Low-wage Work across U.S. Cities: A Multilevel Analysis

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*Abstract* Mobilization by low-wage workers across urban America has yielded changes in local labor market regulation with significant implications for low-wage workers. Our study adds to the growing body of empirical evidence documenting these effects. In particular, we find that workers living in metropolitan regions with higher local minimum wage policies and higher union densities are at a lower risk of low-wage employment than their counterparts in other cities, controlling for individual attributes, occupation, and education. A $1.00 increase in the metropolitan minimum wage tends to decrease the likelihood of low-wage employment for statistically comparable workers by 7%. We model union density as a random effect (varying slope) and find that the size of the negative effect of union density varies across metropolitan labor markets. Union density may matter more for low-wage workers in places where unions are organizing in low-wage service sectors, such as Los Angeles, Honolulu, and Houston.

*Keywords:* low-wage work, wages, unionization, local labor markets, labor market institutions, metropolitan inequality

As U.S. cities slowly emerge from the aftermath of the Great Recession, the prospects of working people in urban America look grim. The vagaries of a jobless recovery have generated economic growth and high returns for Wall Street while urban neighborhoods remain devastated by foreclosure and high unemployment (Mishel, Bivens, Gould, and Shierholz, 2012; Wyly and Ponder, 2011). Weak job growth has kept employment below pre-recessionary levels with middle-class jobs slowest to rebound, and employment gains have been concentrated in low-wage work (Bernhardt, 2012). Dubbed the "low-wage recovery," new opportunities for middle-class employment remain scarce, reinforcing the pattern of growing economic inequality and downgraded work (Doussard, 2013).

Although pundits and scholars alike often describe low-wage work as a feature of the new economy—and now of the new post-recessionary economy—surprisingly little research examines the urban patterning of low-wage work or the geographic variation of the low-wage workforce. Studies have been primarily case study in nature, usually examining jobs within a particular industry or segment, such as day labor (Appelbaum, Bernhardt, and Murnane, 2003; Melendez, Visser, Theodore, and Valenzuela, 2014). References to the relationship between city structure and low-wage employment often are embedded within studies that examine wage or income inequality. These analyses shed important light on the nature of the changing labor market that frame the ontological category of low-wage work, yet the political claimsmaking around low-wage work in the U.S. that has emerged over the last fifteen years beg an empirical accounting of its urban geography.

The salient locus of the urban derives from both theory and practice: from economic geographers' attention to regional and urban economies and from increased political activity focused on the plight of low-wage workers at the city level. Scholars of urban regional economies tout the health of a metropolitan economy as dependent upon growth and productivity, often driven by core high-road industries that generate positive externalities (e.g., Glaeser and Saiz, 2003; Scott and Storper 1987). Low-wage employment lurks as a contextual byproduct in these accounts. But politics matters as well, and for low-wage workers urban politics has become a paramount arena of engagement and redress. From the rise of the living wage movement in the 1990s to the recent wave of strikes by fast food workers across U.S. cities, most political and policy activity aimed at low-wage work has been local. On the policy side, this urban mobilization has led to various forms of local labor market regulation, including living wage, worker retention, and paid sick day ordinances. The culmination of local labor market regulation has been the passage of city minimum wage ordinances, such as the $15/hour wage ordinance adopted by Seattle in 2014. San Francisco was the first city to enact a comprehensive coverage city minimum wage in 2004, pegging the city minimum to inflation and covering all workers, regardless of industry or establishment size.

This study evaluates the distribution of low-wage work across U.S. cities in 2013, focusing on the characteristics of local labor markets that may mediate the likelihood of low-wage employment. We illustrate the large geographic variation of low-wage work across U.S. cities, examine the demographic characteristics of low-wage workers, and analyze the effects of a city's urban economy and regulatory climate on the probability of low-wage employment. Using a multilevel analysis, we evaluate whether individual

characteristics alone account for low-wage employment, or whether a metropolitan region's characteristics might matter, too. For example, would the same immigrant worker experience the same likelihood of low-wage employment whether she resides in Los Angeles or Chicago or San Francisco? We utilize the Current Population Survey Outgoing Rotation Group file, a data set little used by geographers, that provides a rigorous measure of hourly wages and sizable sample counts across the largest 27 metropolitan labor markets in the U.S. We find that local regulatory effects such as the unionization rate and metropolitan minimum wages significantly reduce the likelihood of low-wage employment for less-educated workers. These same workers in urban economies with larger shares of retail and food services employment—engines of the low-wage economy—are more likely to be employed in low-wage work.

**Background**

Over the last three decades, low-wage work has slowly but steadily comprised a growing share of the American job market (Mason and Salverda, 2010). The growth of low-wage work has become especially salient during the Great Recession and its aftermath, with employment losses concentrated in mid-wage occupations and gains concentrated in lower-wage occupations. Lower-wage occupations grew 2.7 times as fast as mid-wage and higher-wage occupations during the recovery (Bernhardt, 2012).

Recent scholarship on low-wage work provides a robust characterization of low-wage workers: female, minority, and immigrant workers are overrepresented in the low-wage workforce (Thiess, 2012; Capps, Fix, Passel, Ost, and Perez-Lopez, 2003); females comprise more than half of all low-wage workers (Cooper and Essrow, 2013); nearly all

low-wage workers are full-time workers or part-time workers who are involuntarily underemployed (Boushey, Fremstad, Gragg, and Waller, 2007); and a little less than two-thirds of low-wage workers are adults older than the age of 24 (Boushey et al., 2007), with 35.5 percent at least 40 years old (Cooper and Essrow, 2013).

Surprisingly, there is a dearth of scholarship on local labor market variations in low-wage work. Qualitative analyses and studies on particular segments of the low-wage market prevail in the literature. Appelbaum, Bernhardt, and Murnane (2003), for example, present a comprehensive case study analysis that examines firms' cost-cutting strategies and subsequent worker outcomes in 25 industries employing low-wage workers; the varied strategies often, but not inevitably, result in declining quality of work for low-wage workers. Other scholars have looked at the destabilizing presence of the day laborer market within the low-wage economy (Melendez, Visser, Theodore, and Valenzuela, 2014); the vulnerability of low-wage workers to wage theft and other workplace violations (Bernhardt et al., 2009); and the low quality of low-wage work in America as compared to select European countries (Schmitt and Gautie, 2010). While wages are the defining feature of low-wage work, nonstandard working hours and lack of employer-sponsored benefits such as sick leave, health insurance, and retirement have also been found to characterize low-wage work in America (Kazis and Miller, 2001).

How cities have shared in the growth of low-wage work has garnered little attention by scholars. In one exception, Florida (2013) finds that cities with strong energy and knowledge economies experienced the largest growth in high-wage jobs during the recovery, while cities in the Sun Belt and Rust Belt sustained the highest growth in low-wage jobs. Yet economic and urban geographers of various theoretical and

methodological proclivities have noted the increasingly rigid labor segmentation that has become a dominant feature of most U.S. cities. Storper and Scott (2009) state that low-wage workers "are not simply a minor side-effect of the new economy or an accidental adjunct to the creative class. . . . The low-wage segment of the labor market it itself one of the critical foundations of urban life today and hence of current patterns of growth. . . ." (164). Doussard, Peck, and Theodore (2009) also document this pattern of growing polarization, as well as illustrate the distinctive pattern of uneven growth at the metropolitan level in cities such as Chicago.

Two primary explanations of low-wage work inhere within the urban economy literature. The first explanation emphasizes industries and skills, the second labor market institutions. A number of specific arguments fall within the first explanatory domain, but all share a focus on the industrial and corollary skill composition of an urban economy. Generally, industry change stimulated by some combination of technological change and global competitive advantage creates a strong demand for higher skilled workers—among labor economists, this is the "skill-based technical change" explanation of growing wage inequality (e.g., Autor, Katz, and Krueger, 1998). Standardization and offshoring eliminate middle-skill jobs, and lower-skill jobs persist in industries of non-tradable and geographically sticky products, such as personal services and construction. In these more neoclassical supply and demand accounts, wages are fully determined by skills.

Among urban geographers, the global city thesis is the most dominant version of this demand-supply explanation (Sassen, 1991). Command-and-control centers of the global economy concentrate finance and producer services, drawing in highly skilled and highly paid workers. The growth of low-wage employment emerges as an outcome

driven by the consumer demands of this labor market elite for personal maintenance services, such as housekeepers, dry cleaners, nannies, and restaurants. Polarization characterizes the occupational and wage distribution of these urban economies. Other versions of this demand-supply skill explanation also highlight the "complementarity" of high- and low-skilled workers (Autor, Katz, and Kearney, 2008). By contrast, research emphasizing the relationship between the concentration of highly skilled workers and urban growth (Glaeser and Saiz, 2003) and to higher metropolitan wages (Florida, Mellander, Stolarick, and Ross, 2012) pays little heed to low-wage employment.

The second explanation of low-wage work emphasizes changes in labor market institutions as an explanation for the growth of low-wage employment. Labor economists point to the decreasing real value of the minimum wage and declining unionization as fundamental contributors to growing wage inequality and low wages (Card, Lemieux, and Riddell, 2003; DiNardo, Fortin, and Lemieux, 1996; Mishel et al., 2012). These processes are central to accounts of the post-Fordist shift in labor markets in geography that emphasize the changing bargain between capital and labor, to capital's advantage (Doussard et al., 2009; Peck, 1996).

Geographers emphasize the local instantiation of labor market institutions, such as unions, but also the larger set of spatially embedded norms, practices, and regulatory policies that influence labor market processes and inhere at a local level (Hanson and Pratt, 1992; Peck, 1996). Empirical research illustrates the regional variation in wage structures that reflects geographic differences in labor market norms (Odland and Ellis, 2001), as well as geographically variable patterns in racial-wage inequality that are partly explained by local racial institutional effects, such as incarceration and public

employment (Parks, 2012). Sociologists also have explored the regional variation in labor market norms, finding that higher levels of union density reduce wage inequality (Western and Rosenfeld, 2011).

*Labor market regulation from below*

The growing mobilization of urban low-wage workers who are demanding, and winning, policy responses at the local level serves as a prime motivation for our study. Living wage, paid sick leave, worker retention, and city minimum wage ordinances comprise a constellation of regulatory policy at the local level that have given rise to a new and politically robust form of labor market regulation from below (Lester, 2011; Luce, 2004; Wells, 2002). The emergence and success of these local political efforts to regulate the local labor market introduce a new dimension of analysis for geographers interested in local labor market institutions—the local regulatory climate.

Motivated by the deteriorating conditions of work and remuneration experienced by low-wage workers, these political movements have sought to secure policy that improves both wages and working conditions in the low-wage labor market. City minimum wage ordinances, first passed in San Francisco in 2004 and most recently in Seattle in 2014, represent the epitome of such political efforts (Jacobs and Reich, 2014; Johnson, 2014). City minimum wages go far beyond living wage ordinances in that they apply to all workers, not just city contractors (Sonn, 2006). These ordinances raise the wage floor of a local labor market, theoretically shrinking the size of the local low-wage labor market.

The map of local labor market regulation is quickly evolving. Mobilization among low-wage workers in urban areas continues, as evidenced in strikes and walk-outs among unorganized low-wage workers in fast-food and retail jobs beginning in 2012 and continuing through 2014. City minimum wage ordinances are under consideration in several U.S. cities, including Oakland and Chicago (Eichelberger, 2014; Levin, 2014; CBS Chicago, 2014). It seems that local labor markets and their regulatory conditions have never been so important to workers in the U.S. since before New Deal labor and employment legislation. With little movement at the federal level to raise the minimum wage, reform labor law, or improve the working conditions of low-wage jobs, the employment prospects of workers at the lower end of the labor market are increasingly defined by local labor market conditions, norms, and policies.

This proposition drives our analysis. We analyze the probability of low-wage employment across different metropolitan labor markets, accounting for differences in industrial mix and regulatory climate, measured by union density and minimum wage. Using a multilevel modeling approach, we control for worker-level characteristics, including occupation, in order to rigorously assess if and how local regulatory context mediates an individual's likelihood of low-wage employment. Is low-wage employment solely determined by individual characteristics? Or would the same low-skilled worker have a better chance of escaping low-wage employment depending upon her city of residence? In short, does geography matter? We take up this question after presenting a descriptive portrait of low-wage workers at the national level in the following section.

**Data and analytic strategy**

Data come from the Center for Economic and Policy Research Uniform Extracts of the U.S. Current Population Survey (CPS) Outgoing Rotation Group (ORG) file (Center for Economic and Policy Research, 2014). The CPS provides timely earnings data, including an hourly wage measure, as well as information on commission, tips, and overtime. The CPS is also the only large public data set that gathers information on union membership and coverage. Researchers are most familiar with the March Supplement of the CPS, yet its small sample size limits geographic analysis below the state level. Compared to the March CPS, the CPS-ORG file provides more robust sample sizes because this file pools earner study respondents across a single year. Respondents are asked earning questions when they rotate out of the survey, in either their fourth or eighth month of survey participation (one-fourth the survey sample that month). Because a respondent is in for four months, then out for eight months, and then in again for a final four months, respondents in the CPS-ORG are not repeated within a 12-month period. The size of the CPS-ORG pooled over twelve months is equivalent to three months of full CPS files.

We use data from the calendar year 2013. All descriptive results are calculated using the CPS sampling weights. We do not use weighted data in the models. The sample includes civilian wage and salary workers, ages 18 to 64, who worked for pay the week prior to the survey and who were not self-employed. We utilize the CEPR's wage variable that accounts for workers' earned income from overtime, commission, and tips (Schmitt 2003). Following Schmitt (2003), we drop respondents with hourly wages of less than $1 and more than $100.

We utilize the smallest geographic area available in the CPS, the Core-Based

Statistical Area. We do this to account best for metropolitan regulatory effects. For example, we use the San Francisco-Oakland-Freemont CBSA rather than the San Jose-San Francisco-Oakland CMSA. We select the largest 27 CBSAs for which there are at least 1,000 observations. Although the CPS utilizes the standard FIPS codes and attendant geographies, there are some differences (noted as "partial" in Table 1). For example, the CPS does not survey respondents in the counties of DeKalb, IL; Jasper, IN; and Kenosha, WI, of the Chicago-Naperville-Joliet, IL-IN-WI CMSA.

Scholars across disciplines, including economics, public health, and sociology, have made use of the metro level structure of the CPS data (Adams & Neumark, 2004; Hirsh and Macpherson 2003; Subramanian, Acevedo-Garcia, and Osypuk, 2005; Western and Rosenfeld, 2011). Use of the CPS among geographers is rare, likely because the smaller sample size of the March CPS hinders fine geographic disaggregation. For researchers interested in earnings and other primary labor market experiences, however, the CPS ORG provides a robust sample when restricted to the largest metropolitan regions. We adopt a multilevel modeling strategy that pools data across the largest metro areas in order to capitalize on the variance available across metro levels in the data.

Lastly, the CPS is a multistage stratified sample but the BLS does not publish information about its stratification method necessary to produce sampling weights, especially for level 2 variables (e.g., the selection of respondent *i* conditional on the selection of metro *j*). Some researchers have begun examining this issue in depth (e.g., Davern, Jones, Lepkowski, Davidson, and Blewett, 2007). We do know that that the stratification approach depends upon metro areas and metro size (respondents are selected conditional on metro residence and respondents in smaller metros are more

likely to be sampled). Because our multilevel model includes variables for both, we have some confidence that our approach reduces standard error bias in the absence of sampling weights. Even if sampling weights were available, we may still prefer unweighted data. Both metro area and metro size are substantive independent variables of interest, and these variables would likely determine sampling weights. In such situations, estimated coefficients would be biased with smaller standard errors than for coefficients estimated with unweighted data (Winship and Radbill 1994).

*Defining low-wage work*

Generally, wage-based definitions of low-wage work utilize either the basic-income or distributional threshold approach. Under the former, the threshold for a low wage is the minimum income a full-time worker must earn to purchase basic needs including food, shelter, health care, and child care (Boushey et al., 2007). The basic-income approach is widely used but is limited by the variable nature of basic needs and their costs and the difficulty of establishing a uniform standard measure.

Another approach sets a threshold value that defines low-wage work in reference to a quantile of either the total wage distribution (i.e., the bottom third of all wages) or of the median wage. This approach allows for comparison of the number of low-wage jobs over time as well as comparison between wages. The OECD, among others, uses the median wage as a base for comparison, defining low-wage work as below two-thirds the national median. Following Boushey et al. (2007), we define low-wage work as work that pays less than two-thirds of the median male hourly wage in 2013, or $12.82. Taking the median wage for males as a base limits the strong effects of gender wage inequality on

the wage distribution, including the higher proportion of women at the lower end of the wage distribution.

*Demographic characteristics of low-wage workers*

We begin by running a simple logistic regression on the entire U.S. workforce to illustrate the demographic and occupational characteristics associated with low-wage employment (Table 1). The binary dependent variable measures whether the respondent's hourly wage is $12.82/hour or less (low-wage or not). In 2013, 31.4% of all workers in the U.S. were in the low-wage workforce with hourly earnings of $12.82/hour or less. In line with previous research, we find that being female, a worker of color, less educated, and an immigrant all increase a worker's likelihood of low-wage employment (base comparison is male, white, native-born, high school graduate, not a union member, employed in a production occupation). All associations are highly statistically significant. Workers with less than a high school diploma are more than twice as likely as high school graduates to be employed in a low-wage job (coefficients are reported as odds ratios for all models). Increasing levels of schooling increasingly decrease one's chance of low-wage employment. Compared to race and nativity, gender is a stronger determinant of low-wage employment: women are 84% more likely than their male counterparts to be employed in a low-wage job. Blacks are 48% and Latinos 21% more likely than whites to hold a low-wage job. Immigrants are 33% more likely than the native-born. Older workers are less likely than younger workers. Although the diminishing effect of age is statistically significant, it is negligible in size (the age-squared term). Being a union member has a sizable negative effect on a worker's

likelihood of low-wage employment. The effect size is on par with that of having a college degree over a high school diploma.

We include occupation dummies to capture effects across industries and occupations that may influence low-wage employment. Occupations provide a parsimonious proxy of these effects because CPS occupations are highly correlated with industry, especially in industries with the highest concentrations of low-wage jobs. For example, 85% of all food preparation and serving occupation are located within food services. Sales occupations are similarly concentrated in retail trade. Production occupations are the base category. All occupations, with the exception of office support, are statistically significantly more or less likely than production occupations to pay low wages. Strikingly, food preparation and serving occupations are more than three times more likely than production jobs to pay low wages. Building and grounds maintenance and personal care services are more than two times as likely. Other low-wage occupations include sales jobs and healthcare support occupations. Results for occupations that are typically understood as higher skilled jobs are as expected: management, business and finance, architecture, legal, and computer occupations are all less likely to result in low-wage employment compared to production occupations. Education occupations are slightly more likely to be low-wage, likely a result of lower-wage preschool workers as well as large numbers of educational support jobs included within this occupational category. We ran this and following models with food preparation and serving occupations as the base category and found no differences in expected directional or significance test results (not reported).

*Metropolitan labor market effects*

We employ a multilevel modeling strategy to test for local labor market effects on the probability of low-wage employment. Multilevel models accommodate individual-level effects (within-metro variation) when estimating metropolitan-level effects (across-metro variation) and generate accurate tests of significance for clustered data (nesting violates the OLS assumption of independent violations)(Gelman and Hill, 2007). Labor market outcomes in the U.S. are heavily dependent upon individual worker characteristics, so accurate accounting of these effects is paramount to discern whether a region is different simply because its population is different or because the region contributes its own contextual influence above and beyond the aggregation of its population attributes (Parks, 2012). The fact that wages are higher in cities with more highly paid (i.e., highly skilled) workers is a tautological claim when using aggregate measures (e.g., Florida et al., 2012). The question of whether a concentration of highly skilled workers generates externalities, such as higher returns to skill, is a question that demands a different analytic strategy. Multilevel models are one such strategy.

The latter example additionally illustrates how multilevel models can be used to test for and estimate random effects—whether a particular effect varies over the data. In our case, we examine whether metropolitan effects vary across cities. For example, does the effect of union density on the probability of low-wage employment differ across labor markets, beyond differences in levels of union density?

The two levels of our multilevel logistic regression are as follows. At the individual-level, we specify the following logistic regression:

$$\Pr\left(y_i = 1\right) = \text{logit}^{-1}\left(X_{ij}\beta + \alpha_{j[i]}\right), \text{for } i = 1, \ldots, \eta,$$

where Pr $(y_i = 1)$ is the probability that worker *i* is employed in a low-wage job, *X* is the matrix of individual-level predictors and *j*[*i*] indexes the metropolitan region where worker *i* resides. The second level of the model is the regression of the metropolitan labor market coefficients:

$$\alpha_j \sim N(U_j\gamma, \sigma_\alpha^2), \text{ for } j = 1, \ldots, 27,$$

where *U* is the matrix of metro-level predictors, $\gamma$ is the vector of coefficients for the metro-level regression, and $\sigma_\alpha$ is the standard deviation of the unexplained group-level errors.

Our primary measures of local regulatory context are union density and the metropolitan minimum wage. Union density is the percentage of workers employed in a job covered by a union contract. A large body of empirical work demonstrates the effect of union membership and union density on wages and inequality, raising the former and reducing the latter (see Western and Rosenfeld, 2011, for a comprehensive review). Theoretically, unions have the strongest distributional impacts in sectors where they are organizing workers. Thus, the union threat effect is described as generating higher wages for nonunion workers because employers will raise wages to steer off unionization drives, thereby equalizing the wage distribution (Kahn and Curme, 1987).

We hypothesize that union effects may vary across labor markets, particularly for low-wage employment. Not only are levels of union density across regions highly variable, the composition of the unionized workforce and the organizing activities of local unions vary greatly across metropolitan regions. The union sector in Detroit remains heavily concentrated in manufacturing and is dominated by older industrial unions such as the United Auto Workers. By contrast, the labor movements in cities such as Los

Angeles and Las Vegas represent workers primarily in the service sector and are dominated by unions such as the Service Employees International Union and Unite Here, a union representing hotel and restaurant employees. Politically, these unions are most actively organizing workers in the lower paid echelons of the service industry, such as janitorial, home health care, and hospitality services. We discuss our empirical test of this hypothesis in the results section.

The metropolitan minimum wage is either the state minimum wage or a weighted average of state minimum wage rates if a metropolitan region covers more than one state (Department of Labor, 2014). In the case of San Francisco, we apply the city minimum of $10.55 to all workers in the San Francisco and Oakland metropolitan region because no geographic information on city residence is available in the CPS. This overestimates the coverage of the San Francisco city minimum wage. However, commuters from the East Bay into San Francisco are captured with this strategy, and the San Francisco city minimum likely exerts regional spillover effects that pull wages in the East Bay up. As a measure of regulatory climate, this strategy captures the current climate in Oakland city politics, likely the next U.S. to adopt a city minimum wage, currently proposed at $12.25/hour, higher than San Francisco's (Levin, 2014).

We include two measures of industrial mix that are of most theoretical significance for low-wage workers. Strong historical evidence points to manufacturing as a marker of fordist labor and industrial relations operating in local markets. Manufacturing jobs are more likely to be unionized, and the concentration of manufacturing jobs can signal norms and practices that spill beyond its union and sectoral boundaries (Western and Rosenfeld, 2011). But deindustrialization and restructuring have

driven a wedge between the politics and the economy of manufacturing. Spatial relocation strategies to nonunion, lower wage regions have driven wage rates in manufacturing down (Bluestone and Harrison, 1982; Angel and Mitchell, 1991). The growth of light, nondurable manufacturing in urban economies such as Los Angeles illustrates the fragmented, nonunion, low-wage nature of post-fordist manufacturing that depends primarily on female, immigrant, and temporary workers (Peck and Theodore, 2001; Scott, 1996). Thus, the directional effect of manufacturing on low-wage employment is an empirical question to be tested.

We include the percentage of food service and retail industrial employment as a proxy for the low-wage industrial composition of a metropolitan economy. These industries have the highest concentration of low-wage jobs at the national level (71% and 52%, respectively). A number of studies illustrate the flexible labor practices that pervade these industries, contributing to their high levels of low-wage employment (Lane, Moss, Salzman, and Tilly, 2003; Carre, Tilly, van Kalveren, and Voss-Dahm, 2010). Larger shares of these industries within an urban economy may signal fewer employment opportunities for workers, pushing them into these industries and thereby increasing their probability of low-wage employment. Further, these industries' larger footprint within an urban economy may influence employment and pay practices in other sectors, ratcheting down wages and working conditions across the lower end of the labor market, thus expanding the low-wage economy and increasing workers' probability of low-wage employment.

We include the percent of the labor force that is foreign-born to test for hypothesized crowding effects that may result from high numbers of immigrants who are

trapped at the lower end of the labor market, or the secondary sector, because of low skills, poor English language ability, or unrecognized credentials. Again, this is an effect measured above and beyond the individual-level effect of being an immigrant.

Lastly, we include unemployment rate as a measure of short-term economic strength. A slack labor market may increase the probability of low-wage employment because competition for scarce jobs will increase. We include the log of population size (taken from the Census 2012 estimates for each metropolitan region) as a control and as a known sampling design stratification variable.

*Metropolitan descriptive statistics*

Low-wage workers are more concentrated in the largest 27 metropolitan areas included in our sample than in the U.S. economy as a whole. Of all low-wage workers in the U.S., 40.2% live in these 27 metropolitan areas, compared to 31.4% across the U.S. The average size of the low-wage workforce across these 27 metropolitan regions is 28%, but relative levels of low-wage employment vary greatly among these cities. Table 2 provides descriptive statistics on the metropolitan variables we examine. Cities are ranked by the relative size of their low-wage employment shares. Los Angeles tops the list with the largest low-wage workforce: 34% of its workers are employed in low-wage work. The Inland Empire of Southern California (Riverside-San Bernardino-Ontario) is the next largest at 34.4%. The Dallas-Fort Worth and Houston regions are third and fourth at 34% and 33.9%, respectively. The Washington D.C.-Arlington VA region has the lowest level of low-wage employment at 20.5%. The greater Seattle region is the second lowest at 20.9%, and San Francisco-Oakland third lowest at 21.6%. The greater

New York region falls near the middle of the urban distribution at 27.2% (fourteenth in rank). The Chicagoland region's share of low-wage employment is 29.6% (tenth in rank).

The metropolitan minimum wage ranges from a high of $10.55 (San Francisco-Oakland) to a low of $7.25, the federal minimum wage. Sixteen of the 27 metropolitan regions have minimum wage rates higher than the federal rate. Unionization rates vary greatly among these urban labor markets. The Riverside-San Bernardino-Ontario region has the highest rate of unionization at 24.9% (measured by union coverage, not strictly union membership), likely driven by its relatively high levels of public employment, especially when compared to Los Angeles. Honolulu is next at 23.4% (another region with high levels of public employment), followed by New York at 21.7% and Seattle-Tacoma at 18.7%. The regions with the lowest rates of unionization include Houston (5.9%), Phoenix (6.2%), and Atlanta (6.8%). Figure 1 plots union density against low-wage employment, weighting each metropolitan circle by the number of low-wage workers. This descriptive analysis illustrates an inverse relationship between higher levels of unionization and lower levels of low-wage employment.

*Model results*

We begin by running a multilevel model with only worker-level characteristics as a comparison to our simple logistic regression run on the national sample (Table 1, Model 2). Because this is a multilevel model, these results account for grouping at the metropolitan area. The results generally mirror the national-level logistic results, albeit with slight shifts in the size of the coefficients that reflect the urban inflection of labor market patterns. Women in these urban markets are slightly less likely to be employed in

low-wage work compared to at the national level (OR of 1.58 compared to 1.843). The effect of race and nativity are more pronounced, capturing the greater disadvantage experienced by immigrants and workers of color that characterizes urban labor markets. Immigrants, for example, are 57% more likely to be employed in low-wage work compared to native-born workers in these urban markets, compared to 33% at the national level.

Occupation continues to exert significant and sizeable effects, but the coefficients shrink for most all occupations (farming and fishing is so small within urban economies that we ignore these coefficients). Workers in food preparation and serving occupations continue to experience the highest likelihood of low-wage work (2.5 times more likely than production workers), but the odds ratio is smaller than for the national sample (OR=3.2). The same urban trend is evident for personal service and building and grounds maintenance occupations.

In a first test of geographic difference, we run a simple model with only metropolitan dummies to evaluate if there is significance variance in the probability of low-wage employment across labor markets (Table 3, Model 3). The model is empty at the individual level. The different coefficient values indicate that the probability varies significantly across regions and that the probability of low-wage employment is statistically significantly lower in every metropolitan region compared to Los Angeles. The odds ratios range from a low of 0.45 for DC to a high of 0.97 for Riverside-San Bernardino: low-wage employment is 55% less likely in DC and 3% less likely in Riverside than in L.A. These results accord with the trends illustrated in the metropolitan descriptive statistics.

Next, we combine individual-level characteristics with the set of metropolitan dummies (Table 3, Model 4). A comparison of the variance components between Model 2 and Model 4 indicates that metropolitan characteristics explain additional variance beyond individual worker characteristics and occupation. The worker-level coefficients remain stable and largely unchanged, but we can discern additional explanation that metropolitan context contributes. The probability of low-wage employment differs statistically from Los Angeles in 13 of the metropolitan regions and is not negative in all cases as we might expect from the descriptive statistics. Statistically identical production workers are 30% more likely to experience low-wage employment in Detroit than Los Angeles, a statistically strong effect ($p<.001$). They are also more likely to experience low-wage employment in Atlanta, although this is a statistically weaker effect ($p<.10$).

Labor markets that are better for workers at the lower end of the labor market than Los Angeles are considerably better, measured by coefficient size and significance. The likelihood of low-wage employment is 34% lower in San Francisco and Washington DC ($p<.001$), 27% lower in Seattle ($p<.001$), 23% in San Diego ($p<.05$), 22% in Minneapolis-St. Paul ($p<.001$), 17% in Denver ($p<.01$), and 14% in New York ($p<.01$) for statistically comparable workers.

The metropolitan dummies absorb all variation at the metropolitan level of analysis, so although they are helpful in ascertaining where workers do better, they cannot tell us why some metropolitan labor markets decrease a worker's likelihood of low-wage employment. Model 5 (Table 4) begins to get at what accounts for these differences by including level-2 measures of metropolitan context.

Population size, the unemployment rate, union density, metropolitan minimum wage, percentage manufacturing, and percentage food and retail all contribute statistically significant explanations of variance. Higher rates of unemployment are positively related with low-wage employment—a 1% increase in unemployment increases the likelihood of low-wage employment by nearly 3%. Higher rates of manufacturing tend to increase the likelihood of low-wage work—a 1% increase in manufacturing is associated with a nearly 2% increase in low-wage employment. Led by Los Angeles, the new urban manufacturing exemplifies the separation of politics from sector: Ford's own sector is no longer fordist. Rather, manufacturing is now primarily associated with lower wages and flexible labor practices. These new labor and employment relations are strong enough to influence labor and pay practices beyond the confines of manufacturing. The negative effects of the lowest wage industries in the U.S. are even stronger. A 1% increase in food and retail employment tends to increase the likelihood of low-wage employment by 3.4%.

The results indicate strong local regulatory effects on the likelihood of low-wage employment. Raising the metropolitan minimum wage by $1.00 tends to reduce the probability of low-wage employment by 7.3%. Whereas an increase of $1.00 may have been simply theoretical in the past, the increases now under consideration in cities such as Oakland and Chicago far exceed this $1.00 margin. Seattle's recently passed $15.00 minimum wage would virtually wipe out low-wage employment under our empirical threshold of $12.82/hour.

Union density may generate different effects across labor markets, particularly for low-wage employment. We empirically test this hypothesis by running the multilevel

logistic regression with union density as a random effect, evaluating whether the slope (or coefficient) on union density varies significantly across metropolitan labor markets. The standard deviation for coefficients on union density is 0.017632—almost five standard errors (0.003709) from zero—indicating that significant city-to-city variation in the slope coefficients on union density exists. A test of model fit indicates that this model fits no better but no worse than the model without random slopes for union density. However, we prefer this model as the test for random variation is significant, and we believe this model reveals a more complete description of the underlying data, e.g., the pattern revealed in Fig. 1.

Figure 2 graphically illustrates the varying effect of union density across metropolitan labor markets. We plot the full metropolitan effect of union density (the fixed plus random coefficient term exponentiated as an odds ratio) for each metropolitan region across average unionization levels. The variation across labor markets ranges from -3.6% for Minneapolis-St. Paul to 1.2% for Chicago. A 1% increase in union density in Minneapolis-St. Paul corresponds to about a 3.6% decrease in a worker's likelihood of low-wage employment. Union density is slightly positive in Chicago and Las Vegas, but is negative in every other city (below the OR line of 1.0). That is, an increase in union density tends to decrease the likelihood of low-wage employment, controlling for individual and other metropolitan characteristics.

The differences between Chicago, New York, and Los Angeles may indicate the different orientations of each region's labor movement. Chicago's labor movement provides minimal, if any, protection from low-wage work for more disadvantaged workers, whereas unionization in both Los Angeles and New York beneficially lowers

the likelihood of low-wage employment. The effect is strongest in Los Angeles, ground-zero for the regeneration of labor in the U.S. as the movement of immigrants, women, and workers of color in the low-wage service sector (Milkman, 2006). The stronger negative effect of union density on low-wage employment for cities with relatively low levels of unionization such as Houston (-2.9%) and Miami (1.2%) indicates the pronounced effect of any union activity in such contexts. Lastly, the fitted line and cases such as Honolulu (2.8% increase) illustrate the increasingly strong effect of unionization at higher levels of union density.

We also ran a set of models (not reported) testing for the metropolitan effect of percent college graduates in the labor force. This variable is highly correlated with unemployment and the size of the immigrant workforce, and we could not estimate a model including all three. When we forced estimation by limiting the number of maximum likelihood iterations, results differed erratically from previous model specifications. Strong multicollinearity or other mediating relationships seem to exist between these three variables. When the sizes of the college graduate and immigrant workforces are included separately, both were statistically insignificant although college grads continued to cause problems for model specification (it is most highly correlated with unemployment) whereas percentage of foreign-born workers did not alter coefficient estimates in unexpected ways. Because we believe unemployment to be a theoretically stronger explanation of low-wage employment (e.g., indicator of general tightness or slackness in the labor market), our model specification strategy privileged this variable above percentage of college graduates. We include a measure of immigrant workforce in our final model and find an insignificant relationship between the size of a metropolitan

market's immigrant workforce and the likelihood of low-wage employment. The size of the immigrant workforce does not determine a worker's probability of low-wage employment.

Our empirical question differs from that taken up by scholars examining the direct wage and employment effects of minimum and living wage increases (Adams and Neumark, 2004; Lester, 2011; Pollin, Brenner, Wicks-Lim, and Luce, 2008). We evaluate whether local minimum wages influence the probability of low-wage employment. The hypothesized mechanism is one of raising the floor of all wages high enough to tip many heretofore low-wage jobs over the $12.82 threshold, a spillover effect of the metropolitan minimum wage on jobs slightly above those that would be directly impacted. Our results indicate the very strong influence of the labor market-wide effect of metropolitan minimum wages. We would perhaps see even stronger effects if we were to examine the change in wages below the $12.82/hour threshold—wages within the low-wage economy.

**Conclusion**

As the low-wage recovery continues unabated, urban economic growth offers little conciliation to workers struggling to make ends meet while maintaining hope for economic mobility. We find that workers in some cities have more cause for hope than others. Although certain individuals are more likely to find themselves among the ranks of the low-wage workforce, their relative risk of low-wage employment depends upon where they live, above and beyond who they are, their education level, or even their occupation.

We found significant variation in the likelihood of low-wage employment across the largest 27 metropolitan areas in the U.S. Statistically identical production workers are 30% more likely to experience low-wage employment in Detroit than Los Angeles and 34% less likely in San Francisco and Washington DC (lowest likelihood). The metropolitan areas with the next lowest likelihood of low-wage employment are: Seattle (27% lower), San Diego (23% lower), Minneapolis-St. Paul (22% lower), Denver (17% lower), and New York (14% lower).

When we examine the specific aspects of a metropolitan labor market that may account for these differences, we find strong evidence that the regulatory context of the local metropolitan labor market matters significantly and substantively. Both higher metropolitan minimum wages and higher levels of union density reduce the likelihood of low-wage employment. The influence of unionization on local norms and pay structures beyond the confines of the union sector likely accounts for this effect by raising the wage floor and pushing the wage distribution upward such that wages for jobs in typically low-wage sectors are pushed over the low-wage threshold. A similar effect is evident for metropolitan minimum wages. None of the minimum wage rates in the metropolitan regions we analyzed are above our low-wage threshold of $12.82/hour. Thus, metropolitan minimum wages not only directly raise wages for workers at the lowest percentiles of the wage distribution, they push wages up for all workers in the lower quantile of the wage distribution, tipping many above the $12.82/hour threshold.

Lastly, we find that the effect of union density varies across metropolitan areas, i.e., the strength of the effect of union density differs across labor markets. Although union density decreases the likelihood of low-wage employment for statistically

comparable workers across all metropolitan markets, the effect is stronger in some metropolitan areas. This is not simply due to different levels of union density, but rather the difference in the effectiveness of different urban labor movements for low-wage workers. The variability of this effect may also illustrate the changing nature of unionization and sectoral trends. Union density seems to matter more for low-wage workers in places where unions are organizing in low-wage service sectors, such as Los Angeles and Houston.

Mobilization from below by urban workers and low-income communities has yielded changes in local labor market regulation with significant implications for low-wage work. These regulatory mechanisms, whether labor market institutions such as unions or policies such as city minimum wages, raise standards in local labor markets and reduce the likelihood of low-wage employment. The momentum of urban labor market regulatory efforts continues to crest. Cities across the U.S. are either considering local minimum wage legislation or have campaigns underway calling for such legislation (Covert and Peck, 2014). The influence of these political efforts from below is pushing beyond the confines of urban politics: President Obama and even the International Monetary Fund have called for the U.S. to raise the federal minimum wage (Schroeder, 2014). Yet these efforts remind us that all politics—and all labor markets—are local.

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Table 1. Results for Models 1 and 2

| | Model 1 Logistic regression across U.S. | | | Model 2 Multilevel logit on 27 Metros | | |
|---|---|---|---|---|---|---|
| | OR | Std Err | P>\|z\| | OR | Std Err | P>\|z\| |
| Female | 1.843** | 0.029 | 0.000 | 1.580** | 0.026 | 0.000 |
| Race (base: White) | | | | | | |
| Black | 1.482** | 0.033 | 0.000 | 1.511** | 0.035 | 0.000 |
| Hispanic | 1.210** | 0.027 | 0.000 | 1.264** | 0.035 | 0.000 |
| Asian | 1.021 | 0.034 | 0.525 | 1.138** | 0.046 | 0.005 |
| Other | 1.320** | 0.077 | 0.000 | 1.279^ | 0.147 | 0.093 |
| Foreign Born | 1.325** | 0.032 | 0.000 | 1.574** | 0.032 | 0.000 |
| Education (base: High School Grad) | | | | | | |
| Less than high school | 2.389** | 0.066 | 0.000 | 2.385** | 0.042 | 0.000 |
| Some college | 0.809** | 0.013 | 0.000 | 0.844** | 0.029 | 0.000 |
| College | 0.325** | 0.007 | 0.000 | 0.341** | 0.036 | 0.000 |
| Advanced | 0.199** | 0.008 | 0.000 | 0.228** | 0.058 | 0.000 |
| Age | 0.770** | 0.003 | 0.000 | 0.760** | 0.006 | 0.000 |
| Age^2 | 1.003** | 0.000 | 0.000 | 1.003** | 0.000 | 0.000 |
| Union Member | 0.391** | 0.010 | 0.000 | 0.437** | 0.040 | 0.000 |
| Occupation (base: production occupations) | | | | | | |
| Management | 0.359** | 0.014 | 0.000 | 0.274** | 0.069 | 0.000 |
| Business & Financial Operations | 0.298** | 0.015 | 0.000 | 0.247** | 0.084 | 0.000 |
| Computer & Mathematical | 0.202** | 0.015 | 0.000 | 0.166** | 0.117 | 0.000 |
| Architecture & Engineering | 0.216** | 0.019 | 0.000 | 0.176** | 0.149 | 0.000 |
| Life, Physical, & Social Science | 0.438** | 0.042 | 0.000 | 0.417** | 0.156 | 0.000 |
| Community & Social Services | 0.758** | 0.048 | 0.000 | 0.648** | 0.110 | 0.000 |
| Legal | 0.261** | 0.027 | 0.000 | 0.195** | 0.174 | 0.000 |
| Education, Training, & Library | 1.213** | 0.049 | 0.000 | 0.926 | 0.070 | 0.271 |
| Arts, Design, Entertain., Sports, Media | 0.706** | 0.044 | 0.000 | 0.482** | 0.102 | 0.000 |
| Healthcare Practitioners & Technical | 0.289** | 0.014 | 0.000 | 0.251** | 0.083 | 0.000 |
| Healthcare Support | 1.667** | 0.074 | 0.000 | 1.459** | 0.074 | 0.000 |
| Protective Service | 1.308** | 0.065 | 0.000 | 1.278** | 0.082 | 0.003 |
| Food Preparation & Serving Related | 3.193** | 0.115 | 0.000 | 2.455** | 0.059 | 0.000 |
| Building/Grounds Cleaning/Maintenance | 2.724** | 0.106 | 0.000 | 2.094** | 0.065 | 0.000 |
| Personal Care & Service | 2.718** | 0.118 | 0.000 | 2.368** | 0.068 | 0.000 |
| Sales & Related Occupations | 1.687** | 0.051 | 0.000 | 1.430** | 0.054 | 0.000 |
| Office & Administrative Support | 0.970 | 0.029 | 0.312 | 0.817** | 0.053 | 0.000 |
| Farming, Fishing, Forestry | 3.519** | 0.277 | 0.000 | 5.030** | 0.258 | 0.000 |
| Construction & Extraction | 0.547** | 0.022 | 0.000 | 0.542** | 0.068 | 0.000 |
| Installation, Maintenance, & Repair | 0.481** | 0.022 | 0.000 | 0.465** | 0.080 | 0.000 |
| Production | (Base) | | | | | |
| Transport & Materials Moving | 1.456** | 0.048 | 0.000 | 1.306** | 0.059 | 0.000 |
| _cons | 122.790** | 9.455 | 0.000 | 153.260** | 0.134 | 0.000 |
| Observations | 154,081 | | | 61,160 | | |
| Wald Chi2 | 31526.62** | | 0.000 | 12398.21** | | 0.000 |

Table 2. Metropolitan Descriptive Statistics

| | % Low-Wage | Median Hourly Wage | Unemp | % Union | % Manuf | % Retail & Food Service | % Foreign Born | Weighted Metro Min Wage | Population Estimate 2013 |
|---|---|---|---|---|---|---|---|---|---|
| Los Angeles-Long Beach-Santa Ana, CA | 34.9% | $16.83 | 9.1% | 14.0% | 11.7% | 16.6% | 40.9% | $8.00 | 13,100,000 |
| Riverside-San Bernardino-Ontario, CA | 34.4% | $16.02 | 12.5% | 24.9% | 9.1% | 17.8% | 26.0% | $8.00 | 4,380,878 |
| Dallas-Fort Worth-Arlington, TX (partial) | 34.0% | $17.00 | 6.1% | 7.0% | 9.7% | 16.5% | 23.0% | $7.25 | 6,810,913 |
| Houston-Baytown-Sugar Land, TX | 33.9% | $17.00 | 5.9% | 5.9% | 10.4% | 16.5% | 28.3% | $7.25 | 6,313,158 |
| Miami-Fort Lauderdale-Miami Beach, FL | 33.2% | $17.31 | 7.2% | 6.8% | 4.9% | 17.4% | 45.9% | $7.79 | 5,828,191 |
| Las Vegas-Paradise, NV | 32.8% | $16.22 | 10.5% | 18.2% | 2.9% | 19.9% | 26.6% | $7.50 | 2,027,868 |
| Atlanta-Sandy Springs-Marietta, GA (partial) | 31.2% | $17.50 | 7.4% | 6.8% | 8.0% | 16.7% | 17.5% | $7.25 | 5,522,942 |
| Detroit-Warren-Livonia,MI | 30.9% | $18.12 | 8.7% | 17.4% | 18.5% | 17.2% | 13.8% | $7.40 | 4,294,983 |
| Phoenix-Mesa-Scottsdale, AZ | 29.7% | $17.58 | 5.8% | 6.2% | 6.6% | 17.7% | 16.8% | $7.80 | 4,398,762 |
| Chicago-Naperville-Joliet, IL-IN-WI (partial) | 29.6% | $18.87 | 8.4% | 16.6% | 11.2% | 16.7% | 20.8% | $8.20 | 9,537,289 |
| NY-New Jersey-Long Island, NY-NJ-PA (partial) | 27.2% | $19.96 | 7.5% | 21.7% | 6.4% | 15.5% | 35.4% | $7.25 | 19,900,000 |
| Omaha-Council Bluffs, NE-IA | 27.2% | $18.50 | 3.7% | 8.9% | 9.6% | 15.9% | 8.6% | $7.25 | 895,151 |
| St. Louis, MO-IL (partial) | 27.0% | $18.27 | 5.7% | 12.7% | 10.3% | 18.2% | 4.4% | $7.45 | 2,801,056 |
| Baltimore-Towson, MD | 26.6% | $20.33 | 6.1% | 11.7% | 6.4% | 17.2% | 13.7% | $7.25 | 2,770,738 |
| Honolulu, HI | 26.5% | $19.00 | 4.2% | 23.4% | 3.0% | 19.0% | 21.5% | $7.25 | 983,429 |
| Providence-Fall River-Warwick, RI-MA | 26.2% | $19.14 | 8.5% | 18.2% | 11.3% | 18.2% | 14.9% | $7.73 | 1,604,291 |
| San Diego-Carlsbad-San Marcos, CA | 25.8% | $19.50 | 5.8% | 15.9% | 7.7% | 16.9% | 27.2% | $8.00 | 3,211,252 |
| Portland-Vancouver-Beaverton, OR-WA (partial) | 25.7% | $19.00 | 7.3% | 14.7% | 14.0% | 16.4% | 13.8% | $8.99 | 2,314,554 |
| Kansas City, MO-KS (partial) | 25.5% | $18.64 | 5.5% | 9.3% | 10.1% | 15.8% | 9.1% | $7.25 | 2,054,473 |
| Denver-Aurora, CO | 24.7% | $19.90 | 5.5% | 10.7% | 5.8% | 17.5% | 13.9% | $7.78 | 2,697,476 |
| Philadelphia-Camden-Wilmington, PA-NJ-DE | 24.6% | $20.00 | 7.5% | 13.6% | 8.7% | 15.9% | 11.7% | $7.25 | 6,034,678 |
| Minneapolis-St Paul-Bloomington, MN-WI (partial) | 22.9% | $20.37 | 5.0% | 14.3% | 13.4% | 16.4% | 11.4% | $7.25 | 3,459,146 |
| Hartford-West Hartford-East Hartford, CT | 21.8% | $21.63 | 7.0% | 14.8% | 13.5% | 13.4% | 15.4% | $8.25 | 1,215,211 |
| Boston-Cambridge-Quincy, MA-NH | 21.7% | $22.50 | 5.7% | 12.3% | 11.3% | 16.1% | 14.1% | $7.25 | 4,684,299 |
| San Francisco-Oakland-Freemont, CA | 21.6% | $23.44 | 6.4% | 16.6% | 7.7% | 14.5% | 35.4% | $10.55 | 4,516,276 |
| Seattle-Tacoma-Bellevue, WA | 20.9% | $20.98 | 5.6% | 18.7% | 11.5% | 18.3% | 17.0% | $9.19 | 3,610,105 |
| Washington-Arlington-Alexandria, DC-VA-MD | 20.5% | $24.04 | 6.3% | 11.2% | 2.1% | 11.9% | 23.8% | $7.70 | 5,949,859 |

Table 3. Models 3 and 4 (level-1 coefficients for Model 4 not reported)

| | Model 3 Multilevel logit on 27 Metros | | | | Model 4 Multilevel logit on 27 Metros | | | |
|---|---|---|---|---|---|---|---|---|
| Base: Los Angeles-Long Beach-Santa Ana, CA | OR | | StdErr | P>\|z\| | OR | | Std Err | P>\|z\| |
| Atlanta-Sandy Springs-Marietta, GA (partial) | 0.815 | ** | 0.059 | 0.000 | 1.144 | ^ | 0.073 | 0.067 |
| Baltimore-Towson, MD | 0.639 | ** | 0.061 | 0.000 | 0.905 | | 0.076 | 0.192 |
| Chicago-Naperville-Joliet, IL-IN-WI (partial) | 0.776 | ** | 0.048 | 0.000 | 1.087 | | 0.060 | 0.165 |
| Dallas-Fort Worth-Arlington, TX (partial) | 0.961 | | 0.053 | 0.456 | 1.057 | | 0.065 | 0.399 |
| Denver-Aurora, CO | 0.597 | ** | 0.061 | 0.000 | 0.825 | * | 0.076 | 0.011 |
| Detroit-Warren-Livonia,MI | 0.795 | ** | 0.065 | 0.000 | 1.295 | ** | 0.081 | 0.001 |
| Honolulu, HI | 0.670 | ** | 0.062 | 0.000 | 0.849 | * | 0.081 | 0.044 |
| Houston-Baytown-Sugar Land, TX | 0.932 | | 0.056 | 0.214 | 0.931 | | 0.069 | 0.301 |
| Kansas City, MO-KS (partial) | 0.616 | ** | 0.072 | 0.000 | 0.883 | | 0.091 | 0.170 |
| Las Vegas-Paradise, NV | 0.886 | ^ | 0.062 | 0.052 | 0.823 | * | 0.075 | 0.010 |
| Miami-Fort Lauderdale-Miami Beach, FL | 0.913 | | 0.060 | 0.128 | 1.017 | | 0.072 | 0.819 |
| Minneapolis-St Paul, MN-WI (partial) | 0.542 | ** | 0.056 | 0.000 | 0.783 | ** | 0.069 | 0.000 |
| NY-New Jersey-Long Island, NY-NJ-PA (partial) | 0.667 | ** | 0.042 | 0.000 | 0.858 | ** | 0.052 | 0.003 |
| Omaha-Council Bluffs, NE-IA | 0.678 | ** | 0.071 | 0.000 | 1.000 | | 0.088 | 0.999 |
| Philadelphia-Camden-Wilmington, PA-NJ-DE | 0.603 | ** | 0.050 | 0.000 | 0.941 | | 0.063 | 0.339 |
| Phoenix-Mesa-Scottsdale, AZ | 0.789 | ** | 0.071 | 0.001 | 0.974 | | 0.087 | 0.766 |
| Portland-Vancouver-Beaverton, OR-WA (partial) | 0.616 | ** | 0.073 | 0.000 | 0.900 | | 0.089 | 0.237 |
| Riverside-San Bernardino-Ontario, CA | 0.974 | | 0.067 | 0.695 | 0.978 | | 0.082 | 0.787 |
| St. Louis, MO-IL (partial) | 0.673 | ** | 0.072 | 0.000 | 1.092 | | 0.088 | 0.320 |
| San Diego-Carlsbad-San Marcos, CA | 0.637 | ** | 0.078 | 0.000 | 0.767 | ** | 0.096 | 0.006 |
| San Francisco-Oakland-Freemont, CA | 0.495 | ** | 0.068 | 0.000 | 0.663 | ** | 0.083 | 0.000 |
| Seattle-Tacoma-Bellevue, WA | 0.487 | ** | 0.069 | 0.000 | 0.733 | ** | 0.086 | 0.000 |
| Washington-Arlington-Alexandria, DC-VA-MD | 0.450 | ** | 0.046 | 0.000 | 0.662 | ** | 0.059 | 0.000 |
| Boston-Cambridge-Quincy, MA-NH | 0.500 | ** | 0.056 | 0.000 | 0.872 | * | 0.069 | 0.048 |
| Hartford-West Hartford-East Hartford, CT | 0.500 | ** | 0.079 | 0.000 | 0.829 | ^ | 0.097 | 0.054 |
| Providence-Fall River-Warwick, RI-MA | 0.657 | ** | 0.054 | 0.000 | 0.953 | | 0.068 | 0.483 |
| _cons | 0.522 | ** | 0.031 | 0.000 | 168.942 | ** | 0.137 | 0.000 |
| | | | | | | | | |
| Observations | 61,160 | | | | 61,160 | | | |
| Wald Chi2 | 704.6 | ** | | 0.000 | 12622.2 | ** | | 0.000 |

Table 4. Results for Model 5 (level-1 coefficients for Model 5 not reported)

| | Model 5 Multilevel logit on 27 Metros | | | |
|---|---|---|---|---|
| | OR | | Rob Std Err | P>\|z\| |
| Metro Demographics | | | | |
| % Unemployment | 1.028 | * | 0.014 | 0.048 |
| % Unionization | 0.988 | ** | 0.003 | 0.000 |
| Weighted Metro Min. Wage_2013 | 0.927 | ** | 0.026 | 0.004 |
| % Manufacturing Jobs | 1.017 | * | 0.007 | 0.024 |
| % Food Service & Retail Jobs | 1.034 | * | 0.014 | 0.015 |
| % College Grad in Labor force | | | | |
| % Foreign Born in Labor force | 1.000 | | 0.003 | 0.972 |
| Log Population Size | 1.030 | | 0.031 | 0.336 |
| _cons | 83.433 | ** | 0.681 | 0.000 |
| | | | | |
| Observations | 61,160 | | | |
| Wald Chi2 | 12465.40 | ** | | 0.000 |
| Random-effects Parameters | | | | |
| sd(unionization_rate) | 0.02 | ** | 0.003709 | 4.754 |
| sd(_cons) | 0.126871 | ** | 0.037901 | 3.347 |
| var(unionization_rate) | 0.000311 | * | 0.000131 | 2.377 |
| var(_cons) | 0.016096 | ^ | 0.009617 | 1.674 |
| cov(unionization_rate,_cons) | -0.002237 | * | 0.001087 | 2.058 |

Fig. 1 Percentage of metropolitan employment that is low-wage plotted against metropolitan union density. Circles are weighted by the count of low-wage workers in each metropolitan labor market.

Fig. 2 Estimated coefficient on union density for each metropolitan labor market plotted against metropolitan union density for the varying-intercept, varying-slope model. A nonparametric regression line fitted to the estimates is overlain for convenience.