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# U.S. Border Enforcement and Mexican Immigrant Location Choice

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**Abstract** We provide the first evidence on the causal effect of border enforcement on the full spatial distribution of Mexican immigrants to the United States. We address the endogeneity of border enforcement with an instrumental variables strategy based on administrative delays in budgetary allocations for border security. We find that 1,000 additional Border Patrol officers assigned to prevent unauthorized migrants from entering a U.S. state decreases that state's share of Mexican immigrants by 21.9 %. Our estimates imply that if border enforcement had not changed from 1994 to 2011, the shares of Mexican immigrants locating in California and Texas would each be 8 percentage points greater, with all other states' shares lower or unchanged.

**Keywords** Unauthorized immigration · Border enforcement · Mexico · Residential location choice

## Introduction

Since the early 1990s, Mexican immigrants to the United States have increasingly chosen nontraditional locations—that is, destinations other than those with historically high Mexican density, such as California and Texas (Card and Lewis 2007; Singer 2004, 2009). The reasons for this diffusion of the Mexican migrant population are complex and varied but are not yet well quantified. A hypothesis that Massey et al. (2002) advanced is that increased border enforcement in traditional migrant crossing

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areas has led migrants to choose alternative border crossing routes and in turn to choose nontraditional destinations. According to this view, an unintended consequence of strengthened border enforcement is a change in traditional settlement patterns among Mexican immigrants. In fact, between 1980 and 2010, the share of Mexican immigrants in California and Texas—the two states where border enforcement increases were most concentrated—fell from 80 % to 58 %. Of course, enforcement is not the only potential driver of location choice. Economic opportunities (Cadena 2013), interior enforcement policies (Watson 2013), and social factors (Bartel 1989) also play important roles.

To our knowledge, however, no causal analysis of the effect of border enforcement on the diffusion of Mexican migrants has been conducted. Indeed, the hypothesis is difficult to evaluate because of data limitations (crossing locations of Mexican immigrants currently in the United States are not available) and the endogeneity of border enforcement (the level of enforcement is likely responsive to illegal crossing behavior). This article quantifies the causal effect of border enforcement on immigrant location choice. We overcome the measurement problem by constructing an index that reflects the number of border agents assigned to disrupt the flow of unauthorized immigrants to each of the 50 U.S. states and the District of Columbia. Our index combines data on enforcement intensity across sectors of the southern border and over time with the historical destination choice of immigrants, drawing on methods developed in the literature (Borger et al. 2012; Pugatch and Yang 2011). We address the endogeneity of the enforcement index to contemporaneous migration flows by relying on administrative delay in enforcement budget allocations. Because budgetary decisions for border resources must be made two years in advance, lagged values of our enforcement index provide identifying variation for the effect of enforcement on immigrant location choice.

We find that increases in border enforcement decreased the share of Mexican immigrants across U.S. destinations. Specifically, we find that every 1,000 additional Border Patrol officers assigned to prevent unauthorized migrants from entering a state decreases that state's national share of Mexican immigrants by 21.9 %. These results are stable across subgroups, with stronger effects for working-age males and noncitizens, and null effects for immigrants who are less likely to be border crossers. Our estimates imply that if border enforcement had not changed from 1994 to 2011, the shares of Mexican immigrants locating in California and Texas would each be 8 percentage points greater, with all other states' shares lower or unchanged.

This study is motivated by the falling share of Mexican immigrants in traditional destinations and the coincident increase in border enforcement displayed in Fig. 1. The concentration of Mexican immigrants in a handful of traditional destinations began to decline in the 1990s, with states in the Southeast, Great Plains, and Midwest experiencing the fastest growth in Mexican immigration between 1994 and 2011.[1] At the same time, control of the southern U.S. border increased substantially, prompting Massey et al. (2002) to hypothesize a causal relationship between these trends. "The massive buildup of enforcement resources in southern California, El Paso, and around other ports of entry," they wrote, "diverted the migratory flows away from traditional points of destination" (Massey et al. 2002:127).

[1] These patterns have been amply documented in other work (Card and Lewis 2007; Hall 2013; Hall and Stringfield 2014; Singer 2004, 2009; Zuniga and Hernandez-Leon 2005).

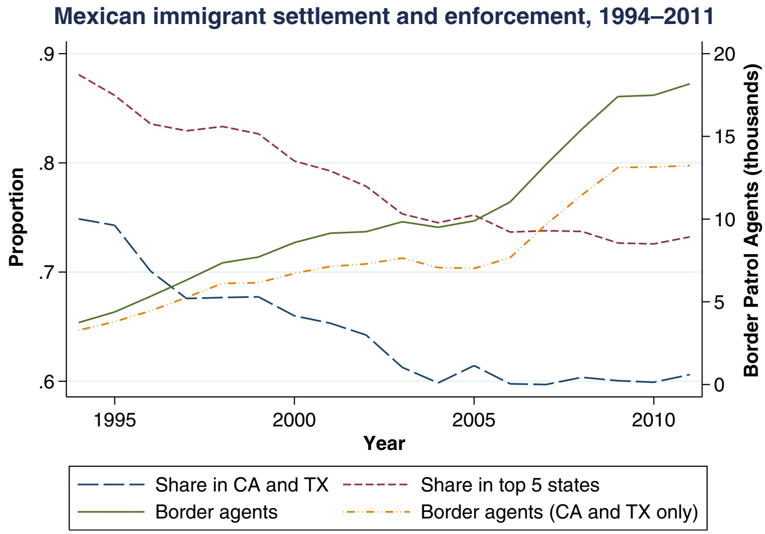


**Fig. 1** Mexican immigrant diffusion and border enforcement. CA and TX exclude only the border sectors that are wholly outside those states. *Source:* CPS and DHS

Gaining a better understanding of the effect of border enforcement on Mexican immigrant settlement patterns should be of major interest to policymakers. Immigrants play an important role in equilibrating local labor markets (Borjas 2001; Cadena and Kovak 2013), and their large share of the workforce has prompted renewed calls for national immigration reform in recent years. State legislatures have entered the immigration policymaking arena in the absence of federal reform, and evidence suggests that state policies themselves are driven by rapid inflows of new immigrant populations (Boushey and Luedtke 2011; Hopkins 2010). Because Mexicans constitute the largest immigrant group in the United States and have a high propensity to enter the United States without authorization, their location decisions hold particular importance. Moreover, attempts to thwart unauthorized immigration come at considerable expense, with the U.S. Customs and Border Protection (CBP) budget for 2012 at nearly $12 billion (U.S. Department of Homeland Security 2013). The role that border enforcement plays in Mexican immigrant locations is thus important at both the national and local levels.

Before proceeding with exposition of our methodological approach, we briefly place this article in the context of two broad literatures: one on impacts of border enforcement, and the other related to immigrant location choice. The influence of border enforcement on aggregate migration flows is the subject of considerable previous research (see, e.g., Angelucci 2012; Cornelius 2001; Gathmann 2008; Hanson and Spilimbergo 1999; Kossoudji 1992; Orrenius 2004; Reyes et al. 2002). Increases in border enforcement alter migrant crossing locations (Cornelius 2001; Massey et al. 2002; Sorensen and Carrion-Flores 2007) and increase migration costs (Orrenius 2004; Roberts et al. 2010). Although apprehensions at the border are apparently correlated with increases in enforcement (Orrenius 2004), it is unclear that illegal immigration is correlated with enforcement, in part because measuring attempted crossing is difficult. However, research has indicated that one unintended effect of increased border enforcement may be to increase the length of stays in the United States by discouraging immigrants currently located in the United States from engaging in return and circular migration (Reyes et al. 2002).

Although these studies and many others have examined migration decisions to and within the United States, none (to our knowledge) have evaluated the causal role of border enforcement on the full spatial distribution of immigrants. The closest antecedents to this study are Alves Peña (2009) and Lessem (2012), both of which developed models in which border enforcement may influence Mexican immigrant residential locations, rather than just aggregate flows. Alves Peña (2009) found that location choice is negatively related to linewatch hours, but the analysis was limited to agricultural workers in four U.S. states. Lessem (2012) found that increased border enforcement is a strong deterrent to migration. Her study did not explicitly address the role of border enforcement in immigrant location decisions within the United States, but the structural model that she developed could be used for this purpose. Crucially, neither study accounted for the endogenous response of border policy to migration flows. In contrast, our article uses large-scale, nationally representative data on all Mexican immigrants to the United States, and isolates plausibly exogenous variation in border enforcement.

Why might border enforcement influence the location of immigrants within the United States in addition to altering the magnitude of overall migration flows? As arguably the most mobile demographic group in the United States, potential immigrants consider several factors when choosing whether to migrate and where to reside, including the presence of others from their home communities (Bartel 1989), local employment opportunities (Cadena 2013, 2014), state immigration policies (Amuedo-Dorantes and Pozo 2014; Bohn et al. 2011; Watson 2013), and migration costs (Chiquiar and Hanson 2005; Orrenius 1999). The link between enforcement and location choice is most closely related to the latter.

Conceptually, it is useful to think of migrants reaching a destination as the outcome of two related choices.[2] The first choice is border crossing location, which carries a fixed cost that depends on enforcement. The second choice is the destination, with a cost that varies according to distance from the border crossing, with distance defined broadly to include direct travel costs, forgone earnings during travel, and the availability of preexisting networks to assist with arrival and employment at the destination. For any destination, migrants will choose the cost-minimizing combination of crossing location and distance. When enforcement at a crossing location is low, nearby destinations may attract migrants even if their economic opportunities and other amenities are poor. When enforcement at a crossing location is high, the benefits of locating nearby must be commensurately greater to attract migrants, making alternative crossing locations and destinations farther from the border more attractive.

Migrant networks, which play an important role in connecting sending communities in Mexico to destinations in the United States (Munshi 2003), mediate this process. Although jobs present a strong pull for potential Mexican migrants to the United States, social networks guide both the decision to migrate and the destination choice by lowering the (broadly defined) costs of migration (Massey et al. 1987, 1994). This social process does not necessarily govern migration in traditional locations only, but social capital and the strength of social networks may drive migration to "new destinations" as well (Hernandez-Leon and Zuniga 2000, 2003; Kritz and Nogle 1994). If increased

[2] We are grateful to the anonymous referees for helping to clarify our thinking about these issues.

enforcement disrupts a crossing location traditionally used by a sending community, the lack of a network to reach the intended destination via an alternate crossing route may lead potential migrants to delay their journey (Borger et al. 2012). Migrants whose networks have better knowledge of these alternative crossing routes and a different set of preferred destinations may take their place, altering the composition of migrants in the United States (Angelucci 2012; Ibarrraran and Lubotsky 2007; Leach and Bean 2008; Lozano and Lopez 2013; Orrenius and Zavodny 2005) and their spatial distribution.

The propensity for return migration may also change differentially across destinations because of border enforcement: circular migrants who anticipate more difficult round trips between the United States and Mexico may choose instead to remain in the United States (Reyes et al. 2002). The effects on prospective immigrants and return migrants work in opposite directions, making the role of border enforcement in immigrant location choice theoretically ambiguous. This multiplicity of channels underscores the need for a theoretical framework and rigorous empirical analysis. We present a migration choice model that formalizes this argument and connects it with our empirical analysis in the following section. Our focus is on consistent estimation of the total effect of border enforcement on the distribution of the immigrant population across destinations. We leave the question of the spatial dimension of selection in response to border enforcement to future work.

## Model and Methodology

Suppose, as in Sjaastad (1962) and Borjas (1987), that a migrant chooses to reside in the location that offers the highest utility net of migration costs. We adapt their models to a random utility framework, following closely the exposition of Scanlon et al. (2002) and Cadena (2013) while placing emphasis on the role of border enforcement in the migrant's location decision. Conditional on migrating, the value function for immigrant $i$ locating in United States destination $k$ in period $t$ is

$$V_{ikt} = \theta e_{kt} + \mathbf{X}_{kt}\beta + \varepsilon_{ikt}, \tag{1}$$

where $e$ is the enforcement intensity associated with locating at the destination in that period, $\mathbf{X}$ is a vector of controls capturing the economic opportunities and other observable characteristics of a destination relevant to location choice, and $\varepsilon$ is the error term. (The controls $\mathbf{X}$ do not carry an $i$ subscript because we will conduct the analysis using destination-level aggregates.) The immigrant chooses destination $k$ if $V_{ikt} \geq V_{ijt}$ for all $j \neq k$. Enforcement affects immigrant location choice by altering the costs of reaching a destination, as described in the Introduction.

We formalize this argument in Online Resource 1, where we present a more detailed model of migrant location choice. We extend the standard migration choice model to allow for multiple destination choices, reachable through multiple border crossings.[3] Each border crossing–destination pair implies an enforcement cost and a distance cost, where distance is defined broadly to include direct travel costs, ease of travel through

[3] The model includes the choice to remain in the source country as the outside option. However, this choice will be unobserved when using U.S. data, so we focus attention on the case where a migrant is choosing among locations in the destination country.

preexisting networks associated with the destination, forgone earnings during travel, and probability of apprehension in the interior. Increased enforcement at the border crossing nearest a particular U.S. destination forces migrants to choose between a shorter route that is now riskier or a longer route for which they may lack knowledge or networks. Either option raises migration costs, making alternative destinations more attractive. The model predicts shifts to destinations with lower migration costs, including more distant destinations not previously considered. This may occur because the same individuals shift their intended destination, or because one group of potential migrants postpones their journey while another group whose intended destination has lower relative migration costs takes their place. Because either explanation will lead to the spatial reallocation of migrants hypothesized by Massey et al. (2002), we leave this question aside for now but return to it in the Conclusion section. Moreover, the model shows that migrant locations can change in response to enforcement because of divergence alone, deterrence alone, or some combination. Our estimates report only the total effect without quantifying the relative contributions of deterrence and divergence.

Although the formulation in Eq. (1) is straightforward, several challenges arise immediately. First, it is not obvious how to measure the level of enforcement $e$ faced by potential migrants to a destination $k$, particularly for destinations in the interior. Second, even if enforcement can be measured for a destination, such enforcement is likely endogenous to immigrant location decisions. For instance, if the government responds to a rapid influx of unauthorized immigrants at a border crossing by increasing enforcement, then enforcement intensity $e$ will be correlated with the error term, preventing us from consistently estimating $\theta$. We address the first of these challenges before returning to a discussion of how we use Eq. (1) as the basis of our empirical specification. We close the section with a description of an instrumental variables strategy that addresses the second concern.

Consider the problem of measuring enforcement faced by a prospective migrant to destination $k$. No large-scale, nationally representative data set exists that provides information on the current U.S. locations of Mexican immigrants and their point of entry.[4] Even if such a data set were available, it is not clear that enforcement at the migrant's point of entry is the proper measure of enforcement that s/he faced. Migrants have a choice among crossing locations and could be influenced by enforcement at alternative locations as well. To address this issue, we build on methods developed by Pugatch and Yang (2011) and Borger et al. (2012) to construct a new measure of border enforcement intensity. We combine data on the historical border crossing and destination patterns of Mexican immigrants to the United States with current measures of border policy to assign a border enforcement index to U.S. locations.

The U.S. CBP splits the southern border with Mexico into nine sectors, with each sector responsible for preventing unauthorized crossings of people and goods in its territory. The CBP adjusts enforcement intensity across sectors to meet perceived security needs, leading to variation in enforcement across sectors and over time. This variation will not affect the desirability of locating in all U.S. destinations equally. For example, suppose that we observed that prior to our sample period, all migrants to

[4] The leading panel data sets with crossing locations of Mexican migrants—the Mexican Migration Project (MMP) and Survey on Migration at the Northern Border (EMIF-N)—do not meet these criteria. The MMP is not nationally representative, and the EMIF-N surveys migrants on the Mexican side of the border, making it difficult to connect crossing locations to the current spatial distribution of migrants. We use the EMIF-N for data on historical crossing locations, however.

Missouri crossed the border through one of two sectors: the Rio Grande Valley (eastern Texas) and the Laredo Sectors. Suppose further that 10 % of migrants crossing through the Rio Grande Valley Sector located in Missouri, and Missouri's share of the Laredo Sector migrant crossers was 5 %. Then a natural measure of border enforcement intensity for Missouri would be to assign 10 % of the Rio Grande Valley Sector and 5 % of the Laredo Sector's enforcement to Missouri, with all other sectors contributing 0. This sector-weighted average of enforcement intensity leads to the following U.S. location-specific enforcement index:

$$e_{kt} = \sum_{s=1}^{9} \omega_{ks} e_{st}, \tag{2}$$

where $\omega_{ks}$ is the share of immigrants who cross at border sector $s$ who locate in destination $k$, and $e_{st}$ is enforcement intensity at sector $s$ at time $t$. We use the number of Border Patrol agents (in thousands) as our enforcement measure so that the index $e_{kt}$ may be interpreted as Border Patrol agents assigned to prevent unauthorized immigration to location $k$ at time $t$.[5] Importantly, the weights used to construct the index are predetermined with respect to enforcement levels so that contemporaneous enforcement patterns do not cause the observed immigrant destination choices.[6] The weights also allow the enforcement index to reflect the importance of migrant networks by placing greater weight on the destinations that are most popular among migrants who cross at a particular sector. If such networks help determine migrant destinations, then these weights will be better predictors of how migrants adapt to changing enforcement patterns than an alternative weight based only on distance, a hypothesis that we will check later. Identifying variation for the effect of border enforcement on immigrant location choice therefore comes from three sources: (1) spatial variation in border enforcement across sectors; (2) time series variation in border enforcement within sectors; and (3) cross-sectional variation in the propensity of immigrants to follow particular routes from border crossings to U.S. destinations.

Return now to Eq. (1), the migrant's value function for locating in a particular destination. Let $\varepsilon_{ujt} = \eta_{kt} + u_{ikt}$ so that the error may be decomposed into a destination- and time-specific component $\eta$ and an idiosyncratic component $u$ that we assume to be independent and identically distributed (i.i.d.) type I extreme value. Then the share of immigrants choosing destination $k$ at time $t$, denoted $\pi_{kt}$, may be expressed as follows:

$$\pi_{kt} = \frac{\exp(\theta e_{kt} + \mathbf{X}_{kt}\beta + \eta_{kt})}{\sum_j \exp\left(\theta e_{jt} + \mathbf{X}_{jt}\beta + \eta_{jt}\right)}. \tag{3}$$

This is the familiar multinomial logit formula with an unobserved destination- and time- specific component $\eta$ included. Letting the sample share of immigrants $S$ differ from the population share $\pi$ by a multiplicative error $\nu$ (assumed uncorrelated with $\pi$) and taking logs yields

$$\log(S_{kt}) = \theta e_{kt} + \mathbf{X}_{kt}\beta + \eta_{kt} - \log(D_t) + \nu_{kt}, \tag{4}$$

[5] Consequently, summing the index by sector across all destinations (i.e., $\sum_k \omega_{ks} e_{st}$) equals total enforcement in that sector.

[6] Our results are robust to using weights based on time-varying crossing location, where this time variation remains predetermined with respect to enforcement. Results are available in Online Resource 1.

where $D_t = \sum_j \exp(\theta e_{jt} + \mathbf{X}_{jt}\beta + \eta_{jt})$, with the subscript acknowledging that this term is identical across all destinations at time $t$. Assume that $\eta_{kt}$ may be further decomposed into time-invariant and time-varying components as $\eta_{kt} = \zeta_k + \phi_{kt}$.Taking first differences of $S$ yields

$$\Delta\log(S_{kt}) = \theta\Delta e_{kt} + \Delta\mathbf{X}_{kt}\beta - \Delta\log(D_t) + \Delta\phi_{kt} + \Delta\nu_{kt}. \tag{5}$$

An empirical specification based on this first-differenced equation offers several benefits relative to multinomial choice estimation. First, it allows for linear estimation with easily interpretable coefficients: the coefficient of interest θ is the *ceteris paribus* effect of a one-unit change in enforcement intensity on the percentage change in the share of immigrants choosing a destination. Second, the specification allows for straightforward incorporation of factors common to all destinations within a time period through the inclusion of period fixed effects, which estimate $\Delta\log(D_t)$. Third, the specification also controls for permanent attributes of a location—such as climate, amenities, and the role of durable immigrant networks—through the term $\zeta$, which differences out of the equation.

A remaining concern, however, is correlation between the destination- and time-specific innovation $\Delta\phi_{kt}$ and changes in enforcement intensity. If border officials respond to shocks that increase the share of immigrants choosing a location by increasing enforcement intensity, then our estimates of θ will be upwardly biased. We address this issue by instrumenting for $\Delta e_{kt}$ with enforcement lagged two periods. As Borger et al. (2012) noted, administrative delays in CBP budget approval led to two-year lags between initial requests and realized outlays. To set its budget, CBP implements a process known as the Operational Requirements Based Budget Program (ORBBP), in which Border Patrol sectors request resources to enforce immigration and customs laws based on an assessment of current needs.[7] This assessment is based on all available information at the time of the request, including data maintained by CBP on current enforcement levels and apprehensions of undocumented migrants. The ORBBP occurs annually, but the lag between initial requests and resource allocation exceeds one year.

Although budget allocations determined through the ORBBP follow a fairly rigid process, the U.S. Department of Homeland Security (DHS) may also address unexpected border enforcement needs through a "surge" of agents or other resources to particular border sectors. Because these additional resources may be contemporaneously correlated with immigrant flows, we are concerned about inconsistent estimates obtained through ordinary least squares (OLS). However, initial budget requests are based on an assessment of enforcement needs before such unexpected shocks are realized. If these initial requests are uncorrelated with the change in unobserved factors realized two years later, then the identifying assumption that $e_{k,t-2}$ is uncorrelated with $\Delta\phi_{kt}$ will hold. This approach also mirrors one that has been used in the labor supply literature (e.g., Ziliak 1997).

The choice of control variables to include in **X** is also important to isolate the role of border enforcement from other factors influencing immigrant location choice. We include a host of destination-specific controls for economic conditions most relevant to prospective immigrants: unemployment rates, hourly wages, GDP per capita,

[7] We base this section on discussions with former U.S. Department of Homeland Security officials.

manufacturing output, agricultural output, construction output, and new housing permits. The economic sectors are chosen because of the high concentration of Mexican immigrants employed in these industries. Moreover, including new housing permits separately from current output helps to capture the role of economic expectations in immigrant location decisions.

We also include measures of state-level legislation aimed at immigrants, which has proliferated since 2004. Arguably in response to increasing unauthorized immigrant populations and federal inaction on comprehensive policy reform, state legislatures have enacted hundreds of laws between 2004 and the present. Most immigrant-related state laws are intended to deter employment or restrict services to unauthorized immigrants; a few have been shown to be effective deterrents, at least to immigrant location choice, if not to the law's stated intent (Bohn et al. 2011). Because policymakers see both border enforcement and state-level legislation as important deterrents to unauthorized immigration, including data on this legislation is critical to isolate the role of border enforcement in immigrant location decisions.

## Data

To conduct the analysis, we need data on population shares of Mexican immigrants by U.S. destination; enforcement intensity by Border Patrol sector; choices of border crossings and destinations by migrants to construct the weights used in the enforcement index; and destination-specific control variables. We describe the sources of these data, with additional details in Online Resource 1.

The main source for population data is the U.S. Current Population Survey (CPS), 1994–2011. We classify immigrants by place of birth; "natives" are those born in the United States. We also combine the 2000 U.S. Census and American Community Survey (ACS) 2001–2011 into an alternate data set to check the consistency of the CPS results. We work with state-level aggregates derived from these sources.[8] Relative to the U.S. Census/ACS, the CPS provides a longer time series, including a set of years (1994–1999) with notable fluctuations in border enforcement. These features lead us to prefer using the CPS despite the larger sample sizes available in the U.S. Census/ACS, although we use both data sources to check for consistent results.[9]

Data on border enforcement are from the DHS, which reports the number of Border Patrol agents employed by the CBP annually in each sector of the southern U.S. border. The primary activity of border agents is linewatch hours (Simanski and Sapp 2013; U.S. Department of Homeland Security 2003), which is a more direct measure of enforcement intensity used in related studies (Angelucci 2012; Gathmann 2008; Hanson and Spilimbergo 1999; Hanson et al. 2002; Lessem 2012; Orrenius 1999, 2004). However, DHS stopped reporting linewatch hours in mid-2004, and denied our

[8] We prefer state to other levels of geographic aggregation, such as the metropolitan statistical area (MSA), because state-years will contain fewer cells with zero immigrants than will alternative geographic units. Passel and Cohn (2010) cautioned against using the CPS and ACS for MSA-level analysis when focusing on unauthorized immigrants. States also leave more scope to control for changing economic conditions because of greater data availability. A few states nonetheless have zero immigrants in a few years of the CPS. We drop these observations and check for robustness to this choice.

[9] See Online Resource 1 for additional details on the consistency and comparability of these data sources.

repeated Freedom of Information Act requests to obtain updated linewatch data. A comparison of agents and linewatch hours in years of overlap suggests that the two are highly correlated, and assures us that an enforcement index based on agents captures enforcement intensity (see Fig. S4 in Online Resource 1).

Data on border crossing patterns used to construct weighted enforcement are from the Survey on Migration at the Northern Border (EMIF-N) , a survey of migrants along the U.S.-Mexico border conducted by the Mexican government annually since 1993. The EMIF-N asks emigrants to the United States and returnees from the United States about border crossing location and state of U.S. residence on previous trips. We use the survey to construct, for each Border Patrol sector, the probability of entering each U.S. state. To do so, we assign each survey respondent to a border sector and a U.S. state according to the crossing point and place of main U.S. residence on his/her last trip to the United States. To mitigate recall bias, we drop any respondents whose last trip was more than 10 years prior to the interview date. The shares of migrants whose last trip to the United States occurred between 1983 and 1993 at border crossing $s$ with last U.S. residence in state $k$ are used to construct the crossing probabilities $\omega$ that appear in Eq. (2).[10] The EMIF-N is also the data source for prices paid to migrant smugglers (coyotes) used in later robustness checks.

Control variables on state-level economic conditions are from U.S. government sources (for details, see Online Resource 1, Section B). Data on state-level legislation aimed at immigrants was compiled from quarterly reports on all state laws related to immigrants from the National Conference of State Legislatures (NCSL).[11] We include as a control the presence of any deterrent state law related to employment or enforcement. This variable will equal 1 when signed into law and 0 in subsequent years because our empirical specification is in first differences.

Figure 1 combines these data sources to describe the phenomenon that motivates our inquiry. Between 1980 and 2010, the share of Mexican immigrants residing in traditional destinations fell from 90 % to 76 %, concomitant with steadily increasing enforcement at the southern border. Figure 2 presents data broadly consistent with the hypothesis that these enforcement increases diverted immigrant flows. Panel a shows border enforcement in selected sectors (only a subset is shown, for clarity), including a substantial increase in enforcement in the San Diego Border Patrol Sector in the mid-1990s that leveled off later in the decade. The Rio Grande Valley (eastern Texas) Sector also experienced an increase throughout the period, ending on a similar level as the San Diego Sector. The sharpest increase was in the Tucson Sector, however. Panel b shows the share of unauthorized Mexican immigrants crossing at each sector. After remaining flat for most of the period 1980–1995, the San Diego Sector began to lose shares beginning in the mid-1990s, while the Rio Grande Valley Sector ended the period at a similar level as its historical average.

[10] Given the availability of data on an immigrant's crossing location and U.S. destination in the EMIF-N, one might reasonably ask why we do not use the EMIF-N to construct our outcome measures in addition to the enforcement weights. We prefer using the CPS (and Census/ACS) for the outcome data because the much larger sample sizes (more than 1.5 million annually in the CPS compared with approximately 15,000 in the EMIF-N) will lead to more accurate measures of population shares. A similar argument applies to the Mexican Migration Project (MMP), which covers only selected Mexican communities, in addition to its relatively smaller sample. See also footnote 4.

[11] National Conference of State Legislatures state laws on immigration are available online (http://www.ncsl.org/research/immigration/state-laws-related-to-immigration-and-immigrants.aspx). Also see Online Resource 1, Section B.

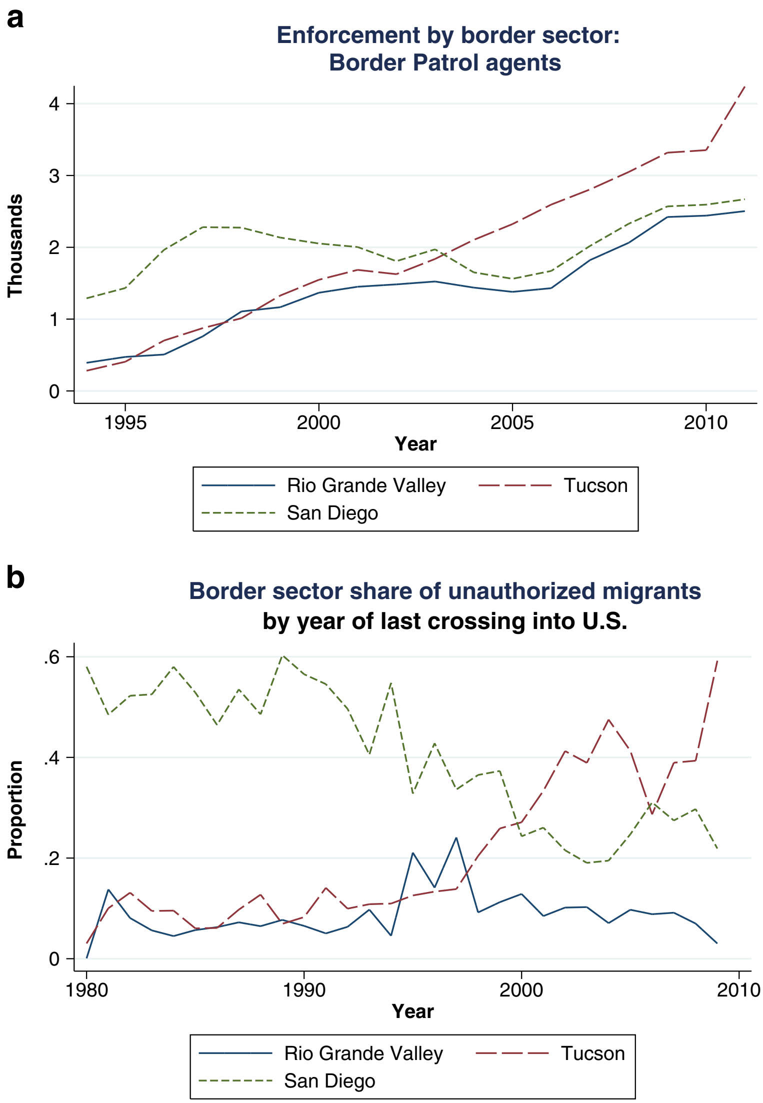


**Fig. 2** Border enforcement and Mexican migrant crossing patterns. *Sources:* DHS for panel a; EMIF, 1993–2009, for panel b

The Tucson Sector share increased considerably over the same period as the decline in the San Diego Sector. Although the evidence is circumstantial, Figs. 1 and 2 show a clear shift in enforcement and crossing activity from the traditional gateways on the western and eastern edges of the border toward the center.

In this article, we seek to determine whether these patterns also led to changes in the residential locations of Mexican immigrants. If changes in border enforcement during our sample period led immigrants to change their crossing patterns but not their destinations, then we would expect to see a weaker link between crossing locations and destinations over time. After all, travel costs within the United States are relatively low after a migrant has successfully crossed the border. In fact, connections between border crossings and destinations are remarkably durable, even as enforcement patterns have drastically changed. The correlation coefficient between the probability of choosing to locate in a U.S. state, conditional on border sector, during 1983–1993 (the period of our enforcement index weights) and 2011 is .91, consistent with the literature on the importance of networks in migrant decisions (Hernandez-Leon and Zuniga 2000, 2003; Kritz and Nogle 1994; Massey et al. 1987, 1994; Munshi 2003). If migrants changed their crossing locations in response to enforcement but not their destinations conditional on crossing, then border enforcement led to changes in immigrant locations. Although this simple correlation is not a substitute for a formal analysis, it suggests that our premise matches basic patterns in the data.

The index we use to measure the border enforcement intensity faced by potential migrants to each U.S. state consists of two components: (1) enforcement intensity by Border Patrol sector, and (2) weights representing the propensity of immigrants crossing at a sector to locate in a particular U.S. state. We have already presented data on the first component. Figure 3 shows data on the second component, in the form of maps showing Mexican immigrant destinations for selected border crossings. Panel a shows the locations chosen from 1983 to 1993 by migrants crossing in the Rio Grande Valley Sector (eastern Texas, with representative city Brownsville circled). Unsurprisingly, the modal destination is Texas, with southeastern states also popular. Panels b and c show the analogous maps for the El Paso (western Texas and New Mexico) and San Diego Sectors. As in panel a, immigrants crossing in these sectors choose destinations that are geographically proximate. This variation in U.S. destinations, conditional on border crossing location, allows us to transform the variation in enforcement across Border Patrol sectors into state-specific measures of border enforcement intensity.

Panel d of Fig. 3 shows the resulting enforcement index for a particular state, Arizona. The solid line shows the enforcement index, which may be interpreted as the number of Border Patrol agents assigned to prevent unauthorized immigrants from entering Arizona. Enforcement in the Rio Grande Valley, San Diego, and Tucson Sectors are also plotted. As shown in the graph, the correlation between Arizona's enforcement index and enforcement intensity in the Tucson Sector is much higher than that for the other sectors. This is the result we would expect if enforcement in the Tucson Sector is more relevant for potential migrants to Arizona than enforcement in the other sectors.[12]

We close this section by presenting summary statistics in Table 1 on the panel of U.S. states used in the analysis. The mean Mexican immigrant share is (approximately) 2 %, which is a mechanical result of the sample size of 50 states and the District of Columbia; we omit reporting shares of other population groups for this reason. The next several rows show average levels of various subpopulations (sample sizes vary because of state-year cells with zero shares, in accordance with the sample used in the regression analysis). The average state has 184,480 Mexican immigrants, compared with 4.1 million natives. Levels of other subpopulations are mostly as expected. The average level of the enforcement index is 0.21, indicating 210 Border Patrol agents assigned to prevent unauthorized immigration to an average state annually. An alternate enforcement index that replaces Border Patrol agents with apprehensions of unauthorized migrants in Eq. (2) shows 21,200 apprehensions

[12] Arizona is chosen as an illustrative example, rather than a representative one. Interior states will not show as strong a correlation between their enforcement index and that of a particular border sector.

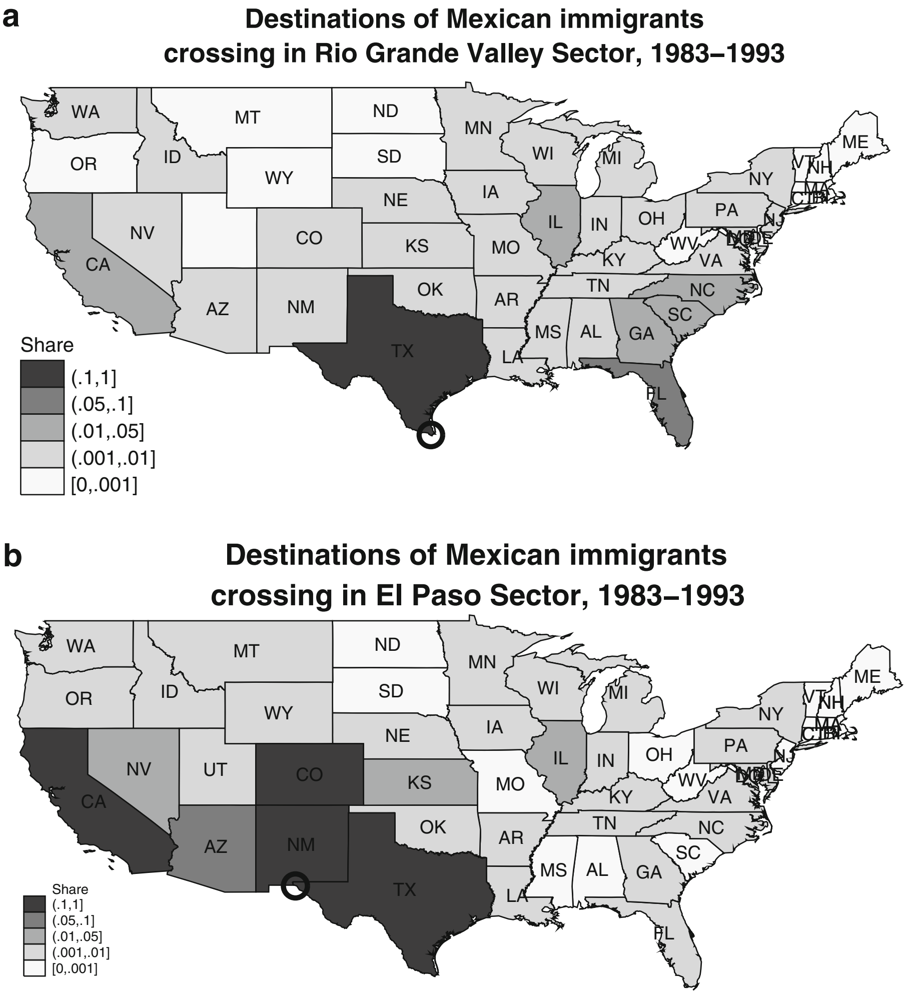


**Fig. 3** Mexican migrant crossing patterns and enforcement index. Panels a–c show the share of migrants at the border crossing choosing to reside in a state, as indicated in the legend. *Sources:* EMIF, 1993–2003, for panels a–c; DHS and EMIF for panel d

intended for an average state per year. The final rows of the table show border-sector specific enforcement, measured by number of agents. There is considerable variation across sectors, with the San Diego, Tucson, and El Paso Sectors assigned the largest numbers of agents.

## Results

### Main Results

In estimating Eq. (3), we include in the vector of controls (**X**) unemployment rates, hourly wages, (log) GDP per capita, (log) manufacturing output,

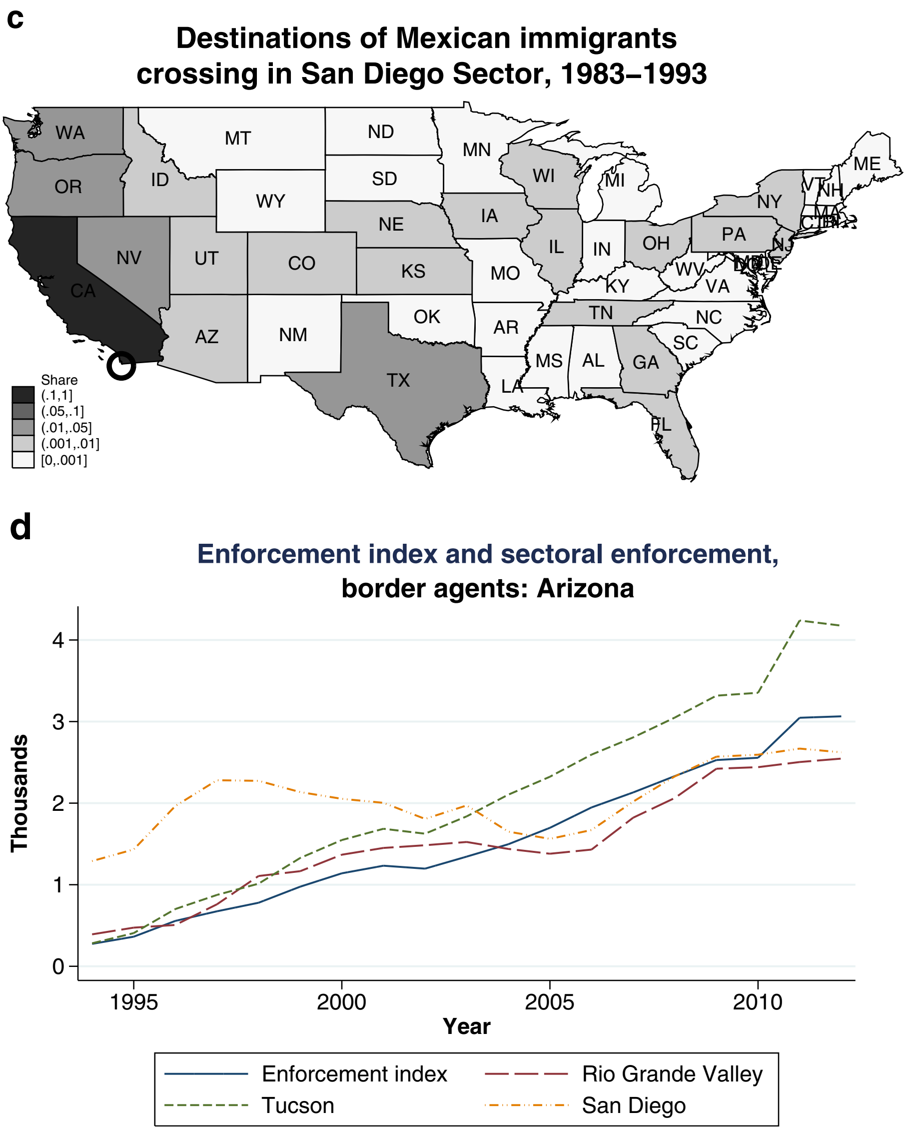


**Fig. 3** (continued)

(log) agricultural output, (log) construction output, (log) new housing permits, and an indicator for passage of any punitive legislation aimed at immigrants, all in first differences. Unemployment rates and hourly wages are specific to the subpopulation whose population shares are under analysis. We also include a constant and year fixed effects. We cluster standard errors by state.

Before discussing results of estimating Eq. (3), we present in Table 2 the results of the first stage, in which we regress the first difference of the enforcement index on its

**Table 1** Summary statistics

| Variable | *N* | Mean | SD |
|---|---|---|---|
| Mexican Immigrants | | | |
| Share | 840 | 0.02 | 0.07 |
| All | 840 | 194,931 | 631,422 |
| Males 16–50 | 810 | 73,869 | 217,852 |
| Naturalized citizen | 790 | 47,152 | 164,748 |
| Not naturalized citizen | 816 | 154,978 | 481,895 |
| Abroad last year | 220 | 1,674 | 3,951 |
| Not abroad last year | 840 | 194,396 | 630,026 |
| Internal migrant | 280 | 834 | 2,058 |
| Not internal migrant | 839 | 194,814 | 631,187 |
| Other Population Groups | | | |
| Non-Mexican immigrant | 867 | 466,837 | 941,395 |
| Natives | 867 | 4,876,839 | 4,898,778 |
| Puerto Rican | 743 | 29,911 | 70,619 |
| Central American | 836 | 47,433 | 121,534 |
| Enforcement | | | |
| Index (agents) | 840 | 0.21 | 0.75 |
| Index (apprehensions) | 840 | 21.2 | 76.9 |
| Weights (state-sector pairs) | 423 | 0.02 | 0.10 |
| Border Patrol agents (sector-years) | | | |
| All sector-years | 153 | 1,177 | 810 |
| Rio Grande Valley (TX) | 17 | 1,491 | 615 |
| Laredo (TX) | 17 | 1,035 | 496 |
| Del Rio (TX) | 17 | 963 | 421 |
| Big Bend (TX) | 17 | 298 | 194 |
| El Paso (TX and NM) | 17 | 1,529 | 747 |
| Tucson (AZ) | 17 | 2,048 | 1,065 |
| Yuma (AZ) | 17 | 492 | 303 |
| El Centro (CA) | 17 | 682 | 337 |
| San Diego (CA) | 17 | 2,058 | 365 |

*Notes:* The table shows summary statistics from U.S. state-years 1995–2011 (unless otherwise indicated). Population data are from the Current Population Survey. Enforcement index = $\Sigma_s$ Pr(U.S. destination | cross at border sector $s$) × enforcement at sector $s$, where enforcement is in thousands of Border Patrol agents or apprehensions of unauthorized migrants. The index may be interpreted as the amount of enforcement dedicated to preventing the arrival of unauthorized migrants at destination. Enforcement weights (crossing probabilities) are calculated from EMIF-N crossings during 1983–1993. Border Patrol agents and apprehensions are from the U.S. Department of Homeland Security.

second lag, with the same set of controls as described earlier. In column 1, the coefficient on the instrument is 0.073, indicating that every 1,000 Border Patrol agents assigned to a state two years ago corresponds to an increase of 73 agents in the past year.[13] The coefficient is precisely estimated, with an $F$ statistic of 74.2. In column 2, we restrict attention to the years 2000–2011, corresponding to the period of the U.S. Census/ACS sample. The coefficient falls slightly to 0.068, with an $F$ statistic of 59.

An alternate instrument that replaces Border Patrol agents with apprehensions of unauthorized migrants (in thousands) also produces strong first-stage results. Column 3 shows that the coefficient for the enforcement index based on apprehensions is 0.57, representing the additional agents assigned to a state for every 1,000 apprehensions two years earlier. In column 5, a second candidate instrument that replaces border agents with the median coyote price to cross the border sector is also strong. Estimates are robust for both alternate instruments using the U.S. Census/ACS sample, with all $F$ statistics greater than 25. We prefer using the agent-based instrument for the second-stage analysis, however, because it is the source of variation most directly related to border policy and because it provides a stronger first stage. We later check the robustness of results to using the alternative instruments.

Table 3 presents the main results from estimation of Eq. (3), with OLS results in panel A and IV results in panel B.[14] Column 1 uses a state's share of all Mexican immigrants located in the United States as the dependent variable. The OLS coefficient of –0.176 indicates that an increase of 1,000 Border Patrol agents assigned to a state is correlated with a 17.6 % decrease in a state's share of Mexican immigrants. In an average state with a 2 % share, this would reduce the share to 1.65 %. The corresponding IV coefficient in panel B is –0.219, meaning that a one-unit increase in the enforcement index leads to a 21.9 % decrease in a state's Mexican immigrant share. The larger magnitude of the IV coefficient is as we would expect if OLS coefficients are upwardly biased because enforcement responds to immigrant inflows. Both coefficients are statistically significant at 1 %.[15]

In subsequent columns of Table 3, we focus on subpopulations of Mexican immigrants to look for differential responses to border enforcement. In column 2, we restrict attention to males aged 16–50 with a high school education or less, a group with a high propensity to migrate. The IV coefficient is nearly 1.5 times both the corresponding OLS coefficient and that for all Mexican immigrants in column 1. The larger magnitude is as expected if this group is more likely to be affected by border enforcement.[16]

[13] For ease of exposition, the enforcement index based on Border Patrol agents is specified in thousands.

[14] First-stage $F$ statistics reported in Table 3 do not correspond exactly to those in column 1 of Table 2 because estimation samples vary as a result of state-years with a zero population share, for which the log population share is undefined. Cells with a zero share also explain the uneven sample sizes across columns. We check the sensitivity of results to exclusion of these observations in Table S4 in Online Resource 1.

[15] To give a better sense of the magnitudes of our estimates, the average change in the enforcement index is 0.017, representing an annual increase of 17 Border Patrol agents assigned to a state. Multiplying this figure by our IV estimate of –0.219 results in a predicted annual decline of 0.37 % in an average state's Mexican immigrant share. In the average state with a 2 % Mexican immigrant share, this will result in a decline to 1.99 % in one year, or a decline to 1.88 % when compounded over the 17 years of our sample. Our IV estimate implies an elasticity of –0.04 (standard error 0.016) when evaluated at the mean level of enforcement ($-0.219 \times 0.21 = -0.04$); standard error found by the delta method.

[16] In Online Resource 1, we investigate the response of unauthorized immigrants. U.S. government surveys do not ask about immigrants' legal status. Instead, we use state-level estimates of unauthorized immigrants from Warren and Warren (2013), multiplied by the proportion of immigrants who are Mexican (according to the state-year cell of the CPS panel) to obtain an estimate of a state's share of unauthorized Mexican immigrants. Our estimates imply that 34 % of Mexican immigrants are unauthorized, which is low compared with Hanson's (2006:870) estimate of 56 %. Given this discrepancy, the results of this supplemental analysis should be taken with a grain of salt. Nonetheless, we find that the response of unauthorized Mexican immigrants to border enforcement is greater in magnitude than for all Mexican immigrants, as expected.

**Table 2** First stage

| | Enforcement Index, First Difference | | | | | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Agent Index, $t$ – 2 | 0.073 | 0.068 | | | | |
| | (0.009)** | (0.009)** | | | | |
| Apprehension Index, $t$ – 2 | | | 0.570 | 0.604 | | |
| | | | (0.095)** | (0.120)** | | |
| Coyote Price Index, $t$ – 2 | | | | | 0.164 | 0.150 |
| | | | | | (0.026)** | (0.029)** |
| Observations | 863 | 612 | 914 | 612 | 918 | 612 |
| $R^2$ | .38 | .37 | .39 | .33 | .50 | .51 |
| First-Stage $F$ Statistic | 74.2 | 59.0 | 36.3 | 25.4 | 40.9 | 27.2 |
| Data Source | CPS | ACS | CPS | ACS | CPS | ACS |

Notes: The table shows regressions of first difference of enforcement index on its second lag. The sample is U.S. state-years (including the District of Columbia) from the Current Population Survey, 1995–2011 in column 1, and U.S. Census 2000 and the American Community Survey 2001–2011 in column 2. Enforcement index = $\Sigma_s$ Pr(U.S. destination | cross at border sector $s$) × enforcement at sector $s$. The index may be interpreted as the amount of enforcement dedicated to preventing the arrival of unauthorized migrants at destination, where enforcement is number of Border Patrol agents or apprehensions of unauthorized migrants (in thousands), as indicated. Crossing probabilities and coyote price are calculated from EMIF-N crossings during 1983–1993. Coyote price is the median price at the border sector in real U.S. dollars, base 1983. Border Patrol agents are from the U.S. Department of Homeland Security. All regressions include year fixed effects and the following controls (in first differences): Mexican immigrant unemployment rate, Mexican immigrant hourly wage, log GDP per capita, log agricultural GDP, log manufacturing GDP, log construction GDP, log new housing permits, and a dummy variable for passage of any punitive immigration legislation. Robust standard errors, shown in parentheses, are clustered by state.

**$p < .01$

Despite the lack of information about immigrant legal status in the data, data on U.S. citizenship can be used to identify a subgroup of immigrants with certain legal status. These immigrants are not at risk of deportation and therefore should not respond to border enforcement in the same manner as noncitizens. Splitting the sample of those born in Mexico into naturalized citizens and noncitizens in columns 3 and 4, we find that naturalized citizens are not responsive to border enforcement when deciding where to reside in the United States, but noncitizens are.

Mexican immigration to the United States is characterized by high rates of circular migration, with migrants cycling back and forth between countries with some regularity (Rendon and Cuecuecha 2010). For migrants currently at a U.S. destination, greater enforcement increases the cost of return migration to Mexico by making it more difficult to engage in circular migration. This increases the incentives for migrants to remain at their U.S. destination when border enforcement tightens (Angelucci 2012; Kossoudji 2002). Although U.S. government surveys do not ask directly about circular migration, they do ask for a respondent's migration status one year ago. Mexican immigrants who report being abroad last year were presumably residing in Mexico, and are likely reentrants or newly arrived migrants to the United States, compared with those who report residing in the United States the previous year. We split Mexican immigrants into

**Table 3** Mexican immigrant population shares and border enforcement

| | All | Males 16–50 HS or Less | Naturalized Citizen | | Abroad Last Year | | Internal Migrant | |
|---|---|---|---|---|---|---|---|---|
| | | | Yes | No | Yes | No | Yes | No |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Panel A: OLS | | | | | | | | |
| Enforcement index | –0.176 | –0.204 | –0.075 | –0.209 | –0.3 | –0.175 | –0.571 | –0.175 |
| | (0.058)** | (0.057)** | (0.050) | (0.056)** | (0.272) | (0.058)** | (0.350) | (0.058)** |
| $R^2$ | .04 | .03 | .02 | .03 | .07 | .04 | .12 | .04 |
| Panel B: IV | | | | | | | | |
| Enforcement index | –0.219 | –0.316 | –0.047 | –0.304 | –0.003 | –0.214 | 0.205 | –0.218 |
| | (0.075)** | (0.078)** | (0.066) | (0.090)** | (0.316) | (0.074)** | (0.292) | (0.073)** |
| Observations | 839 | 809 | 789 | 815 | 220 | 839 | 280 | 838 |
| First-stage $F$ statistic | 75.6 | 75.6 | 75.4 | 75.7 | 49.4 | 75.6 | 53.8 | 75.7 |

*Notes:* The table shows regressions of the log population share on enforcement index, in first differences. The sample is U.S. state-years (including the District of Columbia) from the Current Population Survey, 1995–2011. Population share is the state's share of Mexican immigrants within each category indicated. Enforcement index = $\Sigma_s$ Pr(U.S. destination | cross at border sector $s$) × enforcement at sector $s$, where enforcement is thousands of Border Patrol agents. The index may be interpreted as the amount of enforcement dedicated to preventing the arrival of unauthorized migrants at destination. The instrument in IV specifications is the second lag of enforcement index. Crossing probabilities are calculated from EMIF-N crossings during 1983–1993. Border Patrol agents are from the U.S. Department of Homeland Security. All regressions include year fixed effects and the following controls (in first differences): Mexican immigrant unemployment rate, Mexican immigrant hourly wage, log GDP per capita, log agricultural GDP, log manufacturing GDP, log construction GDP, log new housing permits, and a dummy variable for passage of any punitive immigration legislation. Robust standard errors, shown in parentheses, are clustered by state.

$^{**}p < .01$

groups by their migration status one year ago in columns 5 and 6 of Table 3. In column 5, the IV coefficient for those residing abroad one year ago is nearly zero, while the coefficients for those not abroad one year ago are similar to those for the full sample. The results are consistent with border enforcement leading to postponement of return migration to Mexico, rather than deterring reentry or new migration to the United States.

If migrants are responding to local shocks other than border enforcement, then they may switch locations within the United States rather than change their entry or exit decision. Although our specification controls for many shocks at the state level, the data allow us to further test these responses by classifying Mexican immigrants as internal migrants if they resided in a different U.S. state one year ago, and a noninternal migrant otherwise. We expect border enforcement to exert a greater influence on noninternal migrants. The IV coefficient for noninternal migrants in column 8 of Table 3 is almost identical to that for the full sample in column 1. In column 7, however, the IV coefficient for internal migrants is positive but not statistically significant. These findings show that the effect of border enforcement on location choice is driven by movements across the border, not between U.S. states. This differential response helps alleviate concerns that the enforcement index is correlated with a more general, but unobserved, adverse environment for all Mexican immigrants at a location.

In Table 4, we repeat the specification of Eq. (3), using additional population groups. In these regressions, we replace controls for the Mexican immigrant unemployment rate and hourly wage with those for the relevant subpopulation, but all other covariates are unchanged. In column 1, we analyze shares of all non-Mexican immigrants. The OLS and IV coefficients are statistically indistinguishable from zero, in contrast to our earlier findings for Mexican immigrants.

Column 2 of Table 4 shows the response of natives to border enforcement. Although fear of deportation should not lead natives to respond to border enforcement, they might nonetheless respond indirectly through the effect of border enforcement on the location decisions of other groups. The results show that this is the case, with the IV coefficient positive and significant. This result is consistent with natives engaging in wage arbitrage as immigrants relocate from high to low enforcement-intensity states.[17] The response is relatively mild, however: the IV coefficient of 0.037 implies that an increase of 1,000 Border Patrol agents assigned to a state increases the native population share by 3.7 %, or from 2 % to 2.07 % in the average state.

In column 3 of Table 4, we examine the response of Puerto Ricans, who provide a useful falsification test for our main results because of their linguistic and cultural similarities with Mexicans and their U.S. citizenship. We find no statistically significant movements of Puerto Ricans in response to border enforcement, as might be expected. Column 4 shows the response of Central Americans, with the IV coefficient on the enforcement index of –0.201 significant at 5 %. This is an interesting result, suggesting that Central Americans respond to border enforcement in similar fashion as Mexicans, which is consistent with anecdotal evidence of relatively large flows of unauthorized Central Americans into the United States through Mexico and the southern U.S. border.

## Robustness Checks

In the Data section, we discuss the reasons why we preferred the longer panel based on the CPS relative to the shorter Census/ACS panel. However, the Census/ACS panel provides larger sample sizes than CPS, and thus is a better source for the years over which the panels overlap. Table S3 in Online Resource 1 shows results analogous to Table 3 using the Census/ACS panel, which covers the years 2000–2011. The results are quite similar to those from the CPS panel. Additionally, we check whether including state-year observations with a zero population share leads to different results. To do so, we add one person to all state-year subpopulations and recalculate the population shares so that the log share is defined for all cells. The results, presented in Table S4 in Online Resource 1, are similar to Table 3.

Table 5 presents a series of robustness checks using the main CPS sample with Mexican immigrant share as the dependent variable. In column 1, we replace the enforcement index

[17] Another possibility, suggested by an anonymous referee, is that border enforcement induces Mexican immigrants to self-identify as natives in surveys, leading to a spurious increase in the native share. Using native white non-Hispanic share as the dependent variable, we obtain an IV coefficient of –0.06 (with standard error of 0.19), consistent with this explanation. Other explanations are also plausible. For instance, native Hispanics may relocate as a result of inflows of immigrants because employers may see the newly arrived Mexican immigrants as close substitutes.

**Table 4** Population shares and border enforcement

| | Non-Mexican Immigrant (1) | Native (2) | Puerto Rican (3) | Central American (4) |
|---|---|---|---|---|
| Panel A: OLS | | | | |
| Enforcement index | 0.017 | 0.017 | 0.004 | –0.095 |
| | (0.030) | (0.014) | (0.068) | (0.070) |
| $R^2$ | .04 | .05 | .03 | .04 |
| Panel B: IV | | | | |
| Enforcement index | –0.03 | 0.037 | 0.017 | –0.201 |
| | (0.044) | (0.017)* | (0.105) | (0.084)* |
| Observations | 863 | 863 | 739 | 832 |
| First-stage $F$ statistic | 75.6 | 75.2 | 75.5 | 75.7 |

*Notes:* The table shows regressions of the log population share on enforcement index, in first differences. The sample is U.S. state-years (including the District of Columbia) from the Current Population Survey, 1995–2011. Population share is the state's share within each category indicated. Enforcement index = $\Sigma_s$ Pr(U.S. destination | cross at border sector $s$) × enforcement at sector $s$, where enforcement is thousands of Border Patrol agents. The index may be interpreted as the amount of enforcement dedicated to preventing the arrival of unauthorized migrants at destination. The instrument in IV specifications is the second lag of enforcement index. Crossing probabilities are calculated from EMIF-N crossings during 1983–1993. Border Patrol agents are from the U.S. Department of Homeland Security. All regressions include year fixed effects and the following controls (in first differences): unemployment rate of indicated group, hourly wage of indicated group, log GDP per capita, log agricultural GDP, log manufacturing GDP, log construction GDP, log new housing permits, and a dummy variable for passage of any punitive immigration legislation. Robust standard errors, shown in parentheses, are clustered by state.

*$p < .05$

weights based on migrant crossing patterns with the inverse square root of the distance between border sector and U.S. state. If the original weights reflect the importance of durable migrant networks in location choices, then weighting border enforcement with distance alone should weaken the relationship between enforcement and immigrant locations. The results confirm this hypothesis, with the coefficient on the distance-weighted enforcement index smaller in magnitude and less precise than the main results.

In columns 2–4 of Table 5, we present results using alternative instruments. Columns 2 and 3 use the second lags of the median coyote price and number of apprehensions at the border sector, respectively, to instrument for the first difference of the enforcement index. Each of these measures may better reflect the information used by border sectors when making budget requests than the number of Border Patrol agents. IV estimates using each of these alternate instruments are larger in magnitude than the main results, and are significant at 1 %. An additional concern with the instruments presented thus far is that if unobserved shocks to migrant flows are persistent, then the second lag of each enforcement measure may be correlated with the second-stage error term, making our IV estimates inconsistent. To address this concern, column 4 of Table 5 uses the third lag of border agents as the instrument, which should be less correlated with contemporaneous shocks. The border enforcement coefficient remains precisely estimated and of similar magnitude as the main results.

Another concern about our methodology is the ability of the enforcement index to isolate variation in border policy for a geographic unit as small as a state. This concern is

**Table 5** Robustness checks

| | Distance Weights | Alternate Instruments: Coyote Price | Apprehensions | $t-3$ | Including Enforcement in Other States | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Enforcement Index | –0.145 | –0.220 | –0.408 | –0.238 | –0.200 | –0.373 |
| | (0.075)$^{\dagger}$ | (0.070)** | (0.077)** | (0.072)** | (0.070)** | (0.061)** |
| Neighbor States' Enforcement | | | | | –0.202 | |
| | | | | | (0.092)* | |
| Other States' Enforcement | | | | | | –0.0005 |
| | | | | | | (0.0003) |
| Observations | 840 | 840 | 840 | 792 | 839 | 840 |
| First-Stage *F* Statistic | 118.0 | 37.7 | 32.7 | 57.2 | 77.5 | 31.5 |

*Notes:* The table shows regressions of the log Mexican immigration share on enforcement index, in first differences. The sample is U.S. state-years (including the District of Columbia) from the Current Population Survey, 1995–2011. All regressions are identical to the baseline IV specification, with the following modifications: in column 1, the inverse square root of distance is used as the enforcement index weight (based on the principal city within the border sector and centroid of the U.S. state; the distance to border states is set to 0 if the crossing is within state). In column 2, the instrument is based on the second lag of the median coyote price (in thousand USD, base 1983). In column 3, the instrument is based on the second lag of apprehensions (in thousands). In column 4, agent instrument, $t-3$. Column 5 includes the sum of enforcement in neighboring states, weighted by 1990 population. "Neighboring state" refers to shared border, with the exception of Alaska (WA is defined as neighbor) and Hawaii (CA is defined as neighbor). Column 6 includes the sum of the enforcement index in all other states, weighted by the inverse distance. Weights are normalized to sum to 1. The distance between the states is defined as the distance between centroids. Robust standard errors, shown in parentheses, are clustered by state.

$^{\dagger}p < .10$; $^{*}p < .05$; $^{**}p < .01$

particularly important for interior states: is it really possible to distinguish between border enforcement directed to, say, Georgia and Alabama? We test this possibility by including measures of other states' border enforcement as additional controls in columns 5 and 6 of Table 5. Column 5 includes a population-weighted average of the enforcement index of neighboring states. The coefficient on own-state enforcement shrinks in magnitude slightly to –0.2, but remains precisely estimated. Neighboring states' enforcement also matters to a nearly identical degree, with a coefficient of –0.202, precise at 5 %. The result suggests that migrants consider border enforcement directed to a region, not merely a state, when deciding on location. Column 6 uses the inverse distance-weighted average of all other states' enforcement. The own-state enforcement coefficient grows to –0.373, significant at 1 %, while the coefficient on other states' enforcement is also negative, although not significant.[18]

[18] Online Resource 1 presents results using aggregated data from U.S. Census Department divisions, plus a Mexican border division. Point estimates are substantially smaller in magnitude than their state-level counterparts, suggesting that substitution of migrant destinations within regions in response to enforcement is important. However, 95 % confidence intervals from the state- and division-level analyses generally overlap.

Tables S5–S9 in Online Resource 1 present more robustness checks using additional alternative instruments, enforcement indices, and subsamples. In sum, our main results are highly robust to a variety of concerns.

## Discussion: How Much Does Border Enforcement Matter for Location Choice?

Given the robustness of our results, it would be instructive to determine the extent to which U.S. border enforcement accounted for the spatial diffusion of Mexican immigrants during the sample period. To quantify the effect of border enforcement, we compare actual state shares of the Mexican immigrant population to those implied by our estimates under a counterfactual of no change in enforcement. To calculate these counterfactual population shares, we subtract our baseline estimate of the border enforcement effect (the IV coefficient reported in Table 3, column 1, multiplied by observed changes in border enforcement during the sample period) from actual changes in population shares. Details of the calculation appear in Online Resource 1.

Table 6 presents results of this exercise. Columns 1 and 2 show each state's observed share of the Mexican immigrant population at the beginning and end of the sample period. Column 3 shows the end-period share if border enforcement had not changed over the same period. Taking the first state in the list (Alabama) as an example, we observe that between 1994 and 2011, its share of the Mexican immigrant population rose more than tenfold, from 0.05 % to 0.51 %. This change is indicative of the diffusion of Mexican immigrants to southeastern states. In column 3, we see that our estimates imply that if border enforcement had not changed since 1994, Alabama's share would be only 0.26 %. The corresponding 0.25 percentage point discrepancy reported in column 4 indicates that border enforcement played an important role in the increased presence of Mexican immigrants in Alabama during the sample period.

Similar insights appear throughout Table 6. Of particular note are our estimates for the southern border states. We find that Mexican immigrant shares in California and Texas would be considerably higher if border enforcement had remained static, by more than 8 percentage points in each case. Conversely, immigrant shares in Arizona and New Mexico would be lower, consistent with the Massey et al. (2002) hypothesis of enforcement in high-traffic areas of the border leading to increasing crossing and settlement in border areas with less historical traffic. In fact, our estimates imply that all states would have a lower (or unchanged) share of Mexican immigrants if enforcement had not changed, with the exceptions of California and Texas.

The maps presented in Fig. 4 help to visualize the results presented in Table 6. Panel (a) shows the empirical change in Mexican immigrant shares, and panel (b) presents our estimates from Table 6, column 4. The map shows that the Mexican immigrant population would not have diffused as extensively across the country if enforcement had remained unchanged.[19]

[19] Online Resource 1 presents additional counterfactual results using a range of enforcement coefficient estimates, including using the apprehensions instrument (Table 5, column 3) and halving the preferred point estimate from Table 3, column 1. Although the magnitudes of counterfactual immigrant shares change, the main results remain: California and Texas would have gained migrant share, with all other states losing or experiencing no change.

**Table 6** State shares of Mexican immigrants, 1994–2011 (percentages)

| State | Observed 1994 (1) | Observed 2011 (2) | Counterfactual 2011 (3) | Change (3) – (2) |
|---|---|---|---|---|
| Alabama | 0.05 | 0.51 | 0.26 | –0.25 |
| Alaska | 0.03 | 0.04 | 0.02 | –0.02 |
| Arizona | 4.50 | 5.03 | 4.72 | –0.30 |
| Arkansas | 0.15 | 0.29 | 0.15 | –0.14 |
| California | 54.02 | 38.16 | 46.40 | 8.24 |
| Colorado | 0.57 | 1.78 | 1.02 | –0.76 |
| Connecticut | 0.00 | 0.18 | 0.12 | –0.05 |
| Delaware | 0.06 | 0.19 | 0.10 | –0.09 |
| District of Columbia | 0.02 | 0.06 | 0.03 | –0.03 |
| Florida | 2.17 | 1.95 | 1.09 | –0.86 |
| Georgia | 1.28 | 1.95 | 1.02 | –0.92 |
| Hawaii | 0.01 | 0.08 | 0.04 | –0.04 |
| Idaho | 0.36 | 0.57 | 0.30 | –0.27 |
| Illinois | 6.53 | 5.56 | 3.02 | –2.54 |
| Indiana | 0.09 | 0.82 | 0.42 | –0.40 |
| Iowa | 0.11 | 0.54 | 0.28 | –0.26 |
| Kansas | 0.12 | 0.63 | 0.33 | –0.30 |
| Kentucky | 0.03 | 0.32 | 0.16 | –0.15 |
| Louisiana | 0.07 | 0.21 | 0.11 | –0.10 |
| Maine | 0.00 | 0.01 | 0.00 | –0.01 |
| Maryland | 0.09 | 0.43 | 0.22 | –0.21 |
| Massachusetts | 0.17 | 0.05 | 0.03 | –0.02 |
| Michigan | 0.12 | 0.87 | 0.45 | –0.42 |
| Minnesota | 0.28 | 0.48 | 0.25 | –0.23 |
| Mississippi | 0.03 | 0.12 | 0.06 | –0.06 |
| Missouri | 0.11 | 0.25 | 0.13 | –0.12 |
| Montana | 0.01 | 0.03 | 0.02 | –0.01 |
| Nebraska | 0.13 | 0.54 | 0.28 | –0.26 |
| Nevada | 1.14 | 1.93 | 1.02 | –0.91 |
| New Hampshire | 0.00 | 0.00 | 0.00 | 0.00 |
| New Jersey | 0.37 | 1.80 | 0.92 | –0.88 |
| New Mexico | 1.08 | 0.99 | 0.56 | –0.43 |
| New York | 1.33 | 1.84 | 0.95 | –0.89 |
| North Carolina | 0.53 | 2.03 | 1.09 | –0.94 |
| North Dakota | 0.01 | 0.00 | 0.00 | 0.00 |
| Ohio | 0.11 | 0.34 | 0.18 | –0.17 |
| Oklahoma | 0.56 | 0.44 | 0.23 | –0.21 |
| Oregon | 1.09 | 0.74 | 0.39 | –0.35 |
| Pennsylvania | 0.19 | 0.38 | 0.20 | –0.19 |

**Table 6** (continued)

| State | Observed 1994 (1) | Observed 2011 (2) | Counterfactual 2011 (3) | Change (3) – (2) |
|---|---|---|---|---|
| Rhode Island | 0.02 | 0.04 | 0.02 | –0.02 |
| South Carolina | 0.06 | 0.54 | 0.28 | –0.26 |
| South Dakota | 0.00 | 0.03 | 0.01 | –0.01 |
| Tennessee | 0.03 | 0.65 | 0.34 | –0.32 |
| Texas | 20.85 | 22.45 | 30.63 | 8.17 |
| Utah | 0.32 | 0.64 | 0.33 | –0.31 |
| Vermont | 0.00 | 0.00 | 0.00 | 0.00 |
| Virginia | 0.12 | 0.85 | 0.44 | –0.41 |
| Washington | 0.47 | 1.84 | 0.97 | –0.88 |
| West Virginia | 0.00 | 0.01 | 0.00 | –0.01 |
| Wisconsin | 0.56 | 0.79 | 0.41 | –0.38 |
| Wyoming | 0.05 | 0.05 | 0.03 | –0.02 |

*Notes:* The table shows the state shares of Mexican immigrants in 1994 (column 1), 2011 (column 2), and 2011 (column 3) under the counterfactual in which border enforcement did not change. The final column shows the percentage point change in the state share as a result of enforcement, found by subtracting column 3 from column 2. Column 3 is found by using the estimated IV coefficient on the enforcement index from column 1 of Table 3 to determine the predicted change in the population share in each state resulting from changes in enforcement. This change is then used to predict each state's population share in 2011 as though no change in enforcement occurred. Details are provided in Online Resource 1.

*Data source:* CPS, 1994–2011.

Although the estimates in this section stem from an empirical specification derived from a theory of immigrant location choice, several caveats are in order. First, we do not quantify the effects of border enforcement on aggregate flows between Mexico and the United States, but focus only on the spatial distribution of immigrants across states. Second, the empirical specification embeds policy and economic variables, such as state-level legislation targeted to immigrants and conditions in industries with large concentrations of immigrant workers, that would also likely change in response to any changes in border enforcement. Nonetheless, we think this exercise is instructive to gauge the relative importance of border enforcement in the diffusion of Mexican immigrants to new U.S. destinations in the past two decades.

## Conclusion

To our knowledge, no causal analysis has previously been conducted on the impact of border enforcement on immigrant location choice. Using an instrumental variables approach to overcome the potentially endogenous response of border enforcement to migrant flows, we find evidence that increases in border enforcement decrease the share of Mexican immigrants, on the order of a 21.9 % decrease in share for every 1,000 additional Border Patrol officers. These results are stable across subgroups, with

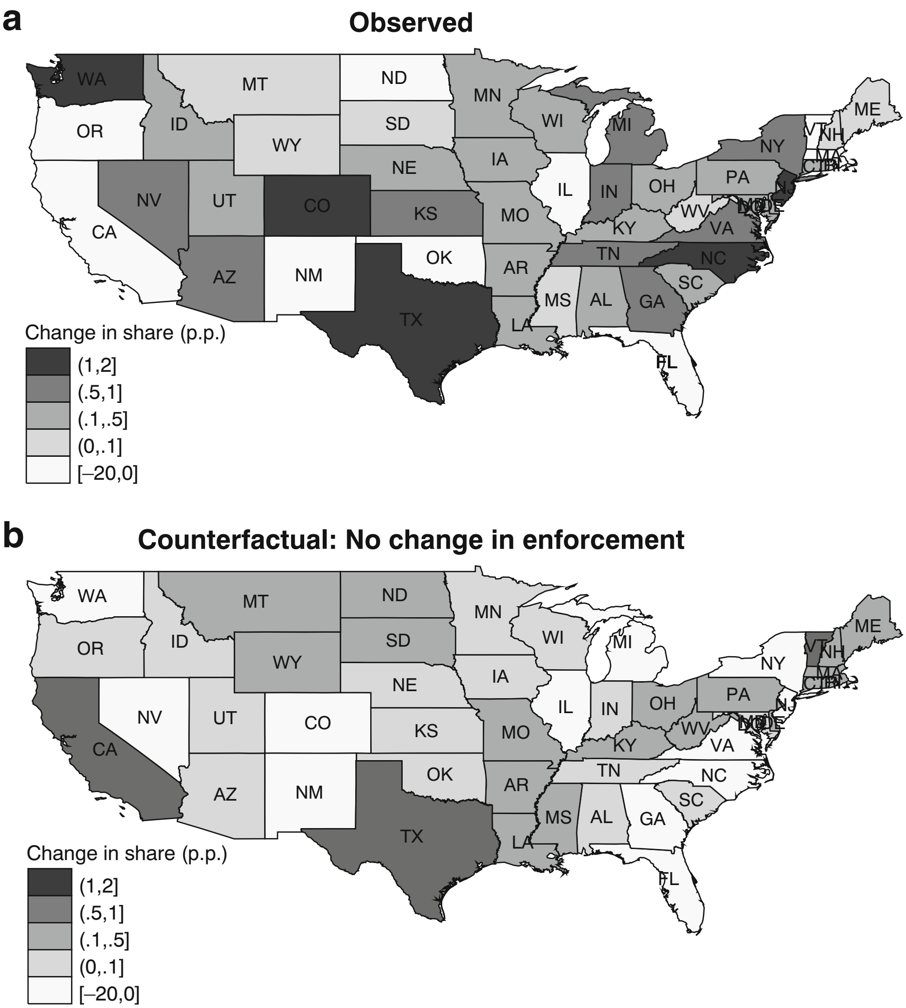


**Fig. 4** Change in state shares of Mexican immigrants, 1994–2011: Observed (panel a) and counterfactual (panel b). The figure groups states by the percentage point change in the Mexican immigrant share, as indicated in the legend. *Source:* CPS

stronger effects for working-age males and noncitizens, and null effects for immigrants less likely to be border crossers. Our estimates imply that California and Texas lost shares of Mexican immigrants to other parts of the country due to border enforcement, consistent with the hypothesis of Massey et al. (2002).

Our data cannot distinguish directly whether border enforcement leads the same individuals to shift their intended destination, or instead leads one group of potential migrants to postpone their journey while another group whose intended destination has lower relative migration costs takes their place. However, important auxiliary evidence for the latter interpretation comes from Borger et al. (2012), who used an identification strategy similar to ours to find that border enforcement, weighted by preexisting migration channels between Mexican states and border sectors, leads potential migrants to remain in Mexico. If the same individuals merely shifted their destination in response to border enforcement, this would not be the case. Disentangling these channels and uncovering their implications for migrant composition is an important topic for future research.

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