0.610	-0.335
	(0.424)	(0.161)	(0.088)	(0.323)	(0.440)	(0.371)
N	147,459	102,193	122,786	147,459	102,193	122,786
Mean of Dep. Var.	0.302	0.057	0.026	0.702	0.757	0.836
Estimation Framework	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D
Weighted	No	No	No	No	No	No
Individual Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
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 (6), 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 (6), 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 between $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	-0.044*	-0.033	-0.058*	-0.043*	-0.043*	-0.053*	-0.043*
	(0.019)	(0.021)	(0.022)	(0.019)	(0.021)	(0.022)	(0.019)
Bound x Post 2	-0.066**	-0.051*	-0.085**	-0.063**	-0.063**	-0.077**	-0.063**
	(0.020)	(0.019)	(0.029)	(0.019)	(0.019)	(0.026)	(0.021)
N	147,459	147,459	147,459	147,459	121,763	124,698	144,499
Mean of Dep. Var.	0.702	0.702	0.702	0.702	0.706	0.699	0.702
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	-0.039	-0.025	-0.038	-0.040	-0.035	-0.047	-0.041
	(0.026)	(0.028)	(0.026)	(0.026)	(0.026)	(0.029)	(0.026)
Bound x Post 2 x Target	-0.068**	-0.048*	-0.067**	-0.073**	-0.062*	-0.074**	-0.066**
	(0.022)	(0.023)	(0.022)	(0.022)	(0.024)	(0.024)	(0.022)
N	270,245	270,245	270,245	270,245	223,669	231,943	264,499
Mean of Dep. Var.	0.763	0.763	0.763	0.763	0.767	0.763	0.763
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 (6), 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 (6), 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 (8), 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 of Binding Minimum Wage Increases across Age Groups**

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable			Employed			
Bound x Post 1	-0.027**	-0.004	-0.004	-0.021*	-0.001	0.001
	(0.009)	(0.006)	(0.005)	(0.009)	(0.006)	(0.006)
Bound x Post 2	-0.023*	0.001	-0.001	-0.024*	0.007	0.006
	(0.010)	(0.006)	(0.006)	(0.010)	(0.006)	(0.007)
Housing Price Index	0.215***	0.101***	0.062*			
	(0.057)	(0.027)	(0.023)			
N	894,384	3,515,120	2,890,905	564,164	2,230,574	1,852,824
Mean of Dep. Var.	0.45	0.79	0.72	0.45	0.80	0.73
Estimation Framework	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D
State and Time Effects	Yes	Yes	Yes	Yes	Yes	Yes
Sample	Full	Full	Full	Matched	Matched	Matched
Age Group	Youngest	Middle	Oldest	Youngest	Middle	Oldest

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 (9), 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 (9), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. Across columns 1 through 3, the samples comprise all individuals ages 16 to 64 who were surveyed in the full Current Population Survey during the relevant years. The sample in column 1 consists of all individuals between ages 16 and 21. The sample in column 2 consists of individuals between ages 22 and 45. The sample in column 3 consists of individuals between ages 46 and 64. The samples in columns 4 through 6 are restricted to individuals in states with housing declines that could be matched to within $20,000. The sample extends from January 2006 through December 2012. The reported standard error estimates allow for state level correlation clusters across the estimation errors.

Table 5: **Effects of Binding Minimum Wage Increases across Skill Groups**

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable	Employed					
Bound x Post 1	-0.037**	-0.010	-0.004	-0.031*	-0.004	-0.001
	(0.011)	(0.008)	(0.005)	(0.015)	(0.010)	(0.005)
Bound x Post 2	-0.037**	-0.006	0.001	-0.037**	-0.003	0.007
	(0.011)	(0.008)	(0.005)	(0.013)	(0.009)	(0.005)
Housing Price Index	0.280*	0.195***	0.063**			
	(0.107)	(0.035)	(0.020)			
N	580,248	1,814,333	4,905,828	358,272	1,143,767	3,145,523
Mean of Dep. Var.	0.38	0.72	0.76	0.38	0.72	0.77
Estimation Framework	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D
State and Time Effects	Yes	Yes	Yes	Yes	Yes	Yes
Sample	Full	Full	Full	Matched	Matched	Matched
Skill Group	Lowest	Middle	Highest	Lowest	Middle	Highest

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 (9), 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 (9), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. Across columns 1 through 3, the samples comprise all individuals ages 16 to 64 who were surveyed in the full Current Population Survey during the relevant years. The sample in column 1 consists of all individuals between ages 16 and 30 with less than a high school education. The sample in column 2 consists of individuals between ages 31 and 45 with less than a high school education and between ages 16 and 30 with a high school education or more. The sample in column 3 includes all remaining individuals between ages 16 and 64. The samples in columns 4 through 6 are restricted to individuals in states with housing declines that could be matched to within 5 log points. The sample extends from January 2006 through December 2012. The reported standard error estimates allow for state level correlation clusters across the estimation errors.

Table 6: **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	-0.044*	0.015+	0.059**	-92.087*	-117.957	0.013
	(0.019)	(0.009)	(0.019)	(36.474)	(85.300)	(0.012)
Bound x Post 2	-0.066**	0.017+	0.082***	-144.042**	-273.063*	0.017
	(0.020)	(0.010)	(0.021)	(44.748)	(119.576)	(0.012)
N	147,459	147,459	147,459	147,459	147,459	147,459
Mean of Dep. Var.	0.702	0.125	0.422	748.459	4,190.870	0.277
Panel B:	Triple Difference Specifications					
Bound x Post 1 x Target	-0.039	0.008	0.047*	-105.029+	-94.301	0.008
	(0.026)	(0.011)	(0.023)	(55.231)	(112.504)	(0.017)
Bound x Post 2 x Target	-0.068**	0.014	0.082***	-174.255*	-276.503*	0.015
	(0.022)	(0.012)	(0.022)	(73.107)	(133.386)	(0.018)
N	270,245	270,245	270,245	270,245	270,245	270,245
Mean of Dep. Var.	0.763	0.089	0.326	995.515	4,249.402	0.230

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 (6), 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 (6), 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 (8), 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 7: **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	-0.021	-0.012	-0.004	-0.007	-0.002
	(0.011)	(0.015)	(0.012)	(0.007)	(0.008)	(0.009)
Bound x Post 2	-0.047***	-0.047**	-0.042*	-0.017+	-0.009	-0.020
	(0.013)	(0.015)	(0.017)	(0.009)	(0.009)	(0.013)
N	147,459	60,507	86,952	147,459	60,507	86,952
Mean of Dep. Var.	0.206	0.154	0.237	0.068	0.032	0.089
Panel B:			Triple Difference Specifications			
Bound x Post 1 x Target	-0.003	-0.034	0.016	-0.013	-0.023+	-0.007
	(0.016)	(0.026)	(0.022)	(0.011)	(0.012)	(0.014)
Bound x Post 2 x Target	-0.054*	-0.081**	-0.038	-0.015	-0.010	-0.019
	(0.024)	(0.029)	(0.030)	(0.015)	(0.016)	(0.018)
N	270,245	111,341	158,904	270,245	111,341	158,904
Mean of Dep. Var.	0.206	0.154	0.237	0.068	0.032	0.089
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 (6), 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 (6), 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 (8), 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 Materials
(Intended For Online Publication Only)

Appendix 1: Supplemental Tables and Figures

Dynamic Employment Estimates across Age and Skill Groups

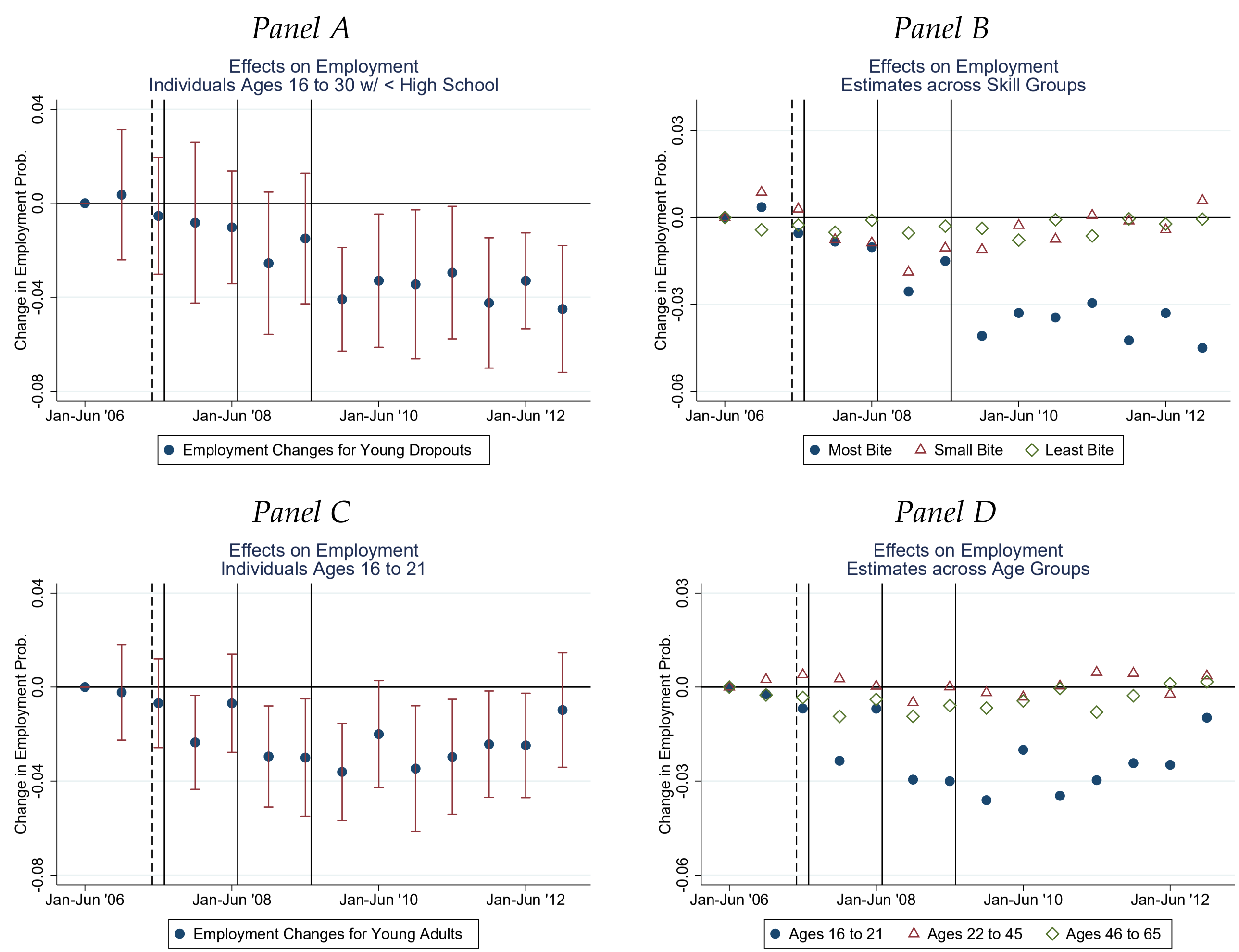

Figure A.1: **Dynamic Regression Estimates:** The figure reports dynamic estimates of differentially binding minimum wage increases' effects on the employment of low-skilled workers relative to higher skilled workers. Each marker in the panels is an estimate of a coefficient of the form $\beta_{p(t)}$ from equation (9), where each time period $p(t)$ is a 6 month interval and period $p = 0$, relative to which subsequent changes are estimated, extends from January to June of 2006. In Panel B, the "Most Bite," "Small Bite," and "Least Bite" division of the sample is the same as that in the main text's analysis of variation in the minimum wage's effects across skill groups. The dashed vertical line indicates the May 2007 passage of the federal minimum wage increases, while the solid vertical lines indicate the implementation of the July 2007, July 2008, and July 2009 increases in the federal minimum wage.

Table A.1: **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	-0.160***	-0.034**	-0.008*	-0.000
	(0.001)	(0.021)	(0.010)	(0.004)	(0.001)
Bound x Post 2	0.001	-0.163***	-0.042***	-0.005	0.000
	(0.001)	(0.024)	(0.009)	(0.005)	(0.001)
N	523,086	147,459	102,193	122,786	1,076,148
Mean of Dep. Var.	0.000	0.302	0.057	0.026	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	-0.009+	-0.044*	0.004	-0.008	0.002
	(0.005)	(0.019)	(0.021)	(0.012)	(0.004)
Bound x Post 2	-0.022*	-0.066**	-0.026	-0.002	-0.003
	(0.009)	(0.020)	(0.021)	(0.013)	(0.005)
N	523,086	147,459	102,193	122,786	1,076,148
Mean of Dep. Var.	0.241	0.702	0.757	0.836	0.936
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-$8.50	$8.50-$10.00	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 (6), where the relevant $p(t)$ corresponds with the period beginning in July 2009 and extending through July 2010. The estimates in row 2 are of the coefficient $\beta_{p(t)}$ from equation (6), where the relevant $p(t)$ corresponds with the period beginning one year after the July 2009 increase in the federal minimum wage. 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. Standard errors are clustered at the state level.

Table A.2: **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+	0.015+	0.014	0.014	0.019*	0.017	0.014
	(0.009)	(0.009)	(0.011)	(0.008)	(0.008)	(0.010)	(0.009)
Bound x Post 2	0.017+	0.018+	0.012	0.012	0.020*	0.014	0.015
	(0.010)	(0.010)	(0.016)	(0.009)	(0.009)	(0.011)	(0.010)
N	147,459	147,459	147,459	147,459	121,763	124,698	144,499
Mean of Dep. Var.	0.125	0.125	0.125	0.125	0.125	0.121	0.126
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	0.008	0.009	0.008	0.007	0.008	0.014	0.007
	(0.011)	(0.011)	(0.011)	(0.011)	(0.011)	(0.012)	(0.011)
Bound x Post 2 x Target	0.014	0.014	0.014	0.010	0.012	0.015	0.011
	(0.012)	(0.013)	(0.012)	(0.012)	(0.013)	(0.015)	(0.012)
N	270,245	270,245	270,245	270,245	249,365	254,704	267,459
Mean of Dep. Var.	0.089	0.089	0.089	0.089	0.093	0.092	0.090
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 (6), 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 (6), 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 (8), 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.3: **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	0.059**	0.049*	0.071***	0.056**	0.062**	0.070**	0.057**
	(0.019)	(0.020)	(0.019)	(0.019)	(0.020)	(0.020)	(0.019)
Bound x Post 2	0.082***	0.068**	0.098***	0.076***	0.083***	0.091**	0.078***
	(0.021)	(0.021)	(0.025)	(0.021)	(0.018)	(0.027)	(0.021)
N	147,459	147,459	147,459	147,459	121,763	124,698	144,499
Mean of Dep. Var.	0.422	0.422	0.422	0.422	0.419	0.422	0.424
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*	0.034	0.047*	0.047+	0.043+	0.060*	0.049*
	(0.023)	(0.026)	(0.023)	(0.024)	(0.023)	(0.025)	(0.023)
Bound x Post 2 x Target	0.082***	0.062*	0.081***	0.083**	0.079**	0.090***	0.077**
	(0.022)	(0.024)	(0.022)	(0.024)	(0.023)	(0.025)	(0.023)
N	270,245	270,245	270,245	270,245	249,365	254,704	267,459
Mean of Dep. Var.	0.326	0.326	0.326	0.326	0.334	0.333	0.327
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 (6), 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 (6), 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 (8), 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.4: **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	-92.087*	-81.512*	-81.505	-83.528*	-95.688*	-132.883**	-88.845*
	(36.474)	(36.007)	(51.270)	(37.526)	(37.780)	(42.963)	(37.056)
Bound x Post 2	-144.042**	-128.897**	-103.013	-116.174*	-149.823***	-189.034**	-135.469**
	(44.748)	(42.904)	(81.985)	(45.446)	(36.653)	(54.550)	(45.570)
N	147,459	147,459	147,459	147,459	121,763	124,698	144,499
Mean of Dep. Var.	748.459	748.459	748.459	748.459	755.041	753.408	744.122
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	-105.029+	-82.630	-105.980+	-107.071+	-111.376*	-123.873+	-116.252*
	(55.231)	(56.567)	(55.513)	(56.414)	(54.396)	(64.723)	(55.462)
Bound x Post 2 x Target	-174.255*	-141.453+	-175.021*	-175.867*	-193.630*	-211.826*	-176.081*
	(73.107)	(71.392)	(73.827)	(72.953)	(79.356)	(85.291)	(74.685)
N	270,245	270,245	270,245	270,245	249,365	254,704	267,459
Mean of Dep. Var.	995.515	995.515	995.515	995.515	971.849	977.143	991.559
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 (6), 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 (6), 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 (8), 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.5: **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	-0.013	-0.007	-0.014	-0.011	-0.022	-0.017
	(0.011)	(0.011)	(0.017)	(0.011)	(0.012)	(0.013)	(0.011)
Bound x Post 2	-0.047***	-0.043**	-0.021	-0.041**	-0.051***	-0.053**	-0.046***
	(0.013)	(0.013)	(0.025)	(0.013)	(0.014)	(0.015)	(0.013)
N	147,459	147,459	147,459	147,459	121,763	124,698	144,499
Mean of Dep. Var.	0.206	0.206	0.206	0.206	0.203	0.222	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	-0.003	0.001	-0.003	-0.007	-0.024	-0.004	-0.009
	(0.016)	(0.015)	(0.017)	(0.017)	(0.018)	(0.019)	(0.017)
Bound x Post 2 x Target	-0.054*	-0.048*	-0.054*	-0.061**	-0.073**	-0.048	-0.059*
	(0.024)	(0.021)	(0.024)	(0.023)	(0.027)	(0.030)	(0.024)
N	270,245	270,245	270,245	270,245	249,365	254,704	267,459
Mean of Dep. Var.	0.282	0.282	0.282	0.282	0.277	0.272	0.282
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 (6), 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 (6), 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 (8), 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.6: **Intensive Margin Effects**

Dependent Variable	(1) Employed	(2) Hours	(3) Full Time	(4) Part Time	(5) Own Business	(6) School/ Training
Panel A:	Difference-in-Differences Specifications					
Bound x Post 1	-0.044*	-0.696	-0.006	-0.038+	0.004*	-0.017
	(0.019)	(0.596)	(0.013)	(0.022)	(0.002)	(0.012)
Bound x Post 2	-0.066**	-1.629*	-0.025	-0.041	0.002	0.001
	(0.020)	(0.786)	(0.020)	(0.025)	(0.001)	(0.016)
N	147,459	147,459	147,459	147,459	147,459	147,459
Mean of Dep. Var.	0.702	24.129	0.268	0.434	0.003	0.311
Panel B:	Triple Difference Specifications					
Bound x Post 1 x Target	-0.039	-0.515	-0.008	-0.031	0.006**	-0.031*
	(0.026)	(0.844)	(0.020)	(0.031)	(0.002)	(0.014)
Bound x Post 2 x Target	-0.068**	-1.888*	-0.031	-0.037	0.003+	-0.023
	(0.022)	(0.914)	(0.024)	(0.029)	(0.002)	(0.016)
N	270,245	270,245	270,245	270,245	270,245	270,245
Mean of Dep. Var.	0.763	27.037	0.344	0.419	0.003	0.261

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 (6), 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 (6), 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 (8), 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 2: Further Analysis of the Minimum Wage's Effect on Employment

This appendix presents further analysis of the minimum wage's effect on employment. We begin with a presentation of further analysis of our baseline result's robustness, with emphasis on the potential relevance of alternative strategies for controlling for heterogeneity in macroeconomic conditions. We next present estimates in which we replace our treatment group with groups selected using the demographic and industrial proxies used regularly in the literature. The latter analysis facilitates a comparison of our approach with alternative research designs.

Appendix 2.1: Further Checks on the Robustness of Our Baseline Estimates

Appendix Table A.7 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 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 eight and a half percentage point declines in the employment of low-skilled workers.

Appendix Table A.8 presents estimates in which we restrict our SIPP analysis samples to states that could be matched on the basis of their housing declines. The matching criterion is the same as that applied to the samples in columns 4 through 6 of tables 4 and 5, which presented our CPS analysis. The specifications are the same set of specifications presented in table 3's robustness analysis. The estimates in table A.8 and 3 are quite similar, providing evidence that our SIPP analysis is robust to the adoption of our matching framework.

Appendix 2.2: 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 (6) 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. Appendix Figure A.2 and Table A.12 characterize the bite of binding minimum wage increases on the wage distributions of groups of workers that are commonly analyzed in the literature. Figure A.2's Panels A and B display the wage distributions of teenagers and food service workers. As summarized in Table A.12, 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 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.

The histograms in Figures 8 and A.2 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 A.2's panel C, comparison samples drawn based on industries will tend to contain many much higher skilled, and thus less directly comparable, individuals. Figure A.2'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.

Column 5 of Appendix Table A.12 reports our estimate that binding minimum wage increases reduced teenager and food service workers' medium-run employment by 4 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). Tables A.13 and A.14 present similar analyses of the probability of working without pay and having no earnings.

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.[36] We estimate that the wage distribution of our target sample was 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 A.15 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 A.15 we present estimates of the minimum wage's effect on the probability that any given individual is employed in the food service sector.

[36]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 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.

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.[37]

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 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 A.15 suggest that, during the period we study, food-service employers had significant scope for substituting between low- and high-skilled workers.

Second, the results in Table A.15 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, substitution between low and high-skilled workers 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.

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

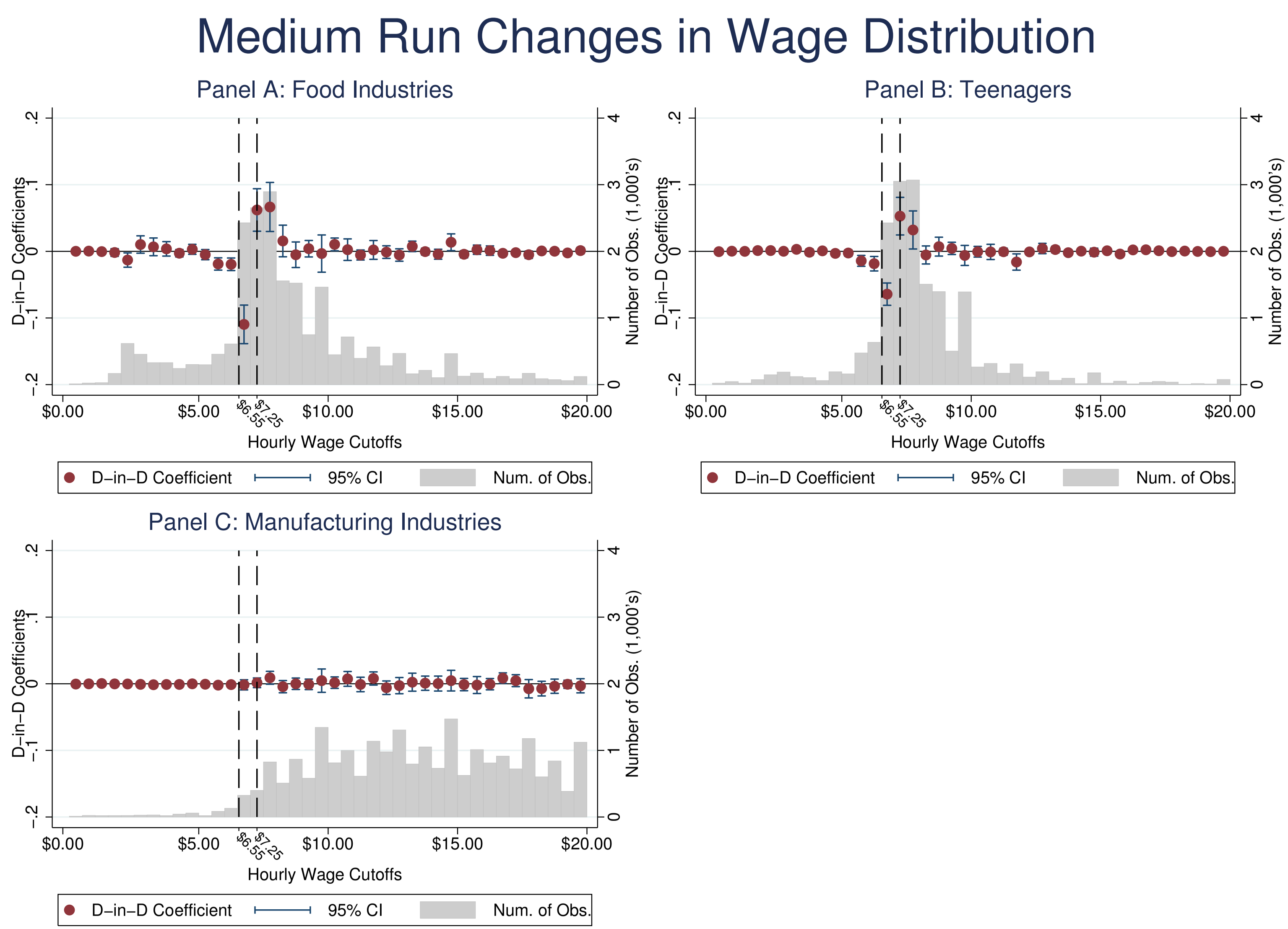

Figure A.2: **Estimated Effects of the Minimum Wage on Hourly Wage Distributions:** The figure reports estimates of binding minimum wage increase's medium run effects on the wage distributions of three groups of low-skilled earners. More specifically, each dot is an estimate of the coefficient $\beta_p(t)$ from equation (6), 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.7: **Further Robustness of the Estimated Employment Effects**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent Variable				Employed			
Bound x Post 1	-0.044*	-0.033	-0.034	-0.041+	-0.049*	-0.059**	-0.059*
	(0.019)	(0.021)	(0.021)	(0.024)	(0.021)	(0.022)	(0.023)
Bound x Post 2	-0.066**	-0.051*	-0.053**	-0.059*	-0.068**	-0.086**	-0.085**
	(0.020)	(0.019)	(0.020)	(0.022)	(0.021)	(0.028)	(0.029)
Housing Price Index	0.755*				0.652+	1.749*	1.785**
	(0.323)				(0.355)	(0.724)	(0.636)
State Employment Rate			0.489	0.401	0.121		0.700
			(0.706)	(0.739)	(0.757)		(0.859)
State Inc. Per Cap. (1000s)			0.004	0.004	0.002		0.005
			(0.004)	(0.004)	(0.003)		(0.004)
Stimulus Per Cap. (1000s)				-0.033	-0.032		0.010
				(0.037)	(0.037)		(0.034)
Predicted State Income				0.499	0.228		-0.774
				(0.529)	(0.536)		(1.426)
Predicted State Employment				-0.597	-0.759		-0.709
				(0.707)	(0.680)		(0.770)
N	147,459	147,459	147,459	147,459	147,459	146,256	146,256
Mean of Dep. Var.	0.702	0.702	0.702	0.702	0.702	0.702	0.702
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 (6), 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 (6), 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.8: **Further Robustness of the Estimated Employment Effects: Matched Sample**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent Variable				Employed			
Panel A:				Difference-in-Differences Specifications			
Bound x Post 1	-0.064**	-0.061**	-0.068**	-0.065***	-0.067**	-0.054*	-0.065***
	(0.018)	(0.018)	(0.023)	(0.017)	(0.021)	(0.020)	(0.018)
Bound x Post 2	-0.073**	-0.069**	-0.069+	-0.072***	-0.080***	-0.068*	-0.075**
	(0.021)	(0.020)	(0.035)	(0.020)	(0.020)	(0.026)	(0.021)
N	98,932	98,932	98,932	98,932	81,425	90,925	98,664
Mean of Dep. Var.	0.705	0.705	0.705	0.705	0.708	0.707	0.705
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	-0.070**	-0.066*	-0.069*	-0.072**	-0.070*	-0.054+	-0.070**
	(0.025)	(0.025)	(0.025)	(0.026)	(0.028)	(0.027)	(0.025)
Bound x Post 2 x Target	-0.082**	-0.078**	-0.081**	-0.089**	-0.087**	-0.070*	-0.084**
	(0.024)	(0.023)	(0.024)	(0.025)	(0.028)	(0.028)	(0.024)
N	170,020	170,020	170,020	170,020	140,833	157,766	169,400
Mean of Dep. Var.	0.764	0.764	0.764	0.764	0.770	0.767	0.764
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. 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 (6), 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 (6), 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). All samples are restricted to individuals in states with housing declines that could be matched to within $20,000. Standard errors are clustered at the state level.

Table A.9: **Summary Statistics on the Magnitudes of Declines in the FHFA Median House Price Index**

	(1) Unbound States	(2) Bound States
	Median House Price Decline (Millions of Dollars)	
Full Sample	0.0720	0.0247
	(0.0666)	(0.0538)
Matched within 5K	0.0297	0.0274
	(0.0403)	(0.0412)
Matched within 10K	0.0332	0.0297
	(0.0397)	(0.0429)
Matched within 20K	0.0385	0.0272
	(0.0371)	(0.0423)
Matched within 5 percent	0.0369	0.0241
	(0.0379)	(0.0394)
No. of States (Full Sample)	24	27

Note: This table reports summary statistics on the magnitudes of declines in states' all-transactions FHFA median house prices indices. Changes are calculated as the average in 2006 minus the average in 2012. Row 1 reports the means of these changes for the full samples of fully and partially bound states. Subsequent rows report means for samples of states that have been restricted based on the quality of nearest neighbor matches. For row 2 the criterion was that the nearest neighbor match deliver a match for which the difference in the states' housing declines was less than $5,000. Row 3 is similar, with a threshold of $10,000, row 4 with a threshold of $20,000, and row 5 with a threshold of 5 log points.

Table A.10: **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	-0.148***	-0.037*	-0.002	-0.061**	-0.006	0.008
	(0.028)	(0.014)	(0.005)	(0.018)	(0.022)	(0.017)
Bound x Post 2	-0.143***	-0.048***	0.000	-0.069**	-0.048*	0.008
	(0.028)	(0.011)	(0.006)	(0.020)	(0.020)	(0.014)
Housing Price Index						
N	98,932	64,848	71,088	98,932	64,848	71,088
Mean of Dep. Var.	0.315	0.065	0.031	0.705	0.771	0.848
Estimation Framework	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D
Weighted	No	No	No	No	No	No
Individual Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
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 (6), 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 (6), 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 between \$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. All samples are restricted to individuals in states with housing declines that could be matched to within \$20,000. Standard errors are clustered at the state level.

Table A.11: **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	-0.143***	-0.035*	-0.004	-0.064**	-0.005	0.009
	(0.025)	(0.014)	(0.005)	(0.018)	(0.023)	(0.017)
Bound x Post 2	-0.137***	-0.045***	-0.002	-0.073**	-0.047*	0.009
	(0.025)	(0.010)	(0.006)	(0.021)	(0.022)	(0.014)
Housing Price Index	-0.861	-0.364+	0.303+	0.511	-0.080	-0.136
	(0.523)	(0.190)	(0.166)	(0.386)	(0.522)	(0.404)
N	98,932	64,848	71,088	98,932	64,848	71,088
Mean of Dep. Var.	0.315	0.065	0.031	0.705	0.771	0.848
Estimation Framework	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D	D-in-D
Weighted	No	No	No	No	No	No
Individual Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
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 (6), 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 (6), 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 between $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. All samples are restricted to individuals in states with housing declines that could be matched to within $20,000. Standard errors are clustered at the state level.

Table A.12: **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	-0.160***	-0.082***	-0.005	-0.044*	-0.024+	0.004
	(0.021)	(0.014)	(0.004)	(0.019)	(0.013)	(0.009)
Bound x Post 2	-0.163***	-0.089***	-0.007	-0.066**	-0.044***	0.005
	(0.024)	(0.016)	(0.005)	(0.020)	(0.012)	(0.011)
N	147,459	275,130	166,662	147,459	275,130	166,662
Mean of Dep. Var.	0.302	0.103	0.014	0.702	0.514	0.917
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 (6), 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 (6), 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 between \$5.15 and \$7.25. In columns 4-6, the dependent variable is an indicator for whether an individual was 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.13: **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	-0.160***	-0.082***	-0.005	0.015+	0.002	0.002
	(0.021)	(0.014)	(0.004)	(0.009)	(0.004)	(0.005)
Bound x Post 2	-0.163***	-0.089***	-0.007	0.017+	0.002	-0.004
	(0.024)	(0.016)	(0.005)	(0.010)	(0.004)	(0.005)
N	147,459	275,130	166,662	147,459	275,130	166,662
Mean of Dep. Var.	0.302	0.103	0.014	0.125	0.039	0.025
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 (6), 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 (6), 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 between \$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.14: **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	-0.160***	-0.082***	-0.005	0.059**	0.026*	-0.001
	(0.021)	(0.014)	(0.004)	(0.019)	(0.012)	(0.010)
Bound x Post 2	-0.163***	-0.089***	-0.007	0.082***	0.046***	-0.010
	(0.024)	(0.016)	(0.005)	(0.021)	(0.012)	(0.012)
N	147,459	275,130	166,662	147,459	275,130	166,662
Mean of Dep. Var.	0.302	0.103	0.014	0.422	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	Manufac-turing	Under $7.50	Food Service and Teens	Manufac-turing

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 (6), 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 (6), 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 between $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.15: **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	1,971,672	147,459	1,824,213
Mean of Dep. Var.	0.047	0.216	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 Population	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 probability of working in the food service sector. More specifically, the estimates in row 1 are of the coefficient $\beta_{p(t)}$ from equation (6), 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 (6), 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.