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THE WAGE CONSEQUENCES OF ENCLAVE RESIDENCE:

Evidence from the 1990 and 1980 Censuses*

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ABSTRACT

In the last century Irish, German, Italian, Jewish and other immigrant groups have tended to congregate in ethnic enclaves upon their arrival in the U.S. Over time members of these groups moved away from enclaves and assimilated into the mainstream society and economy. Today, Hispanics appear to be following this same pattern. While these broad patterns of movement are well documented, little is known about immigrants' labor market experience inside the enclave that might motivate them.

This paper uses the 1980 and 1990 5% Census samples to investigate the wage consequences of enclave residence for Hispanic males. There are three principal findings. First, there are important citywide and Hispanic-specific demand shocks that influence where Hispanics choose to settle and affect their wage rates. Any examination of Hispanics' wages should control for these factors. Second, on average Hispanics choose to live in cities with a large existing stock of Hispanics. There is heterogeneity, however, across Hispanics sub-populations; the lesser skilled (i.e., immigrants and the English deficient) aggregates are drawn to enclaves, while natives and the English proficient tend to leave high-Hispanic cities. Third across a number of specifications, I find that enclave residence reduces the relative (compared to white natives) wages of Hispanics. In the least restrictive specification, I instrument for the 1980 to 1990 change in Hispanics' citywide share of the labor force with the 1980 share and find that a 1-percentage point increase in the share of Hispanics in a city reduces the Hispanic citywide relative wage by approximately 1%. This result holds across broadly defined Hispanic sub-populations, although the evidence is strongest among the least skilled.

At different times in United States history, Irish, German, Italian, Jewish, and Chinese immigrants have clustered in ethnic enclaves within large cities. Today Hispanics, one of the fastest growing segments of the U.S. population reside disproportionately in a few large cities. These ethnic enclaves offer their residents greater social interactions with people who share similar cultures, businesses that supply ethnic products and services, and a haven from the English language. Over time enclave residents invest in skills that are rewarded in the U.S. labor market, including the acquisition of English and U.S.-specific vocational skills (Hashmi 1987). Eventually they move out of the enclave and assimilate into the mainstream society and economy (Wilson and Portes 1980; Bartel 1989).

While these broad trends of movement into and out of enclaves are well documented, little research has been done on labor market experiences inside the enclave that might motivate these movements.[1] To my knowledge, there has only been one multi-city examination of the wage consequences of residence in a language enclave; it found that the wages of English deficient Hispanics are positively correlated with the share of Hispanics in a city (McManus 1990). The implication is that residing in an enclave reduces the penalties associated with weak English abilities. This study was based on a single cross-section and hence relied on variation in the share of Hispanics across cities to identify the enclave effect. This identification strategy is subject to at least two major criticisms: (1) in the presence of capital and labor mobility, wage differentials should not exist in the long run; and (2) citywide or Hispanic-specific labor demand factors are likely to influence where Hispanics choose to settle <u>and</u> to affect their wage rates, which is a classic simultaneity problem and is likely to bias the estimates upwards. The former criticism is mitigated by research that indicates that small labor demand shocks in U.S. states can persist for as long as a decade (Blanchard and Katz 1995). It is with the latter criticism in mind that this paper uses data from the 1980 and 1990 5% Census samples to empirically examine the relationship between Hispanics' wages and enclave residence.

Beginning with a simple, local production function, I show that the effect of living in an

[1] A separate but related question concerns immigrants' effects on host countries. See Friedberg and Hunt (1995) for a review of the substantial literature on this topic and Borjas (1994) for a broader discussion of the economics of immigration.

enclave on hourly wages can be estimated from correlations between the fraction of Hispanics in a city and Hispanics' citywide wage. The citywide wages are calculated by regressing the individual-level data against the standard determinants of wages and a wide variety of background variables (i.e., country of birth, years since arrival in U.S., etc.) to remove the influence of observable factors. The regression-adjusted citywide mean wage measures are captured by a full set of city-specific indicators. In order to allow these citywide wage measures to vary by broadly defined skill groups, separate regressions were fitted for a number of Hispanic labor aggregates (e.g., natives, immigrants, the English proficient, and the English deficient) and white natives.

The regression-adjusted means are then regressed against a number of city-level variables, including a measure of enclave status — the density of Hispanics. Of course, a simple cross-sectional correlation of this type is likely to be subject to the simultaneity problems mentioned above. For example, Hispanics are likely to choose to live in cities with high citywide or Hispanic-specific wages. In fact, I find evidence that these cross-sectional or "between" estimates are prone to substantial misspecification.

As an initial remedy to this misspecification problem, I remove permanent and transitory factors that affect the citywide wage level. This is accomplished by calculating the difference of Hispanics' and whites' average regression-adjusted citywide wages; and, this relative wage measure should be purged of factors that equally affect the wages of whites and Hispanics within a city. I then regress this differenced wage measure against the density of Hispanics to obtain relative enclave effects in 1980 and 1990 cross-sectional regressions. These estimates suggest that living in an enclave reduced Hispanics' relative wages, but only modestly.

It is possible that this differenced cross-sectional estimation procedure is misspecified due to permanent and/or transitory factors that are specific to Hispanics in particular cities. For example, Hispanics may choose to live in a city with industries that disproportionately employ Hispanics. To control for permanent factors of this type, I regress the change in the differenced wage measure between censuses against the change in Hispanic density. This approach also produces negative estimates of the enclave effects, particularly for lesser skilled Hispanics (i.e., recent immigrants and the English deficient). Moreover, these estimates are of a greater magnitude than the differenced cross-sectional ones, suggesting that the estimates from this more restrictive approach were biased upwards.

To address any remaining misspecification problems, I instrument for the change in a city's share of Hispanics between 1980 and 1990 with the 1980 share and its square.[2] This may be a valid procedure to identify the effect of Hispanic density on Hispanics' wages, because I find that Hispanics largely ignored economic conditions and chose their destination based on the existing stock of Hispanics.[3] Estimates from this IV procedure imply that a 1-percentage point increase in the citywide share of Hispanics is associated with a 1-% reduction in Hispanics' relative wages. This appears to hold across broadly defined Hispanic sub-populations, although the evidence is strongest among the least skilled (i.e., recent immigrants and the English deficient).

The paper proceeds as follows. Section 1 describes the data, provides a working definition of enclaves, presents descriptive characteristics about whites and Hispanics who live in high-Hispanic cities, and locates them in the U.S. Section 2 presents a simple theoretical model that uses a local production function to describe the effects of imperfectly substitutable labor aggregates in geographically distinct labor markets. This model underlies the subsequent empirical analysis. In Section 3, I operationalize the model developed in the previous section and discuss more generally the potential sources of misspecification inherent in the model and my efforts to purge the estimates of them. Section 4 presents the empirical results, and Section 5 concludes the paper.

1. Data Description and Implementation Issues

1.1 Data Sources

An "ideal" data source for examining the wage effects of Hispanic enclaves might include detailed longitudinal information on Hispanics' wages and location of residence.[4] These data would

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[2] This empirical strategy is similar in spirit to ones used to examine the impacts of inflows of immigrants on natives' wages (Altonji and Card 1991; Schoeni 1996).

[3] Bartel (1989) first identified this pattern in the data, and I show that it still holds in the 1980s.

[4] The Panel Study of Income Dynamics (PSID) recently began a supplementary survey that added Hispanics to their universe of respondents. Although it may eventually prove useful, the sample size is currently too small and there has not yet been enough movement between cities to carefully explore the relationship between enclaves and

permit precise comparisons of the changes of individual Hispanics' wage rates in cities with varying population shares of Hispanics. In the absence of such a data set, this paper uses the individual data from the 1980 and 1990 5% Census samples; so, the labor market information refers to 1979 and 1989. The sample is limited to white native and Hispanic men between the ages of 16 and 64 who were: in the labor force (employed or unemployed) when surveyed; employed for pay during 1979 (1989); and residents of one of the Census Bureau's Standard Metropolitan Statistical Areas (SMSAs).[5] Respondents' hourly wage was calculated by dividing total earnings by the product of weeks worked and usual hours per week. The sample was restricted to a 10% random sample of all white natives and 100 percent of all Hispanics from the Censuses. These selection criteria yielded 322,207 observations in 1980 and 393,309 in 1990.

1.2 Defining Enclaves

Enclaves are frequently defined as clusters of people that share a common background (Borjas 1995; McManus 1990). Given this definition, at least two issues arise in an empirical examination of enclaves. The first is how to translate the informal notion of a "cluster" into an observable geographic unit. Some observers have defined enclaves at the level of a neighborhood, suggesting that the proper unit of observation might be a census tract or a small group of tracts. This definition is problematical for a study that attempts to draw inferences on the determinants of wages across labor markets, because these fine geographic units are likely to arbitrarily divide unified labor markets.

To avoid this division of labor markets, I define enclaves at the SMSA level, which are drawn such that they have "a high degree of economic and social integration" (U.S. Bureau of the Census 1993).[6] This definition of an enclave may be larger than the typical conception of one, but it avoids what I consider to be the more serious problem of splitting labor markets.[7] Using the two

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Hispanics' wages.

[5] In 1990 the Census Bureau changed the name of large urban areas from SMSAs to Metropolitan Statistical Areas (MSAs). For clarity, these geographical areas are henceforth referred to as SMSAs for both Census years.

[6] A county based definition of enclaves might also avoid dividing labor markets. McManus (1990) found that a county-based definition of enclaves yielded results that were similar to a SMSA-based one.

[7] Card (1997), Altonji and Card (1991), LaLonde and Topel (1991) and others have all defined labor markets at the

censuses, I was able to identify 300 of the 318 SMSAs in 1980, 324 of the 335 in 1990, and to match 280 across censuses. The Data Appendix describes the procedures used to identify the SMSAs in each year and to match SMSAs across them.

The second issue centers on determining who shares a "common background." In the case of German or Italian enclaves, where speakers of that language come from a single county, this is a straightforward task. Hispanics, on other the hand, come from many places (e.g., Mexico, Cuba, Puerto Rico, Central America, etc.) but share a common language. Borjas (1994) argues that pooling people from different countries into a single category may lead to "aggregation bias" that makes it difficult to discern the effect on a person from a particular country. I rely, however, on the language-based classification adopted by the Census Bureau and assign people descended from Spanish-speaking countries to the relatively crude classification of "Hispanic." This assignment rule may miss some of the country-specific elements of enclaves, but allows for an accurate description of the economic impacts of Spanish language enclaves.

1.3 Hispanics' Disproportionate Location in A Few Cities

Table 1 presents Hispanics' share of the male, urban labor force in 1980 and 1990 and underscores their substantial and growing presence in the U.S. economy. Notice first that blacks, Asians, and other groups are included in the calculation of the labor force, so the percentages do not add up to one hundred. In 1980 Hispanics accounted for 7.1%, but by 1990 their share had increased by approximately 34% to 9.5%. This 2.4% increase was due to a 2.8 percentage point increase in the share of Hispanic immigrants and a .4 percentage point decline in Hispanic natives. English deficient Hispanics' share of the labor force increased from 1.5% to 2.2% in this period.[8] These gains in Hispanics' share of the labor force came at the expense of whites whose share declined from 77.2% to 66.8%.

An important difference between Hispanics and white natives is their distribution across

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SMSA level.

[8] The Census questionnaire provides limited measures of English ability. Following Chiswick (1991) and Carliner (1995), respondents are designated English proficient if they report that they speak English "very well" or English is the only language they speak at home. All other respondents are labeled English deficient.

cities. In both 1980 and 1990, Hispanics were disproportionately located in a few large cities and whites were almost evenly spread across U.S. cities. The skewed geographic concentration of Hispanics in 1980 is illustrated in Figures 1a, which graphs the cumulative fraction of Hispanic males and recent Hispanic immigrants against the cumulative fraction of whites in the same cities. If these three groups were evenly located across the country, the cumulative fraction lines would lie along the 45-degree line, which is pictured in Figure 1a. Along the three cumulative fraction lines, the cities are ordered by their share of the national total of urban dwelling Hispanic males in the labor force. For example, the Los Angeles-Long Beach SMSA had the largest Hispanic population in 1980, so it is the first point on each line. A convenient way of summarizing this skewed distribution is by noting that nine SMSAs accounted for 51% of all urban dwelling Hispanic males and 66% of recent[9] Hispanic male immigrants in 1980.[10] Yet, these same cities only accounted for 20% of total (white and Hispanic) urban, working aged males in this period. Figure 1b presents the analogous graph for 1990. In this year, eleven cities housed 54% of Hispanic males and 63% of recent Hispanic immigrant men but contained only 20% of urban working aged men.[11]

To explore the differences between Hispanic enclave cities and low-Hispanic cities, I separately ordered the 1980 and 1990 SMSAs by the percentage of Hispanics in the male labor force.[12] The cities were then divided into four equal groups; Group 1 SMSAs have the lowest share of Hispanics and Group 4 the highest. The first panel of Table 2 presents summary data for Hispanic males. Hispanics in Group 1 SMSAs were more likely to be English proficient than their Group 4 counterparts in both years and were better educated. In both surveys, the percentage of Hispanics who had moved to their current state in the last 5 years was significantly greater in Group

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[9] Recent immigrants are defined as people who have arrived in the U.S. in the previous 15 years.

[10] In 1980, the nine SMSAs with the largest populations of Hispanic males and their shares of the total, urban, male, Hispanic population were: Los-Angeles-Long Beach, CA (16.6%), New York, NY (9.6%), Chicago, IL (4.9%), Miami, FL (4.7%), Houston, TX (3.8%), San Antonio, TX (3.6%), San Francisco, CA (3.0%), Riverside-San Bernardino-Ontario, CA (2.6%), and Anaheim, CA (2.5%).

[11] In 1990, these eleven SMSAs and their respective shares of the total urban, male, Hispanic population were: Los-Angeles-Long Beach, CA (17.6%), New York, NY (7.8%), Miami, FL (4.7%), Chicago, IL (3.8%), Houston, TX (3.7%), Riverside-San Bernardino, CA (3.5%), Anaheim, CA (3.4%), San Antonio (2.8%), San Diego (2.7%), Dallas-Ft. Worth, TX (2.1%), and San Jose, CA (1.8%).

[12] The population of the labor force is the weight adjusted number of working aged men of any ethnicity in the labor force, where the weight comes from the Census files.

1 cities, implying that Hispanics living outside of enclaves are more mobile. Additionally, the hourly wage of Group 4 Hispanics was 9% (1%) less than that of Group 1 Hispanics in 1990 (1980). In sum, Hispanics in enclave cities were less educated, more likely to be English deficient, more likely to have lived in the same location longer, and earned less than those in low-Hispanic SMSAs.

The middle panel of Table 2 illustrates that whites residing in enclave cities differ from those in low-Hispanic ones. Whites were slightly better educated (.7 years in both 1990 and 1980) in Group 4 than in Group 1 cities, but had less potential experience. The whites in 1990 Group 4 SMSAs were more likely to have changed states in the previous 5 years than their counterparts in Group 1 SMSAs. This difference in mobility, however, was not evident in 1980. The most striking difference is that the hourly wage of white natives was 18% (11%) higher in 1990 (1980) in Group 4 cities than in Group 1 ones. In addition to having high concentrations of Hispanics, Group 4 cities appear to be high wage locales for whites.

The bottom panel of Table 2 details the relationship between some observable city factors and the share of Hispanics in a SMSA. The weighted number of people in the sample from a given city rises with the percentage of Hispanics, from less than 85,000 (85,000) in Group 1 to more than 250,000 (235,000) in Group 4 in 1990 (1980). In 1980 whites accounted for 91% of the male working aged population in Group 1 SMSAs but only 71% in Group 4 ones; Hispanics comprised less than 1% in Group 1 cities and almost 18% in Group 4 ones. A similar pattern held in 1990. Additionally, Group 4 cities have slightly smaller unemployment rates in both census years.

The summary statistics presented above indicate that neither Hispanics nor whites randomly locate across the U.S. In studying enclaves' effect on Hispanics' wages, it will therefore be important to adjust for other factors that might influence both wages and location, such as years of education, time in the labor force, language ability, as well as observed and unobserved city characteristics.

1.4 Where are the High-Hispanic Cities?

Table 2 showed that the characteristics of high-Hispanic cities, and of their residents differ, from low-Hispanic ones. To highlight where these enclave cities are located, Figures 2a and 2b graphically display the percentage of Hispanics in each SMSA in the 1980 and 1990 Censuses, respectively. As Table 1 highlighted, Hispanics accounted for 7.1% (1980) and 9.5% (1990) of the total male, SMSA labor forces. In these figures, cities are shaded three different colors based on the share of Hispanics within their boundaries: light gray indicates less than 1% of the white and Hispanic, male labor force, gray for 1%-10%, and greater than 10% is shown with black. The cities located in states that border Mexico and those that are closest to Cuba (i.e., California, Arizona, New Mexico, Texas, and Florida) consistently have the highest shares of Hispanics. However, a number of cities outside these states also qualify as high-Hispanic SMSAs. In particular, many of the major cities in other Western states, the Rust Belt and the Northeast are popular locations for Hispanics. For example by 1990, Hispanics' share of the labor force exceeded 10% in Seattle, Denver, Chicago, New York, Baltimore, and Washington, DC. In contrast, there were still relatively few cities in the South where the share of Hispanics exceeded 1% of the population in 1990.

Figure 2c presents a graphical description of the change in male Hispanics' share of the labor force in the 1980s. In this figure, light gray identifies cities where Hispanics' share declined, gray indicates a change between 0% and 2%, and black signifies an increase of more than 2%. Of the 280 SMSAs pictured, 72 experienced a decline, 155 were in the middle group, and the remaining 53 were in the top group. The three largest declines were in cities near the Mexican border: Laredo, TX (-6.4%, from 90.2% to 83.8%), Albuquerque, NM (-2.7%), and McAllen-Pharr-Edinburg, TX (-1.4%). The three largest increases were in Miami, FL (12.8%), the city that absorbed the majority of the Marielitos, Los Angeles-Long Beach, CA (8.2%), and Visalia-Tulare-Poterville, CA (7.9%).

Visual inspection of Figures 2a and 2c suggests that the cities with high Hispanic shares in 1980 were the ones that experienced the largest increases in Hispanic shares between the two Censuses. I demonstrate this relationship more formally later, but the pattern is important because it suggests that Hispanics were drawn to enclave cities. If this attraction to enclave cities is independent of wage rates, except through enclaves' effect on wages, it is valid to use the 1980 share of Hispanics as an "instrument" for the change in the share of Hispanics in a wage regression.

2. A Framework for Estimation

A natural framework for analyzing the effect of differential supplies of various Hispanic skill groups on Hispanics' wages in a local labor market is to treat each city c as a geographically distinct, competitive economy with a single output good (Y_c). Assume that each city's period-specific output Y_{ct} is a concave function of the human capital supplied by non-Hispanic white natives, different Hispanic skill groups, and non-labor inputs (e.g., capital);

$$Y_{ct} = F[(\eta_{ct})\, p(\exp(L_{ct1}),\ldots,\exp(L_{ctk})), (\alpha_{ct})\, v(\exp(N_{ct1})\ldots,\exp(N_{ctm}))]. \qquad (1)$$

For this examination of Hispanic enclaves, I assume that the shares of human capital in p(·) are confined to a number of Hispanic skill groups or aggregates (j=1,...,k-1) and whites (j=k).[13] The function v(·) defines the effects of capital and other resources that are necessary for production but are exogenous to the analysis. I assume weak separability of the local production function and thereby avoid calculating local capital stocks whose existing estimates are notoriously imprecise (Hojvat-Gallin, LaLonde, and Topel 1995). The parameters η_{ct} and α_{ct} are exogenous shocks to demand for the inputs that are both locale-specific and factor neutral. For example, an increase in η_{ct} might represent a local labor demand shift that is common to all labor aggregates in city c.

To derive the wage of group j workers in locale c, it is assumed that each member of labor aggregate j supplies their aggregate's average amount of human capital. Without loss of generality, I introduce a loading factor ω_{cjt}, which has a weighted average of one across all labor aggregates within a city. The advantage of including the loading factor is that it allows labor demand shocks to differentially affect each aggregate's wage. Formally, the partial derivative of the production function with respect to L_{ctj} yields

$$\partial Y_{ct} / \partial L_{ctj} = W_{ctj} = F_1(\cdot)\, (\eta_{ct})(\omega_{cjt})\, \exp(L_{ctj})\, p_j\, (\exp(L_{ct1}),\ldots,\exp(L_{ctk})) \qquad (2)$$

where subscripts denote partial derivatives with respect to the relevant argument. Application of the

[13] A native is defined as any person born in the United States. Non-Hispanic white immigrants are excluded from the sample, thus non-Hispanic white natives are henceforth referred to as whites. Section 4 describes my efforts to divide Hispanics into the k – 1 labor aggregates.

weak separability assumption of the production function in (1) means that capital and other inputs enter (2) only through $F_1(\cdot)$. Shifts in these inputs therefore equally affect each labor aggregate's wage leaving relative salaries unchanged.

Taking the natural logarithm of equation (2) leads to the following expression for the period t wage rate of group j in locale c:

$$W_{ctj} = \ln[F_1(\cdot)(\eta_{ct})] + \ln[(\omega_{cjt})] + \ln[\exp(L_{ctj})p_j(\exp(L_{ct1}),\ldots,\exp(L_{ctk})]. \qquad (3)$$

Note that in this model the wage rate is determined by three factors: a common city-specific component; a labor aggregate and city-specific component; and the relative population shares of the various skill groups. The form of this function implies that deviations of each group's wage from the period t city average depend only on the labor aggregate-specific factors and the effects of the share of each labor aggregate.

Equation (3) is the basis for the empirical analysis in this paper. In particular assume that the term $\ln[F_1(\cdot)(\eta_{ct})]$ can be decomposed as:

$$\ln[F_1(\cdot)(\eta_{ct})] = \nu_c + \nu_{ct},$$

where ν_c represents fixed city factors and ν_{ct} city-level shocks. Similarly assume $\ln[(\omega_{cjt})]$ can be written as:

$$\ln[(\omega_{cjt})] = \nu_{cj} + \nu_{cjt},$$

where ν_{cj} captures fixed labor aggregate- and city-specific factors and ν_{cjt} represents labor aggregate- and city-specific shocks. In order to make the last term tractable, I take a first order Taylor series approximation by expanding around the United States averages L_{ci}^*. Thus (3) can be rewritten as a simple regression model of the form

$$w_{cjt} = \sum_i L_{cit}\lambda_{ji} + u_{cjt}, \qquad (4)$$

where

$$u_{cjt} = \nu_c + \nu_{ct} + \nu_{cj} + \nu_{cjt}$$

is an unobserved error. Thus the error term is comprised of permanent and transitory city-specific factors and permanent and transitory aggregate-specific factors. This specification of the error term underscores that if any of these four types of factors raise wages <u>and</u> attract in-migrants to a city, the estimated λ_{ji}s (i.e, the enclave effects) will be biased upwards. The empirical analysis attempts to obtain estimated enclave effects that are free of these potential biases.

The form of equation (4) has an additional feature that merits mention. Specifically, it permits a Hispanic aggregate's share of the population to affect its own wage rate and the wage rate of other Hispanic aggregates. This is an attractive feature, because it is possible that an increase in the share of low (or high) skilled Hispanics differentially affects the wages of low and high skilled Hispanics.

To recap, positive estimates of the λ_{ji}s, or "enclave effects," would imply that enclave residence raises Hispanics' wages and that aggregates j and i are complements in production. Such a result would be evidence that enclaves reduce the economic costs of assimilating into the U.S. economy. Negative enclave effects, in contrast, imply that aggregates j and i are substitutes and exert downward pressure on each other's salaries.

3. Identification Strategy

Having established the basic regression equation in the previous section, this section presents a few operational details and discusses the efforts to eliminate the potential sources of misspecification highlighted in equation (4).

3.1 Computing Labor Aggregate-Specific Wages

In order to estimate equation (4), it is necessary to calculate period-specific citywide wages for the Hispanic labor aggregates and white natives. Previous research (McManus 1990) has suggested that there is heterogeneity in the estimated enclave effects, so I divided Hispanics into six different labor aggregates: all Hispanics, natives, immigrants who arrived in the previous 15 years, immigrants who arrived more than 15 years ago, the English deficient, and the English proficient. Moreover, I allow individuals within a labor aggregate to supply different amounts of human capital and, in turn, to have heterogeneous wages. Specifically, an individual's stock of human capital, l, is assumed to be determined by his characteristics, X. Thus, person h's human capital, in labor aggregate j, in city c in period t, can be modeled as $l_{cjth} = \exp(X'_{cjth}\theta_{jt} + \varepsilon_{cjth})$. Consequently, I estimate the citywide wages by substituting l_{cjth} into (3) and fitting:

$$w_{cjth} = \Omega_{cjt} + X'_{cjth}\theta_{jt} + \varepsilon_{cjth}, \tag{5}$$

where the vector w is the natural logarithm of the hourly wage. The vector, Ω_{cjt}, contains a full set of labor aggregate by SMSA by year indicator variables. The vector X_{cjth} contains a number of variables that might be correlated with both wages and the percentage of Hispanics in a city. It includes years of education, potential experience (age - 6 - years of education), and potential experience squared.[14] X_{cjth} also includes indicator variables for birth in the U.S., several categories of the number of years since arrival for immigrants (Chiswick, 1978; Borjas, 1985), whether an immigrant received the majority of their schooling in the U.S. (Friedberg, 1991), deficient English language skills (McManus, Gould, and Welch, 1983), U.S. citizenship status, enrollment in school, country of birth (Borjas, 1994), and whether or not the person lived in the same SMSA five years earlier.[15] The last indicator is intended to control for otherwise unobserved motivation or ability that is correlated with changing cities. The subscripts on the vector θ_{jt} indicate that these parameters are allowed to vary at the labor aggregate by year level.

Consequently, the estimated parameters $\hat{\Omega}_{cjt}$ are the period-, labor aggregate-specific mean SMSA wages, or w_{cjt} from equation (4), after individual differences in <u>observable</u> characteristics have been purged. The use of regression adjusted means has two advantages over simply computing the aggregate-specific arithmetic mean: first, it should eliminate bias in the estimated enclave effects that results from correlation between the fraction of Hispanics and <u>observable</u> factors; and second, it should reduce the sampling variation associated with citywide wage measures.

In the second step of the analysis, I return to equation (4) and regress the estimated cell means against the city's period-specific enclave status and city level controls. Specifically, I fit

$$\hat{\Omega}_{cjt} = V'_{ct}\pi_{jt} + \sum_i L_{cit}\lambda_{ji} + u_{cjt}, \tag{4'}$$

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[14] The 1990 Census reports education in 17 discrete categories. Park (1994) provides a mapping to convert these categories to a continuous years of education variable.

[15] To avoid perfect collinearity, it is necessary to exclude some of the indicator variables in the vector X_{cjth} for some of the labor aggregates. The English deficient variable, for instance, is excluded from the fitting of (5) for the English proficient and deficient aggregates.

where

$$u_{cjt} = \nu_c + \nu_{ct} + \nu_{cj} + \nu_{cjt}$$

and $\hat{\Omega}_{cjt}$ is the adjusted hourly wage for labor aggregate j in city c and period t, which was estimated in equation (5). The vector V_{ct} contains the city unemployment rate and the natural logarithm of the number of male labor market participants, both calculated from the Census files. These two variables should control for local economic conditions and the higher price levels larger cities (e.g., Los Angeles and New York).

Two problems with the standard errors are addressed. First, the dependent variable in (4') is an estimate, not a population value. Since the number of individuals in the sample varies across SMSAs, the dependent variables are estimated with differing levels of precision. To remedy this problem of heteroskedasticity, I use weights of $(N_{jt})^{1/2}$ in the second stage where N_{jt} is the number of observations in labor aggregate j at time t. Second, the two-stage approach solves the problem of correlated random effects.[16] For expository purposes, I fit a one-stage regression and found that the standard errors were approximately 1.2-2 times larger in the two-stage approach.

3.2 An Illustration of Misspecification Bias

The specification of the error term in (4') highlighted that there are at least four categories of factors that may jointly determine the fraction of Hispanics in a city and Hispanics' citywide wage. Hispanics, for example, may be more likely to live in cities that have permanently higher wages or that have experienced positive labor demand shocks. In either of these cases, the estimated enclave effects would be biased, because they would capture the effect on wages of both the unobserved variable(s) and the concentration of Hispanics.

As an illustration of the bias associated with unobserved city factors, I regressed the 1990 regression-adjusted SMSA mean wage of Hispanic natives against the fraction of Hispanic natives. The first panel of Table 3 reports the results and Figure 3a, which is derived from the "residual

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[16] When aggregate variable (i.e., percentage of Hispanics) are used in an individual-level regression, the estimated standard errors will generally be significantly downward biased, due to correlation in the errors terms of people with the same group (i.e., city). In the two-stage approach, the dependent and independent variables are at the same level and the standard errors fully account for this correlation. See Moulton (1986 and 1990).

regression" technique, graphically illustrates them.[17] Together they indicate that a 1-percentage point increase in the share of Hispanic natives reduces the average wage of Hispanic natives by .7%.[18] These results *seem* to suggest that an increase in the share of Hispanic natives reduces their wages.

Yet, the share of Hispanic natives in a city is not the only city-level variable that affects wages. To investigate the source of this result, I regressed whites' 1990 regression-adjusted SMSA mean wage against the fraction of Hispanic natives. Whites natives are chosen, because Spanish language enclaves are a Hispanic phenomenon and, as such, they should not directly affect whites' wages. However, whites' wages are a function of factors that affect the overall level of wages in a city.[19] The results are shown in the middle panel of Table 3 and Figure 3b and taken at face value they imply that there is a substantial *negative* enclave effect for *whites*, which is almost identical to the effect on Hispanic natives. This result underscores the importance of unobserved, citywide factors in determining wage levels and Hispanics' location decisions.[20]

One means to purge the estimated enclave effects of these citywide unobservables is to use white natives as a "control" group. In this case, the second stage regression becomes

$$\hat{\Omega}_{cj\text{-}kt} = \hat{\Omega}_{cjt} - \hat{\Omega}_{ckt} = (\pi_{jt} - \pi_{kt}) X_{ct} + \sum_i (\lambda_{ji} - \lambda_{ki}) L_{cit} + u_{cj\text{-}kt}, \qquad (6)$$

where

$$u_{cj\text{-}kt} = \nu_{cj} - \nu_{ck} + \nu_{cjt} - \nu_{ckt} = \nu_{cj\text{-}k} + \nu_{cj\text{-}kt}.$$

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[17] Figures 3a-3c plot the residuals (on the y-axis) from the regression of the citywide wage measures against all the explanatory variables in equation (4'), except the Share of Hispanic Natives, against the residuals (on the x-axis) from the regression of the Share of Hispanic Natives against all the other explanatory variables in (4'). The regression line in these figures is derived from the bivariate regression of these residuals. See Goldberger (1991) for a detailed description of the "residual regression" technique.

[18] I was concerned that in the context of fitting a weighted regression, these findings might be the result of one of the "outlier" cities having a large Hispanic population. It does not appear that this is the case. Specifically, the cities with the largest negative enclave effects are generally near the U.S./Mexico border like Corpus Christi, TX, Laredo, TX, and Albuquerque, NM. Although these towns all have a high percentage of Hispanics, their population of Hispanics in levels ranks them in the middle of the cities under consideration. Consequently, their influence on the regression line is limited. Los Angeles, New York, Miami, and Chicago and the other cities with the largest number of Hispanics are generally in the thick mass of points.

[19] The use of whites to control for these citywide factors is indirectly supported by Lalonde and Topel (1991) who found that the effects of immigration are smallest on white natives and greatest for recent immigrant groups. Altonji and Card (1991), Card (1997), and Borjas, et. al (1997) have similar findings.

[20] Previous work on Hispanic enclaves restricted the sample to Hispanics and implicitly assumed that unobserved, citywide determinants of wages were orthogonal to the share of Hispanics in a city (McManus, 1990).

Hispanic labor aggregates are denoted by j and white natives by k.[21] Notice that ν_c and ν_{ct}, the permanent and transitory city factors from (4'), are differenced out and cannot bias the parameters of interest. Thus equation (6) estimates the *relative* enclave effect, or the effect on Hispanics minus the effect on whites.

Returning to the example of whether Hispanic natives benefit from or are penalized by enclave residence, I fit equation (6). The bottom panel of Table 3 and Figure 3c together illustrate that once differences in citywide wage levels are removed, the evidence for an enclave effect has disappeared. This example has illustrated that there are important city-level factors that equally affect the wages of Hispanics and whites. Throughout the rest of the paper, this differenced wage measure is used to control for these potential sources of misspecification.

4. Empirical Estimates of Enclave Effects

This section presents estimates of the enclave effects from three different regression approaches: differenced cross-sectional, difference in differences, and difference in differences instrumental variables. The differenced cross-sectional specification is the fitting of equation (6), while the latter two techniques are extensions of this equation that are more formally introduced in this section. Each of the techniques requires different assumptions to produce consistent enclave effects; the differenced cross-section requires the most restrictive and the difference in differences instrumental variables the least.

An issue that equally applies to the three estimation techniques is how to best divide Hispanics into different labor aggregates. For instance, Hispanics could be divided into aggregates based on birth, language ability, or education, to name just a few. The placement of Hispanics in a greater number of distinct aggregates, however, creates at least two problems. First, these finer groupings of Hispanics might belong to the same labor aggregate, which confuses the interpretation of the enclave effects. Second, the greater the number of enclave effects, the more imprecise their

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[21] Equation (6) is weighted by $(N_{t,j}^{-1}+N_{t,k}^{-1})^{-1/2}$, where j represents the relevant Hispanic aggregate and k indicates whites.

estimation. This imprecision reflects a vain effort to estimate an increasing number of parameters from observations on a limited number of SMSAs. After some experimentation, I fit two specifications for all six aggregates. The first specification limits the enclave variable to Hispanics share of the labor force. A second specification used the shares of the Hispanic English deficient and Hispanic English proficient for the language aggregates and the shares of Hispanic natives and Hispanic immigrants for the other aggregates.

4.1 Differenced Cross-sectional Estimates

In the differenced cross-sectional approach, equation (6) is separately estimated for 1980 and 1990. This approach yields consistent estimates of the enclave effects if neither fixed nor transitory city by labor aggregate factors are correlated with the differenced hourly wage and the share of Hispanics. Recall from Table 2, however, that Hispanics and whites have different rates of mobility both in and out of high-Hispanic cities. If Hispanics' location decisions are more responsive to factors, such as positive labor demand shocks or other aggregate-specific factors, this specification will produce inconsistent estimates of the enclave effects.

Table 4 presents the differenced cross-sectional estimates. Each panel contains estimates from four separate regressions where the dependent variable is always the relative regression-adjusted, SMSA-specific mean wage for the relevant labor aggregate, as in equation (6). The estimates in columns (1) and (2) are from 1980 and columns (3) and (4) display the 1990 ones. In examining these estimates, it is useful to bear in mind that Hispanics average share of the labor force in these two years was 7.1% and 9.6%, respectively.

Columns (1) and (2) of the top panel indicate that an increase in the share of Hispanics had an inconsequential effect on Hispanics' relative wage in 1980. Although there is no effect in the aggregate, the other five panels indicate that there is considerable heterogeneity across the different Hispanic sub-populations. In particular, the estimated enclave effects are negative and statistically significant for the lesser skilled recent (i.e. arrived in the last 15 years) immigrant and English deficient aggregates. A 1-percentage point increase in the share of Hispanics is associated with approximately a .3% decline in the relative hourly wage of these two aggregates. In contrast, the enclave effects are generally very small and usually statistically indistinguishable from zero for the higher skilled aggregates (i.e., natives, immigrants who have been in the US more than 15 years, and the English proficient).

The 1990 differenced cross-sectional results are presented in columns (3) and (4). In the context of the sampling errors, these estimates are very similar to the 1980 ones. For example, these estimates also imply that an increase in the share of Hispanics is correlated with lower wages for both recent immigrants and the English deficient. A 1-percentage point increase in the share of Hispanics is associated with an estimated .4% reduction in the English deficient aggregate's relative wages. Just as in 1980, this finding is largely accounted for by differences in the share of the English deficient across SMSAs. One result that differs from 1980 is that an increase in the share of Hispanic natives modestly raises the relative wages of Hispanic immigrants who have been in the U.S. for more than fifteen years. Taken together, the differenced cross-sectional estimates provide little evidence for the "wage-raising" view of enclaves.

4.2 Difference in Differences (DD) Estimates

If there are unobserved factors that differentially affect the wages of whites and Hispanics and are correlated with Hispanic density, the differenced cross-sectional results will be biased. A city's industrial composition might be an example of a factor that favors (or hurts) Hispanics. To control for these factors, the DD approach subtracts the 1980 differenced cross-sectional equation from the 1990 one. In other words, the change in the relative wage is regressed against the change in Hispanic intensity between 1990 and 1980:

$$\Delta \hat{\Omega}_{cj\text{-}kt} = \hat{\Omega}_{cj\text{-}k,1990} - \hat{\Omega}_{cj\text{-}k,1980} = \pi_{j\text{-}k}(\Delta X_{ct}) + \sum_i \lambda_{j\text{-}ki}(\Delta L_{cit}) + \Delta u_{cj\text{-}kt}, \qquad (7)$$

where

$$\Delta u_{cj\text{-}kt} = \nu_{ck\text{-}j1990} - \nu_{ck\text{-}j1980}.$$

The advantage of equation (7), compared to (6), is that any permanent labor aggregate by city factors (i.e. $\nu_{ck\text{-}j}$) are differenced out and are no longer a potential source of misspecification. An inherent problem with the DD approach, however, is that the labor force share variables are derived from the Census 5/100 samples; they are not population measures. Consequently, enclave status is measured with error, which will attenuate the estimated enclave effects. The measurement problem is also present in the differenced cross-section, but it is worse in the DD approach where the share Hispanic variables result from the difference of two mismeasured variables.

Columns (1) and (2) of Table 5 present the estimated enclave effects from the DD identification approach. The table is designed to resemble Table 4, so it contains aggregate-specific estimates of the enclave effects (i.e., $\lambda_{j\text{-}ki}$) from the estimation of (7). The relative decrease in the signal to noise ratio of the share variables is reflected in the substantially larger standard errors. Despite this increase in noise, the estimates indicate that an increase in the share of Hispanics is correlated with a statistically significant decrease in the relative wages of 4 of the 6 labor aggregates. Moreover, all six of the estimated effects are negative and of a greater magnitude than in either of the differenced cross-sections. In particular, the estimates suggest that a 1-percentage point increase in the share of Hispanics is associated with .4% decline in Hispanics' relative wages. Since Hispanics' share of the male labor force increased by approximately 2.5% in this period, this translates into a 1% decline in Hispanics' relative average hourly wage. Although the sampling errors are large, the estimated enclave effects continue to be of a greater magnitude for the less skilled aggregates.

4.3 Difference in Differences Instrumental Variables (DDIV) Estimates

Comparison of the differenced cross-sectional results to the DD ones suggest that permanent city by labor aggregate factors may have biased the differenced cross-sectional results upwards. The DD estimates, however, are still subject to at least two other potential sources of misspecification. First, wage changes and Hispanics' location decisions may be jointly determined by transitory labor aggregate by city demand shocks, $\nu_{ck\text{-}jt}$. For instance if Hispanics are better able to foresee future labor aggregate-specific demand shocks, they might migrate to (away from) cities with rising (declining) wages. Since plant openings and closings are often announced in advance and Hispanics are more mobile than whites, these shocks may be a source of endogeneity. Second, the mismeasured Hispanic share variables may bias the estimated enclave effects towards zero. In light of these possibilities, the causal effect of enclave size can only be identified if there is an exogenous determinant of enclave size.

The tendency of newly-arriving Hispanic immigrants (Hispanic natives) to move to (away) from cities established by earlier immigrants suggests one such determinant.[22] In particular, a Hispanic aggregate's share of the labor force in 1980 and its square are used to predict the change in that aggregate's share between 1980 and 1990 in an instrumental variables version of equation (7). The intuition is that changes in Hispanic density are composed of both labor aggregate-specific demand shocks and the desirability (or undesirability) of living with people who share a language background. The IV estimation procedure attempts to identify the enclave effects by purging the component of the change in Hispanic density that is determined by labor demand shocks. This estimation procedure will provide consistent estimates of the enclave effects if two conditions are met: the instruments are correlated with the change in the share Hispanic aggregate measures; and they are uncorrelated with the error in the wage equation.

Appendix Table 1 presents the "first-stage" results that suggest that the first condition is met; over the relevant range of 1980 shares, Hispanics are on average drawn to high-Hispanic cities. This relationship is even stronger for immigrants and the English deficient. In contrast, the table indicates that Hispanic natives are modestly repelled from cities with high shares of natives and the English proficient tend to leave cities whose population of English proficient is greater than 20% of the male labor force. Across all of the aggregates the relationship between the 1980 level and the 1980 to 1990 change is quite strong.[23] In particular, the F-statistic on the instruments (i.e., the 1980 level and its square) in the first-stage equations were 38 (Share Hispanics), 25 (Share Hispanic Natives), 240 (Share Hispanic Immigrants), 15 (Share Hispanic Eng. Deficient), and 86 (Share Hispanic Eng. Proficient).[24] The second condition that the fitted value from the first stage be

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[22] Bartel (1989) first documented that the existing stock of immigrants with the same background are a substantially more important determinant of immigrants' location decisions than are local economic conditions. Altonji and Card (1991) and Card (1997) have used city-specific levels to instrument for city-specific inflows of immigrants.

[23] To reduce the degree of finite sample bias introduced by IV, it is important that the instruments explain a large share of the change in Hispanic aggregates' intensity (Bound, Jaeger, and Baker 1994).

[24] The importance (relative to other factors) of a city's enclave status in determining changes in the share of Hispanics is further underscored by the unimportance of the citywide unemployment rate in the first-stage models. In particular, the citywide unemployment rate is statistically insignificant in all of the models, except for the most skilled aggregate, the English proficient.

uncorrelated with the unobserved components of the wage process is not directly testable.

Columns (3) and (4) of Table 5 present the estimation results from the application of the DDIV procedure. The table also includes Wald statistics from a Hausman test for misspecification of the share Hispanic variables (Hausman 1978). The Hausman test's null hypothesis is that the DD model is correctly specified and hence produces consistent and efficient parameter estimates. Thus when the Hausman null is rejected, DDIV is an appropriate technique to obtain consistent parameter estimates as long as the instruments are uncorrelated with the unobserved components of the wage process.

The results in column (3) of Table 5 reveal that an increase in the share of Hispanics is correlated with substantial declines in the relative wages of all Hispanic labor aggregates. The parameter estimates on the Share Hispanics variable are within one standard deviation of -1 for all six aggregates, indicating that a 1-percentage point increase in the share Hispanics is correlated with a 1% decline in Hispanics relative wages. The view that enclaves raise the wages of the English deficient is undermined by the estimated enclave effects, which are negative and almost identical for both the English proficient and the English deficient aggregates. Another striking regularity across the six aggregates is that the DDIV coefficients on the Share Hispanics variables are of a greater magnitude than the DD ones for all six aggregates. Together the estimated enclave effects suggest that a 1-percentage point increase in the share of Hispanics is associated with a 1% decline in the relative wages of all Hispanic aggregates.

The DDIV estimates in column (4) do not have such a clear pattern. The estimated effects of the Share Hispanic Immigrant and Share English Deficient variables are relatively stable when compared to the DD ones in column (2). However, the Share Hispanic Natives and the Share English Proficient variables occasionally change signs and generally are much less stable across the two columns. These patterns partially reflect the relative "strength" of the relevant first-stage equations. This differential "strength" can also be seen in the change in the sampling errors across the approaches. For example, the standard errors on the share Hispanic immigrant variable only increase by 10-20% from column (2) (i.e., DD) to column (4) (i.e., DDIV), but they almost double for the Share Hispanic Natives variable. In general, the imprecision of the estimates in column (4) makes it difficult to draw any firm conclusions about the separate effects of changes in the shares of these finer groupings of Hispanics.

The Hausman test statistics have an interesting pattern. They indicate that the DD estimates are misspecified for the higher skilled aggregates (i.e., natives and English proficient) but not for the other aggregates. These differences in the test statistics across aggregates suggest that measurement error is not the principal source of the misspecification; if it were, the Hausman test should have rejected the DD model for all aggregates. A more plausible explanation is that higher skilled Hispanics aggregates are better able to foresee and migrate in response to aggregate and city-specific demand shocks. In other words, upon news of an impending increase (decline) in earnings opportunities in another (their) city, members of higher skilled Hispanic aggregates are better able to move to (from) that SMSA than are the lesser skilled Hispanics. This ease of mobility seems to bias the DD estimates of the enclave effects upward for these aggregates.

It is worth mentioning that these models have contained an informal "over-identification" test. Specifically, all of the specifications included the share of the white natives as an explanatory variable; and, the coefficient on this variable was always small and statistically indistinguishable from zero. This is not direct evidence that the share Hispanic variables are capturing the effects of enclave residence, but if this parameter had been an important determinant of Hispanics' relative wages it would have lessened my confidence in the reliability of the estimated enclave effects.

4.4 Are the DD and DDIV Results Robust to Controlling for State-Specific Trends?

It is possible that the enclave effects estimated in Table 5 do <u>not</u> result from variation in the share of Hispanics across cities but are the consequence of unobserved state-level factors. In this section, I allow for unrestricted state-specific trends in Hispanics' relative wages by including a separate indicator for each state. The inclusion of these indicators will absorb any labor aggregate-specific shock that is common to all SMSAs within a state. Consequently, the identification of the enclave effects will come from SMSA-specific changes in Hispanic density, controlling for the statewide change in density.

In addition to providing a robustness check for the DD and DDIV estimates, the state indicators provide an informal test of the validity of the instruments. In particular, the instruments are intended to purge any labor-aggregate by city shocks (i.e., $\Delta u_{cj\text{-}k}$ from equation (7)) to demand. A necessary, albeit not sufficient, condition for the instruments to be "working" is that they also purge any labor aggregate by state shocks to demand. Consequently, the estimated enclave effects from the DDIV specification should be robust to the inclusion of the state by time indicators. The same reasoning does not hold for the DD estimates, which may be still be misspecified due to labor-aggregate by city shocks.

Before turning to the results, two caveats to the estimation of these additional parameters merit attention. First, if enclaves are best defined at the state level, this identification technique will remove valid variation. Second, there are only 280 SMSA observations and the inclusion of the additional regressors opens the possibility that the model will be "over-fit."

Table 6 presents the DD and DDIV estimation results from these heavily parameterized specifications. A comparison of the results in this table and those in Table 5 suggests two different conclusions. For Hispanics as a whole, both immigrant aggregates, and the English deficient, the estimated effect of the Share Hispanics variable in both the DD and DDIV approaches are very similar with and without the state indicators.[25] The differences that do exist are generally insubstantial in the context of the standard errors. Moreover, an F-test on the state indicators fails to reject that they are jointly equal to zero for each of these specifications at the 5% significance level. On the whole, these results are supportive of the estimates in Table 5.

The results for the higher skilled Hispanic aggregates lead to a less sanguine conclusion. Most strikingly, the coefficient on the share Hispanics variable is quite sensitive to the inclusion of the state indicators: it is –1.516 (without them, in Table 5) and –0.019 (with them, in Table 6). Similarly, the estimated enclave effects for the English proficient aggregate, which overlaps with the native one, decline in magnitude by approximately half when the state indicators are included. It seems that the estimated enclave effects for these aggregates may not solely reflect differences in the share of Hispanics; there appear to be important statewide Hispanic native-

[25] It is noteworthy that even within a state, the 1980 shares of Hispanic aggregates are important determinants of the change in the share. The F-tests on the instruments (see the note at the bottom of Table 6) in the "first-stages" reveal that they still do a substantial proportion of the "work."

specific factors that affect both their wages and the density of Hispanics. These findings imply that the instruments do not purge these factors and hence are not "working" properly for the higher skilled aggregates, which undermines my confidence in the reliability of the estimates for these aggregates.

As a whole, the results presented in this sub-section are supportive of the earlier findings that enclave residence reduces the relative wages of Hispanics, particularly the lesser skilled. However, the estimation of the extra parameters in this section, is asking a lot from the data, as can be seen by the large sampling errors. Consequently, the findings from Table 6 should primarily be viewed as suggestive.

4.5 Implications

Taken at face value, the preponderance of the evidence presented in this paper indicates that enclave residence is associated with significant wage declines. In the context of mobile labor, why aren't these differences arbitraged away? The "compensating differentials" literature provides a convenient framework to explain these results.[26] As applied to enclaves, this literature's insight is that Hispanics who choose to live in enclaves must derive the same level of utility from their location choice as those who live outside enclaves. The implication is that the persistence of higher wages outside of the enclave must be "compensation" for some benefits that accrue to enclave residents. These benefits may be non-pecuniary, cultural ones associated with living among people who share a common background, or, as Cranston et al. (1996) have suggested, they may take the form of lower migration costs associated with moving to a city with an existing stock of immigrants. Whatever the exact form of the benefits, the compensation allows for an equilibrium where Hispanics live inside and outside enclaves.

This paper has attempted to purge the confounding differences in labor demand across cities to estimate the "causal" effect of the share of Hispanics in a city on Hispanics' wages. However as is always the case with a non-experimental design, there is a form of unobserved heterogeneity that can also explain the findings without a causal interpretation. Consequently

[26] See Rosen (1986) for a review of the compensating or equalizing differences literature.

even though all observable productivity characteristics have been "regressed" out, the possibility remains that Hispanics with low (high) levels of unobserved ability systematically select into (out) of enclaves. If this non-random selection occurs, the estimated effects presented in this paper would at least partially reflect the lower abilities of Hispanics residing in enclaves and not the "causal" effect of enclave residence.

5. Conclusions

This paper has examined the wage consequences of enclave residence for Hispanics by using the 1980 and 1990 5% Census samples. There are three principal findings. First, there are important citywide and Hispanic-specific demand shocks that influence where Hispanics choose to settle and affect their wage rates. Any examination of Hispanics' wages should control for these factors.[27] Second, on average Hispanics' choose to live in cities with a large existing stock of Hispanics. There is heterogeneity, however, across Hispanics sub-populations; the lesser skilled (i.e., immigrants and the English deficient) aggregates are drawn to enclaves, while natives and the English proficient tend to leave high-Hispanic cities. Third across a number of specifications, I find that enclave residence reduces the relative (compared to white natives) wages of Hispanics. In the least restrictive (and preferred) specification, a 1-percentage point increase in the share of Hispanics in a city reduces the average Hispanic's relative wage by approximately 1%.[28] This appears to hold across broadly defined Hispanic sub-populations, although the evidence is strongest among the least skilled (i.e., recent immigrants and the English deficient). The differences between this finding and the previous literatures' can largely be explained by this paper's efforts to purge the influence of unobserved labor demand factors.

Using a number of different sources of identifying variation, this paper has documented a robust empirical regularity; Hispanics' relative wages are negatively correlated with the share of

[27] This is consistent with the immigration literature, which has found that wage levels are generally higher in cities with high immigrant inflow rates (Card 1997; Schoeni 1996).

[28] Evans (1985) and Chiswick and Miller (1996) found that immigrants who form enclaves in Australia are slower to learn English than immigrants who settle outside the enclave. This finding suggests that in addition to the short-run losses established in this paper, enclave residence may also damage long-term earnings prospects

Hispanics in a city. Future research should concentrate on more firmly determining the basis of this relationship. In particular, it would be important to differentiate between the compensating differentials and low unobserved ability explanations. Since this paper and others provide evidence that some Hispanics choose to leave enclaves, a fruitful direction may be to examine more closely the dynamics that lead to this decision.

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Data Appendix

a. Assigning SMSA Codes in the 1980 and 1990 Censuses

I define enclaves at the SMSA level. Most respondents who live in a SMSA are assigned a SMSA code. The 5/100 samples contain a question on SMSA of residence and for large SMSAs the proper code is disclosed. The confidentiality rules, however, require that the finest geographic unit identified in the 5/100 samples are Public Use Microdata Areas (PUMAs) which are arbitrary geographic divisions that contain no less than 100,000 people. These arbitrary divisions are troublesome for at least two reasons: SMSAs with less than 100,000 people cannot directly be identified; and PUMAs often cross SMSA boundaries. In the case of these "mixed" PUMAs, a SMSA code is not directly assigned.

I used the 1980 and 1990 *County Group Equivalency* files to assign these "mixed" PUMAs an SMSA code. From these files, one can identify the SMSA that contributed the largest fraction of the population to each PUMA. If over 50 percent of the PUMA's population was attributable to a single SMSA, I assigned all PUMA residents to the majority SMSA. If less than 50 percent of a PUMA's population lived in a single SMSA, all respondents in that PUMA were excluded from the analysis. The computer code for this assignment is available upon request.

b. Matching SMSAs across the Censuses

Since the paper relies on comparing changes within a city between 1980 and 1990, it is important that the definition of a city remain stable across the two samples. SMSAs are redefined after each census, generally resulting in the addition or, less frequently, the subtraction of counties. Also new SMSAs might be created and in rare cases they might be deleted.

To obtain a time invariant definition, I redefined 1990 SMSAs to match the 1980 boundaries. Again, I relied on the *County Group Equivalency* files to identify the PUMAs that contained the affected counties in the 1990 Census. If the counties in question comprised more than half of the PUMA's population, I assigned all respondents to the pertinent SMSA. If greater than 10% of a SMSA's 1990 population was affected by the boundary changes and was unrecoverable from the County Equivalency files, I dropped the city from the analysis. Dayton and Springfield, Ohio were the only such cities. This computer code is also available upon request.

Table 1: White and Hispanics Labor Aggregates' Percentage of Male, Urban, Labor Force

	1980	1990	Change
Whites	77.2%	66.8%	-10.4%
All Hispanics	7.1%	9.5%	2.4%
Hispanic Natives	4.2%	3.8%	-0.4%
Hispanic Immigrants	2.9%	5.7%	2.8%
Hispanic English Deficient	1.5%	2.2%	0.7%
Hispanic English Proficient	5.6%	7.2%	1.6%

Notes: These percentages are calculated from the sample of 280 SMSAs that can be matched across the 1980 and 1990 Censuses. Calculations from the 300 SMSAs in the 1980 sample and the 324 SMSAs in the 1990 sample yield almost identical estimates of the labor force percentages.

Table 2: Mean Characteristics of Hispanics, White Natives, and SMSAs, Across SMSAs Grouped by Percentage of Hispanics, 1980 and 1990

	1980				1990			
	1st	2nd	3rd	4th	1st	2nd	3rd	4th
Hispanics								
Share English Deficient	.027	.052	.110	.181	.060	.079	.145	.218
Hourly Wage	6.91	7.73	7.23	6.83	11.48	10.30	10.67	10.47
Potential Experience	14.61	13.86	14.90	16.20	14.40	14.78	15.27	17.11
Year of Education	11.99	11.93	11.18	10.18	12.81	12.33	11.50	10.63
Changed States in Last 5 Years	.322	.301	.316	.282	.330	.324	.306	.168
Self Employed	.030	.015	.023	.017	.023	.025	.026	.024
White Natives								
Hourly Wage	7.79	8.45	8.53	8.63	13.27	13.35	14.50	15.69
Potential Experience	17.10	16.60	16.98	16.30	18.01	17.07	17.26	17.34
Years of Education	12.39	12.72	12.76	13.10	13.07	13.40	13.45	13.82
Changed States in Last 5 Years	.302	.305	.308	.298	.093	.127	.165	.156
Self Employed	.039	.040	.046	.048	.052	.052	.050	.064
SMSA Characteristics								
Population of Working Aged Men	83,831	125,328	136,894	235,935	83,647	128,520	161,911	254,376
Share White	.913	.851	.842	.706	.808	.773	.769	.608
Share Hispanic	.005	.012	.026	.176	.005	.014	.038	.213
Share Hispanic Immigrant	.004	.009	.019	.121	.002	.005	.020	.100
Unemployment Rate	.067	.065	.062	.058	.060	.059	.059	.057

Note 1: In both 1980 and 1990, the SMSAs are ordered by the percentage of Hispanics: 1st refers to the quarter of the SMSAs with the smallest share of Hispanics; 2nd the second smallest share, etc. In calculating the SMSA averages, each city received equal weight. In other words, the averages are not adjusted for city population. There were 300 SMSAs in 1980 and 324 in 1990.
Note 2: All the means for white natives and Hispanics are calculated from men aged 16-64 who worked for pay in the year preceding the Census. The hourly wages was determined by dividing total earnings by the product of weeks worked and usual hours per week. The SMSA characteristics were calculated from working aged men of all races and ethnicities.

Table 3: Enclave Effects or Unobserved City Factors?

	1990 Cross-Section
Hispanic Natives' Wage	
Share Hispanic Natives	-.695***
	(.090)
Share Hispanic Immigrants	.356***
	(.101)
White Natives' Wage	
Share Hispanic Natives	-.693***
	(.088)
Share Hispanic Immigrants	.363***
	(.098)
Hispanic Natives' Wage Minus White Natives' Wage	
Share Hispanic Natives	-.002
	(.052)
Share Hispanic Immigrants	-.007
	(.059)

Notes: Each panel is a separate regression. Entries are estimated regression coefficients of share Hispanic aggregate variables in models for regression adjusted, city wage. The equations also include the fraction of whites in the male, labor force, the SMSA unemployment rate of men ages 16-64, and the natural logarithm of the number of men ages 16-64 both in and out of the labor force. Sample includes data from 324 SMSAs. Standard errors are in parentheses.

* $p<.10$ (two-tailed test), ** $p<.05$ (two-tailed test), *** $p<.01$(two-tailed test)

Table 4: Estimated Differenced Cross-Section Regression Models for Hispanic Aggregates' Regression Adjusted City-Wide Wage

	1980 Differenced Cross-Section		1990 Differenced Cross-Section	
	(1)	(2)	(3)	(4)
Hispanics				
Share Hispanics	-.078		-.081**	
	(.060)		(.040)	
Share Hispanic Natives		-.070		.038
		(.066)		(.051)
Share Hispanic Immigrants		-.097		-.204***
		(.086)		(.052)
Hispanic Natives				
Share Hispanics	.126*		-.004	
	(.070)		(.044)	
Share Hispanic Natives		.141*		-.002
		(.078)		(.052)
Share Hispanic Immigrants		.084		-.007
		(.117)		(.059)
Hispanic Imm's 0-15				
Share Hispanics	-.338***		-.253***	
	(.095)		(.060)	
Share Hispanic Natives		-.393***		-.047
		(.105)		(.086)
Share Hispanic Immigrants		-.244**		-.424***
		(.121)		(.078)
Hispanic Imm's >15				
Share Hispanics	-.022		.131**	
	(.115)		(.059)	
Share Hispanic Natives		-.016		.348***
		(.120)		(.076)
Share Hispanic Immigrants		-.039		-.087
		(.146)		(.076)
Hispanic Eng. Deficient				
Share Hispanics	-.293**		-.398***	
	(.121)		(.068)	
Share Hispanic Eng. Deficient		-.679**		-.912***
		(.339)		(.185)
Share Hispanic Eng. Proficient		-.174		-.122
		(.156)		(.114)
Hispanic Eng. Proficient				
Share Hispanics	-.077		-.022	
	(.060)		(.039)	
Share Hispanic Eng. Deficient		-.501***		-.257**
		(.189)		(.110)
Share Hispanic Eng. Proficient		.038		.076
		(.077)		(.058)

Notes: Entries are estimated regression coefficients of share Hispanic aggregate variables in models for relative, regression adjusted, city wage. Regressions are estimated separately for each labor aggregate (listed in bold). All models include the fraction of whites in the male, labor force, the SMSA unemployment rate of men ages 16-64, and the natural logarithm of men ages 16-64 who are in or out of the labor force. Sample includes data from 300 SMSAs in 1980 and 324 in 1990. Standard errors in parentheses.
* p<.10 (two-tailed test), ** p<.05 (two-tailed test), *** p<.01(two-tailed test)

Table 5: Estimated Differences in Differences and Differences in Differences IV Regression Models for Hispanic Aggregates' Regression Adjusted City-Wide Mean Wage

	Diff in Diff		Diff in Diff IV	
	(1)	(2)	(3)	(4)
Hispanics				
Share Hispanics	-.418**		-1.162***	
	(.169)		(.375)	
Share Hispanic Natives		-.296		.251
		(.326)		(.678)
Share Hispanic Immigrants		-.430**		-.338*
		(.172)		(.197)
Hausman Statistic			3.49*	0.64
Hispanic Natives				
Share Hispanics	-.111		-1.516***	
	(.218)		(.591)	
Share Hispanic Natives		-1.267***		.577
		(.398)		(.797)
Share Hispanic Immigrants		-.006		.188
		(.216)		(.261)
Hausman Statistic			6.49**	-15.95***
Hispanic Imm's 0-15				
Share Hispanics	-.471*		-.666	
	(.252)		(.428)	
Share Hispanic Natives		-.184		-.173
		(.461)		(.835)
Share Hispanic Immigrants		-.515**		-.536*
		(.259)		(.294)
Hausman Statistic			0.25	-0.28
Hispanic Imm's >15				
Share Hispanics	-.318		-.826	
	(.302)		(.554)	
Share Hispanic Natives		.215		-.887
		(.530)		(.892)
Share Hispanic Immigrants		-.420		-.400
		(.313)		(.364)
Hausman Statistic			1.02	-0.74
Hispanic Eng. Deficient				
Share Hispanics	-.571*		-1.161**	
	(.292)		(.558)	
Share Hispanic Eng. Deficient		-.287		-.281
		(.482)		(.862)
Share Hispanic Eng. Proficient		-.903*		-.592
		(.537)		(.926)
Hausman Statistic			1.13	-0.22
Hispanic Eng. Proficient				
Share Hispanics	-.297*		-1.050***	
	(.174)		(.396)	
Share Hispanic Eng. Deficient		-.072		.847
		(.300)		(.525)
Share Hispanic Eng. Proficient		-.612		-1.788***
		(.384)		(.696)
Hausman Statistic			3.84**	3.42

Notes: Entries are estimated regression coefficients of share Hispanic aggregate variables in models for relative, regression adjusted, city wage. Regressions are estimated separately for each labor aggregate (listed in bold). The F-statistic on the instruments (i.e., the 1980 level and its square) in the relevant first-stage equations of columns (3) and (4) were 38 (Share Hispanics), 25 (Share Hispanic Natives), 240 (Share Hispanic Immigrants), 15 (Share Hispanic Eng. Deficient), and 86 (Share Hispanic Eng. Proficient). I also present a Wald test statistic based on the Hausman test to determine whether the DD models in columns (1) and (2) are misspecified. All models include the fraction of whites in the male, labor force, the SMSA unemployment rate of men ages 16-64, and the natural logarithm of the number of men ages 16-64 both in and out of the labor force. Sample includes data from 280 SMSAs. Standard errors in parentheses.
* p<.10 (two-tailed test), ** p<.05 (two-tailed test), *** p<.01(two-tailed test)

Table 6: Are the Results Robust to the Inclusion of State Indicators?

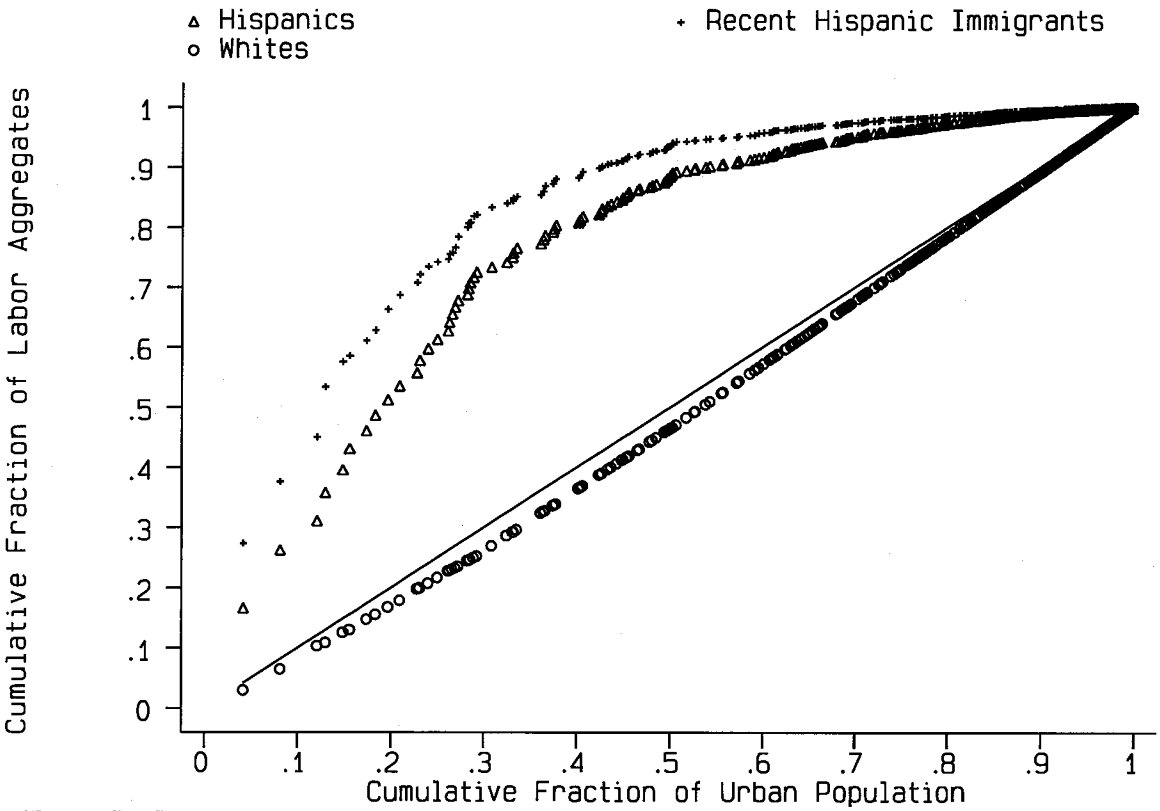
	Diff in Diff		Diff in Diff IV	
	(1)	(2)	(3)	(4)
Hispanics				
Share Hispanics	-0.560***		-0.637	
	(.202)		(.438)	
Share Hispanic Natives		-1.051**		-0.387
		(.466)		(.754)
Share Hispanic Immigrants		-0.525***		-0.343
		(.205)		(.246)
F-Statistic for State FEs	1.41*	1.44*	1.44*	1.36*
Hispanic Natives				
Share Hispanics	-0.156		-0.019	
	(.235)		(.590)	
Share Hispanic Natives		-1.490***		-1.264*
		(.482)		(.745)
Share Hispanic Immigrants		-0.008		0.287
		(.235)		(.287)
F-Statistic for State FEs	2.45***	2.43***	2.59***	2.57***
Hispanic Imm's 0-15				
Share Hispanics	-0.691		-0.704	
	(.317)		(.563)	
Share Hispanic Natives		-0.466		-0.380
		(.692)		(.952)
Share Hispanic Immigrants		-0.709**		-0.739*
		(.322)		(.390)
F-Statistic for State FEs	1.18	1.16	1.29	1.28
Hispanic Imm's >15				
Share Hispanics	-0.382		-1.028	
	(.396)		(.697)	
Share Hispanic Natives		-0.922		0.369
		(.801)		(1.086)
Share Hispanic Immigrants		-0.338		-0.296
		(.400)		(.486)
F-Statistic for State FEs	0.83	0.81	0.86	0.71
Hispanic Eng. Deficient				
Share Hispanics	-0.418		-0.759	
	(.381)		(.782)	
Share Hispanic Eng. Deficient		-0.159		0.246
		(.750)		(.995)
Share Hispanic Eng. Proficient		-0.446		-0.049
		(.389)		(.498)
F-Statistic for State FEs	0.91	0.91	0.90	0.94
Hispanic Eng. Proficient				
Share Hispanics	-0.504		-0.489	
	(.207)		(.450)	
Share Hispanic Eng. Deficient		-1.280***		-0.720
		(.467)		(.747)
Share Hispanic Eng. Proficient		-0.442		-0.267
		(.209)		(.249)
F-Statistic for State FEs	1.38*	(.209)	1.40*	1.36*

Notes: Entries are estimated regression coefficients of share Hispanic aggregate variables in models for relative, regression adjusted, city wage. Regressions are estimated separately for each labor aggregate (listed in bold). I also present the F-statistic for the test that the state fixed effects are jointly zero. The F-statistic on the instruments (i.e., the 1980 level and its square) in the first-stage equations (with the state fixed effects included) were 31 (Share Hispanics), 42 (Share Hispanic Natives), 210 (Share Hispanic Immigrants), 27 (Share Hispanic Eng. Deficient), and 14 (Share Hispanic Eng. Proficient). All models include the fraction of whites in the male, labor force, the SMSA unemployment rate of men ages 16-64, and the natural logarithm of the number of men ages 16-64 both in and out of the labor force. Sample includes data from 280 SMSAs. Standard errors in parentheses.
* $p < .10$ (two-tailed test), ** $p < .05$ (two-tailed test), *** $p < .01$ (two-tailed test)

Appendix Table 1: First Stage Results, Estimated Regression Models for 1990 to 1980 Change in Share of Hispanic Aggregates

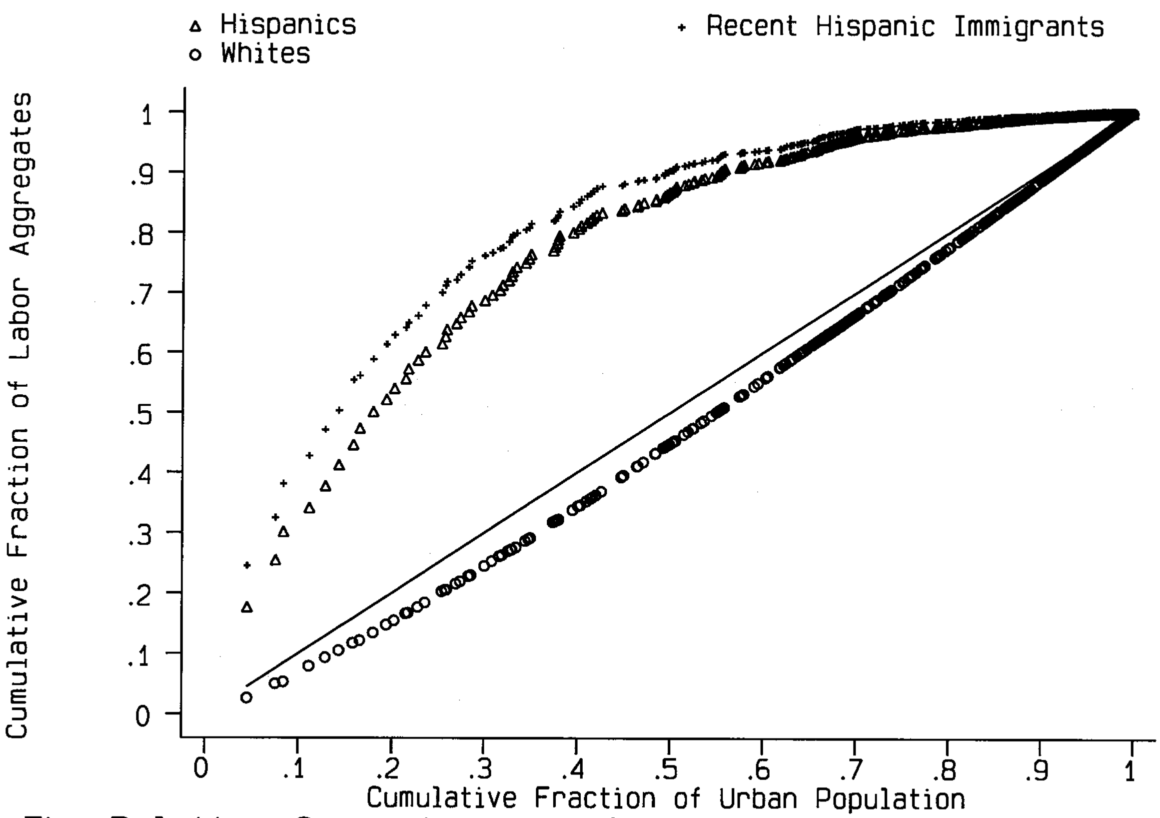
	Regression Results
Change in Share Hispanics	
1980 Share Hispanics	0.251***
	(.029)
1980 Level Squared	-0.350***
	(.049)
F-Statistic for Instruments	38.2***
R squared	0.58
Change in Share Hispanic Natives	
1980 Share Hispanic Natives	-0.090***
	(.026)
1980 Level Squared	0.048
	(.059)
F-Statistic for Instruments	24.6***
R squared	0.18
Change in Share Hispanic Imm's	
1980 Share Hispanic Imm's	0.720***
	(.044)
1980 Level Squared	-1.441***
	(.157)
F-Statistic for Instruments	240.2***
R squared	0.80
Change in Share Hispanic Eng. Deficient	
1980 Share Hispanic Eng. Deficient	0.165***
	(.036)
1980 Level Squared	-.218***
	(.067)
F-Statistic for Instruments	15.3***
R squared	0.55
Change in Share Hispanic Eng. Proficient	
1980 Share Hispanic Eng. Proficient	.616***
	(.047)
1980 Level Squared	-3.596***
	(.300)
F-Statistic for Instruments	85.9***
R squared	0.61

Figure 1a



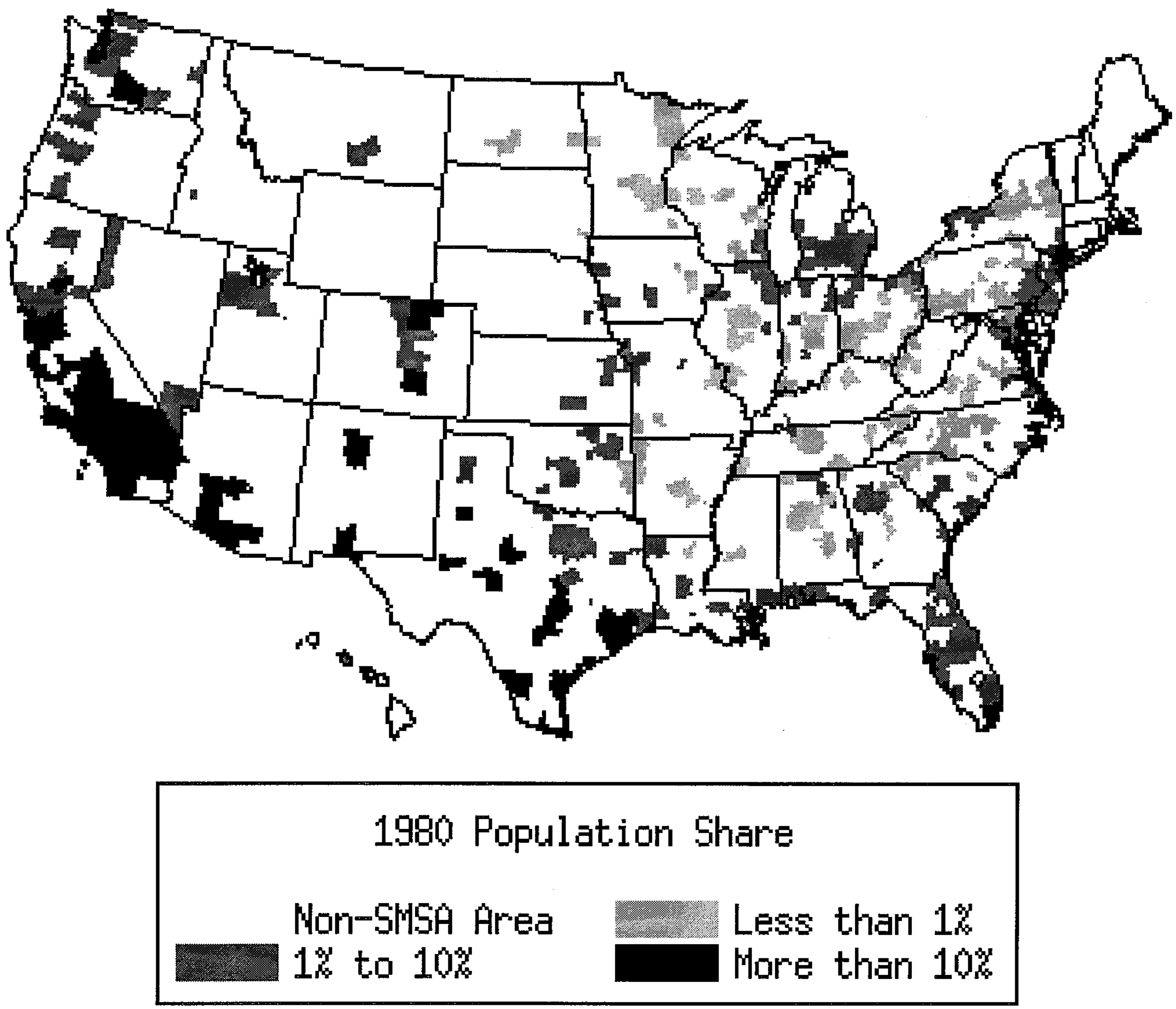
The Relative Concentration of Hispanics in Cities, 1980

Figure 1b



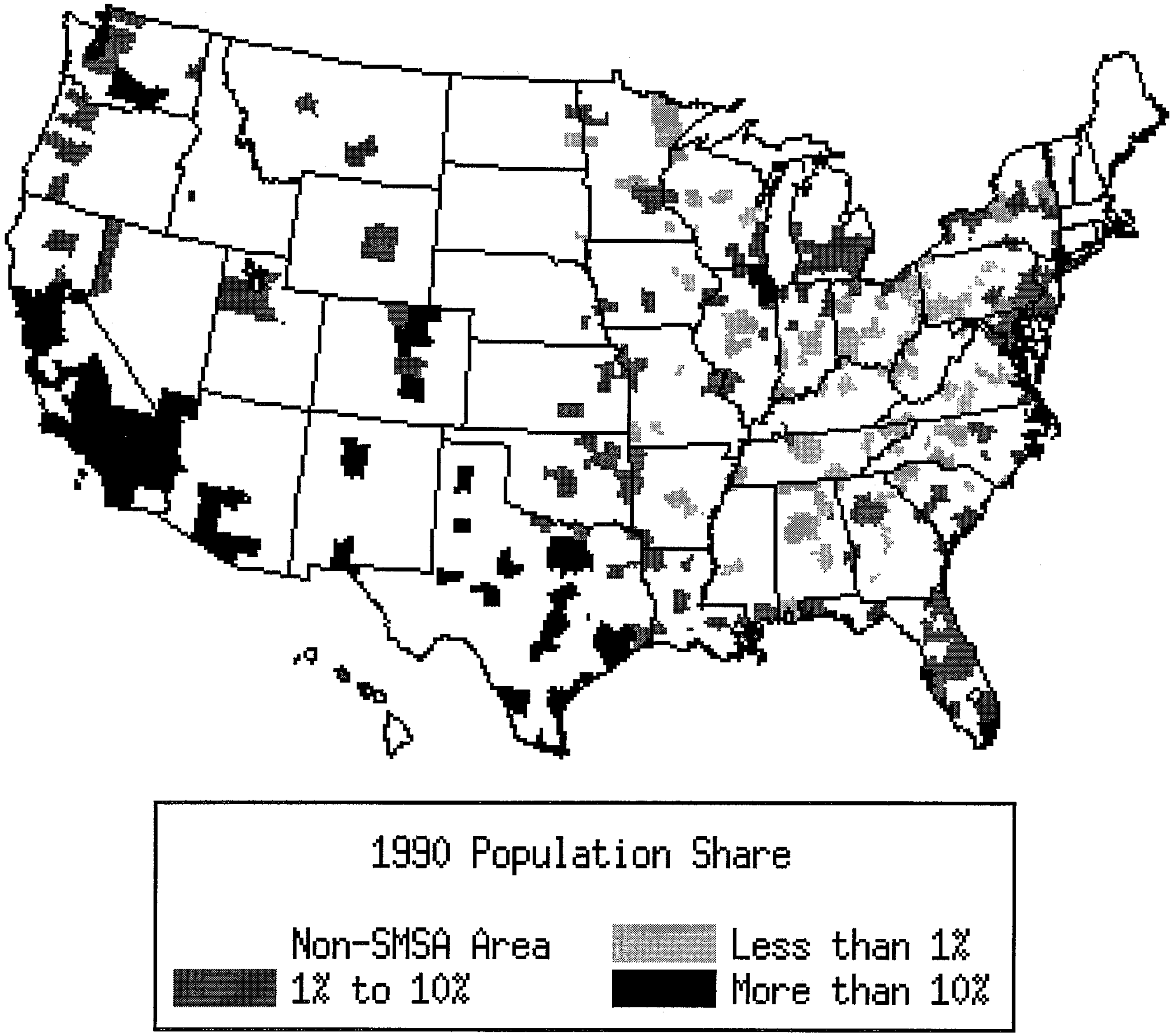
The Relative Concentration of Hispanics in Cities, 1990

Figure 2a: Hispanics' Share of Male Labor Force by SMSA, 1980



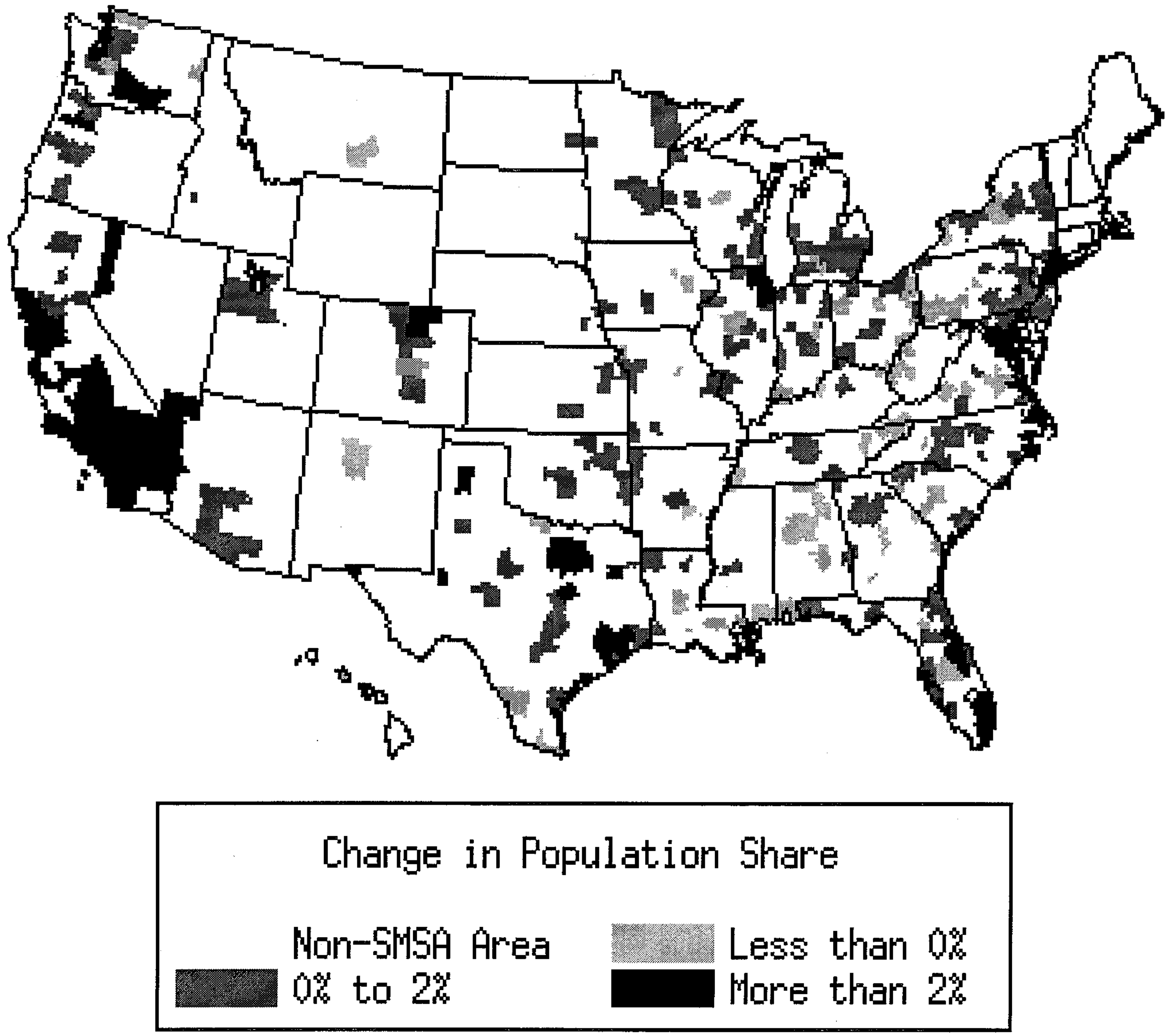
Source: Author's Calculations from 1980 5% Census samples

Figure 2b: Hispanics' Share of Male Labor Force by SMSA, 1990



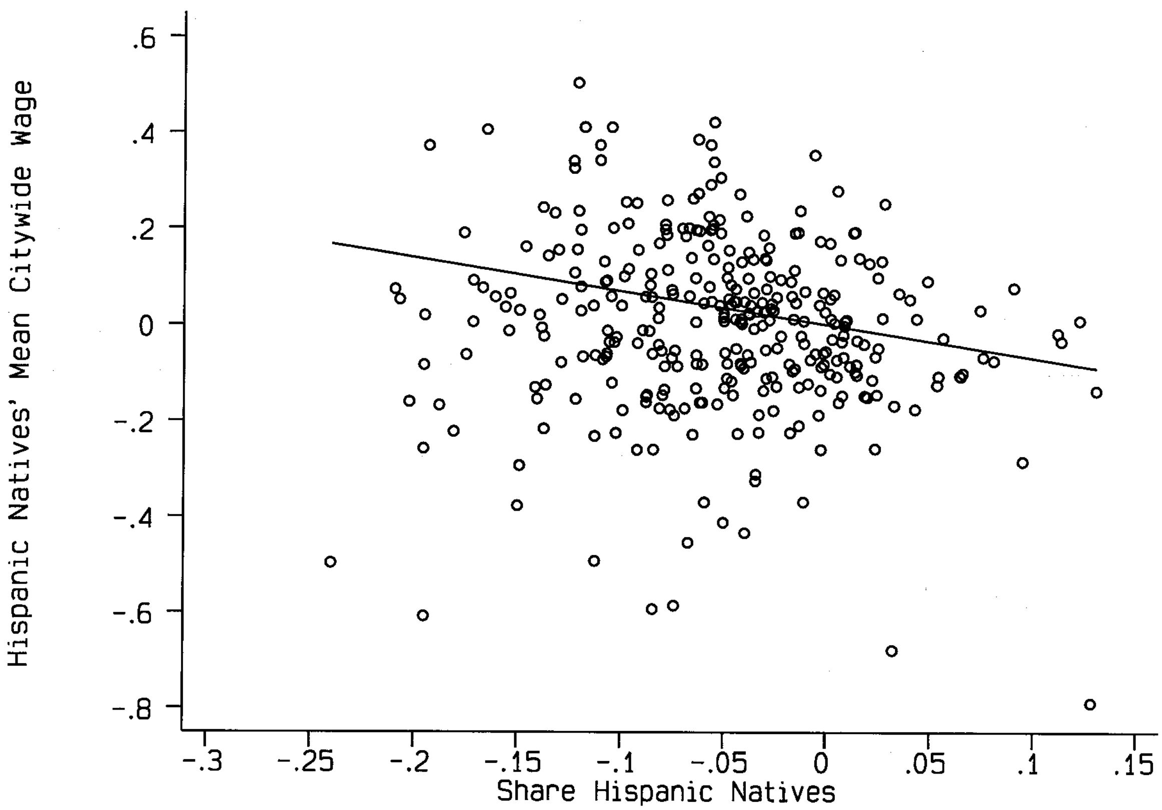
Source: Author's Calculations from 1990 5% Census samples

Figure 2c: Change in Hispanics' Share of Male Labor Force by SMSA, 1990 - 1980



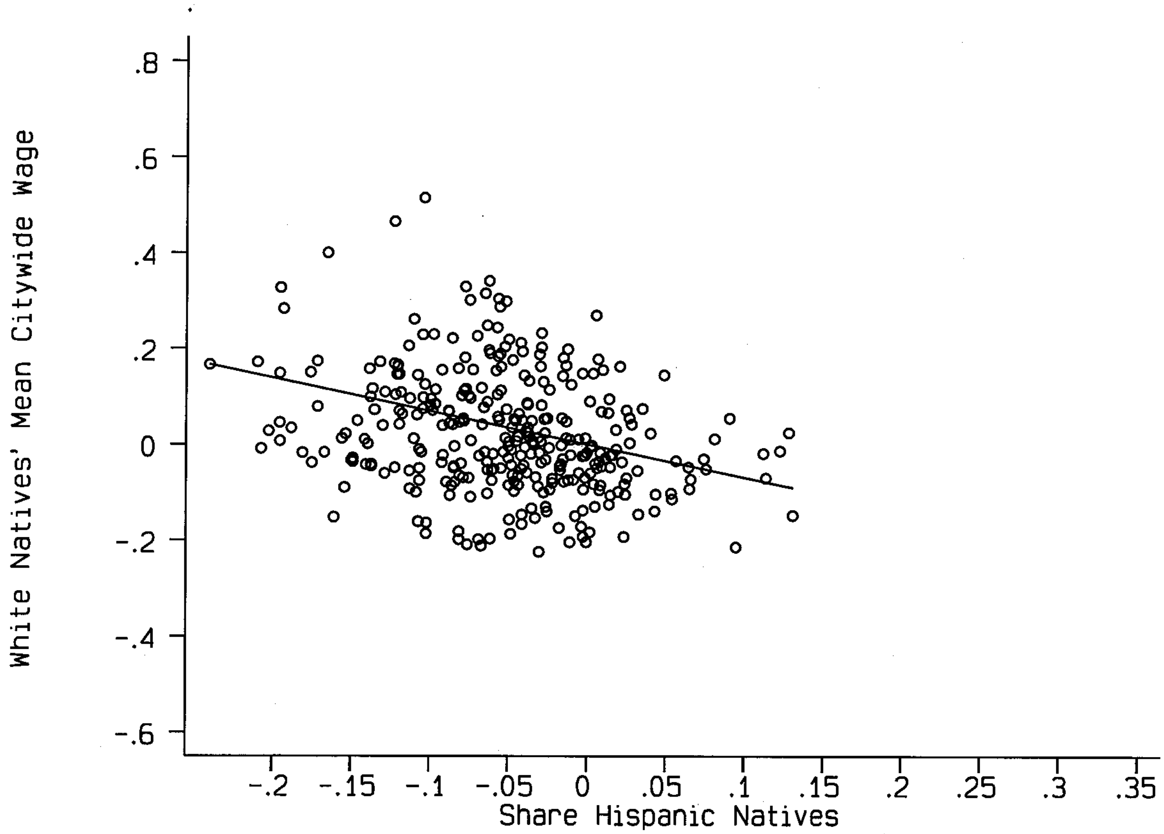
Source: Author's Calculations from 1980 and 1990 5% Census samples

Figure 3a: Correlation Between Hispanic Natives' Regression Adjusted Citywide Wage and Share Hispanic Natives



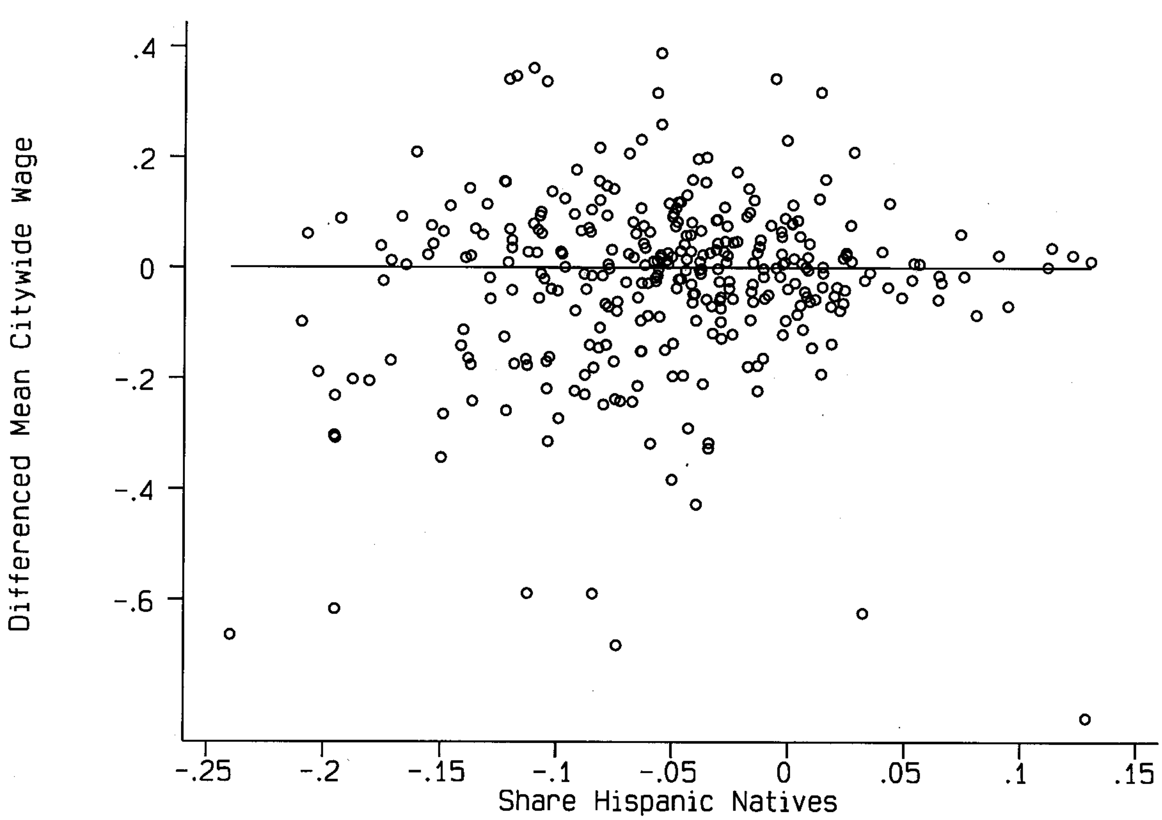
Note: See footnote 17 for a description of the technique underlying this graph.

Figure 3b: Correlation Between Whites' Regression Adjusted Citywide Wage and Share Hispanic Natives



Note: See footnote 17 for a description of the technique underlying this graph.

Figure 3c: Correlation Between Differenced (Hispanic Natives' Minus Whites')
Regression Adjusted Citywide Wages and Share Hispanic Natives



Note: See footnote 17 for a description of the technique underlying this graph.