American Economic Review: Papers & Proceedings 2015, 105(5): 220–225
http://dx.doi.org/10.1257/aer.p20151043

The Long-Run Effect of Mexican Immigration on Crime in US Cities: Evidence from Variation in Mexican Fertility Rates†

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Over the past 30 years, crime rates in cities across the United States initially increased and then declined precipitously, in many cases, reaching historic lows. At the same time, the share of the foreign born among the US population has increased rapidly, with the foreign-born Mexican share of the population quadrupling since 1980. The majority of the increase in immigration has taken place since 1990 and coincides with the largest decline in US crime rates since crime has been reliably measured.

Research suggests that immigration has either played no role in this historic decline in crime (Butcher and Piehl 1998; MacDonald, Hipp, and Gill 2013; Chalfin 2014) or has possibly contributed importantly to the decline (Ousey and Kubrin 2009; MacDonald and Saunders 2012). In particular, researchers have pointed to weak cross-sectional relationships between immigrant concentrations and crime at the neighborhood level and small and often negative correlations between changes in a city's immigrant share and changes in a city's crime rate over time. A recent exception to the entirety of the extant literature is that of Spenkuch (2014) who, in a careful analysis, studies the relationship between immigration and crime at the county level and concludes that there is a positive relationship between immigration, particularly Hispanic immigration, and instrumental crimes such as robbery and burglary.

Despite the recent surge of academic interest in this topic, the literature remains unsatisfying in several ways. First, the available literature rarely disaggregates the effects of immigration on crime by nationality. As Mexican immigrants comprise over one-third of all immigrants to the United States and over half of all undocumented immigrants (Passel and Cohn 2009), assessing the effects of Mexican immigration on crime is of particular relevance. Second, prior literature has examined only the effect of immigration on crimes reported to police. To the extent that immigrants are less likely to report crimes, an alternative explanation for a negative relationship between immigration and crime in the data is that immigration drives down crime reporting (Butcher and Piehl 1998). To address this issue, I provide an auxiliary analysis of the effect of immigration on the rate at which crimes are reported to police, using MSA (metropolitan statistical area)-level data from the National Crime Victimization Survey (NCVS). This analysis suggests that differences in crime reporting rates are unlikely to explain negative correlations between immigration and crime in the extant literature. Finally, regression-based estimates of the effect of immigration on crime can only be ascribed a causal interpretation under stringent assumptions regarding the inability or unwillingness of migrants to adjust the timing and destination of their arrival in the United States in response to social and economic conditions in US destinations. I describe a novel identification strategy that plausibly addresses this issue.

I. Identification Strategy

The primary strategy that has been used to isolate quasi-random variation in immigrant flows is to instrument for recent flows of country-specific immigration with immigrant flows that are predicted by the national flow of migrants to the United States and the location decisions of past migrants.[1] The approach

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† Go to http://dx.doi.org/10.1257/aer.p20151043 to visit the article page for additional materials and author disclosure statement.

[1] This instrumental variable was pioneered by Altonji and Card (1991) in their seminal treatment of the cross-city

relies on the empirical regularity that migrants tend to take advantage of preexisting social networks and sojourn to US destinations that host an existing community from their country of origin. Accordingly the instrumental variable arising from this framework has come to be known as the "network" instrument. Formally, the network instrument, Z_{it}^{n}, can be written as

$$(1) \qquad Z_{it}^{n} = \sum_{m=1}^{M} \Delta MIG_{mt} \times \lambda_{im},$$

where ΔMIG_{mt} is the change in the number of immigrants from source region m who are living in the United States in year t relative to some base year and λ_{im} is a matrix of weights that capture the strength of migration networks between the M source regions and each US destination. Z_{it}^{n}, is the interaction of these two terms summed over the M source regions and is the predicted change in the number of migrants living in city i in year t.

To the extent that the lagged values of the stock of the foreign-born population do not directly affect contemporary crime rates, the network instrument is valid. However, if migrants are drawn to a particular city due to the existence of certain characteristics related to crime (e.g., the city's industrial mix) in 1960, to the extent that these characteristics persist, today's migrants may be pulled to a city for similar reasons thus invalidating the instrument. Another way to see this is to consider that while the migration weights in (1), λ_{im}, are not time-varying, the change in the stock of immigrants in each US city, ΔMIG_{mt}, will be a function of both events that are unfolding in source regions and contemporaneous conditions in network-linked US destinations.

In this section, I propose a decomposition of Z_{it}^{n} that can be used to isolate "pull" from "push" variation using long-differenced data.[2] The idea is that ΔMIG_{mt} can be apportioned into a part that is explained by past fertility rates and a part that is attributable to the time-varying conditional probability of immigration. Let N_{mt} be the available supply of Mexican citizens living in Mexican state m who are eligible to migrate to the United States in year t and let p_t be the annual conditional probability of migration among the originally eligible migrants. N_{mt} is a function of the number of lagged Mexican state-specific births (where the length of the lag will correspond with the ages of likely migrants) and the number of deaths in each cohort among the N_{mt} individuals. The annual number of Mexicans from each state who *actually* migrate to the United States is given by

$$(2) \qquad \Delta MIG_{mt} = N_{mt} \times p_t.$$

While N_{mt} is a function of fertility and mortality conditions in Mexico many years ago, p_t is a function of contemporary conditions in Mexico as well as traditional destinations in the United States. For example, p_t might rise due to a currency crisis in Mexico or due to favorable employment conditions in Los Angeles or Chicago.[3] Thus p_t creates a potential problem for the network instrument. Recognizing this, there is promise in removing p_t from the equation, instead focusing only on the size of available migration cohorts. Let

$$(3) \qquad N_{mt} = \sum_{t_{\min}}^{t_{\max}} B_{mt},$$

where B_{mt} is the number of births in a Mexican state-year. Further let $t_{\min}$ and $t_{\max}$ be the minimum and maximum years of birth of eligible migrants, where we assume that migrants will be between the ages of 17 and 52 upon leaving Mexico.[4] Summing the number of births over

effect of immigration on the wages and employment of natives.

[2] See Chalfin and Levy (2013) for further details.

[3] Likewise p_t might change if migration is sensitive to the size or conditions of illicit labor markets or to the scope of local immigration enforcement.

[4] Though this window reflects the age-range in which migration is most common, its precise upper bound is chosen for reasons of data availability. Natality data by state exist dating back to 1928 in Mexican government almanacs, defining the upper bound of the age window for the 1980 sample at 52 years old. For consistency, this upper bound is retained for the other years. This is not a serious concern because birth cohorts earlier than 1928 would have passed prime migration age by the late 1960s, prior to which most migration was seasonal and thus did not contribute to large-scale growth in the Mexican migrant population share. I have run all analyses shifting the window's lower and upper bounds one and two years earlier, shifting the lower bound one and two years later, and shifting only the upper bound one and two years earlier. None of these changes had a material effect on any of the results presented.

all eligible birth cohorts yields an estimate of the number of migrants eligible to migrate from each state in each year.

Substituting (3) into (2) and (2) into (1) and rearranging yields an alternative specification of the network instrument:

$$Z_{it}^{n} = \sum_{m=1}^{M} \left[\lambda_{im} \times p_t \sum_{t_{\min}}^{t_{\max}} B_{mt} \right]. \tag{4}$$

Finally, recalling that p_t is implausibly exogenous, I rewrite (4) excluding p_t:

$$Z_{it}^{b} = \sum_{m=1}^{M} \left[\lambda_{im} \sum_{t_{\min}}^{t_{\max}} B_{mt} \right]. \tag{5}$$

This is the "births" instrument. In practice, I divide the quantity in (5) by the 1980 estimate of each MSA's population in order to meaningfully scale the variable. Hence Z_{it}^{b} is simply the number of Mexican births that are predicted to end up in each US MSA in a given year under the assumption that the entire cohort migrates, deflated by MSA population. The instrument is used to predict changes in an MSA's Mexican immigrant share.

II. Data

Using data from three US censuses (1980, 1990, and 2000), I begin with a sample consisting of 92 MSAs with a sufficient presence of Mexican immigrants to allow reliable estimation. Data on state-specific births are drawn from tabulations of registered births and male-to-female birth ratios in Mexican states included in statistical almanacs produced by the Mexican government's *Instituto Nacional de Estadistica, Geografia e Informatica* (INEGI). Data on each MSA's population, its Mexican immigrant share, and relevant control variables are derived from 5 percent samples of the US census accessed using IPUMS.[5] Data on index crimes reported to police come from the Federal Bureau of Investigation's Uniform Crime Reporting (UCR) program and were aggregated from the agency to the MSA level in order to accord with the available migration data.[6] Finally, to construct a set of weights that capture migration patterns linking Mexican sending states to US destinations, I use migrant-level survey data from the Mexican Migration Project (MMP). The MMP surveys Mexican households in known sending regions and includes data on when migrants embarked on their first journeys to the United States, each migrant's state of birth, and the US MSA in which that migrant subsequently settled. In order to construct weights that are predetermined with respect to the study period, exclude data on journeys undertaken after 1979.

III. Results

I begin with a discussion of the first stage relationship between growth in an MSA's foreign-born Mexican population share and the change in the immigrant share that is predicted on the basis of the size of lagged Mexican birth cohorts (the births instrument). Table 1 presents regression estimates of this relationship. Columns 1 through 5 report a coefficient and standard error arising from a least squares regression of the change in the Mexican immigrant share on the births instrument, conditional on a variety of control variables.[7] Because all variables in the model are differenced, the model uses only within-city variation.

Column 1 reports estimates that condition only on a decade fixed effect. In column 2, I add time-varying covariates and, in column 3, the decade fixed effect is replaced by interacted region × decade fixed effects. These interacted fixed effects are powerful and account for, among other things, the emergence of new immigrant destinations, at least at the level of the region. Finally, columns 4 and 5 report the

[5] I focus on the share of Mexican immigrant adults among each MSA's population non-institutionalized population. Relevant control variables include the educational attainment of each MSA's population (< HS, High School, Some College, and College+), the age composition of the population (0–14, 15–24, 25–39, 40–54, and 55+), gender composition, race and ethnicity (black, US-born Hispanic, and non-Mexican immigrants), and the employment-to-population ratio. In addition, I control for changes in the lagged number of US births to Mexican-born parents in each MSA.

[6] Crime data were available for the seven standard index crimes—murder, forcible rape, robbery, aggravated assault, burglary, larceny, and motor vehicle theft. In an auxiliary analysis, MSA-level data were obtained by the National Crime Victimization Survey (NCVS) which surveys potential crime victims to learn about victimizations that were reported to police as well as those which remain unreported.

[7] Standard errors are clustered at the MSA level and are robust to within-MSA dependence.

Table 1—First Stage Models: Proportion of Migration-Eligible Births in Network-Linked US Metropolitan Statistical Areas

	Foreign-born Mexican adults					Other subgroups		
						Other foreign born	US-born Hispanics	Mexican-born children
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Births instrument	0.168***	0.080***	0.064***	0.062***	0.053***	−0.033	0.036	0.000
	(0.018)	(0.021)	(0.018)	(0.018)	(0.018)	(0.029)	(0.022)	(0.005)
F-statistic	90.3***	14.6***	12.6***	11.6***	8.7***	1.3	2.7	0.0
Year effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Covariates	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Region × year effects	No	No	Yes	Yes	Yes	Yes	Yes	Yes
Los Angeles excluded	No	No	No	Yes	No	No	No	No
Chicago excluded	No	No	No	No	Yes	No	No	No

Notes: Columns 1 through 5 report estimates from a least squares regression of the decadal change in an MSA's foreign-born Mexican population share on the decadal change in the predicted foreign-born Mexican share using the number of eligible lagged births in network-linked Mexican states. Columns 1 and 2 condition only on a year fixed effect. Column 2 adds a vector of covariates while column 3 adds interaction region × year fixed effects. Columns 4 and 5 remove from the panel, Los Angeles and Chicago, respectively. Columns 6 through 8 test whether lagged Mexican births predict the decadal change in foreign-born from countries other than Mexico (column 6), US-born Hispanics (column 7), and Mexican-born children who are too young to have reached prime migration age. All models are weighted accordingly using 1980 MSA population and cluster-robust standard errors are reported in parentheses below the coefficient estimates. The *F*-statistic is the square of the *t*-statistic on the birth cohort instrument and is a sufficient statistic to assess the strength of the first stage relationship between Mexican immigration and the birth cohort instrument. The partial R^2 is the percentage of the variation in MSA-specific migration that is explained by the births instrument, conditional on the covariates in the model. The sample size is 184 city-years.

***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

first stage coefficients having excluded each of Los Angeles and Chicago from the sample. All regressions are weighted using each MSA's 1980 population.

Prior to assessing the strength of the first stage relationship, it is worth noting that the coefficients have a useful interpretation. Since both the instrument and the endogenous regressor are scaled by the MSA's population, the coefficient can be interpreted as the estimated probability that a migration-eligible birth can be found in a network-linked US destination. Referring to column 3, between 1980–2000, approximately 6 percent of individuals born in Mexico in the eligible window could ultimately be found in a network-linked US destination. In all of the specifications, the *F*-statistic on the excluded instrument meets standard criteria for instrument relevance which is impressive given the size of the sample ($N = 184$ MSA-years) and that models 3–5 condition on region × decade fixed effects. Columns 6–8 provide a series of falsification tests that bolster the validity of the identification strategy. In particular, I show that while the births instrument is a strong predictor of Mexican migration to network-linked US destinations, it does not predict a change in migration from other source countries, internal (within US) migration of US-born Hispanic citizens, or the migration of Mexican children. This is crucial to establish as an association between lagged Mexican births and any of these variables would call into question whether the exclusion restriction on the instrument is, in practice, met.

Table 2 presents substantive results with respect to per capita crime and is divided into two panels. Panel A presents least squares coefficients and standard errors from a regression of the natural log of the number of crimes per MSA resident on the Mexican immigrant share. Panel B reports the corresponding 2SLS estimates using the births instrument. Overall,

Table 2—Least Squares and 2SLS Estimates of the Effect of Mexican Immigration on Crimes Reported to Police

Violent crimes				Property crimes		
Murder (1)	Rape (2)	Robbery (3)	Aggravated assault (4)	Burglary (5)	Larceny (6)	Motor vehicle theft (7)
Panel A. Least squares estimates						
0.075***	0.002	0.095***	0.037	0.025	0.000	0.042
(0.028)	(0.029)	(0.025)	(0.029)	(0.019)	(0.023)	(0.028)
Panel B. 2SLS Estimates [births instrument]						
−0.022	−0.131**	0.049	0.197***	−0.092	−0.105**	−0.149*
(0.094)	(0.062)	(0.094)	(0.075)	(0.057)	(0.047)	(0.081)

Notes: In panel A, each column reports estimates from a least squares regression of the decadal change log crimes per capita on the decadal change in the foreign-born Mexican population share. Panel B reports 2SLS estimates where the birth cohort measure is used as an instrument for decadal change in the foreign-born Mexican population. All models condition on a set of control variables capturing demographic changes in each MSA as well as interacted region × year fixed effects. All models are weighted by 1980 MSA population. I report cluster-robust standard errors in parentheses below the coefficient estimates.

***Significant at the 1 percent level.

**Significant at the 5 percent level.

*Significant at the 10 percent level.

the pattern of the least squares results is not reflective of particularly large associations between Mexican immigrant settlement and MSA-level crime rates. In contrast, the 2SLS results lead to a dramatically different impression. Indeed, immigration can be shown to have a protective effect on several crime types. In particular, a 1 percentage point increase in the immigrant share leads to a 13 percent reduction in rape, an 11 percent reduction in larceny, and a 15 percent reduction in motor vehicle theft ($p < 0.10$).[8] On the other hand, the 2SLS coefficient in the aggravated assault model is positive and significant indicating that Mexican immigration appears to increase the rate of assault. For an MSA that received a large influx of Mexican immigrants over the course of a decade—for example, a 3–5 percentage point increase in the immigrant share—the effects on the crime rate can be very large. Given that the estimated share of Mexican immigrants in the US population increased from approximately 1.7 percent in 1990 to 3.5 percent in 2000, holding other factors constant, point estimates derived from the 2SLS procedure imply that Mexican immigration may explain as much as half of the decline in property crimes over this time period.

[8] The coefficient on burglary is also relatively large and negative though it is not precisely estimated.

While these results may sound implausible, to the extent that other factors may have been driving crime upwards over this time period, these estimates are best thought of as upper bounds with respect to the protective role of Mexican immigration.

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