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RICHARD T. ELY LECTURE

Immigration and Inequality

By David Card*

Immigration has long been a controversial topic among economists.[1] The issue nearly disappeared in the 1960s, but over the past three decades professional interest has picked up as immigrant inflows have surged. The new immigration has attracted attention in part because of its sheer size—approximately 1.25 million people per year over the first half of this decade—and in part because of its composition.[2] A third or more of the new arrivals are undocumented immigrants from Mexico and Central America with low education and limited English skills (Jeffrey Passel 2005). These immigrants presumably compete for the same jobs held by the least-skilled native workers, contributing to a trifecta of economic factors—technology, trade, and immigration—that are thought to have led to a rise in skill differentials in the US economy since the late 1970s.

This paper presents an overview and synthesis of research on the connection between immigration and wage inequality, focusing on the evidence derived from comparisons across major US cities. The appeal of this research design is illustrated in Table 1, which presents recent data on immigrant densities, education outcomes, and mean salaries for workers in 12 of the nation's largest cities.[3] The immigrant share of the working-age population in these cities ranges from about 10 percent in Philadelphia and Detroit to nearly 50 percent in Los Angeles. These differences are correlated with large differences in the relative shares of different skill groups. Figure 1, for example, presents a scatter plot of the share of high school dropouts in the working-age population of the biggest 124 cities against the corresponding immigrant share. On average, each percentage point rise in the immigrant share is associated with a 0.2 percent rise in the relative share of dropouts.[4] Consistent with this general pattern, the dropout share in Los Angeles is over twice as large as the share in Philadelphia or Detroit. My view is that there is something to be learned about labor markets in general, and the effects of immigration in particular, by studying how wages and other outcomes respond to this variation.

Nevertheless, as emphasized by George Borjas, Richard Freeman, and Lawrence Katz (1997), cross-city comparisons are far from a panacea. Natives and immigrants can move between cities, and depending on how sensitive these flows are to differences in local wages, naive cross-city comparisons may reveal a lot or little about the underlying technological parameters that determine the effects of immigration on native opportunities. In fact, Borjas, Freeman, and Katz (1997) argue that mobility rates are so sensitive to relative wages that

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[1] The founding president of the American Economic Association, Francis A. Walker, wrote a well-known article arguing in favor of restricting immigration (Walker 1896). Walker believed that a particular problem was the declining "quality" of the new immigrants. He wrote: "Fifty, even thirty years ago, there was a rightful presumption regarding the average immigrant that he was among the most enterprising, thrifty, alert, adventurous, and courageous of the community from which he came. It required no small energy, prudence, forethought, and pains to conduct the inquiries relating to his migration, to accumulate the necessary means, and to find his way across the Atlantic. To-day the presumption is completely reversed." Walker's view was disputed by Paul H. Douglas (1919).

[2] See US Department of Commerce (2006).

[3] Throughout this paper, I identify "cities" with Metropolitan Statistical Areas (MSAs), or in the case of larger urban agglomerations, with the constituent Primary Metropolitan Statistical Areas (PMSAs).

[4] The linear regression coefficient is 0.21 with a standard error of 0.02; the R^2 is 0.38.

Table 1—Immigrant Presence, Education, and Earnings in Large US Cities

	Working-age population (thousands)	Share of US population (percent)	Percent immigrants	Education distribution: Less than high school	Education distribution: College or more	Mean salary
All US	174,870	100.0	16	14	28	39,806
Larger cities (124)	116,748	66.8	21	14	31	43,102
Rest of country	58,122	33.2	7	15	21	32,988
Los Angeles	5,828	3.3	48	24	27	41,260
New York City	5,687	3.3	44	17	35	49,613
Chicago	5,114	2.9	25	14	34	46,522
Washington, DC	3,359	1.9	25	10	45	56,076
Atlanta	3,055	1.7	17	13	34	44,110
Philadelphia	3,017	1.7	11	11	33	46,149
Houston	2,904	1.7	31	22	27	41,688
Detroit	2,634	1.5	11	12	27	43,937
Dallas	2,516	1.4	26	20	30	42,545
Phoenix	2,348	1.3	22	17	26	40,949
Riverside County, CA	2,266	1.3	31	23	17	37,409
Boston	2,055	1.2	22	9	46	52,584

Notes: Based on tabulation of 2005 and 2006 American Community Surveys. Working-age population includes people age 18 or older with 1–45 years of potential experience. Cities are Primary Metropolitan Statistical Areas (PMSAs).

intercity comparisons are essentially uninformative. Instead, they propose the use of aggregate time series models to estimate the underlying parameters of interest.

I will make the case that appropriately identified parameters derived from cross-city and aggregate time series comparisons are in fact mutually consistent. In particular, estimates from both data sources point to three key conclusions. The first is that workers with less than a high school education are *perfect substitutes* for those with a high school education. This conclusion is important because it means that the impact of low-skilled immigration is diffused across a relatively wide segment of the labor market (the roughly 60 percent of the population who are counted as "high school equivalents" workers) rather than concentrated among the much smaller dropout population (only 14 percent of the population). The second conclusion is that "high school equivalent" and "college-equivalent" workers are imperfect substitutes, with an elasticity of substitution that appears to be similar whether the relevant labor market is defined as the nation as a whole, or an individual city.[5] A third conclusion is that within broad education classes, immigrant and native workers appear to be imperfect substitutes, with a large but finite elasticity of substitution. As was first pointed out by Gianmarco Ottaviano and Giovanni Peri (2006), if immigrants and natives in the same skill category are imperfect substitutes, the competitive effects of additional immigrant inflows are concentrated among immigrants themselves, lessening the impacts on natives.

When the demand side of the national labor market is parameterized with these three assumptions, and capital is assumed to be perfectly elastically supplied, the net impact of immigrant inflows over the past two decades on mean wages of different subgroups of native workers is quite small (Ottaviano and Peri 2008). A similar parameterization of the

[5] Franklin Fisher (1969) showed that the conditions for existence of a well-behaved aggregate production function—even when all firms face identical prices for all inputs and outputs—are extremely stringent. In particular, the existence of a labor aggregate such as "high school labor" requires the absence of specialization across firms. Nevertheless, simulations reported in Fisher, Robert Solow, and James Kearl (1977) suggest that economies made up of heterogeneous firms with different constant elasticity of substitution (CES) production functions behave as if they were generated by an aggregate CES. An interesting issue for further research is under what conditions an aggregate economy made up of separated markets—each containing heterogeneous CES firms, and each with potentially different relative prices—behaves as a single aggregate CES, and whether the synthetic parameters at the market and aggregate levels are related in any systematic way.

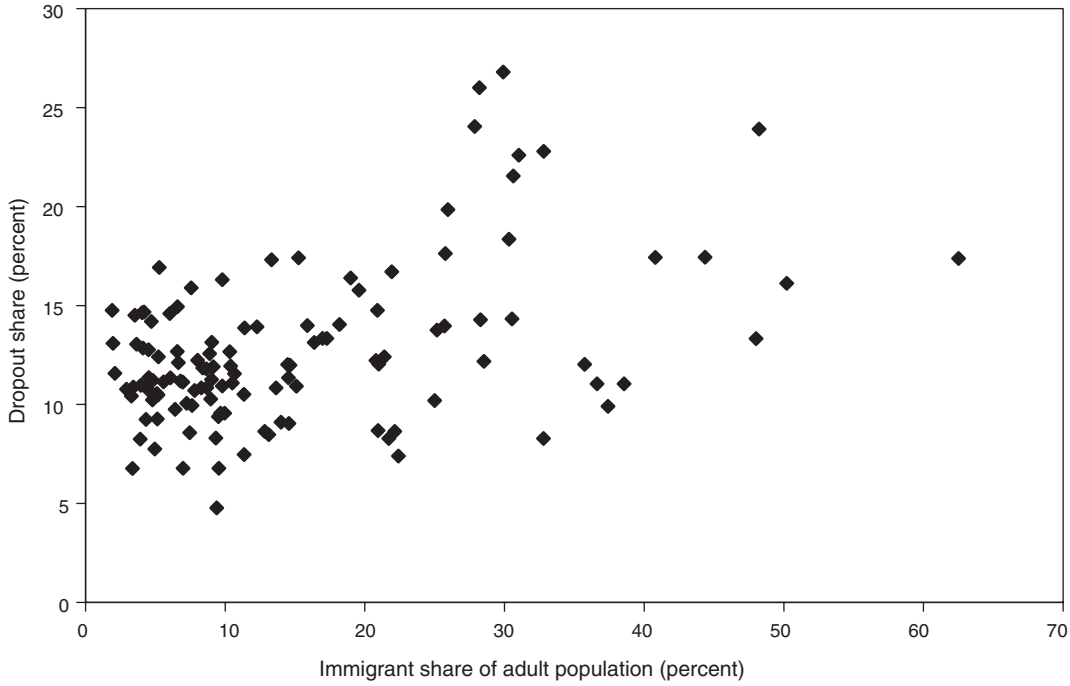

Figure 1. Immigrant Presence and Dropout Share

demand side of the local labor market implies that the effects of immigration on mean wages for different groups of natives in most cities are also small.

Most of the existing research on immigration has focused on between-group inequality. A significant share of the overall rise in US wage inequality is attributable to increases in within-group or residual wage inequality—the variation that remains after controlling for education, age, gender, race, and ethnicity (see Thomas Lemieux 2008 for a recent summary). Empirically, residual wage variation among native workers is significantly higher in cities with more immigrants. The relative level of residual wage inequality for natives in different skill groups is uncorrelated with the relative fraction of immigrants, however, suggesting that immigration has a relatively small causal effect.[6] Taken together with the evidence on between-group wage differentials, I conclude that immigration has had very small impacts on wage inequality among natives.

Nevertheless, when immigrants themselves are counted in the overall population, their effect on inequality is clearly positive. Immigrants tend to be concentrated in the upper and lower tails of the skill distribution. Residual wage inequality among immigrants is also higher than among natives. A simple calculation suggests that the presence of immigrants can explain about 5 percent of the rise in overall wage inequality between 1980 and 2000.

I. Conceptual Framework

A. *Overview*

The main theoretical apparatus used in studies of wage inequality, and in studies of immigration, is a model of the demand side of the labor market.[7] In both literatures the supply side is usually

[6] Lemieux's (2008) comparisons of trends in residual wage inequality by subperiod and within narrow occupations also seem to rule out a major role for immigration.

[7] In the inequality context see, for example, Katz and Kevin Murphy (1992), Murphy and Finis Welch (1992), Katz and David Autor (1999), and Claudia Goldin and Katz (2008), all of which focus exclusively on demand

simplified by assuming that per capita labor supply is perfectly inelastic, although in models of local labor markets the number of workers can potentially vary (e.g., Card 2001). Most often, capital is assumed to be separable from labor inputs so the issue becomes one of specifying the relative demand for different skill groups.

As emphasized in international trade theory, some fraction of the overall response of labor demand to relative wages presumably arises through sectoral adjustments (see Peter Kuhn and Ian Wooten 1991, for an illuminating analysis in the immigration context). In the stark world of the Hecksher-Olin theorem, this channel is so important that relative demand curves are flat: changes in the relative supplies of different skill groups lead to expansions and contractions of different industries with no change in relative wage or relative skill utilization within any particular industry. One might guess that sectoral adjustments are particularly important in understanding derived demand in local labor markets (e.g., at the city or state level). Surprisingly, however, this does not appear to be the case. Ethan Lewis (2003) showed that differences in relative supplies of different education groups across cities are almost entirely absorbed by within-industry changes in utilization.[8] Likewise, Lewis (2008) concludes that sectoral shifts played a relatively small role in the adjustment of the Miami labor market to the Mariel Boatlift. In view of these results it does not appear too unrealistic to adopt a one-sector model of the demand for labor at either the national or local levels.

In a one-sector framework, the properties of the relative demands for different skill groups are derived from the properties of the relevant production function, which I will write in general form as

$$y = f(N_1, N_2, \ldots, N_S), \tag{1}$$

where N_S represents labor input from skill group $s = 1, 2, \ldots, S$. It is standard to assume that f exhibits constant returns to scale, implying that the inverse demand functions for each type of labor are homogenous of degree 0 in the vector of quantities $N_1, N_2, \ldots, N_S$. This has the important implication that a "skill-balanced" inflow of immigrants—one with the same skill distribution as the existing labor force—has no effect on the relative wage structure.[9]

As a point of departure it is helpful to think of the case where all labor types are perfect substitutes, so f has the form

$$f(N_1, N_2, \ldots, N_S) = h(\Sigma_s \, \theta_s \, N_s) \tag{2}$$

for some set of (possibly time-varying) productivity weights θ_s. In this case the relative wage between any two groups, say 1 and 2, is determined by the ratio θ_1/θ_2, and is *independent* of relative supply. Moreover, if capital is freely available at a fixed price, immigration has no effect on the absolute level of wages for any group. While a one-skill model is extremely simple, there is a long tradition in labor and macroeconomics of using such a model, and I believe it represents a useful "null hypothesis."

The most widely used form for the production function is a "two-group CES" in which workers are partitioned into "high school equivalents" (H) and "college equivalents" (C).[10] Within each subgroup, workers are assumed to be perfect substitutes. Thus,

$$f(N_1, N_2, \ldots, N_S) = [\, \alpha_H \, L_{H^\rho} + \alpha_C \, L_{C^\rho}]^{1/\rho}, \tag{3}$$

$$L_H = \sum_{s \in H} \theta_s \, N_s,$$

$$L_C = \sum_{s \in C} \theta_s \, N_s,$$

$$\rho = (\sigma_e - 1)/\sigma_e,$$

side modeling. In the immigration context see George E. Johnson (1980), Jean Grossman (1982), Borjas (2003), and Ottaviano and Peri (2006), all of which likewise focus on the demand side.

[8] Card and Lewis (2007) present a simplified version of Lewis's analysis for the case of Mexican immigrants and reach the same conclusion.

[9] Provided that capital is available at a fixed interest rate, such a "balanced" immigration will have no effect on the level of wages for any group either. Researchers sometimes simulate the effect of immigration assuming that the US capital stock is fixed (e.g., Borjas and Katz 2007). My view is that, over a longer horizon, it is sensible to assume that capital is elastically supplied to the United States. See Ottaviano and Peri (2008) for more discussion.

[10] I am not sure of the origins of this specification, but it dates at least to Freeman (1976). H is assumed to consist of dropouts, people with exactly a high school education, and a fraction (usually ½) of those with 1–3 years of postsecondary education; C is assumed to contain all those with a BA or more, plus the remainder of the "some college" group.

where σ_e is the elasticity of substitution between high school and college labor. Usually, high school dropouts, high school graduates, and some fraction of people with 1 to 3 years of postsecondary education are classified as high school equivalents. College graduates and the remaining fraction of those with some college education are classified as college-equivalents (see, e.g., Katz and Murphy 1992). This simple specification has become a cornerstone of the wage inequality literature (see, e.g., Katz and Murphy 1992; Katz and Autor 1999; Daron Acemoglu 2002).

Estimates based on US time series data up to the early 1990s suggested that with the addition of a linear trend term (representing skill-biased technical change), a model based on (3) could provide a reasonable description of education-related wage differences in the economy, with estimates for $1/\sigma_e$ centered around 0.7. The addition of more recent data tends to lead to smaller estimates of $1/\sigma_e$: for example, Acemoglu (2002) reports a value of 0.56 using data for 1939–1996, while Ottaviano and Peri (2008) present a range of estimates between 0.3 and 0.7. Goldin and Katz (2008) argue that recent estimates are confounded by a slowdown in the pace of skill-biased technical change (i.e., the trend in α_C/α_H) in the early 1990s. Allowing for such a shift, they obtain estimates for $1/\sigma_e$ close to the benchmark 0.7 number.

The assumption embodied in (3) that there are only two distinct skill groups seems relatively strong, and a number of extensions have been proposed. Card and Lemieux (2001) proposed a nested CES structure, allowing for imperfect substitution between different age or experience groups within each of the H and C groups. Their implied estimates of the elasticity of substitution (σ_x) are relatively large (10 or so) and the age structure of the immigrant labor force is not so different from that of natives, so ignoring this form of imperfect substitution makes relatively little difference in the immigration context (see Ottaviano and Peri 2008 for comparisons of various simulations).

A much more important distinction is whether two education classes are enough. Borjas (2003) and Borjas and Katz (2007) assume there are four: dropouts (D), high school graduates (H), people with some college (S), and college graduates (C), implying a specification like

$$(4)\quad f(N_1, N_2, \ldots, N_S) = [\alpha_D L_{D^\rho} + \alpha_H L_{H^\rho} + \alpha_S L_{S^\rho} + \alpha_C L_{C^\rho}]^{1/\rho},$$

where L_D, L_H, L_S, and L_C are CES aggregates of labor inputs of different experience groups within each education class, as in Card and Lemieux (2001). Note that the elasticity of substitution between any two education groups ($\sigma_e = 1/(1-\rho)$) is assumed to be constant. Thus, the inverse relative demand function for any two groups j and k has the simple form

$$(5)\quad \log w_j/w_k = \log \alpha_j/\alpha_k - 1/\sigma_e \log L_j/L_k.$$

In the two-skill case, an equation like (5) holds for the college–high school wage premium, but not for the wage of dropouts relative to high school graduates, which depends only on the relative number of efficiency units of high school–equivalent labor held by the two groups (i.e., the θ_s's in equation (3)).

Though seemingly innocuous, the switch from a two-skills model like (3) to a four-skills model like (4) has extremely important implications for the potential effect of immigration on native wage inequality. This is because the immigrant population has a high fraction of dropouts relative to natives (31 percent versus 11 percent) but a very similar fraction of high school equivalents (63 percent versus 59 percent).[11] In a four-skills model the relatively high fraction of dropouts in the immigrant population distorts the relative share of dropouts and lowers the wage of dropouts relative to other education groups. In a two-skills model, however, the share of dropouts is irrelevant. What matters, instead, is the relative share of high school–equivalent labor, which is unaffected by immigration if immigrants and natives have the same relative shares of high school–equivalent labor.

A second important question is whether immigrants and natives in the same education

[11] In the 2005/2006 American Community Survey, the education distribution for working-age immigrants is 30.5 percent dropouts, 24.2 percent high school graduates, 18.6 percent some college, and 26.7 percent college graduates. The corresponding fractions for natives are 10.8 percent, 30.4 percent, 30.9 percent, and 27.8 percent. I assume that each dropout supplies 0.7 units of high school labor, that one-half of those with some college supply 1.2 units of high school labor, and that the other half supply 0.8 units of college-equivalent labor.

(and experience) class are perfect substitutes (as is assumed in Borjas 2003; Borjas and Katz 2007), or imperfect substitutes (as was implicitly assumed in the seminal studies by Grossman 1982 and Borjas 1987). Ottaviano and Peri (2006, 2008) and Marco Manacorda, Alan Manning, and Jonathan Wadsworth (2006) propose an additional nest to the CES structure to allow for imperfect substitutability. In these studies there are three elasticities of substitution: one between immigrants and natives with the same age and education (which I will denote by σ_I); one within education classes between different experience groups (σ_x); and the third (σ_e) between education groups.

Interestingly, even a modest degree of imperfect substitutability between natives and immigrants makes a significant difference in the implied impacts of immigration on native wages. Loosely speaking, the higher is $1/\sigma_I$, the greater is the concentration of the wage impacts caused by immigrant inflows on immigrants themselves, and the smaller is the spillover effect on natives. For example, in a two-education group model that ignores differences by age (i.e., $1/\sigma_x = 0$), the effect of an inflow of high school–equivalent immigrants on the college–high school wage gap for *natives* depends on the difference $[1/\sigma_e - 1/\sigma_I]$, rather than on $1/\sigma_e$ as in equation (5).[12]

B. *Aggregate Time Series Evidence on the Relevant Elasticities*

Before turning to the main evidence that I review here, drawn from cross-city comparisons, it is useful to briefly summarize the state of the literature based on aggregate time series studies. Focusing first on the issue of the number of education groups, my reading is that recent studies support a relatively simple two-group structure. Goldin and Katz (2008, chap. 8) present an analysis of the determinants of the high school graduate wage premium from 1915 to 2006, allowing the inverse elasticity of substitution between high school graduates and dropouts to vary over time. In a model with a simple linear trend interaction, they find that the inverse elasticity was 0.35 at the beginning of their sample and has steadily trended toward zero. Indeed, their estimates imply that the inverse elasticity reached zero in 1985. Confirmatory evidence is presented by Ottaviano and Peri (2008), who use annual data from 1963 to 2006, and obtain estimates for the inverse elasticity of substitution between dropouts and high school graduates in the range of 0 to 0.04. Ottaviano and Peri (2008) also attempt to estimate the elasticity of substitution between workers with 1–3 years of college, and those with a bachelor's degree or more.[13] Again, they obtain estimates of the inverse substitution elasticity that are small and statistically insignificant (taking account of serial correlation).

These findings shed some light on the apparent difficulty that Borjas (2003) and Borjas and Katz (2007) had in attempting to estimate a single inverse elasticity of substitution among four education groups. Borjas (2003) reports two estimates of $1/\sigma_e$: 0.74 (with a standard error = 0.65) and 0.76 (with standard error = 0.58). Borjas and Katz (2007) report an estimate of 0.41 (standard error = 0.31). I conjecture that an important reason for the imprecision (apart from the relatively small number of time series observations) is that the data do not support the four-skills model. In particular, the wage gap between dropouts and high school graduates appears to be uncorrelated with the relative supply of dropouts. Moreover, judging from the estimation results reported in Katz and Murphy (1992), Acemoglu (2002), and Goldin and Katz (2008), the college–high school wage gap depends on the relative number of people with 12 years *or less* of schooling, not just the number with exactly 12 years of schooling, so again a four-group model is misspecified.

There is somewhat less time series evidence on the elasticity of substitution between immigrants and natives (σ_I). Ottaviano and Peri (2006) originally reported estimates of $1/\sigma_I$ in the range of 0.10 to 0.20. Some details of their analysis have been criticized by Borjas, Jeffrey Grogger, and Gordon H. Hanson (2008), and in their latest work Ottaviano and Peri (2008) present estimates for $1/\sigma_I$

[12] See Manacorda, Manning, and Wadsworth (2006), equation 12.

[13] Strictly speaking, the traditional two-education group model assigns some of the workers with some college to the high school group and the remainder to the college graduate group.

Table 2—Characteristics of Immigrants in 2000

	Working age population (thousands)	Share of all immigrants (percent)	Fraction arrived		Educational attainment			
			After 1980	After 1990	Mean years completed	Dropouts	12–15 years	College or more
Natives	141,272	—	—	—	13.3	14.2	60.6	25.2
Immigrants	23,627	100.0	70.5	39.9	11.6	37.4	38.8	23.8
By country of origin								
Mexico	7,267	30.8	75.1	43.8	8.6	69.8	26.5	3.7
Philippines	1,078	4.6	66.1	31.5	14.1	9.2	43.7	47.0
India	838	3.5	78.4	51.4	15.6	9.6	20.2	70.2
Vietnam	806	3.4	75.3	39.7	11.7	34.6	45.8	19.6
China	715	3.0	82.0	50.1	13.6	24.2	29.2	46.7
El Salvador	698	3.0	85.1	37.0	8.9	65.0	30.6	4.4
Korea	664	2.8	66.4	33.1	14.0	10.6	45.8	43.6
Cuba	586	2.5	52.3	29.1	12.5	30.0	48.3	21.7
Dominican Republic	536	2.3	74.2	38.1	10.8	48.8	41.9	9.3
Canada	517	2.2	47.6	31.9	14.3	8.9	49.8	41.3
Germany	455	1.9	32.6	21.0	13.9	8.3	59.3	32.4
Jamaica	429	1.8	66.7	27.3	12.6	23.8	57.8	18.4
Columbia	400	1.7	71.9	40.5	12.5	24.7	53.3	21.9
Guatemala	400	1.7	84.0	45.9	8.8	64.5	30.4	5.1
Haiti	333	1.4	75.1	34.5	11.8	35.2	51.3	13.5
Poland	310	1.3	74.5	42.3	13.3	16.3	58.2	25.6

Notes: Based on tabulation of 2000 Census. Working age population includes people age 18 or older with 1–45 years of experience.

that range from 0.04 to 0.08. Though small in magnitude, the estimates are quite precise. Manacorda, Manning, and Wadsworth (2006) present estimates based on aggregate UK data that are larger in magnitude (around 0.15) but also relatively precise.

II. Immigrant Settlement Patterns as a Source of Identifying Information

While the aggregate time series evidence is useful, I believe that additional information on the critical parameters of the demand side of the labor market can be gleaned from cross-city comparisons. As noted above, the main issue in interpreting cross-city comparisons is that the supplies of labor in a city may respond to relative wages, leading to a classic identification problem. In this paper I present evidence based on instrumental variables (IV) specifications that use earlier immigrant settlement patterns as a source of identifying information. This section briefly reviews the basis for this strategy.

As a starting point, Table 2 summarizes some of the main characteristics of immigrants from different sending countries observed in the 2000 census. For reference, the first two rows compare natives and all immigrants. As noted in the last section, a striking difference is in the distribution of very low versus "middle" levels of education. Immigrants are relatively clustered at the lowest levels of education, while natives are clustered near the middle. The shares in the upper quartile of the education distribution are more similar. Comparisons across the different rows of the table show that the education distributions vary widely by source country. Immigrants from Mexico—by far the largest source country—are very poorly educated, as are those from El Salvador and Guatemala. Immigrants from the Philippines and India (the second and third largest source countries, respectively) are better-educated than natives.

A well-known fact about immigrants is that they tend to settle in country-specific enclaves. Interesting examples include the clustering of Arab immigrants in Detroit (see Nabeel Abraham and Andrew Shryock 2000), Polish immigrants in Chicago (Dominic Pacyga 1991), and Mexican immigrants in Los Angeles and Chicago. Figure 2 illustrates the pattern using data for Filipino immigrants in the 124 largest

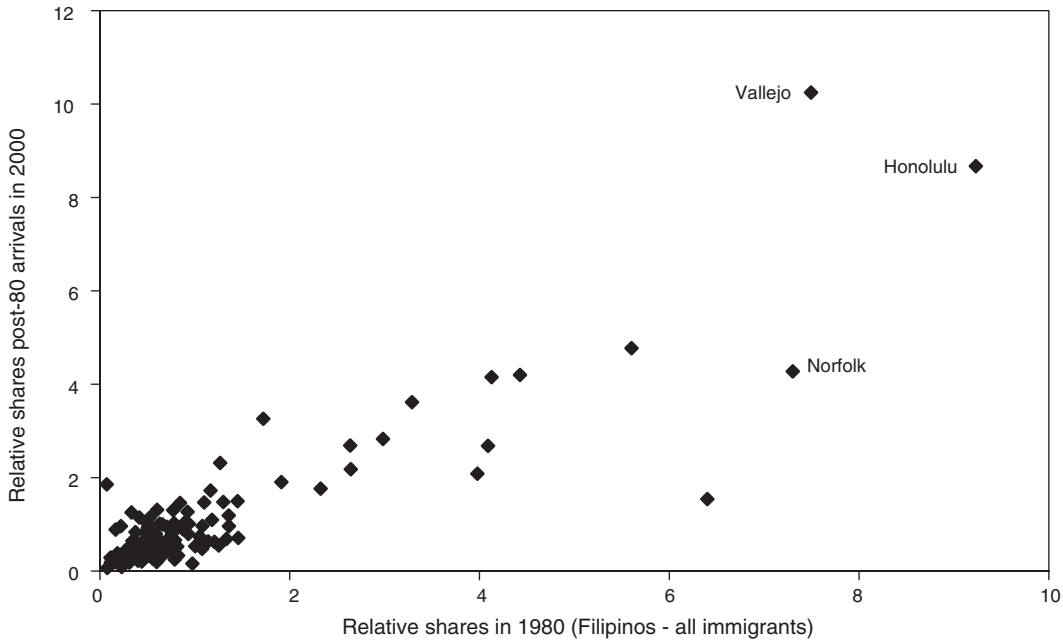

Figure 2. Relative Shares of Filipino Immigrants in Major Cities

US cities. The x-axis of this figure represents the ratio between the fraction of Filipino immigrants who lived in a specific city in 1980 and the average fraction of all immigrants who were living in this city in 1980. The y-axis represents the same relative share, taken over the set of immigrants observed in the 2000 census who arrived after 1980. The graph suggests that cities where Filipino immigrants were clustered in 1980 (relative to the locations of all immigrants) tended to attract an excess share of Filipino immigrants in the following decades. Similar, though on average somewhat weaker, patterns hold across nearly all source countries.[14]

As a result of the tendency for new immigrants to move to the same cities as earlier immigrants from the same country, the number and relative skill distribution of immigrants arriving in a city over a given interval of time is fairly predictable. If M_m immigrants arrive from country m to the United States as a whole (say between 1990 and 2000), and the fraction of earlier immigrants from country m who lived in city j at some previous date (say 1980) is λ_{mj}, then a naive clustering model predicts that $\lambda_{mj} M_m$ immigrants from country m will move to city j. Letting N_m denote the earlier population of immigrants from country m in the United States as a whole, and N_{mj} denote the number living in city j, the predicted inflow rate, as a fraction of the city's population (P_j) at some reference date, is

$$(6) \quad \left(\sum_m \lambda_{mj} M_m\right) / P_j = \sum_m [N_{mj}/P_j]\, M_m/N_m ,$$

which is a weighted average of the national inflow rates from each source country, with weights that depend on the shares of the country's earlier immigrants in city j.[15] If the national

[14] The R^2 for the scatter of points in Figure 2 is 0.77. The mean R^2 across 37 source countries/country groups is 0.36. Other source countries with high enclave tendencies include Mexico, Cuba, and Poland. Immigrants from India, China, and Taiwan show relatively low enclave tendencies.

[15] If N_{mc} and P_j are measured at the same time point, then N_{mj}/P_j is just the fraction of immigrants from source country m in city j at that point. If the two variables are measured at different time points, then N_{mj}/P_j is the fraction of immigrants from country m in city j at the date of measurement for N_{mj}, multiplied by the overall growth rate in the population of city j from that date to the date of measurement of P_j.

inflow rates from each source country are exogenous to conditions in a specific city, then the predicted inflow based on (6) will be exogenous. Assuming that a fraction δ_{ms} of immigrants from country m are in skill group s, the predicted inflow rate of new immigrants in skill group s is

$$(7) \quad \left(\sum_m \lambda_{mj} M_m \delta_{ms}\right) / P_j = \sum_m [N_{mj}/P_j]\, \delta_{ms} M_m / N_m,$$

which again can be interpreted as a weighted average of the skill group–specific inflow rates from each source country. Finally, note that a predicted *relative* inflow rate (for example, of college-equivalent versus high school–equivalent immigrants) can be calculated by taking the ratio of two expressions like the one in equation (7).

In the analysis in the next section, I use predicted relative inflows calculated in this way as instrumental variables for the relative number of dropout workers in a city and the relative number of college-equivalent workers in a city. Given the very large unskilled inflows from Mexico over the 1980s and 1990s, the calculated values of the instruments are significantly correlated with the fraction of Mexican immigrants in a city in 1980, leading to a concern that the instruments may be picking up other features of a city that are also correlated with the initial fraction of Mexicans living there (even if the surge in national inflows of Mexicans is exogenous to conditions in the city). To address this concern, I have refit many of the models excluding Mexican immigrants from the calculation of the instrumental variables. This does not change the general pattern of the results in any of the tables reported below, though the point estimates and precision are sometimes affected.

III. Cross-City Evidence on the Impacts of Immigration on Native Wage Differentials

A. *Data Overview*

I use 1980–2000 census data along with data from the combined 2005 and 2006 American Community Surveys (ACS) to construct a panel data-set of city-level labor market outcomes in 1980, 1990, 2000, and 2005/2006. As noted in the discussion of Table 1, I define "cities" as MSAs or PMSAs as of the 2000 census. I have used information on the changing definitions of MSAs and PMSAs to attempt to match the 2000 boundaries as closely as possible in 1990 and 1980. (The ACS public use files use the same geographic coding as the 2000 census). Some boundary changes cannot be incorporated because the smallest geographic units in the public use census files (the so-called Public Use Micro Areas) are rather large (100,000 people or more) and do not always correspond to the geographic units that define MSAs and PMSAs (counties in most of the country and towns in New England). Such problems are likely to be more serious for relatively small cities. For this reason, and because immigrants live primarily in larger cities, I focus on the 124 largest MSAs or PMSAs in the country as of 2000.

Table 3 presents some summary statistics on the characteristics of native and immigrant men and women in the four sample years. The education, experience, and employment data in the table are constructed from samples that include persons over the age of 18 with between 1 and 45 years of potential experience. The wage outcomes are constructed by dividing annual wage and salary earnings by the product of weeks worked and usual hours per week. To eliminate the influence of outliers, I have "Winsorized" hourly wages in each sample year at a lower value of 0.75 times the federal minimum wage, and an upper value of 50 times the minimum wage. In calculating wages I also exclude people who have positive self-employment income, since there is no information on how they divide their hours of work between self-employment and non-self-employment jobs.

The entries in the first column of Table 3 suggest that both natives and immigrants have become better educated over the past 25 years, with a slightly bigger gain for natives. Natives have also become older (a gain in over two years of potential experience since 1980), while the immigrant population's experience levels have remained constant, reflecting the rapid inflow rate of new, relatively young immigrants over the period. Employment rates of native men have drifted down (see Autor and Mark Duggan 2003), while those of immigrant men have remained constant, and rates for women have

Table 3—Summary Statistics for Samples from 1980, 1990, 2000 Census and 2005/2006 ACS

		Mean years of				Variance (log wage)	
		Education	Experience	Employment rate (%)	Mean wage	Overall	Residual
Native men	1980	12.5	18.8	90.1	25.07	0.385	0.288
	1990	12.9	18.7	88.5	23.72	0.462	0.322
	2000	13.2	20.4	86.8	25.86	0.487	0.353
	2005/6	13.4	21.4	86.2	25.35	0.522	0.361
Native women	1980	12.2	19.6	65.4	16.75	0.317	0.269
	1990	12.8	19.4	74.7	17.05	0.382	0.295
	2000	13.3	20.7	77.1	19.51	0.408	0.313
	2005/6	13.5	21.8	76.8	19.74	0.456	0.335
Immigrant men	1980	11.6	19.1	87.5	24.49	0.444	0.321
	1990	11.4	18.0	86.5	21.73	0.517	0.347
	2000	11.6	18.8	86.5	23.21	0.557	0.390
	2005/6	12.0	19.9	90.6	21.45	0.544	0.352
Immigrant women	1980	11.0	20.6	60.0	17.15	0.343	0.291
	1990	11.2	19.9	65.0	16.94	0.414	0.318
	2000	11.7	20.0	64.8	19.27	0.484	0.367
	2005/6	12.2	20.9	67.2	18.58	0.515	0.356

Notes: Samples include persons age 18 or older with 1–45 years of potential experience. Wages are reported in 2007 dollars. Residual wage variance is based on linear prediction models, fit separately by year, gender, and immigrant status.

risen. Mean real wages of native men have been roughly constant (using the CPI as a price index) while those of immigrant men have fallen. Real wages of both native and immigrant women have risen.

The last two columns of Table 3 show two simple measures of wage inequality: the variance of log hourly wages and variance of residual wage inequality.[16] By these measures, wage inequality has risen for all groups, with a particularly large jump between 1980 and 1990. Hourly wages are measured with some error in the census, so the precise levels of inequality reported in Table 3 may be too high (see Lemieux 2006 for a careful discussion of measurement problems in the context of wage inequality). The trends after 2000 may also be affected by the switch between the census questionnaire, which asks about earnings in the previous calendar year, and the ACS questionnaire, which asks about earnings in the previous 12 months. Presumably these measurement problems affect all four groups, so comparisons in relative trends between immigrants and natives are still informative.

[16] Wage residuals are obtained from a series of linear regression models fit separately by gender, immigrant status, and year. The models for natives include a flexible combination of age, education, and ethnicity variables (a total of 55 covariates). The models for immigrants include these controls as well as dummies for each of 38 countries of origin (or country groups), interactions of the origin dummies with a measure of years in the United States, and interactions of education variables with broad country group dummies (a total of 168 covariates).

B. *Analysis of the Dropout/High School Graduate Wage Gap*

Perhaps the single most important issue for understanding how immigrant inflows have affected native wage structures is the degree of substitutability between high school graduates and dropouts. Cross-city comparisons are potentially useful here because the relative share of dropouts varies widely across cities (see Figure 1). If there is an important degree of imperfect substitution, the relative wage of dropouts should be lower in high-dropout cities. To address the question, I estimate cross-city models of the form

$$(8)\quad r_{Djt} - r_{Hjt} = a + b\,\mathbf{X}_{jt} + c \log\left[S_{Djt}/S_{Hjt}\right] + e_{jt},$$

where r_{Djt} represents the mean residual log wage among *native male* dropout workers in city j and year t, r_{Hjt} represents the mean residual log

wage among *native male* high school graduates, $\mathbf{X}_{jt}$ is a vector of city-level control variables, S_{Djt} and S_{Hjt} represent total hours of work by all dropouts and high school graduates in city j and year t, and e_{jt} is an error term. As in equation (5), the coefficient c is interpretable as the negative of the inverse elasticity of substitution between dropouts and high school graduates, and is expected to be negative or zero in the limiting case of perfect substitutes. Note that wages are measured for native men, but supplies are measured over all workers present in city j. Note, too, that I use residual wages from a model fit to the entire US workforce. Thus, r_{Djt} is interpretable as the mean wage differential for native male dropouts who work in city j, relative to the national labor market, after adjusting for observed characteristics.[17]

A concern with the specification of equation (6) is that there may be unobserved city-specific factors that shift the relative demand for dropout workers, leading to higher relative wages and higher relative employment shares, and confounding the estimation of the inverse substitution elasticity. I use as instrumental variable for the relative supply ratio the log of the ratio of the predicted inflows of dropout and high school immigrants to city j over the previous ten years. Specifically, I calculate expressions like (7) for each city in 2000 using national inflows of immigrants from each of 38 source countries/country groups over the period from 1990 to 2000, and the shares of each group observed in each city in 1980. I use estimates for the skill shares (the δ_{ms} terms in equation (7)) derived from the skill shares of the national pool of immigrants from each country who arrived between 1990 and 2000. I then take the log of the predicted relative inflows for dropouts and high school graduates. This instrument has a relatively strong correlation with the actual ratio of dropout to high school labor supply in 2000, with t-statistics in the first-stage equation on the order of 14 (or 10 if Mexicans are excluded from the construction of the instrument).

Estimation results for various versions of equation (8) are presented in Table 4. An issue for the specification is whether one should allow city fixed effects that capture any permanent city-specific factors that account for differences in the mean (adjusted) wage gap between dropouts and high school graduates. To probe this issue, I include specifications with a lagged dependent variable (estimated for the city in 1990).[18] If there are permanent differences across cities, the coefficient of the lagged dependent variable will be close to one. As control variables, I include in some specifications the log of the city size in 1980 and 1990, the share of college workers in the city in 1980 and 1990, the share of workers employed in manufacturing in 1980 and 1990, and the mean wage residual for all native workers in the city in 1980 and 1990. The last are meant to pick up any potential correlation between the high school/dropout relative wage, and the average level of wages in a city.

All the specifications point to the same conclusion, which is that elasticity of substitution between dropouts and high school graduates is effectively infinite. The conclusion is similar whether controls are added to the model or not, and whether the model is estimated by OLS or IV.[19] Although not shown in the table, IV results based on predicted inflows that exclude Mexicans are also very similar (and similarly precise). Figure 3 illustrates the configuration of the data underlying the IV results. This figure plots the residual wage gap between dropouts and high school graduates against the instrumental variable. Clearly there is wide variation across cities in the predicted relative inflow of dropouts to high school graduates. But this variation is uncorrelated with the residual wage gap for native dropouts.

[17] The average log wage differential between native male dropouts and high school graduates in the census samples was -17.5 percent in 1980, -22.9 percent in 1990, -25.4 percent in 2000, and -26.6 percent in 2005/2006. Much of the change over time can be attributed to changing composition of the two groups: as discussed in Card (2005), the gap adjusted for differences in experience and ethnicity is slightly declining over the 1980–2000 period.

[18] I fit some of these models using the lagged wage for native women as an instrument for the lagged dependent variable, and obtained coefficient estimates very similar to the ones reported here.

[19] The covariates are all individually insignificant except the mean residual wage in 1990, which has a negative coefficient in all models (typical estimate = -0.15, typical standard error = 0.05). The estimates imply that the high school–dropout wage gap is higher in high-wage cities. This may reflect the upward pressure of the minimum wage in low-wage cities.

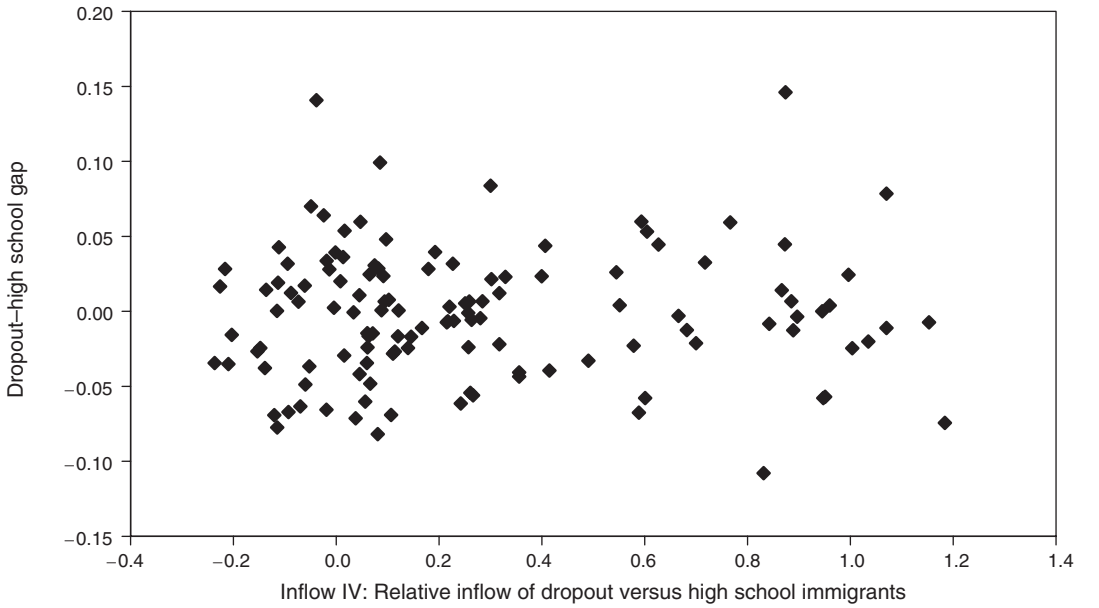

Figure 3. Reduced-Form Relationship: Inflow IV and Dropout Wage Gap

The estimates in Table 4 strongly support the conclusion from recent studies based on aggregate time series data. Both sources of evidence suggest that "dropout labor" is not an independent factor of production. Rather, dropouts are a subset of high school equivalent labor. I have fit many different versions of the models in Table 4, including specifications using native female wages, and pooled male and female wages, and other models that exclude the largest three or ten cities in the country. All of these models yield estimates of the inverse substitution elasticity that are close to zero.[20]

C. *Analysis of the College/High School Wage Gap*

Based on the preceding analysis, I believe it makes sense to use a simple two-skills model, with college-equivalent and high school–equivalent labor types, to study wage differentials at the city level. Following the literature, I fit models of the form

$$(9) \quad r_{Cjt} - r_{Hjt} = a + b\,\mathbf{X}_{jt} + c \log\left[S_{Cjt} / S_{Hjt}\right] + e_{jt},$$

where r_{Cjt} represents the mean residual wage of *native male* workers with exactly 16 years of education in city j and year t, r_{Hjt} represents the mean residual wage for *native male* high school graduates, and S_{Cjt} and S_{hjt} represent the total hours of college-equivalent and high school–equivalent labor employed in city j in year t. In this equation, the coefficient c is interpretable as the negative of the inverse elasticity of substitution between college-equivalent and high school–equivalent labor. I define high school–equivalent hours as the sum of hours worked by high school graduates, plus 0.7 times the hours worked by dropouts (assuming that dropouts have 0.7 efficiency units of high school graduates) plus 1.2 times the hours worked by one-half the people with 1–3 years of postsecondary schooling (assuming that half the people with some college are high school equivalents, and

[20] Very similar findings are reported by Card and Lewis (2007) who look at the high school graduate–dropout wage premium for a slightly larger set of cities in 2000 and between 1990 and 2000. The models presented in Card and Lewis (2007) control for overall employment growth between 1990 and 2000. They find that the high school premium is uncorrelated with employment growth rates in both OLS and IV specifications that use employment levels from 1982–1990 to instrument the 1990–2000 growth rate.

Table 4—Estimated Models for the Relative Wage Gap between Native Male Dropouts and High School Graduates

	Estimated by OLS				Estimated by IV		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Log relative supply of dropout vs. high school	—	0.00 (0.01)	0.01 (0.01)	0.01 (0.01)	0.02 (0.01)	0.02 (0.01)	0.02 (0.01)
Lagged dependent variable	0.29 (0.09)	—	—	0.25 (0.08)	—	—	
Controls for log city size, college share, manufacturing share, and mean wage residuals for all workers in 1980 and 1990	No	No	Yes	Yes	No	Yes	Yes
R^2	0.09	0.01	0.24	0.29	0.01	0.24	0.29
First-stage t-statistic	—	—	—	—	14.82	14.03	14.16

Notes: Standard errors in parentheses. All models are estimated on cross section of 124 larger cities in 2000. Estimates are weighted OLS or IV, using the 1990 population of the city as weight. Dependent variable is the difference between the mean adjusted log wage of high school dropouts, and the mean adjusted wage of high school graduates. Log relative supply measure is based on annual hours of all dropouts and all high school graduates (men and women, natives and immigrants). Instrumental variable for models in columns 5–7 is the log of the ratio of predicted inflows of dropout immigrants and high school graduate immigrants over the 1990–2000 period, based on national inflows of 38 source country groups and shares of each group in a city in 1980.

each has 1.2 efficiency units of high school labor). I define college-equivalent hours as the sum of hours worked by college graduates plus 0.8 times the hours worked by one-half the people with 1–3 years of postsecondary schooling (assuming that half the people with some college are college equivalents, and each has 0.8 efficiency units of college labor).

Table 5 presents estimation results for a variety of alternative specifications. One immediate and important difference between the results in this table and those in Table 4 is that city-specific values of the college–high school wage gap are highly correlated over time, with rather complex dynamics. As shown in column 1 of Table 5, when two lagged values (i.e., the wage gaps in 1990 and 1980) are included, the sum of their coefficients is very close to one, suggesting that an appropriate model may be a first-order autoregression in differences:

$$\Delta\left(r_{Cjt} - r_{Hjt}\right) \approx \beta\,\Delta\left(r_{Cjt-1} - r_{Hjt-1}\right) + \text{other terms},$$

where β is a number like -0.3. Specifications that impose a first-differences structure on both the lagged dependent variables and the relative supply terms are reported in columns 4 and 8.

When dynamics are ignored, and no other covariates are included, the results in column 2 of Table 5 suggest that the college–high school wage gap in a city is positively correlated with the relative supply of college workers. Adding two lags of the dependent variable, lagged relative supply, and controls for city size and the employment share in manufacturing in 1980 and 1990 (column 3) pushes the coefficient on current relative supply into the negative range, but the point estimate is small in absolute value. Imposing a first differences structure on the lagged dependent variables and the relative supply variable (column 4) leads to estimates that are very close to those from the corresponding unrestricted specification in column 3.

One obvious interpretation of these OLS estimates is that the relative supply of college workers is responsive to relative wages, leading to a positive bias in the estimated inverse elasticity of substitution. This conclusion is consistent with existing work on relative migration flows of different education groups (e.g., Gordon Dahl 2002), which suggests that mobility rates of college workers are sensitive to group-specific wage levels.

Table 5—Estimated Models for the Relative Wage Gap between Native Male College and High School Graduates

	Estimated by OLS				Estimated by IV			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Log relative supply of college vs. high school	—	0.08 (0.02)	−0.04 (0.05)	−0.01 (0.05)	−0.42 (0.28)	−0.41 (0.15)	−0.26 (0.12)	-0.28 (0.12)
Log relative supply lagged 10 years	—	—	0.12 (0.08)	0.01 (—)	—	0.60 (0.21)	0.34 (0.17)	0.28 (—)
Log relative supply lagged 20 years	—	—	0.01 (0.04)	—	—	-0.10 (0.07)	0.01 (0.07)	—
Relative wage gap lagged 10 years	0.66 (0.08)	—	0.66 (0.07)	0.68 (0.06)	—	0.46 (0.12)	0.51 (0.10)	0.56 (0.09)
Relative wage gap lagged 20 years	0.29 (0.06)	—	0.26 (0.06)	0.32 (—)	—	0.45 (0.10)	0.35 (0.09)	0.44 (—)
Controls for log city size and mfg. share in 1980 and 1990	No	No	No	Yes	No	No	Yes	No
R^2	0.59	0.09	0.68	0.37	0.02	0.60	0.71	0.24
First-stage t-statistic	—	—	—	—	1.98	4.73	4.87	4.66

Notes: Standard errors in parentheses. All models are estimated on cross section of 124 larger cities in 2000. Estimates are weighted OLS or IV, using the 1990 population of the city as weight. Dependent variable is the difference between the mean adjusted log wage of college graduates and the mean adjusted wage of high school graduates. Log relative supply measure is based on annual hours of all college equivalent and high school quivalent workers (men and women, natives and immigrants). Models in columns 4 and 8 are fit in first difference form. Instrumental variable for models in columns 5–8 is the log of the ratio of predicted inflows of college-equivalent and high school–equivalent immigrants over the 1990–2000 period, based on national inflows of 38 source countries and shares of each group in a city in 1980.

Columns 5–8 of Table 5 present instrumental variables estimation results, using predicted relative inflows of college-equivalent and high school–equivalent immigrants from 1990 to 2000 as an instrument for the relative supply of college-educated workers. This instrument is weaker than the instrument for the relative supply of dropouts used in Table 5, but when lagged supply is added to the specification (columns 6–8) the t-statistic on the instrumental variable in the first stage equation is over 4.5 (see the entries in the bottom row of Table 5). The IV estimates of the inverse elasticity of substitution between college and high school workers range from 0.26 to 0.41—not far below the estimates obtained in many recent aggregate time series studies, and consistent with a value for σ_e in the range of 1.5–2.5.

My interpretation of the results in Table 5 is that relative demand for college versus high school workers at the city level exhibits about the same elasticity with respect to relative wages as relative demand at the national level. However, relative supply is endogenous at the local level, confounding simple observational comparisons across cities (such as the univariate regression model in column 2). This does not mean that cross-city comparisons are uninformative. Rather, it means that researchers have to address the endogeneity problem to obtain interpretable estimates. For this purpose, the enclaving tendency of immigrants, coupled with differences in the education distributions of immigrants from different origin countries, is particularly helpful, although other identification strategies are certainly worth pursuing.[21]

[21] Enrico Moretti (2004) presents an analysis of the effect of the college share on wages of different education groups in a city, using the age structure of a city in 1970 as a predictor of the change in the college share in a city from 1980 to 1990. This is a reasonably powerful instrument. His estimates (Moretti 2004, table 5) imply that a 10 percentage point increase in the college share reduces the college–high school wage premium by about 0.12 log points. This implies an elasticity with respect to the log of the relative supply of

D. *Analysis of the Immigrant/Native Wage Gap*

A third key issue for understanding the impact of immigration is whether immigrants and natives in the same broad skill class are perfect or imperfect substitutes. Here, cross-city comparisons are potentially very useful because there is enormous variation in the relative fraction of immigrants across cities. Assuming there are only two skill groups—high school equivalents and college equivalents—I estimate models separately by skill group of the form

$$(10)\ r_{Mjt} - r_{Njt} = a + b\,\mathbf{X}_{jt} + c \log\,[S_{Mjt}/S_{Njt}] + e_{jt},$$

where r_{Mjt} represents the mean wage residual for *immigrant men* in a particular skill group in city j and year t, r_{Njt} represents the corresponding mean wage residual for *native men*, and S_{Mjt} and S_{Njt} represent the total hours of work by all immigrants and all natives in the skill group (i.e., including men and women). The coefficient c in this equation is interpretable as an estimate of the inverse elasticity of substitution between immigrants and natives in the particular skill group (i.e., $c = -1/\sigma_I$). Again, there is an obvious concern that unobserved factors in a city may lead to both higher wages and higher employment levels of immigrants relative to natives, confounding the estimation of the inverse substitution elasticity. Following the strategy in Tables 4 and 5, I use predicted inflows of immigrants in the particular skill group (high school or college) to city j over the 1990–2000 period (based on immigrant shares in 1980) to instrument for the share of immigrants in 2000.

It is worth noting a potential difference between the IV strategy for the estimation of equation (10) and the corresponding strategies for equations (8) and (9). In the latter cases, the instrumental variable is the *ratio* of predicted immigrant inflows in two skill groups (e.g., high school equivalents versus college equivalents). These ratios are not as strongly correlated with the initial immigrant share in 1980 as the predicted immigrant inflows of college- or high school–equivalent workers used as instruments for the estimation of (10).[22] To the extent that initial immigrant shares are correlated with other unobserved features that affect relative wage differentials in a city, an enclave-based identification strategy may be less attractive for estimating (10) than for estimating (8) or (9).

Estimation results for equation (10) are shown in Table 6. All the specifications include measures of city size, college share, and manufacturing share in 1980 and 1990, as well as the mean wage residuals for all immigrants and all natives living in the city in 1980. The latter are meant to control for any attributes in a city that lead to persistently higher wages for immigrants and/or natives, although the coefficients of these variables are generally small and statistically insignificant. Wage gaps between immigrants and natives in a city do not exhibit the complex dynamics of the college–high school wage premium, and models with a lagged dependent variable have a relatively small autoregressive coefficient (around 0.15). The estimated relative supply effects are also very similar, whether the lagged dependent variable is included or not.

The estimated relative supply coefficients are generally bigger in absolute value when estimated by IV than by OLS, suggesting some endogeneity in relative supply, although both OLS and IV estimates are uniformly negative. The relative supply coefficients are also bigger (in absolute value) for college workers than high school workers. Taken literally, this means that less-educated immigrants and natives are closer to perfect substitutes ($\sigma_I \approx 40$) than are more-educated immigrants and natives ($\sigma_I \approx 17$).

Figure 4 illustrates the reduced-form relationship between the predicted inflow of college-educated immigrants to a city over the 1990–2000 period (normalized by the size of the total city population in 2000) and the residual wage gap between college-equivalent immigrants and college-equivalent natives in a city. The figure illustrates the strong negative relationship (weighted correlation = -0.60) between

college-equivalent workers of about -0.25, which is not too far off the estimates in Table 5.

[22] The correlation of the instrument for the ratio of dropouts to high school graduates with the share of immigrants in a city in 1980 is 0.43. The correlation of the instrument for the ratio of college- to high school–equivalent labor with the 1980 immigrant share is -0.55. By comparison, the correlation of the instruments for the fractions of immigrants in the high school or college workforces with the immigrant share in 1980 is 0.90.

Table 6—Estimated Models for the Relative Wage Gap between Immigrants and Natives within Skill Group

	High school equivalent workers				College equivalent workers			
	Estimated by OLS		Estimated by IV		Estimated by OLS		Estimated by IV	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Log relative supply of immigrants/natives	−0.019 (0.006)	−0.019 (0.006)	−0.023 (0.008)	−0.022 (0.008)	−0.036 (0.011)	−0.029 (0.011)	−0.060 (0.014)	−0.055 (0.015)
Lagged dependent variable	—	0.159 (0.060)	—	0.159 (0.060)	—	0.200 (0.100)	—	0.128 (0.106)
Controls for log city size, college share, and mfg. share in 1980 and 1990, and mean wage residuals for all natives and all immigrants in 1980	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.308	0.349	0.301	0.338	0.444	0.463	0.455	0.471
First-stage t-statistic	—	—	11.67	11.68	—	—	13.76	12.85

Notes: Standard errors in parentheses. All models are estimated on cross section of 124 larger cities in 2000. Estimates are weighted OLS or IV, using the 1990 population of the city as weight. Dependent variable is the difference between the mean adjusted log wage of male immigrants and natives who are classified as "high school equivalent" workers (columns 1–4) or "college equivalent" workers (columns 5–8). Log relative supply measure is based on annual hours of all high school equivalent or college equivalent workers (men and women). Instrumental variable for models in columns 3–4 is the predicted inflow of high school equivalent immigrants between 1990 and 2000, divided by city population in 2000. Instrumental variable for models in columns 7–8 is the predicted inflow of college equivalent immigrants between 1990 and 2000, divided by city population in 2000.

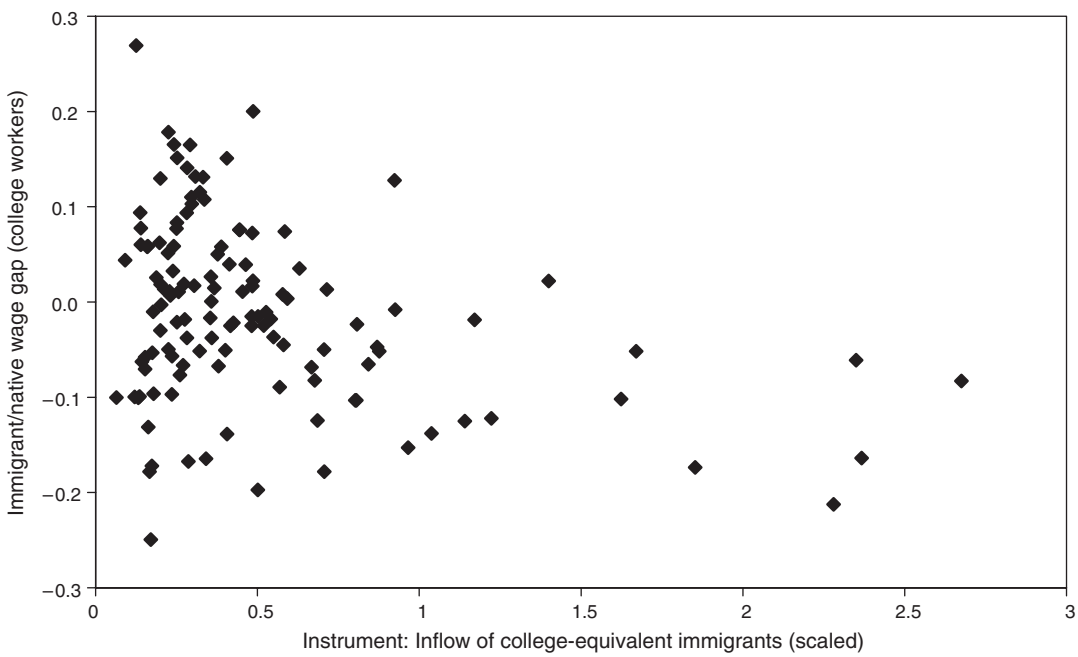

Figure 4. Reduced Form: Inflow IV and Immigrant Wage Gap (*College*)

the instrument and the immigrant/native wage gap that drives the IV results shown in Table 6.

Overall, the estimated inverse elasticities in Table 6 are consistent with the range of estimates for $1/\sigma_I$ obtained by Ottaviano and Peri (2008), which center around -0.04. Specifications similar to the ones in Table 6 fit for both men and women, and for women only, also confirm this conclusion, with estimates for the inverse substitution elasticity in the same range. IV estimates of the inverse elasticity using predicted inflows that exclude Mexican immigrants are slightly larger in absolute value (e.g., the estimate corresponding to the entry in column 4 is -0.031 with a standard error of 0.008), as are estimates from specifications that exclude the control variables.[23] Overall, I conclude that both the time series and cross-city evidence are consistent with a small but detectable degree of imperfect substitution between immigrants and natives.

E. *Within-Group Residual Inequality*

Existing studies of the impact of immigration on the wages of natives have focused on the effect on mean wage differentials between skill groups. While much of the rise in wage inequality over the past two decades has been driven by increases in between-group inequality—especially the gap between high and low education groups—within-group inequality has also risen substantially (Katz and Autor 1999; Lemieux 2006, 2008).

Across major cities the level of within-group or residual wage inequality is strongly correlated with immigrant densities. This is illustrated in Figure 5, which plots the residual variance of wages for college-equivalent and high school–equivalent male workers in each of the largest 124 cities in 2000 against the immigrant share.[24] A 10 percentage point increase in the immigrant share is associated with a 0.025 point rise in the residual variance of high school–equivalent men's wages (standard error = 0.002 points), and a 0.027 point rise in the residual variance of college-equivalent men's wages (standard error = 0.003 points).

To the best of my knowledge, there is no well-developed theory of how the presence of immigrants (or a more diverse workforce in general) affects residual wage inequality among natives. If one assumes that the causal effect of higher immigration depends on the fraction of immigrants in a worker's *own* skill group, however, then an appropriate empirical specification has the form

$$(11)\quad \log\,[V_{Cjt}/V_{Hjt}] = a + b\,\mathbf{X}_{jt} + c\,\log\,[Imm_{Cjt}/Imm_{Hjt}] + e_{jt},$$

where V_{Cjt} represents the residual variance of wages among college-educated workers in city j and year t, V_{Hjt} represents the corresponding variable for high school–educated workers, and Imm_{Cjt} and Imm_{Hjt} denote the fractions of immigrants in the college-equivalent and high school –equivalent workforces. Here the coefficient c is interpretable as the effect of a shift in the log immigrant share in a given skill group on the residual variance of wages for natives in the same skill group.

Table 7 presents estimation results for specifications based on equation (11). The table includes both OLS and IV models that use the ratio of the predicted number of college and high school immigrants moving to the city between 1990 and 2000 as an instrumental variable for the relative fraction of immigrants in the two labor pools.

Two main conclusions emerge from the table. First, the ratio of the residual variances among college and high school workers is fairly persistent over time, with a coefficient on the lagged dependent variable close to 0.5. Second, there is no evidence of an effect of relative immigrant densities on the relative residual variances of college and high school workers. The latter conclusion appears to be quite robust. For example, estimates from IV models based on predicted immigrant inflows excluding Mexicans, and from models for the residual wage variance among both men and women, are quite similar.

[23] The IV coefficient estimates from models without controls are -0.027 (standard error = 0.006) and -0.080 (standard error = 0.010) respectively.

[24] The same patterns holds using women's wages, or men's and women's wages pooled. For example, the correlation between immigrant share and the residual variance of college-equivalent women's wages is 0.62, versus 0.64 for men. The correlation between immigrant share and the residual variance of high school–equivalent women's ages is 0.75, versus 0.77 for men.

Table 7—Estimated Models for the Effect of Immigration on Relative Within-Group Residual Variance for Native Men

	Estimated by OLS				Estimated by IV		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Log relative fraction of immigrants (college/high school)	—	0.00 (0.03)	0.01 (0.03)	-0.03 (0.05)	0.07 (0.05)	0.03 (0.04)	0.06 (0.22)
Lagged log relative fraction of immigrants (1990)	—	—	—	0.04 (0.04)	—	—	-0.02 (0.15)
Lagged dependent variable	0.51 (0.06)	—	0.47 (0.06)	0.45 (0.07)	—	0.48 (0.07)	0.49 (0.11)
Controls for log city size, college share, and manufacturing share in 1980 and 1990	No	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.34	0.21	0.46	0.47	0.21	0.46	0.46
First-stage t-statistic	—	—	—	—	9.46	9.89	2.32

Notes: Standard errors in parentheses. All models are estimated on cross section of 124 larger cities in 2000. Estimates are weighted OLS or IV, using the 1990 population of the city as weight. Dependent variable is the log of the ratio of the residual wage variance for native men with college or more versus exactly 12 years of education. Instrumental variable for models in columns 5–7 is the log of the ratio of predicted inflows of college-equivalent and high school–equivalent immigrants over the 1990–2000 period, based on national inflows of 38 source countries and shares of each group in a city in 1980.

Tentatively, I conclude that the correlations exhibited in Figure 5 are not a *causal* effect of immigration, but rather a consequence of unobserved city-wide factors that are associated with higher immigration and a higher level of residual variance for both skill groups. It should be noted, however, that a "relative" specification such as (11) is not the only possible causal model. If a greater fraction of immigrants in one skill group exerts an equal effect on the residual variances of wages for workers in both groups, then a specification like (11) will fail to estimate the true effect of immigration.

IV. Summary and Discussion

Cross-city and time series comparisons of the effects of relative supplies of different skill groups on relative wages are consistent with three key hypotheses:

- Workers with below high school education are perfect substitutes for those with a high school education;
- High school–equivalent and college-equivalent workers are imperfect substitutes, with an elasticity of substitution on the order of 1.5–2.5;
- Within education groups, immigrants and natives are imperfect substitutes, with an elasticity of substitution on the order of 20.

These hypotheses imply that the structure of relative labor demand at the city or national level is consistent with a simple nested CES production function for labor inputs with only two skill groups (high school and college), each of which is itself an aggregate of immigrant and native labor inputs within the appropriate skill group.

These three assumptions plus an assumption of perfectly elastic capital supplies means that, at the national level, immigration over the past two or three decades has had only minor effects on the mean wage differences between natives in different skill groups (see Ottaviano and Peri 2008), and a negligible effect on between–skill group wage variability. The main explanation for this somewhat surprising conclusion is that in a two-group model, what matters for the structure of wages is the relative fractions of immigrants and natives who are high school–equivalent and college-equivalent workers. As it happens, immigrants are only slightly underrepresented in the college-equivalent group relative to natives (36 percent versus 41 percent). Compared to the distribution among natives

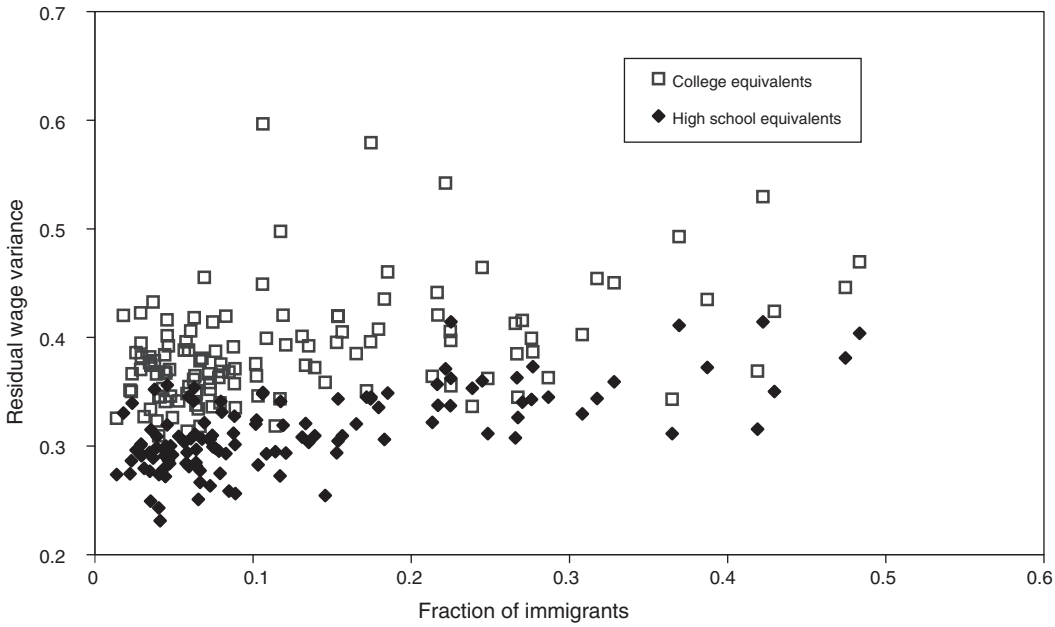

Figure 5. Immigrant Density and Residual Wage Inequality (*Native Men*)

alone, immigrant arrivals have hardly distorted the relative fraction of college-equivalent workers, and have therefore had little impact on the college–high school wage gap, relative to the counterfactual of no immigrant presence in the economy.

The overall impact of immigration on native inequality depends on the effects on between-group differentials and on the effects on within-group inequality. There is little existing theoretical or empirical research on the latter channel. Assuming that any within-group impacts caused by immigrants are concentrated among natives in the same skill group, however, the estimates presented here suggest that immigration has not caused a rise in residual wage inequality among natives.

Overall, my interpretation of the evidence is that immigration has not had much effect on *native* wage inequality in the United States. Nevertheless, because immigrants are clustered at the high and low ends of the education distribution, and because they also tend to have higher residual inequality than natives (see Table 3), wage inequality over all workers in the economy is higher than it would be in the absence of immigration. Table 8 illustrates this point, showing data on wage inequality in 1980 and 2005/2006 for all workers, and for immigrants and natives. In 1980, for example, the variance of log hourly wages across all male workers was 0.390, versus 0.385 among native men. Likewise, the variance among all female workers was 0.318, versus 0.317 among native women. Over the past 25 years, the gap between the variance of wages in the entire workforce and among natives has widened: thus, immigration can be said to have contributed to the rise of inequality in the workforce. The effect is relatively small, however. For men, native inequality rose by 0.137 while inequality among all workers rose by 0.142, a differential of about 4 percent. For women, native inequality rose by 0.139 while overall inequality rose by 0.148, a difference of about 6 percent. These comparisons suggest that the presence of immigration can account for a relatively small share (4–6 percent) of the rise in overall wage inequality over the past 25 years.

While recent research using time series and cross-city comparisons has made significant progress in clarifying the effects of immigration on labor market outcomes for natives, several important issues deserve further attention. First,

Table 8—Summary of Changes in Variance of Log Wages for All Workers and Natives Only

	Variance of log hourly wages			
	All workers	Natives	Immigrants	Percent immigrants
Male workers				
1980	0.390	0.385	0.444	6.9
2005/2006	0.532	0.522	0.544	18.0
Change	0.142	0.137	0.100	11.1
Female workers				
1980	0.318	0.317	0.343	6.7
2005/2006	0.466	0.456	0.515	13.9
Change	0.148	0.139	0.172	7.2

Note: Samples include persons age 18 or older with 1–45 years of potential experience who have positive wage and salary earnings and no self-employment earnings.

given the importance of the degree of substitutability between dropouts and high school graduates, more research on the types of jobs held by the two groups and further evidence on how excesses of dropout labor are absorbed in high-immigrant cities would be useful. Additional insights into the degree of substitution between similar immigrants and natives are also needed. Finally, more research is needed to better understand why a simple two-skill group CES model seems to work so well at both the national and city levels, and how the estimated elasticities of substitution from these two levels of aggregation are related to underlying parameters of firm-level production functions.

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