HISPANICS AND ORGANIZED LABOR IN THE UNITED STATES, 1973-2007*

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ABSTRACT

Past research finds that minority populations in the United States secure unionized employed as part of the process of economic incorporation. Yet little work has systematically tested whether this pattern holds for the nation's largest minority, Hispanics, during recent decades of labor decline. After juxtaposing traditional labor market position theories of unionization with solidaristic accounts, we utilize 1973-2007 Current Population Survey (CPS) data to provide the most comprehensive analysis of Hispanics and organized labor in the United States to date. We disaggregate the Hispanic population by citizenship, nationality, and time since arrival to uncover subpopulation differences in unionization propensities. Additional analyses take advantage of the CPS structure to target individuals who join a union, allowing us to test whether the much-publicized efforts by organized labor to incorporate recent immigrants have resulted in detectable gains. Consistent with solidaristic accounts of labor organization, results suggest that certain Hispanic subpopulations – especially those born in the U.S. and immigrants who have secured citizenship – have higher unionization odds and join unions at higher rates than U.S.-born whites even after controlling for traditional positional accounts of labor organization. However, the large substantive effects of positional variables, such as sector, occupation, and firm size, indicate that organized labor's revival depends on more than any particular group's capacity for collective action.

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INTRODUCTION

The highly publicized (and ultimately successful) organizing campaign of largely Hispanic janitors in Southern California during the late 1980s provided a beleaguered labor movement with a rare source of hope after decades of steady membership losses. Nearing the 20th anniversary of the Justice for Janitors campaign, it is time to evaluate our knowledge of the changing relationships between the labor movement and Hispanics in the contemporary U.S. Despite a rebirth of academic interest in organized labor since that successful organizing drive, we still know surprisingly little about basic patterns of unionization among Hispanics after the dramatic expansion of the Hispanic population following the 1965 Immigration and Nationality Act. Much of sociologists' focus in recent years has been on case studies of successful and unsuccessful organizing drives. This research has generated important new theories about the interrelationships between labor organizing, immigrant incorporation, and Hispanics at the dawn of the 21st Century, and deepened our knowledge of organizing trends in particular locales (California especially). However, the lack of a broad, national-level focus has left gaps in our understanding of how representative such campaigns really are, whether the labor movement has made quantifiable inroads into organizing Hispanic natives and immigrants during the last decades, or whether the factors leading to or impeding unionization among Hispanics and Hispanic subpopulations differ from those of other groups.

These unanswered questions remain important for two main reasons. On the one hand, the extraordinary growth and diversity of the nation's largest minority population has sparked a large body of literature investigating Hispanics' changing labor market outcomes (for recent treatments, see Borjas 2006; Massey 2007: Figure 4.6; Reed and Danziger 2007). In decades past, minority populations – including immigrants and their offspring – have used union employment as a means of ascent to middle-class status (Lichtenstein 2002: 82-85; Milkman 2006: 119-122; Piore 1979: 156-157). Whether or not Hispanics and Hispanic subpopulations are able to secure membership rates similar to others speaks to their ability to rise economically. On the other hand, innovative factions within the battered U.S. labor movement have identified Hispanics – especially Hispanic immigrants – as potential sources for revitalization, given the rapid growth of the Hispanic population and its perceived capacity for collective action. Whether or not solidarity (actual or potential) on the part of Hispanics and certain Hispanic populations can overcome otherwise disadvantageous labor market positions and precarious legal statuses on a scale large enough to help staunch union membership losses remains unclear.

Figure 1 below casts the issues of an expanding Hispanic population and a devastated labor movement in sharp relief. Over the last three and a half decades, as the Hispanic proportion of the working population nearly quadrupled, unionization rates halved. Research on the timing of union decline and Hispanic growth in specific industries goes far toward dispelling any causal link between the two trends in the past (Allen 1994; Milkman 2006: 105-107; Waldinger et al 1998). However, hopes for the labor movement's revival may rest on whether organizers can capitalize on the burgeoning Hispanic population to swell union ranks in the future.

[FIGURE 1 HERE]

Our specific objectives are three-fold: First, using a dataset comprised of several series of Current Population Survey (CPS) data and spanning nearly four decades of growth and diversification in the Hispanic population, we provide the most comprehensive analysis of Hispanics and organized labor in the United States to date. Unlike the few other quantitative analyses on this topic, the thirty-five years of data assembled for the project allow us to determine whether and how the relationship between Hispanics and labor unions has changed since the dramatic increase of the Hispanic population. Given the enormous diversity within the Hispanic population (for an early overview, see Portes and Truelove 1987), we disaggregate models of union membership by subgroup, testing whether the factors affecting unionization differ according to whether one is an immigrant, time since immigration, citizenship, and national origin.

Second, since much of the rebirth of interest in Hispanic and organized labor among sociologists stems from contemporary developments, we create a unique panel dataset to capture the most current trends in individual-level change in union status. The few prior quantitative analyses on labor unions and the Hispanic population rely on membership rates. Such population-based measures combine those who just secured union employment with those who have worked under a union contract for decades. Our series of mini-panels captures within-individual variation in union status over a single year. These panels enable us to see if Hispanic ethnicity and various subgroup characteristics predicts ones odds of gaining a union job, thereby establishing whether the much-publicized efforts of organized labor to incorporate Hispanics and Hispanic immigrants into its fold have resulted in detectable gains during the early years of the 21st Century.

The quantitative analyses carry implications for theoretical accounts of immigrant incorporation and unionization. Our third objective is to clarify whether the relationships between Hispanics and organized labor follow labor market position accounts of unionization, or are suggestive of group-specific processes that would lend credence to solidaristic theories of worker mobilization. Whereas positional theories emphasize the differences in organizing costs attached to *workplace* characteristics, solidaristic accounts stress differences in organizing costs associated with *worker* traits. The rich set of covariates in the CPS datasets allow us to target positional factors that either impede or promote unionization. Unionization differentials that remain after controlling for labor market characteristics are suggestive of solidaristic explanations of unionization.

UNION REPRESENTION AND IMMIGRANT INCORPORATION

Solidaristic theories of unionization emphasize the ways in which group solidarity – ethnic or class-based – on the part of Hispanics and Hispanic subpopulations may raise or lower their costs of organizing. A labor market position account of unionization emphasizes the ways in which relatively stable industrial, geographic, and occupational factors structure unionization costs. These factors pattern unionization across specific groups. Below we discuss the theories and their applications to recent trends in unionization and Hispanics in the United States.

Solidaristic theories of union membership

The onset of severe deunionization in the 1970s coupled with a dramatic increase in the Hispanic population rekindled long-standing tensions between organized labor and the nation's fastest growing minority group. In Los Angeles, unionists were "openly hostile" to Hispanic workers pouring into the local labor market during the 1970s and 1980s (Milkman 2006: 114). Worries about the destabilizing impact of ethnic divisions surely motivated some of the hostility. In the early decades of the 20th Century, AFL craft unions in particular viewed outsiders as threats to solidarity and prevailing wage standards, with labor conflict erupting as levels of immigration increased or with the growth of non-white native populations (Olzak 1989). The implications of this historical legacy suggest that union leaders prefer organizing largely white, native workers. Given that the AFL-CIO reversed its restrictionist stance on immigration only in 2000, one might expect that organizers' fear of the Balkanizing effect of growing ethnic heterogeneity would extend to the close of the 20th Century.

This fear is echoed in recent comparative research on unionization rates and ethnic pluralism (Hechter 2004; Lee 2005). An influx of immigrants into a particular country is theorized to raise the costs of unionization by undermining local solidarity among the working class and ratcheting up competition between workers (Hechter 2004; Lee 2005). Evidence for the effect remains uneven, however. Whereas Lee finds that a percentage point increase in the net migration rate is associated with approximately a half percentage point decrease in unionization (80: Table 2, Model 5), Brady finds a *positive* effect of a country's net migration rate on its level of union membership (85: Table 4, Model 5). While these studies do not focus specifically on the effect of growing ethnic pluralism on *minorities*' unionization rates, one implication of this theory is that organizers worried about the destabilizing impact of status competition may avoid bringing newcomers into labor's fold.

In the context of U.S. immigration patterns, a weaker version of the solidarity thesis presented above suggests that part of the process of economic assimilation for immigrant populations and their offspring involves developing the capacity for class-based collective action in the U.S. (Cohen 1990: 324-325; Piore 1979; Sanchez 1993). Status-based identities lose their force as new populations become incorporated into the labor market and develop a capacity for collective mobilization that transcends ethnic lines. Such a development is unlikely to occur among the most recent arrivals, who must overcome cultural and legal hurdles to convince themselves and labor organizers that they are prime candidates for unionization. The class-based solidarity necessary for unionization is likely to develop over time and across generational divides. Thus we expect to see higher rates of unionization the longer a Hispanic immigrant has been in the U.S., and higher rates of unionization for Hispanics born in the United States.

Likewise, the potential for class-based collective action in the U.S. should be lower for arrivals who maintain strong ties across borders. Immigrants with a history of cycling between the U.S. and their home nation are more likely to compare their work conditions in the U.S. to those experienced in their homeland (Waldinger and Der-Martirosian 2000; Waldinger and Lichter 2003). This cross-national frame of reference is unlikely to translate into mobilization and agitation over unfair pay and practices in the U.S. Among Hispanic subpopulations, Mexicans move back and forth across the two nations' borders at relatively high rates. If such mobility inhibits the generation of labor solidarity in the U.S., we expect lowered odds of unionization for Mexican immigrants.

In the most comprehensive quantitative test of Hispanics and unionization in the U.S. to date, Waldinger and Der-Martirosian (2000) model the likelihood of union membership for immigrants (Hispanics and others) using a pooled sample of CPS-March data that spans the mid-1990s. The authors find that American-born Mexicans hold union jobs at similar rates as whites, while other American-born Hispanics exhibit a higher likelihood of unionization compared to whites after controlling for relevant demographic, geographic, and industry characteristics. Recent Mexican immigrants, however, have a lower likelihood of unionization (see their Figure 2.2), buttressing claims that cross-national identification among recent Mexican arrivals leads to depressed unionization rates in the U.S. Blanchflower, in his investigation of union predictors in the U.S., U.K., and Canada, reports slightly higher probabilities of unionization for Hispanics compared to whites in the U.S. between 1984 and 2002 (2007: see Tables 4 and 5). By the end of his series, however, the Hispanic differences fail to reach statistical significance. Blanchflower does not disaggregate the Hispanic population by immigrant status or nationality, so growing heterogeneity within the Hispanic population resulting from expanding immigration could explain the decline in unionization probabilities, and may provide evidence of the lack of class-based solidarity among recent arrivals.

Solidaristic accounts that stress the destabilizing effects of increased immigration conceptualizes an increase in ethnic heterogeneity as undercutting economic solidarity within the working class. But what if that solidarity does not exist? The decades-long decline in unionization in the U.S. has left many workers unaware of and unfamiliar with the potential benefits of union employment. Other research on Hispanic subgroups in the late-20th Century U.S. suggests a more complex relationship between ethnicity, immigration, and labor solidarity (for a few notable examples, see Fantasia and Voss 2004: 134-150; Lichtenstein 2002: 262-268; Milkman 2002; 2006: 114-116). Many immigrants arrive from countries with a robust and vibrant labor movement, or with past experiences of other forms of collective mobilization (Lichtenstein 2002: 267). These experiences in their homelands may make more *recent* arrivals receptive to union efforts, lowering their costs of organizing (Milkman 2006: 133). Indeed, in a recent survey, Weir finds greater support for unions among non-union, non-citizens in California than non-union citizens (2002: Figure 4.10). Milkman suggests that the "shared experience of stigmatization" may help foster group consciousness in employment settings dominated by Hispanic immigrants (2006: 133); a consciousness that can be capitalized on by enterprising organizers. Thus, it could be that solidarity on the part of newcomers may compensate for the lack of class organization or other experiences with collective mobilization on the part of native-born Americans.

The various solidaristic models of unionization lend themselves to certain competing hypotheses. Controlling for relevant labor market position variables, if union organizers fear the potential destabilizing effects of incorporating Hispanics and Hispanic immigrants into labor's fold, we expect to find depressed unionization rates for Hispanics, especially at the beginning of our series, prior to the AFL-CIO's reversal of its immigration stance. From 1994 on we can distinguish immigrants from non-immigrants in the CPS datasets, as well as time since arrival and country of origin. If the economic incorporation experiences of Hispanics follow the pattern of prior immigrant populations, we expect to find increasing odds of organization the longer a Hispanic immigrant has been in the U.S. We also expect lower unionization probabilities for Mexican migrants, given their higher rates of return migration (Waldinger and Der-Martirosian 2000; Waldinger and Lichter 2003).

The hypotheses presented above presume an existing level of class-solidarity among non-Hispanic, non-immigrant Americans, and a corresponding fear among labor organizers of incorporating Hispanics – especially recent arrivals – into labor's fold. If, on the other hand, immigrants' stronger networks or past experiences with worker mobilization lower their costs of organizing substantially, we would expect their odds of unionization to exceed other groups, especially as unionization rates begin to bottom out among native-born populations. Non-Mexican Hispanic migrants should have higher odds of unionization relative to other groups, due to the large presence of political refugees and labor activists among the population (Milkman 2006: 137; Waldinger et al 1998: 117).

Labor market position model of union membership

Solidarity means little if Hispanics concentrate in unorganizable sectors of the economy. A positional theory of union organizing focuses on the ways in which relatively stable industrial, occupational, and geographical factors affect unionization costs. Hispanic unionization rates will vary according to how Hispanic employment maps onto the labor market positions of the economy that determine the costs of unionization in the institutional context of the modern U.S.

One such feature is high labor costs relative to other factors in production (Western 1994). In industries with high wage bills combined with largely unskilled employment, the barriers to unionization are especially substantial: not only do employers' incentives against unionization rise with relative wage costs, the substitutability of labor lowers workers' bargaining leverage. Smaller firms raise the costs of unionization as well, as monitoring costs are low, thereby negating a potential benefit of union presence (Hirsch and Addison 1986: 61; for evidence of firm size effects on Hispanic immigrants' unionization rates, see Defreitas 1993). Productivity gains spurred by union grievance procedures (and the resulting happier workforce) are likely lower in small firms, adding further to the costs of unionization (Freeman and Medoff 1984; Western 1994: 498).

In modern capitalist economies, these characteristics – high wage bills, unskilled employment, and smaller firms – are found in many service industries, and contribute to the comparatively low unionization rates found in the service sector. By contrast, the capital-intensive manufacturing sector, with large, hierarchically arranged firms, constituted the unionized core of the economy in mid-20th Century America and many other developed nations. That unionization rates have not declined in the U.S. and elsewhere even further with the collapse of manufacturing is due to robust public sector unionization rates. Costs of unionization in the public sector are lower for both employers, due to a looser connection between employment and wage levels (Western 1994), and for union organizers, who can bypass the increasingly onerous demands of a union election drive and exert pressure on political actors to unionize government workers.

Industry *location* affects unionization costs independent of characteristics specific to the industry itself. High density locales reduces costs for organizers, leading to greater unionization rates in metropolitan areas. In the U.S., state-level differences in labor laws and the (related) less union-friendly political environment of the South may depress unionization rates, other factors held constant. Occupation factors into unionization patterns as well. Costs of organizing white collar workers – with well-defined career ladders, higher levels of workplace autonomy, and greater pay – are higher than their blue-collar counterparts, who are less likely to identify with management in battles over representation (Freeman and Medoff 1984; Western 1994).

Positional accounts of union membership imply distinct hypotheses from solidaristic theories. Our data contains direct measures of the labor market position factors most likely to affect unionization rates. By nesting models and introducing key positional variables, we can determine which positional factors dampen or increase Hispanic unionization rates relative to other groups. Controlling for such factors should result in predicted probabilities of unionization equal to other populations. Should Hispanic or Hispanic subpopulation differences remain, we can compare their magnitude to those positional variables emphasized in the traditional unionization literature.

DATA AND METHODS

The CPS is a monthly survey of approximately 60,000 households, containing information on demographic and labor force characteristics. (For details on the structure of the CPS, see Appendix A). To estimate Hispanic unionization probabilities we rely on various series of the CPS, 1973 to 2007. Prior to 1982, union status questions were asked only in the May sample: For all pre-1983 estimates, we utilize data from the CPS-May surveys.[1] Beginning in 1983, union questions were included in basic questionnaire. For the annual estimates we present in Figures 1 and 2, we prefer the merged outgoing rotation group (CPS-MORG) data for the years 1983 to 2007 due to the much larger sample sizes of the CPS-MORG series. The March data include certain key covariates missing from the CPS-MORG, including a question on firm size. Thus for the cross-year models presented in Tables 2-4, we use CPS-March survey data. In all of our analyses, we restrict the data to non-self-employed workers, age 18 to 65.[2]

For the final set of analyses presented in Table 5 we take advantage of the limited longitudinal feature of the CPS to construct 1-year mini-panels using the CPS-MORG data. Because all CPS respondents are interviewed for a period, are then re-interviewed after a hiatus, a portion of any given year's sample is re-interviewed the following year. We capitalize on the survey's structure, and match a respondent's data across these two observation points to identify changes in key characteristics over a one-year period. These mini-panels allow us to isolate the population of respondents who gain union status across a particular year. For our final models, we restrict the sample to the last five years of available data, providing insight into the most recent trends in unionization. For a more thorough discussion of the matching procedure used in our analysis, see Appendix A.[3]

[1] No union questions were asked in 1982. To estimate 1982 unionization rates and probabilities, we average results from 1981 and 1983. For a discussion of union variables in the various CPS series, including changes in question wording beginning in 1977, see Hirsch, Macpherson, and Vroman (2001).
[2] The CPS asks union questions of non-self-employed wage and salary questions only. The vast majority of U.S. workers do not own their own business, and self-employment rates for Hispanics and Hispanic immigrants are lower than the national average (Alba and Nee 2003: 235).
[3] Madiran and Lefgren (1999) provide more technical details on matching procedures with CPS datasets. For a recent sociological analysis using matched CPS mini-panels, see DiMaggio and Bonikowski (2008).

Table 1 below presents sample sizes for the various data sets used in the analysis along with unweighted descriptives of key covariates (a full set of descriptives for all controls used in the models are available from the authors on request). Interestingly, in the early years of our samples Hispanic unionization rates actually exceed those of others by three percentage points. The surveys covering later years – CPS-MORG and CPS-March – suggests a temporal shift as Hispanic unionization rates average slightly below those of non-Hispanics. Table 1 includes transportation industry means as well as averages for two occupational categories due to the strong effects these controls have in subsequent models; a listing of the sixteen additional industry dummies and two other occupational dummies included in the models is available in Appendix B.

[TABLE 1 HERE]

For the main set of analyses, our outcome variable of interest is a binary measure of union status. Given that our dependent variable is dichotomous, we fit a series of logistic regression models for these analyses. For individual *i*,

$$\eta_{ij} = \log \frac{\pi(union)_i}{1-\pi(union)_i} = \alpha + H_i\beta + X_i\gamma + \varepsilon_i$$

where *H* captures whether the respondent is Hispanic, X_i is a set of demographic, geographic, and socioeconomic covariates, and ε_i is residual individual-level variation. We present odds ratios from the logistic regressions due to their ease of interpretability, supplementing key results with estimates of the predicted change in unionization probability given a particular outcome of a covariate of interest.

The main models contain a four-category race/ethnicity measure (white, African-American, Hispanic, and other). The Hispanic subsample analyses disaggregate the four-category measure by time since immigration, nationality, and citizenship. The other demographic controls included in all analyses are sex, education, a dichotomous measure of current marital status, age, and age squared. Research has demonstrated that gender, age, and education potentially alter the costs of unionization, although the effects are inconsistent. Research has found a higher likelihood of membership among older workers (Waldinger and Der-Martirosian 2000), however, other work has shown higher levels of support for unions among younger non-members (Freeman and Medoff 1984: 31). Traditionally, unionists have been largely "'3M' (male, manual, manufacturing)" workers, although the rapid growth of public sector unionization has tempered the gender gap in representation (Visser 2002: 405). The growth of public sector unions relative to once-strong manufacturing unions may have also altered the relationship between education and unionization: Whereas past theory suggests that unions benefit less-educated workers with fewer outside employment prospects (Freeman and Medoff 1984: 98), more recent empirical work has found a positive union-education relationship (Waldinger and Der-Martirosian 2000: Appendix A). To the extent that these demographic factors affect unionization costs, groups that vary by age, gender, and education will have different membership rates.

The geographical controls we introduce include four dummy measures of metropolitan residence, four region dummies, and – in the most restrictive models – state fixed effects. Our full model includes all of the demographic and geographic controls described above, plus a dichotomous sector measure, occupation and industry fixed effects. In the full models we substitute a potential experience (age minus years spent in school) and potential experience squared measure for the age and age squared variables. Year dummies are included in the cross-year models presented in Tables 2 and 3. For supplementary models presented in Tables 2 and 3, we include measures of firm size. We present the controls used for the various models in more detail in Appendix B.

Since many unions push to convert part-time positions to full-time during contract negotiations, in the models presented here we do not include a measure of full-time status. However, given the strong positive relationships between ethnicity, immigrant status, and contingent work arrangements (including part-time status) (Kalleberg, Reskin, and Hudson 2000), and the strong negative relationship between unionization and contingent work arrangements, as a robustness check we include a full-time indicator for our main models; results are substantively similar from the ones presented in the article and available upon request.[4]

The final set of analyses use the MORG mini-panels to isolate the population of union joiners for the 1996-2007 period.[5] Here our dependent variable of interest has four categories defined by the respondent's union status at time 1 and time 2: one can never have been in a union across the two time periods (0,0), one could have left a union (1,0), always been in a union (1,1) or joined a union (0,1). To estimate the probability of joining a union relative to having not been in a union at time 1, we fit a multinominal logistic regression model. For individual i and category j of the dependent variable with J categories, the model takes the general form:

$$\eta_{ij} = \log \frac{\pi_{ij}}{\pi_{iJ}} = \alpha_j + H_i \beta_j + X_i \gamma_j + \varepsilon_{ij},$$

[4] We also test to see whether our model results differ when controlling for whether the respondent moved in the prior year (a variable only available in a subset of the 1983-2007 CPS-March series). Results remain unaffected; truncated models with the move covariate are available upon request. We do not include an income measure due to endogeneity, as unionization has long been associated with higher wages.
[5] We modeled the population of joiners for the entire 1983-2007 period, but only present here the results from more recent years. Full model results are available from the authors on request.

where α_j is a constant, H captures whether the respondent is Hispanic, X_i is a set of demographic, geographic, and socioeconomic covariates, β_j and γ_j are vectors of regression coefficients, for j=1,2,…, J-1 and ε_{ij} is residual individual-level variation. π_{ij} reflects the probability of being in outcome category j for each individual i. This model is analogous to a series of J-1 binary logistic regression equations that contrast each outcome j against a common category J.

The controls for these models largely mirror those previously discussed for the main analysis, with the following exceptions. In these models we use region dummies rather than state fixed effects, owing to computation challenges in estimating the more complex multinomial logit model with year, state, industry, and occupation fixed effects.[6] The mini-panel models also include controls for experiencing a change in industry or occupation between the two time periods, and a specification limited to only those individuals with stable industry and occupation codes across the two periods.[7]

RESULTS

Modeling unionization among Hispanics

In Table 2 below, we begin the multivariate analyses by testing what variables help explain unionization differences between Hispanics and others for a subset of our time period, 1983-2007, using CPS-March files.[8] The first column presents odds ratios

[6] Checks of reduced models using state fixed effects result in substantively similar findings.
[7] A lack of information on specific job moves in the CPS-MORG series prevents us from restricting the sample to only those individuals with the same employer across panels.
[8] We exclude the May years of the survey (1973-1981) for these models for the following reasons: Lack of state identifiers for most states prior to 1977 prevent the inclusion of state fixed-effects, and inconsistencies in the industry and occupation codes between the CPS-May and CPS-March surveys prevent a seamless comparison between survey years. CPS-May data is utilized for the individual year models presented in Figure 2.

from a simple model including only demographic controls and year fixed effects. Without accounting for labor market position variables, Hispanics have 6% *higher* odds of unionization compared to whites. The second column adjusts for positional differences between Hispanics and other groups. The inclusion of this battery of measures more than triples the number of parameters in the model. The inclusion of positional controls results in a non-significant Hispanic odds ratio, suggesting that the positive Hispanic odds ratios in the previous column are due in part to Hispanics' concentration in labor market positions with lower unionization costs.

[TABLE 2 HERE]

The final column in Table 2 includes all the variables of the labor market position model plus a set of firm size dummy variables (firm size measures are available only in a portion of the CPS-March surveys). Odds ratios of unionization for employees in the largest firm size category – 1,000 workers or more – are over four times those for employees working in the smallest category of less than 25 employees, consistent with structural explanations for inter-firm union variation (Hirsch and Addison 1986; Western 1994). The firm size controls do not affect the Hispanic odds ratios. Hispanics as a whole seem no more or less "organizable" than whites; differences in raw unionization rates observed in Table 1 can be attributed to differences in positional characteristics that pattern unionization rates in the contemporary U.S.

The cross-year models presented in Table 2 obscure any time variation in the relationship between Hispanics and unionization. In Figure 2 below we present odds-ratios for individual-year models using CPS-May and CPS-MORG data. The point estimates represent Hispanic odds ratios of unionization and are derived from models run on each year of survey data. The reference category is white workers' odds of unionization. The dark line represents models that only include race/ethnicity identifiers. As shown, the series trends sharply downward over the three and a half decades covered by our data. Whereas in the mid-1970s Hispanics had around 20% higher odds of unionization than whites, by the late 1990s Hispanics' odds of unionization stabilized at about 20% *lower* than those of whites.

To investigate whether the temporal shift reflects positional differences between the populations or results from other factors, we next run labor market position models on each survey year. Models used to generate the series include the same set of covariates shown in Table 2 with the exception of firm size dummies (unavailable in both the CPS-May and CPS-MORG surveys).[9] Two results from the picture stand out: One, for the majority of the series the point estimates hover right near one, indicating little difference in unionization odds between Hispanics and whites once one adjusts for differences in labor market position. Indeed, the Hispanic odds ratio is significant for less than a third of the years in our data (p-values and confidence intervals for the 68 models used to generate Figure 3 available upon request). Two, the trendline shows a slightly downward slope. Beginning in the mid-1990s, Hispanics' odds of unionization fall below one; for five out of the last six years of data, the difference in odds ratios between whites and Hispanics is significant at the .05 level. To investigate this temporal pattern in more detail, we now turn to analyses that disaggregate the Hispanic population by time since immigration, citizenship, and nationality.

[FIGURE 2 HERE]

[9] Due to variable differences between the May and MORG series, inter-year comparisons should be made with some caution. For geographic controls in the pre-1977 May series, we include dummies for the 12 states and District of Columbia that are identified, and dummies for the 10 multi-state groupings.

Modeling unionization among Hispanic subpopulations: 1994-2007

In the tables that follow, we present only the odds ratios for various Hispanic subcategories. Except for the models restricted to immigrants only, the reference category is non-immigrant whites. All models include the full set of covariates from Table 2; coefficients for all control variables are available upon request. The first set of models examines the effect of time since migration on the odds of unionization. A clear cleavage emerges between recent Hispanic arrivals and non-immigrant whites: Hispanic immigrants who settled in the U.S. within the last five years had less than half the odds of unionization compared to non-immigrant whites. Those Hispanic immigrants who have been in the U.S. for more than two decades exhibit no difference in unionization than non-immigrant whites, while Hispanic natives have 20% *higher* odds of belonging to a labor union than their white counterparts.

The time since migration effects evident here are broadly consistent with past research on the topic (Defreitas 1993; Waldinger and Der-Martirosian 2000), and may be suggestive of organizers' worries about the destabilizing impact of recent immigrants, or may reflect uncaptured labor market features unique to recent Hispanic arrivals. One such feature is firm size. Research has shown how recent arrivals often rely on entrenched immigrant networks for finding employment (for a recent overview of this literature, see Alba and Nee 2003: 163-166; for recent empirical work on network effects and Mexican immigrant incorporation, see Aguilera and Massey 2003). These networks often work to steer migrants to ethnic workplaces, often of small size (Aldrich and Waldinger 1990). The second column includes a set of firm size indicators. While the difference in odds ratios between recent Hispanic arrivals and white natives diminishes slightly, the overall story of depressed odds for recent Hispanic arrivals remains. The final column limits the sample to employees in firms with 1,000 or more workers. With this further restriction, except for Hispanic immigrants who have been in the U.S. between five and ten years, Hispanic immigrants no longer differ from non-immigrant whites. This finding suggests that the strong negative odds of unionization in the prior models may be partly due to recent arrivals' concentration in smaller firms, where unionization costs are high.[10]

[TABLE 3 HERE]

Next we subdivide our race/ethnicity and immigrant categories by citizenship status. The lowest odds of unionization are for non-citizen Hispanic immigrants. This remains true (although the effect is slightly diminished) when restricting the sample to employees in the largest firms, perhaps indicative of organizers' resistance to unionizing non-citizen immigrants, or immigrants' worries about the legal ramifications of an organizing drive. The third set of models isolates Mexicans from the rest of the Hispanic population. Prior research suggests that Mexicans have lower odds of unionization than other Hispanic immigrants (Waldinger and Der-Martirosian 2000: 53-54). Additional research suggests that Central Americans often arrive with past experiences of labor militancy and other solidaristic ties likely to boost their odds of unionization (Milkman 2006: 137; Waldinger et al 1998: 117). The labor market position model indicates that Mexican immigrants indeed have lower odds of belonging to a labor union than non-Mexican migrants, but the effect disappears when restricting the sample to individuals in

[10] The non-significant effect for the earliest arrivals remains somewhat puzzling. Sample sizes for certain immigrant subcategories run quite low, and this result may reflect a lack of power in the model.

large firms. The durable fault lines seem to reside along citizenship and time since arrival, not nationality.

The final set of models in Table 3 restricts the sample to immigrants only. Here we test whether Hispanics' odds of unionization differ from other immigrants' odds along the two key dimensions of citizenship and nationality. These models control for years since immigration. For the citizenship analyses, the reference group is white immigrant citizens. For the nationality investigation, the reference group is white immigrants. Controlling for duration in the U.S., no evidence exists to suggest that organizers resist unionizing Hispanics compared to other immigrant populations; indeed, among Hispanic citizens, these models reveal a higher likelihood of being organized. Interestingly, non-citizen Hispanics have similar unionization odds as white immigrant citizens, suggestive that group solidarity can partially overcome barriers to unionization associated with non-citizen status. The nationality investigations reveal no significant differences between Mexican immigrants, non-Mexican Hispanic immigrants (mostly Central Americans), and whites.

The results from Table 3 reveal important subpopulation differences among Hispanics. The question remains how these findings compare in magnitude to the effects of key labor market position determinants of unionization. For the citizenship model in which we control for firm size, the 34% lower odds of unionization among Hispanic non-citizens translates into a predicted probability of belonging to a union 3.6 percentage points lower than the unionization probability of a Hispanic non-immigrant, setting the other covariates at their mean values. Other predictions (not shown; available upon request) reveal a roughly similar negative difference in the probability of unionization between the most recent Hispanic migrants and Hispanic non-immigrants. To put these differences in context, in Table 4 below we present differences in predicted probabilities for some of the major structural covariates in our model. The influence of citizenship status and time since migration on unionization probabilities is dwarfed by the positional characteristics of firm size, sector, and occupation.

[TABLE 4 HERE]

Modeling transitions in union status

Union membership rates combine individuals who have belonged to a union for decades with those who joined a week prior to the survey. To target the latest developments in unionization, we use the 1-year mini-panels created from the CPS-MORG series to isolate the population of union joiners (see the details of constructing this dataset in the earlier data section or Appendix A). In Table 5 below, we again disaggregate the Hispanic population by duration of U.S. residence, citizenship status, and nationality, and model the odds of joining a union relative to never being in a union. We present results from four specifications: one capturing labor market positional factors, a second including a dummy variable to account for any change in industry or occupation across the two observations, a third restricting the analysis to those individuals who remained in the same occupation and industry across the two time periods,[11] and finally a replication of the labor market position specification limited only to the last five years (2002-2007).

[11] Substantive results are similar if we instead limit to only individuals who do change industry or occupation.

The first three sets of models corroborate the findings from Table 3: The odds of joining a union are approximately 22-33% lower for the most recent Hispanic immigrants relative to non-immigrant whites. These depressed odds translate to less than half a percentage point difference in the probability of joining for recent immigrants compared to Hispanic non-immigrants, except in the last model (2002+), where there is a one point difference. Non-immigrants and immigrants living in the U.S. for 20 or more years show 36-46% higher odds of joining a union than native whites. These odds translate into approximately a 1-1.5 percentage point increase in the probability of joining a union. The targeting of recent Hispanic immigrants (many of them undocumented) discussed in the case study literature (for examples see Rudy 2004; Wells 2000) has not yet translated nationally into elevated odds of joining a union among non-citizens, as least compared to white natives.[12] Native-born Hispanics and Hispanic immigrant citizens have roughly 30-43% higher odds of joining a union relative to native whites; these elevated odds translate into a 1-1.3 point increase in the percent joining a union. The next set of models again suggests that there is nothing about *Mexican* immigrants that impede their organization: they show no statistical differences with native whites in joining a union. Non-Mexican Hispanic immigrants, along with Mexican-origin and other Hispanic-origin natives have odds 22-60% higher than native born whites, translating to a 0.7 to 1.7 point difference in the probability of joining a union.

[12] We also model (results available upon request) the odds of leaving a union relative to being in a union at time 1 and 2. Generally, Hispanics (and African-Americans) have a higher likelihood of leaving a union compared to whites due, we suspect, to higher rates of job turnover within the same industry and occupation. Specifically, we suspect that many changes out of union status stem from leaving a unionized firm, and do not reflect decertification campaigns. Unfortunately, indicators of firm change are unavailable in the CPS-MORG data to test this hypothesis.

Results of the first three sets of models suggest duration of residence in the U.S. is the primary hindrance to joining a union among Hispanics and Hispanic immigrants. While non-citizens and Mexican immigrants do not display higher odds of joining a union than native whites, this may simply result from the short duration of stay in the U.S. among these groups. We also examine the influence of citizenship and nationality, controlling for duration of U.S. residence in models limited to immigrants. These results (reported in the final section of Table 5) confirm that net of the influence of duration of residence in the U.S., Hispanic non-citizens and Mexican immigrants are no less likely, and in some cases are more likely, to join a union than other immigrants.

[TABLE 5 HERE]

DISCUSSION AND CONCLUSIONS

Minority populations in the past secured union employment to ascend economically, reshaping the demographic profile of the American middle-class. The results of our analyses suggest that this process of economic assimilation remains similar for Hispanics in the contemporary U.S. Incorporation into organized labor is not immediate, as the odds of unionization increase dramatically with time spent in the U.S. This pattern also has historical precedent, as union queues are often long, and immigrants' first U.S. jobs often rely on informal networks that steer the arrivals toward relatively unorganized niches (Alba and Nee 2003: 232-235; for the role of subcontractors in the hiring of recent Hispanic arrivals, see Massey 2007: 143-146). Economic ascent among immigrant populations and their offspring typically occurs after transitioning into the open labor market, where union jobs are more likely to be found.

These results remain relatively unsurprising, given past research on minority populations and economic incorporation.

What is surprising is the evidence we uncover revealing higher unionization probabilities among certain Hispanic subpopulations. Even in our most restrictive models (last column of Table 3), Hispanic immigrant citizens have much higher odds of unionized employment than native whites, controlling for a battery of relevant labor market position variables. The analyses using mini-panels (Table 5) reveals positive odds of joining a union compared to whites except in the cases of the most recent arrivals, non-citizens, and Mexican immigrants (where the results are non-significant). Among immigrant populations, Hispanics tend to have higher probabilities of being in a union (Table 4), and of joining a union (Table 5). Taken together, these results suggest that solidaristic ties help Hispanics overcome other barriers to unionization, and point to a promising site for labor organizers desperate for new members. To the extent that union membership continues to provide a toe-hold to the middle-class among economically disadvantaged groups, these findings portend economic gains over time and across generations for Hispanic immigrants and their offspring.

The results from the nationality models and models restricted to immigrant populations offer insights into various proposed mechanisms stemming from solidaristic theories of collective action. We find very little evidence to buttress claims suggesting organizers avoid Mexicans, given high rates of return migration among this Hispanic subpopulation. While Mexican immigrants seem to join unions at lower rates than Mexican non-immigrants and non-Mexican Hispanic migrants, their odds of joining rival those of native whites (Table 5). Membership rates among Mexican immigrants trail those of native whites, but the difference disappears after restricting the analysis to workers in large firms (Table 4). Evidence of greater capacity for collective mobilization among non-Mexican Hispanic immigrants is likewise mixed: While membership rates for non-Mexican Hispanics is comparatively high in some of our models, the results are not robust to restricting the sample to large firms. However, the final models in Table 5 do reveal higher odds of joining unions for non-Mexican Hispanic immigrants compared to other immigrant populations, perhaps indicative of past experiences with collective action and labor organization among these (largely Central American) migrants.

While we believe these insights help illuminate our understanding of the dynamics between Hispanic subpopulations, solidarity theories, and labor organization over the past decades, certain shortcomings remain. Similar to past research on the topic, our models lack exact tests that can differentiate between various hypotheses stemming from solidaristic theories of group mobilization. We cannot determine precisely, for example, whether the elevated odds of unionization for Hispanic subpopulations uncovered in our analyses stem from past experiences of labor organization or from organizers' ability to capitalize on ethnic-based solidarities. Definitive statements regarding the precise characteristics specific to Hispanic subpopulations leading to high unionization rates await further in-depth qualitative work.

Regardless of the precise mechanisms, the elevated odds of joining a union for many Hispanic populations should encourage labor organizers. Yet the ability of Hispanics to reshape fundamentally the institutional underpinnings of organized labor appears limited. The substantive effects of immigrant status and ethnicity pale in comparison to positional factors like sector, occupation, and firm size (Table 4). At the individual level, Hispanic ethnicity, immigration status, and time since immigration remain statistically significant indicators of unionization; however, recent growth in these subpopulations has little aggregate effect on the overall unionization rate, given their comparatively small substantive effects combined with the shrunken fraction of all workers in labor unions. It terms of one's union status, it seems to matter much more where you work (public versus private sector; small versus large firm) and what kind of work you do rather than your ethnicity and immigration status.

Capacities for collective mobilization among Hispanics can certainly help labor's revitalization. Yet the results from this analysis suggest that a dramatic turnaround rests on two other developments: first, a thawing in anti-immigration rhetoric and policies to capitalize on greater group solidarity among Hispanic subpopulations. Citizenship represents a durable fault line in unionization probabilities; should pathways to citizenship be further restricted, we can expect unionization rates among non-citizen Hispanics to decrease, threatening their economic gains. Second, the institutional framework that has constrained organizing efforts in the U.S. for over half a century now must change if labor hopes for a turnaround. The substantive effects of the positional variables uncovered in our models highlight the difficulty of organizing high-growth service industries in the private sector. Altering this equation likely requires dramatically changing the legal-political environment surrounding unions and employers in the U.S. Without such a development, unionization rates will remain at historic lows, regardless of any particular group's capacity for collective action.

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Table 1. Descriptives of various CPS data sources.

	May	MORG	March	MORG Matched^
Percentage Hispanic	4.7	8.1	8.6	8.8
Percentage Hispanic immigrant (1994+)	n/a	5.2	4.6	4.5
Percentage Union	23.9	15.6	15.8	15.5
Percentage union Hispanic	26.9	13.5	14.5	13.6
Percentage union non-Hispanic	23.8	15.7	15.9	15.6
Selected demographic controls				
Average age	37.2	38.2	38.3	40.8
Percentage male	57.7	51.4	50.9	50.5
Percentage African-American	9.1	9.4	9.8	8.6
Percentage less than HS	22.7	11.3	11.2	8.35
Percentage HS degree or equivalent	38.1	34.3	33.7	31.8
Percentage some college	21.2	28.2	28.5	29.6
Percentage BA or higher	18.1	26.3	26.7	30.3
Percentage married	70.0	62.4	63.2	66.4
Selected labor market position controls				
Percentage Southern	29.2	29.6	29.2	28.9
Percentage private sector	80.4	82.4	82.0	80.8
Percentage transportation industry	3.6	4.5	4.5	4.7
Percentage professionals / managers	26.3	30.5	29.2	33.6
Percentage production / craft / repair occupations	34.3	25.8	25.2	23.5
Years of sample used	1973-1981	1983-2007	1983-2007	1996-2007
N	413,821	4,078,063	355,857	531,203

Note : Unweighted means presented. Descriptives for all covariates used in the models availabe upon request.

^Matched dataset for immigrants begins in 1996 due to a lack of adequate identifiers in 1994 and 1995.

Table 2. Odds ratios from logistic regressions predicting unionization, 1983-2007.

	Demographics Model		Labor Market Postion Model		Firm Size Model	
Demographic controls:						
Hispanic (reference is white)	1. 06**	(1.02, 1.11)	. 98	(.93, 1.03)	1. 02	(.96, 1.08)
African-American	1. 51***	(1.46, 1.56)	1. 52***	(1.46, 1.59)	1. 41***	(1.35, 1.48)
Other race	. 93**	(.88, .98)	. 73***	(.68, .78)	. 70***	(.65, .76)
Male	1. 61***	(1.58, 1.65)	1. 26***	(1.23, 1.29)	1. 23***	(1.19, 1.27)
Married	1. 09***	(1.06, 1.11)	1. 15***	(1.12, 1.18)	1. 13***	(1.10, 1.17)
Age (experience)	1. 17***	(1.17, 1.18)	1. 08***	(1.08, 1.09)	1. 08***	(1.07, 1.08)
Age^2 ($experienced^2$)	. 99***	(.99, .99)	. 99***	(.99, .99)	. 99***	(.99, .99)
H.S. (reference is H.S. dropout)	1. 46***	(1.41, 1.52)	1. 43***	(1.37, 1.49)	1. 42***	(1.34, 1.50)
Some college	1. 27***	(1.22, 1.32)	1. 38***	(1.32, 1.45)	1. 34***	(1.26, 1.42)
College degree or higher	1. 23***	(1.18, 1.28)	1. 51***	(1.43, 1.59)	1. 42***	(1.34, 1.52)
Labor market position controls:						
Private sector			. 13***	(.13, .14)	. 16***	(.16, .17)
Selected industries (reference is ag./forestry/fisheries)						
Transportation			8. 50***	(6.83, 10.57)	6. 88***	(5.25, 9.02)
Mining			5. 25***	(4.09, 6.73)	3. 94***	(2.89, 5.36)
FIRE			. 89	(.71, 1.13)	. 74*	(.56, .99)
Selected occupations (reference is prof./managerial)						
Production / craft / repair			3. 84***	(3.68, 4.00)	4. 18***	(3.98, 4.39)
Service occupations			1. 24***	(1.20, 1.28)	1. 28***	(1.23, 1.33)
Firm size (1,000+ vs. <25)					4. 33***	(4.11, 4.57)
State and metro effects	No		Yes		Yes	
Year effects	Yes		Yes		Yes	
N	355,857		355,587		282,090	
Number of parameters	34		107		106	
McFadden's R^2	. 04		. 23		. 25	
BIC	296,871		240,896		180,605	

Note : Robust 95% C/Is in parentheses.

*p < .05 **p < .01 ***p < .001

Data come from the March files of the CPS, 1983 - 2007. All models weighted with the appropriate March CPS weights. Three firm size items, thirteen industry dummies, and an occupation control included in the final model but not shown. See Appendix B for further detail on the control variables used in the analysis.

All coefficients from suppressed covariates available upon request.

Table 3. Hispanic immigrant subcategories analysis: odds ratios from logistic regressions predicting unionization, 1994-2007.

		Labor Market Position		Firm Size		1,000+ Firms Only	
1). Years since immigration	Hispanic non-immigrant	1. 20***	(1.10, 1.31)	1. 16**	(1.06, 1.27)	1. 10	(.98, 1.23)
	Hispanic immigrated: 20+ years	1. 05	(.92, 1.19)	1. 10	(.96, 1.26)	1. 11	(.91, 1.36)
	Hispanic immigrated: 10 to 20 years	. 75***	(.64, .88)	. 83*	(.71, .97)	1. 04	(.81, 1.35)
	Hispanic immigrated: 5 to 10 years	. 44***	(.35, .55)	. 51***	(.40, .65)	. 46**	(.28, .73)
	Hispanic immigrated: less than 5 years	. 41***	(.30, .56)	. 50***	(.36, .68)	. 71	(.43, 1.18)
	N	202, 975		199,493		84,194	
	Number of parameters	112		116		112	
	McFadden's R^2	. 23		. 25		. 24	
	BIC	129,120		124,416		65,990	
2). Citizenship	Hispanic non-immigrant	1. 20***	(1.10, 1.30)	1. 16**	(1.06, 1.27)	1. 10	(.98, 1.23)
	Hispanic immigrant citizen	1. 15	(.99, 1.32)	1. 20*	(1.04, 1.39)	1. 29*	(1.06, 1.58)
	Hispanic immigrant non-citizen	. 58***	(.52, .66)	. 66***	(.58, .75)	. 73**	(.59, .90)
	N	202, 976		199,494		84,194	
	Number of parameters	104		108		104	
	McFadden's R^2	. 23		. 25		. 24	
	BIC	129,055		124,342		65,927	
3). Nationality	Hispanic non-immigrant, non-Mexican	1. 29***	(1.14, 1.45)	1. 25***	(1.10, 1.42)	1. 16	(.98, 1.37)
	Hispanic immigrant, non-Mexican	. 79***	(.69, .89)	. 89	(.78, 1.01)	. 93	(.76, 1.14)
	Mexican non-immigrant	1. 14*	(1.02, 1.28)	1. 10	(.98, 1.24)	1. 06	(.92, 1.24)
	Mexican immigrant	. 70***	(.62, .80)	. 77***	(.67, .87)	. 94	(.77, 1.16)
	N	202, 976		199,494		84,194	
	Number of parameters	102		106		102	
	McFadden's R^2	. 23		. 25		. 24	
	BIC	129,158		124,412		65,954	
4). Immigrants only	Hispanic immigrant citizen	1. 46***	(1.20, 1.78)	1. 47***	(1.20, 1.79)	1. 47**	(1.10, 1.96)
(models control for years	Hispanic immigrant non-citizen	. 89	(.72, 1.10)	. 92	(.74, 1.13)	. 88	(.64, 1.21)
since immigration)	N	22,302		21,640		7,476	
	Number of parameters	100		104		98	
	McFadden's R^2	. 21		. 24		. 26	
	BIC	13,517.2		12,911.6		6,129.5	
	Hispanic immigrant non-Mexican	1. 10	(.93, 1.30)	1. 13	(.95, 1.34)	1. 24	(.96, 1.61)
	Hispanic immigrant Mexican	1. 14	(.94, 1.38)	1. 16	(.95, 1.42)	1. 32	(.99, 1.76)
	N	22,302		21,640		7,476	
	Number of parameters	97		101		95	
	McFadden's R^2	. 21		. 23		. 25	
	BIC	13,538		12,924		6,124	

Note : Robust 95% C/Is in parentheses.

*p < .05 **p < .01 ***p < .001

Data come from the March files of the CPS, 1994 - 2007. All models include state, metro, and year effects and all covariates shown in the Full Model from Table 2. Reference category in the years since immmigration, citizenship, and nationality models is non-immigrant whites. Reference category for the immigrants only citizenship models is white immigrant citizens. Reference category for the immigrants only nationality models is white immigrants. All models weighted with the appropriate March CPS weights. Odds ratios from suppressed covariates available upon request.

Table 4. Differences in predicted probability of unionization for key covariates.

	Predicted probability difference
1). Hispanic non-immigrant vs. Hispanic non-citizen	3.6 points
2). Public vs. private sector employment:	21.5 points
3). Production/craft/repair vs. professional/managerial employment:	12.9 points
4). 1,000+ firm vs. <25 workers:	8.4 points

Note : Predictions generated from Citizenship model in Table 4, column 2. Predictions generated using the STATA version 10 prchange command. A full set of predicted probabilities for all covariates in the model is available from the authors on request.

Table 5. Hispanic immigrant subcategories analysis: odds of joining a union (vs. never being in a union) in a one-year period, 1996-2007.

		Labor Market Position	Occ. and Ind. Change	Stable Occ. & Ind. Only	2002+ Only
1). Years since immigration	Hispanic non-immigrant	1. 37*** (1.28, 1.47)	1. 37*** (1.27, 1.47)	1. 36*** (1.24, 1.48)	1. 38*** (1.25, 1.53)
	Hispanic immigrated: 20+ years	1. 46*** (1.30, 1.64)	1. 45*** (1.30, 1.63)	1. 43*** (1.24, 1.66)	1. 36*** (1.16, 1.60)
	Hispanic immigrated: 10 to 20 years	1. 14* (1.01, 1.28)	1. 13* (1.00, 1.28)	1. 14 (0.98, 1.34)	0. 96 (0.79, 1.15)
	Hispanic immigrated: 5 to 10 years	0. 88 (0.74, 1.06)	0. 88 (0.73, 1.06)	0. 80 (0.62, 1.03)	0. 72* (0.54, 0.96)
	Hispanic immigrated: less than 5 years	0. 74** (0.59, 0.92)	0. 73** (0.58, 0.92)	0. 78 (0.58, 1.06)	0. 67* (0.48, 0.93)
	N	531, 202	531, 202	358,351	256,011
	Number of parameters	62	64	62	56
	McFadden's R^2	. 16	. 17	. 18	. 17
	BIC	612,550	606,096	431,017	288,898
2). Citizenship	Hispanic non-immigrant	1. 37*** (1.28, 1.47)	1. 36*** (1.27, 1.46)	1. 35*** (1.24, 1.48)	1. 38*** (1.24,1.52)
	Hispanic immigrant citizen	1. 43*** (1.28, 1.61)	1. 42*** (1.27, 1.60)	1. 40*** (1.20, 1.63)	1. 30** (1.10, 1.54)
	Hispanic immigrant non-citizen	1. 01 (0.92, 1.12)	1. 01 (0.92, 1.11)	1. 02 (0.89, 1.16)	0. 88 (0.76, 1.02)
	N	531, 203	531, 203	358,352	256,012
	Number of parameters	54	56	54	48
	McFadden's R^2	. 16	. 17	. 18	. 17
	BIC	612,309	605,830	430,738	288,655
3). Nationality	Hispanic non-immigrant, non-Mexican	1. 57*** (1.41, 1.74)	1. 55*** (1.40, 1.72)	1. 60*** (1.40, 1.84)	1. 42*** (1.24, 1.62)
	Hispanic immigrant, non-Mexican	1. 40*** (1.26, 1.55)	1. 39*** (1.25, 1.54)	1. 40*** (1.22, 1.61)	1. 27*** (1.11, 1.45)
	Mexican non-immigrant	1. 26*** (1.15, 1.38)	1. 26*** (1.15, 1.38)	1. 22*** (1.09, 1.38)	1. 34*** (1.20, 1.49)
	Mexican immigrant	0. 98 (0.88, 1.08)	0. 98 (0.88, 1.08)	0. 98 (0.86, 1.12)	0. 96 (0.84, 1.08)
	N	531,200	531,200	358,351	350,197
	Number of parameters	52	54	52	48
	McFadden's R^2	. 16	. 17	. 18	. 17
	BIC	612,657	606,181	431,015	399,075
4). Immigrants only	Hispanic immigrant citizen	1. 47*** (1.24, 1.75)	1. 47*** (1.23, 1.75)	1. 46*** (1.17, 1.83)	1. 37* (1.06, 1.77)
(models control for years	Hispanic immigrant non-citizen	1. 23* (1.03, 1.47)	1. 23* (1.03, 1.47)	1. 24 (0.99, 1.57)	1. 11 (0.85, 1.45)
since immigration)	N	55,540	55, 540	36,652	28,248
	Number of parameters	53	55	53	47
	McFadden's R^2	. 14	. 15	. 16	. 15
	BIC	62,810	62,560	42,840	31,089
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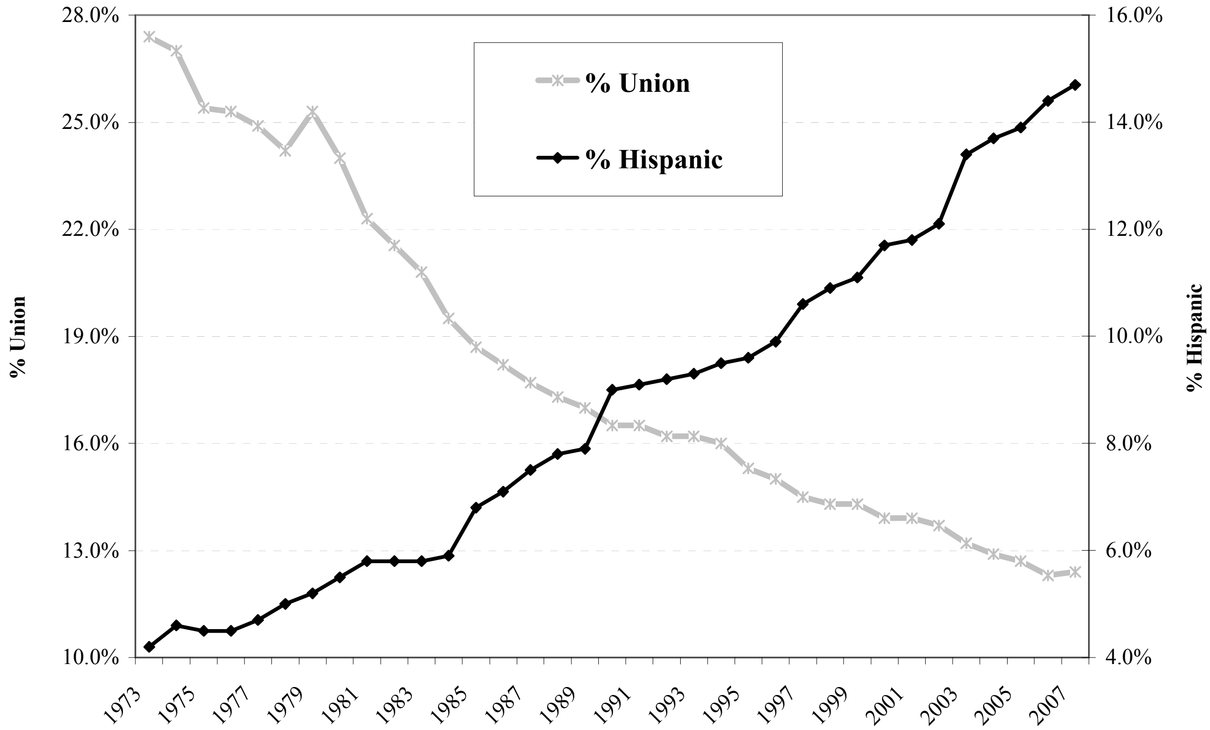
	Hispanic immigrant non-Mexican	1. 42*** (1.23, 1.64)	1. 41*** (1.23, 1.63)	1. 36*** (1.13, 1.64)	1. 30* (1.05, 1.62)
	Hispanic immigrant Mexican	1. 17 (0.99, 1.37)	1. 17* (1.00, 1.37)	1. 14 (0.93, 1.40)	1. 14 (0.91, 1.44)
	N	55,540	55, 540	36,652	28,248
	Number of parameters	50	52	50	44
	McFadden's R^2	. 14	. 15	. 16	. 15
	BIC	62,856	62,608	42,878	31,064

Note : 95% c/i in parentheses.

*p < .05 **p < .01 ***p < .001

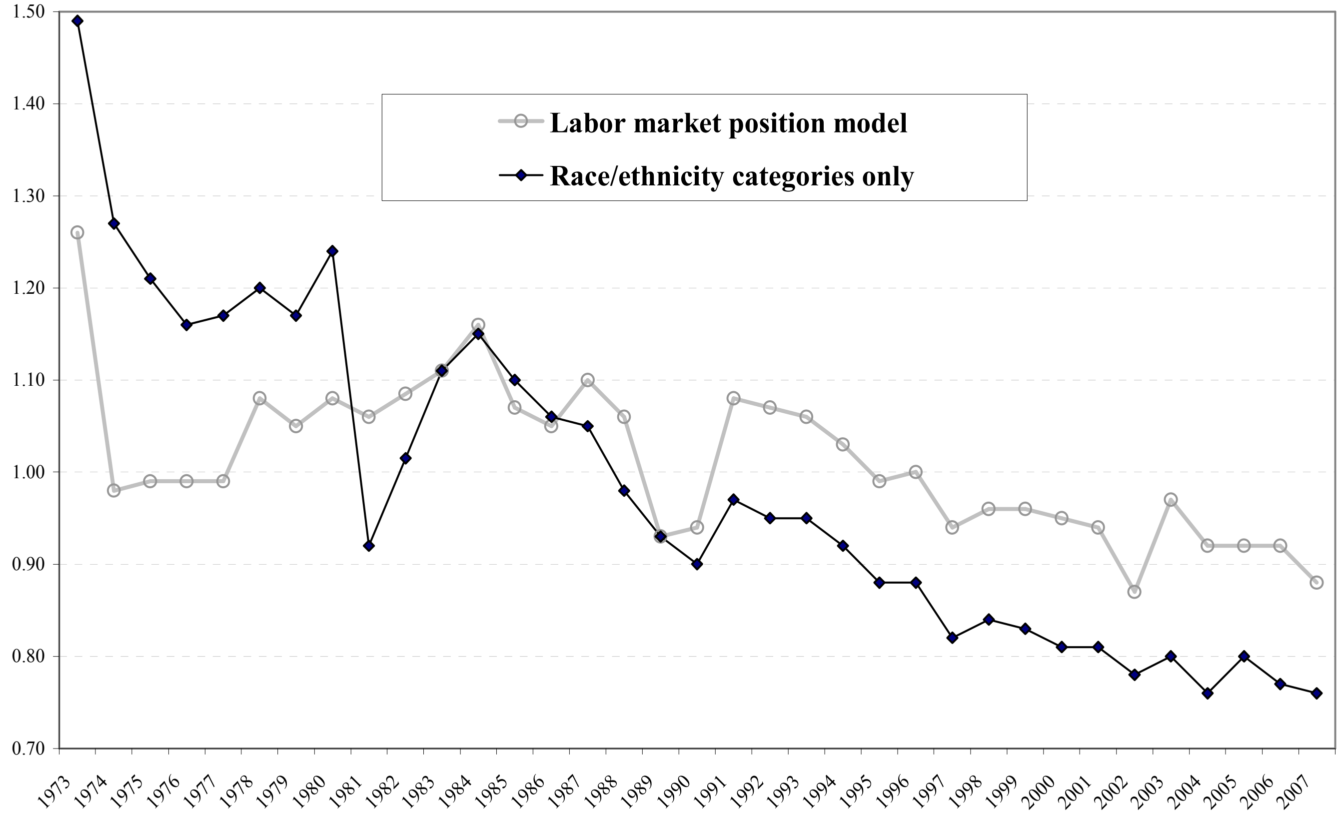
Data come from the matched MORG files, 1996 - 2007. All models include region, metro, and year effects and all covariates shown in the Full Model from Table 3. Reference category in all models is non-immigrant whites. All models weighted with the appropriate final CPS weights. Odds ratios from suppressed covariates available upon request. Reference category is non-immigrant whites in sections 1-3, and immigrant whites in the last set of models.

Figure 1. Declining unionization and the growth of the Hispanic population, 1973-2007.



Note : Data come from the May and MORG series of the Current Population Survey (CPS), various years. Estimates restricted to workers age 18 to 65, excluding the self-employed. No union questions were asked in 1982; we average the 1981 and 1983 rates to produce estimates for 1982. We multiply the 1973 to 1976 unionization rates by 1.094 to reflect changes in CPS wording on the union question (see Hirsch, McPherson, and Vroman 2001: 51). Estimates weighted with the appropriate CPS weights.

Figure 2. Hispanic workers' odds of unionization, 1973-2007.



Note : Point estimates represent the Hispanic odds ratio from annual models predicting union status. Reference category for all models is white. Data come from the May and MORG series of the CPS, various years. No union questions were included in the 1982 surveys; we generate 1982 estimates by averaging the 1981 and 1983 odds ratios. Estimates restricted to workers age 18 to 65, excluding the self-employed.

Appendix A. Creating the CPS-MORG matched data file for mini-panel analysis.

We use the CPS Merged Outgoing Rotation Group data (MORG) files for the years 1983-2007 from the UNICON Corporation. The procedure to create the series of CPS-MORG mini-panels capitalizes on the longitudinal aspect of the CPS rotation group structure. While the CPS is a monthly survey, the sample does not consist of a new household each month. Instead, each household is in the "observation" sample for four consecutive months, out of the sample for the following eight months, and returns to the observation sample for four additional months. Sampled households would typically be observed for eight total months, with the household in observation months four and eight being considered the "outgoing rotation groups" because they are about to leave the observation sample, either for an eight-month hiatus, or permanently. CPS respondents are interviewed for 4 consecutive months ("month in sample" or MIS=1-4), are out of the sample for 8 months, and then return to the interview sample for 4 months (MIS=5-8). Thus, if an individual is first interviewed in January of year 1 (MIS=1), s/he will be re-interviewed one year later, in January of year 2 (MIS=5). The MORG dataset is made up of only the outgoing rotation groups (those leaving the observation sample with MIS=4 and 8); we match individuals' data across MIS=4 and MIS=8. That is, we match a respondent's data across time 1 and time 2 to identify changes in key characteristics in a one-year period. We create these single year mini-panels for all years where time 1 is between 1983-2006 and time 2 is between 1984-2007.

We adapt the matching algorithm and Stata do-files described in Madrian and Lefgren (1999). Our do-files are available on request, and we briefly summarize the logic here. First, we extracted single year MORG data files from UNICON. Next, we cleaned and recoded several variables of interest for the study to ensure consistency across years, saving the results in single-year data files. We eliminated observations outside the working ages of 18-65, and those not employed in both time 1 and time 2. We initiate the matching process by generating separate data files for observations where MIS=4 in year N, and another where MIS=8 in year N+1. These two files are merged using the following identifying variables in the CPS: *state hhid hhnum lineno.*

Because the CPS is a sample of housing units (or addresses) within states, not specific individuals, families or households, new residents are surveyed if an individual or family moves out of a sampled household after the time 1 survey. The naïve matching described above will include a number of erroneous matches, notably in any instance where the original occupant/s moved, died, or did not participate in the survey for any other reason at time 2. To accurately reflect a match of the same individual across time we compare the sex, race, and age of the naïvely-matched individual at time 1 and time 2. Matches that differ on sex or race are considered failed matches, likely reflecting a household with a new resident replacing the initially-surveyed respondent. We also drop any matches where the age between time 1 and time 2 differs by something other than 0, 1, or 2 years. (If an individual's birthday fell near the survey date in either time 1 or time 2, his or her age may differ by 0 years or by 2 years, rather than the expected 1 year). Our naïve match rate across all years averaged 71%, whereas the effective match rate averaged 62% (See Table A1 below for yearly match rates). Differences between our match rates and those reported in other analyses may result from the elimination of cases not meeting the study criteria (e.g. not employed or in the working age population) *before* conducting the matching procedures.

Table A1. Naïve and valid merge rates for matched MORG data.

Year	Naïve merge rate	Valid merge rate
1983	68.7%	61.5%
1986	67.8%	60.1%
1987	65.7%	57.9%
1988	67.1%	59.3%
1989	69.7%	61.7%
1990	68.9%	61.1%
1991	69.8%	62.0%
1992	70.1%	62.2%
1993	69.3%	61.0%
1996	73.6%	65.1%
1997	72.7%	64.0%

1998	73.2%	64.2%
1999	73.5%	64.5%
2000	73.0%	64.5%
2001	72.9%	64.5%
2002	73.3%	62.8%
2003	72.1%	63.6%
2004	65.7%	61.2%
2005	72.7%	63.4%
2006	72.6%	63.0%

The mini-panels are appended into a single analysis file. Failed matches are dropped from this working file. In 1985 and 1995, the CPS revised the geographic identifiers in the public-use CPS files. As a consequence, matching across years with differing definitions of detailed geographic areas was precluded and panels including 1985 and 1995 as either the time 1 or time 2 data were dropped from our analysis.

Appendix B: Covariates used in constructing the odds of unionization models (Table 2).

Demographics Model	**Labor Market Position Model**	**Firm size model**
Race:	All Demographics Model controls, plus:	All Labor Market Position Model controls, plus:
White (reference)		
Black	*Sector:*	*Firm size:*
Hisapanic	Public (reference)	< 25 (reference)
Other	Private	25-99
		100-499
Sex:	*Occupation:*	500-999
Female (reference)	Prof./managerial (reference)	1000+
Male	Production/ craft/ repair	
	Service occupations	
Age & age^2 (continuous)	Farm/forestry/fishery occupations	
Marital Status:	*Industry:*	
Not married (reference)	Ag./forestry/fisheries (reference)	
Married	Mining, construction,	
	manu. durables, manu. non-durables,	
Education:	transportation, communications,	
< HS (reference)	utilities/sanitary svcs., wholesale trade,	
HS graduate	retail trade, FIRE, business repair svcs.,	
Some college	rec./entertainment svcs., prof. svcs.,	
B.A. or higher	government, unclassified	
Year fixed effects:	*Metro status:*	
1983 (reference)	In metro area (reference)	
Dummies for 1984-2007	In rest of SMSA	
	Not in SMSA	
	Missing	
	State fixed effects:	
	Maine (reference)	
	Dummies for other 49 states plus D.C.	
	Age & age^2 replaced with	
	Experience & experience^2	