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Who Stays? Who Goes? Selective Emigration Among the Foreign-Born

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Abstract

We investigate the level and selectivity of emigration from the United States among foreign-born adults. We use the CPS Matching Method (Van Hook et al. 2006) to estimate the probability of emigration among foreign-born adults aged 18–34, 35–64 and 65+ from 1996 to 2009 (N = 92,852). The results suggest higher levels of emigration than used in the production of official population estimates. Also, indicators of economic integration (home ownership, school enrollment, poverty) and social ties in the U.S. (citizenship, having young children, longer duration in the United States) deter emigration. Conversely, having connections with the sending society, such as living apart from a spouse, was associated with emigration, particularly among Mexican men. Health was least strongly related to emigration. Simulations suggest that selective emigration may alter the home ownership and marital status, but not health, composition of immigrant cohorts. The implications for public policy are discussed.

Keywords

Immigration; Emigration; Foreign-born

Introduction

As many as one-third of U.S. immigrants eventually return home (Duleep 1994), 18% within 5 years and 22% within 10 years of arrival (Borjas and Bratsberg 1996). The magnitude of return migration flows clearly suggests that emigration has the capacity to shape the size and characteristics of the foreign-born population solely through the force of attrition (Jasso and Rosenzweig 1990). Changes in immigrant characteristics with increasing time in the country are frequently interpreted as evidence of assimilation, often as a positive sign of the capacity of immigrants to adapt to the U.S. environment. But such patterns may in fact reflect a harsher reality involving the departure of less successful immigrants.

Beyond its implications for the social and economic integration of immigrant cohorts, the level and selectivity of emigration has enormous policy implications. For example, if many immigrants were in fact temporary residents, this would suggest a mismatch between U.S. immigration policy—which emphasizes permanent settlement—and U.S. demand for temporary workers (Massey et al. 2002; Meissner et al. 2006). High rates of emigration may also slow the pace of economic and social integration into U.S. society. For example, some

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evidence suggests that circular migration among Mexican immigrants disrupts children's schooling (Rendall and Torr 2008). Finally, foreign-born emigration is a critical yet poorly understood component of official population estimates, which in turn affect the distribution of public funds, public planning, and business decision-making, form the basis for sampling weights for surveys such as the Current Population Survey and the American Community Survey, and serve as denominators for fertility and mortality rates.

Theoretical perspectives differ about the causes of emigration and, by implication, the characteristics of emigrants compared with stayers. For example, the neo-classical economics perspective views emigration as a "corrective" process whereby less successful immigrants return home. However, the social capital perspective views emigration more neutrally as immigrants' responses to their rights and obligations within social and family networks (Massey 1999). Still another perspective focuses on life course factors such as retirement and poor health as spurring decisions to emigrate (Duleep 1994; Turra and Elo 2008).

This study investigates the relative importance of the factors emphasized by these three theoretical approaches. Despite the lack of official governmental statistics on emigration (Kraly 1998), some studies have made creative use of longitudinal data or administrative records to assess the economic integration (Borjas 1989; Borjas and Bratsberg 1996; Reagan and Olsen 2000), social capital (Massey et al. 2002; Lindstrom and Massey 1994; Constant and Massey 2003), and health (Turra and Elo 2008) of emigrants. However, these studies have each focused on a relatively narrow set of factors; none examined economic integration, social capital, and health characteristics all together. In addition, prior work has focused on foreign-born subgroups that are less likely to emigrate, such as legal immigrants (Jasso and Rosenzweig 1982; Borjas and Bratsberg 1996), those who arrived in the U.S. as children (Reagan and Olsen 2000), or Social Security recipients (Turra and Elo 2008; Duleep 1994). An examination of a wide range of characteristics among all immigrants is important for evaluating the relative strength of the effects of economic integration, social capital, and health/aging on emigration, and for assessing the impact of selective emigration on the composition of the foreign-born population.

In this paper, we estimate emigration of the foreign-born by using an approach we developed earlier based on attrition from the Current Population Survey (Van Hook et al. 2006). We present estimates of emigration rates of foreign-born adults, estimate multivariate models to assess the independent association of economic-, social-, and health-related factors, and then simulate the effects of emigration on the composition of the foreign-born population for selected characteristics. We find that the relative importance of these factors and the direction of their effects vary considerably by age, gender and Mexican origin.

Background

Estimates of Foreign-Born Emigration

For reasons of convenience, we use the terms "foreign-born population" and "immigrants" interchangeably, with both referring to the US-resident foreign-born population (legal or unauthorized). Official statistics have not been collected on emigration by the U.S. government since 1956 (Kraly 1998). Of necessity then, emigration among the foreign-born has been estimated with a variety of indirect demographic methods, the most prominent of which has been the residual method. The residual method estimates emigration by comparing the size of foreign-born cohorts between two decennial censuses after adjusting for mortality. The U.S. Census Bureau currently uses this method to obtain estimates of foreign-born emigration for the population estimates series (U.S. Census Bureau 2010). Residual-based estimates of the annual emigration rate of foreign-born tend to be close to

1.0% (Warren and Peck 1980—1.2%; Ahmed and Robinson 1994—1.2%; Mulder 2003—0.9%). A major weakness of the residual method is its inability to estimate emigration for recently arrived immigrants—i.e., those who arrived between the two censuses. For this group, the earlier census is not available, so emigration rates are not calculated from the data. Rather, in most cases, immigrants who arrived in the decade before the second census are assigned emigration rates that were calculated for longer-term immigrants.

Alternative methods that estimate emigration for recent arrivals directly yield higher estimates than the residual method. For example, Jasso and Rosenzweig (1982) estimated an annual emigration rate of 2.1% for the 1971 immigrant admission cohort based on follow-up rates in the Alien Address Report Program data. In another study, Duleep (1994) used Social Security administrative records matched across years to estimate emigration of legal immigrants. She found that about 30% of all legal immigrants returned home, and 83% of emigrants left in the first 10 years in the country. This implies an annual emigration rate of 2.8% in the first decade of residence, but less than 1% among more settled immigrants. Borjas and Bratsberg (1996) estimated emigration rates for legal immigrants who were admitted from 1971 to 1986 based on the number enumerated in the 1980 Census after adjusting for mortality and unauthorized migration. They estimated that 17.5% emigrated within 5 years of arrival and 21.5% within 10 years, implying annual emigration rates of 3.8% in the first 5 years and 0.8% in the second 5 years of U.S. residence. Van Hook et al. (2006) estimated foreign-born emigration based on attrition from the 1996–2005 Current Population Surveys (i.e., the CPS Matching Method). This work yielded annual emigration rates of 2.9%, with much higher emigration rates for recently arrived immigrants (5.0 and 3.6% for those with 0–4 and 5–9 years of U.S. residence, respectively) than more settled immigrants (2.0% for those with 10+ years of U.S. residence). Finally, Rendall et al. (2009) used Mexican household survey data (ENOE) to estimate emigration. They found that during 2005/2006, approximately 335,093 Mexican-born returned to Mexico from the United States. This represented about 3.7% of the Mexican born living in the United States at the time (5.5% of males and 1.5% of females). They also found that emigration back to Mexico declined during the late 2000s.

Another factor affecting the magnitude of emigration estimates is whether emigration is defined as including temporary trips home. Massey et al. (2002) estimated extremely high annual rates of emigration among Mexican immigrants between 1965 and 1989—between 12.5 and 29% among recent arrivals. These much higher emigration rates are derived through the analysis of life histories documenting the number of trips to the United States by Mexican migrants who have returned to Mexico. The detail in the data permit the identification of each separate trip as contributing in- and out-migration, and thus may reflect gross exits more than net emigrants over time. In general, studies that measure emigration over small time intervals (such as Massey's work) will be more likely to pick up the multiple trips back home by circular migrants and will yield higher emigration rates than those that measure permanent emigration over longer time intervals (such as the residual method, which often uses ten-year intervals). The CPS matching method, which measures emigration over the period of 1 year, falls between these two extremes.

Economic Characteristics and Emigration

Research on the economic characteristics of emigrants has been guided by neoclassical economic choice theory, which predicts that migration occurs when expected earnings in the destination country exceed earnings in the source country net of migration costs. Return migration decisions are theorized to occur in much the same way, except that earnings in the destination country are known with more certainty. This suggests that less economically successful immigrants (the "mistaken" immigrants) would be most likely to return home. Consistent with this idea, Borjas (1989) modeled outmigration on the basis of sample

attrition from a longitudinal sample of scientists and engineers, finding that the least successful scientists and engineers were the most likely to drop out of the sample, and by inference, leave the United States. Reagan and Olsen (2000) examined attrition due to emigration from the 1979 National Longitudinal Survey of Youth, finding that those with lower potential wages in the U.S. were more likely to emigrate. Massey (1987) and Lindstrom and Massey (1994) similarly found that Mexican emigrants were negatively selected (on wages and education). Similar results have been found outside the United States, including Germany (Constant and Massey 2003; Schmidt 1994), Egypt (Bauer and Gang 1998), and Sweden (Edin et al. 2000). Consistent with the economic choice perspective, the costs of migration may also affect whether a person returns home. In his analysis of the Mexican Migration Project data, Riosmena (2004) found that the probability of return migration has declined in recent years, purportedly because increased enforcement of the U.S. Mexico border has driven up costs of circular migration (Massey et al. 2002; Riosmena 2004).

However, the results of some studies also show that those with higher levels of education are more likely to emigrate than those with lower levels (Jasso and Rosenzweig 1988; Reagan and Olsen 2000), and that emigration decisions are unrelated to the generosity of social welfare benefits in the immigrants' state of residence, even though the availability of welfare income should have reduced the economic incentive to emigrate (Reagan and Olsen 2000). Immigrants may not be sensitive to the level of welfare benefits, particularly if they are in the country to work or attend school (Van Hook and Bean 2009). To help explain the apparent contradictory findings about educational attainment, Borjas and Bratsberg (1996) reasoned that emigrants can be positively or negatively selected depending on the selection that characterized the original migration flow. Their analysis suggested that if the high-skilled are more likely to immigrate, then the lower-skilled among the immigrants are more likely to return; conversely, if the original migration flow is negatively selected, then emigrants are likely to be positively selected. Their explanation is that the economic conditions that would produce a positively selected (negatively selected) immigration flow —i.e., higher earnings for high-skilled (low-skilled) workers in the destination relative to the source country—also discourage return migration of high-skilled (low-skilled) workers. Consistent with this idea, Ramos (1992) found that migrants from Puerto Rico were negatively selected, and that return migrants were the most skilled among the migrants.

Social Capital and Emigration

Emigration is also related to social and family ties in source and destination countries. As noted above, recently arrived immigrants are much more likely to emigrate than more settled immigrants (Reagan and Olsen 2000; Jasso and Rosenzweig 1988; Duleep 1994; Borjas and Bratsberg 1996; Van Hook et al. 2006), suggesting that those who are more socially or economically integrated are less likely to return. As demonstrated in Massey's work on Mexican labor migration to the United States (1987), migration often represents the enactment of household-level strategies to send household members abroad, typically young men (Cerrutti and Massey 2001), to earn money before returning home again. Thus among Mexican labor migrants, both immigration and emigration (i.e., circular migration) tends to be concentrated among young men, particularly those with family ties in Mexico (Massey et al. 2002). These patterns have also been found for non-U.S. labor migration flows. Constant and Massey (2003) found that emigration is positively associated with factors such as having a spouse living in another country, remittances, and duration of German residence. Schmidt (1994) also found that immigrants in Germany were more likely to return home if they had a spouse living abroad. Conversely, Jensen and Pederson (2007) found that family ties in the host society (number of children younger than 18 years and whether the individual has a native-born spouse) lowered the likelihood of emigration of immigrants living in Denmark.

Health and Emigration

In comparison to social and economic factors, much less work has examined how health or mortality is related to emigration. One idea, referred to as the "salmon-bias" hypothesis, is that unhealthy immigrants will be more likely to return home to receive care and eventually die among family and friends. Duleep (1994) further suggests that older immigrants—who are less likely to be healthy than younger immigrants—are likely to emigrate upon retirement. Selective migration has been used as an explanation for immigrant and Hispanic mortality advantages (Markides and Eschbach 2005), particularly at older ages (Liao et al. 1998; Palloni and Arias 2004) and among Mexican immigrants, who are more likely to emigrate than Cubans and Puerto Ricans (Palloni and Arias 2004, but see also Abraido-Lanza et al. 1999). More recent evidence appears to confirm these inferences. For example, Palloni and Arias (2004) examined Mexican data on the health of older former U.S. immigrants (i.e., those who returned to Mexico). They found that former immigrants tended to report worse health than Mexican immigrants living in the U.S. Also, in the most convincing study to date, Turra and Elo (2008) used Social Security records to estimate the risk of mortality among foreign-born Social Security recipients age 65+ who remained in the U.S. and those who emigrated. They found that those who emigrated were at higher risk of dying. However, they also found that emigration was not sufficient to explain the mortality advantage of the foreign-born relative to native-born non-Hispanic whites. Hummer et al. (2007) came to the same conclusion in a study of infant mortality, in which they found immigrant advantages in infant survival even at very young infant ages, such as at 1 day old, when it is nearly impossible for a mother to have emigrated with her child prior to the child's death (although they did not consider the possible effects of migration-related selection bias on maternal health).

Despite the emphasis on salmon bias in the literature on immigrant health, a plausible argument may also be made for the opposite pattern for certain immigrant groups, whereby healthy immigrants are more likely to emigrate. Immigrants who become injured or ill may choose to remain in the country if they have access to U.S. medical care, have family in the United States, or are simply too ill to travel. Consistent with these ideas, residential mobility has been observed in several countries to be much higher among young adults than older people (White and Lindstrom 2006). Similarly, emigration rates appear to be higher among working aged immigrants and taper off at older ages (Ahmed and Robinson 1994; Van Hook et al. 2006). In addition, the relationship between health and duration of U.S. residence is consistent with emigration being higher among healthier immigrants. Cho et al. (2004) found that duration of U.S. residence is associated with worse health among Hispanics (particularly Mexicans). Holding other factors constant, if emigration were associated with poor health, this would result in a pattern in which health would appear to improve, not worsen, with increased residence.

The Current Study

In this study, we examine the associations of various indicators of economic integration, social capital (both in the U.S. and abroad), and health with the annual probability of emigration. Although some emigrants subsequently return to the United States (i.e., they are what some call "circular" migrants or sojourners), we focus here on emigration without attempting to make any adjustments for how emigration might be offset by emigrants who return (like nearly all other studies). We expect to find support for all of the theoretical frameworks described above. Nevertheless, the relative contributions of economic integration, social capital, and health are likely to vary across groups depending on immigrants' motivations for coming to the United States. For example, labor migrants may be especially responsive to economic conditions while family reunification immigrants may be more sensitive to the presence or absence of social networks.

Our data (the Current Population Survey) do not provide information on the motivations for migration or conditions of admission. Nevertheless, we focus on differences by age, gender, and Mexican origin because these factors are likely to be roughly aligned with migration experiences. Younger immigrants, especially men, are more likely to be in the labor force, while immigrant women have been documented as being less likely to be in the labor force than men or native women (Lowell, Gelatt, and Batalova 2006). Mexican women have been observed as being more likely to follow husbands or fathers rather than immigrate for their own work opportunities (Parrado and Flippen 2005; Cerrutti and Massey 2001; Pessar 1999). We therefore expect that younger men will be more likely to respond to factors related to economic integration, while women and older immigrants may be more sensitive to indicators of social capital and health.

We also expect to find differences by Mexican origin. Mexican immigrants have been identified as composed primarily of unauthorized circular labor migrants engaged in household-level income generating strategies (Massey 1987) and are likely to be engaged in low-skilled labor (Bean and Stevens 2003). Therefore, Mexican immigrants may be more responsive to social capital (especially social capital in the country of origin) than non-Mexican immigrants.

Methodology

Data and Sample

We used the 1996–2009 March Current Population Surveys (CPS) to estimate the number and characteristics of foreign-born emigrants for the late 1990s and 2000s. The CPS is a monthly survey of over 50,000 households conducted by the Bureau of the Census for the Bureau of Labor Statistics. The sample is scientifically selected to represent the civilian noninstitutional population. The survey has been conducted for more than 50 years, and serves as the primary source of information on labor force trends in the United States. In addition, the CPS is the only ongoing nationally representative survey with questions on nativity, citizenship, and generational status. The March CPS includes a wide range of socioeconomic and demographic variables such as labor market activity, educational attainment, school enrollment, and self-reported health.

The analytical sample used to estimate the number and characteristics of foreign-born emigrants includes all foreign-born adults in the 1996 through 2008 CPS March samples who were eligible to be followed up in the following year. This means that the sample was restricted to those in months-in-sample 1–4 (or 5 in the case of the oversample). The final sample included 92,852 foreign-born adults age 18+ (44,861 men, 47,991 women). Because emigration has been hypothesized to vary across the life course (Duleep 1994), we conducted all analyses separately for young adults (age 18–34; $N = 33,273$), older working-aged adults (age 35–64; $N = 48,599$) and the elderly (age 65+; $N = 10,980$).

Methodology for Estimating and Modeling Emigration

We used the CPS Matching Method to estimate the annual rate of emigration from the U.S. The methodology is presented in greater detail elsewhere (Van Hook et al. 2006). Like other methods for estimating foreign-born emigration (e.g., Reagan and Olsen 2000), the CPS Matching Method uses the rate of sample attrition (with adjustments) as an indicator of emigration. The CPS is a longitudinal survey. It interviews occupants of housing units for 4 consecutive months, and again for 4 additional consecutive months in the following year. If a CPS respondent moves to a new address, he/she is not followed. Rather, the new occupants of the housing unit are interviewed and the original respondent is dropped from the sample (U.S. Census Bureau 2002).

The key insight is that individuals in the March CPS in one particular year (year t) who do not appear in the following year's March CPS (year t + 1) include those who died, internal migrants (who moved to other residences in the U.S.), emigrants who moved out of the country, and a residual group who cannot be matched for other reasons. The rate of attrition (a) can thus be represented as the sum of the rates of internal migration (m), death (d), emigration (e), and attrition for other reasons (r). For the foreign-born (f), the relationship is:

$$a^f = m^f + d^f + e^f + r^f \quad (1)$$

Rearranging terms, we obtain an expression for the rate of emigration:

$$e^f = a^f - m^f - d^f - r^f \quad (2)$$

Attrition (a) is estimated as the proportion not followed-up among those eligible to be followed (described below). The rate of internal migration (m) is based on the CPS question that asks respondents whether he/she lived in a different residence 1 year before, and is estimated as the proportion of movers among those who reported having lived in the United States 1 year before.

The CPS does not contain specific information about mortality (d) or non-response (r). To estimate the proportion of CPS respondents who died between year t and t + 1 (d), we used data from the National Health Interview Survey or NHIS (Palloni and Arias 2004). NHIS respondents are routinely linked to a data base of all deaths in the United States. We used the 1989–1994 NHIS and linked deaths from 1989–1997 to estimate discrete-time hazard models of mortality as a function of age, race/ethnicity, education, marital status, household composition, and self-reported health status, estimated separately by gender, nativity, and Mexican ethnicity. We used the vector of coefficients from these models to generate a predicted probability of mortality among CPS respondents, which when averaged gives the proportion that is likely to have died.

We used two assumptions to estimate the proportion that failed or refused to respond (r). The first was that foreign-born and second generation adults (U.S.-born children of foreign-born parents) have the same non-follow-up probabilities after adjusting for compositional differences in demographic characteristics. Thus:

$$\begin{aligned} r^f &= r^s \\ &= a^s - m^s - d^s - e^s. \end{aligned} \quad (3)$$

To remove the influence of compositional differences and obtain estimates of a^s, m^s, and d^s, we estimated logistic regression models of attrition, internal migration, and mortality for the second generation, and then used the coefficients from these models to generate predicted probabilities for the foreign-born respondents (i.e., plugging in foreign-born respondents' own characteristics into the second generation prediction equation).[1]

The second assumption was that $e^s = 0$. Prior work suggests that emigration among the second generation may be as low as 0.02% per year (Fernandez 1995), if not lower (Gibbs et al. 2003), and probably no higher than 0.2% per year. If we substitute Eq. 3 into Eq. 2 and assume that $e^s = 0$, we obtain:

[1]The logistic regression models include a wide range of predictor variables, including age, sex, ethnic origin, education, marital status, household composition, year, month-in-sample, self-reported health, home ownership, and poverty status.

$$e^f = a^f - m^f - d^f - (a^s - m^s - d^s). \quad (4)$$

One complication is that our estimates of internal migration are based on a retrospective survey question. The population at risk of internal migration—as measured in the CPS in year t + 1—excludes some who were actually at risk of moving internally in year t such as those who died or emigrated in the previous year. The "true" population at risk of moving internally between t and t + 1 (P_t*) is actually $P_t^* = P_{t+1}/(1 - e - d)$, where e and d are the proportions emigrating and dying, respectively. Accordingly, we adjust the internal migration probability as: m* = m(1 – e – d). For second generation adults, among whom e is assumed to be zero, the adjusted internal migration probability reduces to m* = m(1 – d). After replacing m with m* and rearranging terms, Eq. 4 becomes:

$$e^f = \left[a^f - m^f + m^f d^f - d^f - a^s + m^s - m^s d^s + d^s \right] / \left(1 - m^f\right). \quad (5)$$

To obtain emigration rates for the foreign-born altogether and by subgroups, we estimated the components of Eq. 5 separately (first for all foreign-born and then within each subgroup), and then plugged the averaged values into Eq. 5.

The estimated emigration rates cannot be modeled directly since they are not measured for individuals but only for groups. However, we are able to estimate models of sample attrition (which is observed for individuals) while adjusting for sources of attrition other than emigration. If Y_i is the probability of attrition and G_i is the probability of attrition for reasons other than emigration, then a model of emigration can be written as:

$$Y_i - G_i = \mathbf{X_i'B} + \varepsilon_i$$

where (applying Eq. 4)

$$Y_i = a_i^f$$
$$G_i = \left[m_i^f + d_i^f + \left(a_i^s - m_i^s - d_i^s\right) \right]$$

Note that a_i^f is a binary variable indicating attrition, and m_i^f, d_i^f, a_i^s, m_i^s, and d_i^s are all predicted probabilities (based on prediction models as described above). To estimate this model, we carry G_i over to the right-hand side:

$$Y_i = \mathbf{X_i'B} + G_i + \varepsilon_i$$

We initially estimated this model as a linear probability model (with the coefficient of G_i constrained to one). We also estimated nonlinear models in which we made the adjustments for the retrospective measure of internal migration (applying Eq. 5). In yet another set of analyses, we estimated logistic regression models of attrition while including G_i as a control variable. All sets of models produced nearly equivalent results and are available to interested readers upon request. We opted to present the logistic regression models here because they model proportional rather than absolute differences, thereby ensuring that predicted values fall between zero and one (important since our outcome—emigration—tends to be close to zero for many groups).

Measures

Attrition—To determine whether a respondent in the March Supplement to the CPS in year t was not successfully followed up the following year t + 1, we match those eligible for follow-up in the 1996–2008 March CPSs with respondents in the following years' CPSs, 1997–2009. In general, households from rotation groups 1–4 in each year t are matched to rotation groups 5–8 in the following year t + 1. Then, matching individuals in these households are identified. We use the methodology developed by Madrian and Lefgren (1999) for linking cases across CPS files, matching on household identification number and person line number. Because matched cases may not represent the same individual due to coding errors on the person or household identification variables, we also require consistency in sex and age before considering a case a "true" match.[2] We do not require consistency on race or Hispanic origin because the race question changes in 2003 (allowing responses in multiple categories) and because of response inconsistency and variability.

Economic Characteristics—We estimated rates of emigration for foreign-born adults separately by a variety of commonly used indicators of human capital and economic integration. These included educational attainment (less than high school, high school, some post-secondary schooling, college degree or more), employment status (employed fulltime, full-year), school enrollment (enrolled in high school or college), poverty status (under poverty threshold, 100–150% of poverty threshold, and 150% of poverty threshold or higher), and home ownership. We also examined logged income but later dropped this term because of multicolinearity.

Social and Family Ties—We examined marital status, household size, and the presence of children as indicators of the presence and location of family ties. We distinguished among four marital statuses: married-spouse present, married-spouse absent, formerly married (divorced, separated or widowed), and never married. We expected that those who are married but not living with their spouse and never-married immigrants would be particularly likely to emigrate. We also distinguished among those living with pre-school and/or school-aged children (ages 0–4 and 5–17, respectively). We expected that people with children (especially school-aged children) would be less likely to emigrate.

We also examined the relationship between emigration and correlates of assimilation: naturalization status (i.e., whether the immigrant is a U.S. citizen) and duration of U.S. residence (0–4, 5–9, 10–14, and 15+ years in the country). Although new arrivals with fewer than 5 years of US residence are generally not allowed to naturalize (except veterans and those with citizen spouses), the correlation between duration of residence and citizenship status is only moderately high ($r = 0.5$).

Health—We examined the relationship between emigration and two indicators of health status: Age and self-reported health status. Age is strongly associated with health and the risk of mortality. Self-reported health (SRH) is five-point scale that ranges from excellent to poor. Due to limitations in sample size, we combine the "Very Good" and "Good" responses, and the "Fair" and "Poor" responses into another category. SRH has been found to be a very strong predictor of subsequent mortality risk, net of a number of important social, economic, behavioral, and physician-assessed health control variables (Idler and Benyamini 1997). SRH also predicts subsequent mortality strongly among all major racial/ethnic groups in the United States (McGee et al. 1999), including Hispanics, and for both women and men, although the association is somewhat stronger for men across several causes of death (Benjamins et al. 2004). Because health insurance may offset the tendency

[2]For example, a person at year t can be no more than 2 years younger than the matched case in year t + 1.

for ill persons to return home, we also control for the effects of health insurance coverage on emigration (covered by health insurance versus not).

Time Period—To assess the idea that emigration rates have changed in recent years as a consequence of increased border enforcement or the current recession, we examined the relationship between time period and emigration.

Results

Rates of Emigration

We estimate that foreign-born persons emigrated at an average rate of 2.9% per year between 1995 and 2009 (Table 1). In 2000, when the foreign-born population (age 18+) was estimated at 27 million, this rate would imply that about 846 thousand foreign-born adults leave the United States each year. If a newly arrived immigrant cohort consistently emigrated at a rate of 2.9% per year, one-quarter would have emigrated after 10 years and 36% would have left by 15 years.[3] The estimate of 2.9% is consistent with our earlier-published estimates (Van Hook et al. 2006) and nearly identical to Borjas and Bratsberg's estimate that 22% emigrated after 10 years. In addition the emigration estimates for Mexicans are consistent with the levels estimated by Rendall et al. (2009); they estimate a rate of about 3.7% for all Mexicans, 5.5% for men, and 1.5% for women, while we estimate a rate of about 3.0% for all, 4.3% for men, and 1.6% for women. Our estimates tend to be higher than most residual estimates, which generally fall around 1.0%. As noted earlier, this discrepancy may be attributable to the detail of the time interval under study. While most residual estimates assess emigration over the period of 5 or 10 years, our method measures emigration over the period of 1 year, which is more likely to capture the multiple trips home of circular migrants. Thus, although 846 thousand may leave each year, a portion of these people subsequently returned to the United States. Our other work on this (published elsewhere) suggests that the rate of returning emigrants is nevertheless quite small, amounting to about less than 1% per year among all foreign-born.

In general, emigration rates decline with age from 4.2% among adults aged 18–34 to 2.3% among adults older than 35. Among adults aged 18–34, Mexican-origin immigrants were more likely to emigrate than non-Mexican immigrants, and among Mexicans, men were more likely to emigrate than women. Among older adults aged 35–64, non-Mexicans were more likely to emigrate than Mexicans, and among the elderly aged 65+, women were more likely to emigrate than men. Emigration is particularly high among Mexican women aged 65+, but this may not be reliable given the smaller sample sizes among elderly Mexicans.

Selective Emigration

To assess the degree and patterns of emigration selectivity, we estimated logistic regression models of attrition while controlling for the probability of attrition for reasons other than emigration. We estimated the models separately by age group (18–34, 35–64, and 65+) and by Mexican origin and gender among adults aged 18–64. The results are shown in Tables 2 and 3.

The neo-classical economic perspective predicts that those who have made fewer economic investments in the U.S. economy or who are less likely to be successful will be the most likely to return home. However, educational attainment may also be associated with emigration if opportunities for high skilled labor are better in the sending country. The results generally support this perspective. Among working-aged adults aged 35–64,

[3]The proportion of a cohort that emigrated (p_x) after x years with an annual emigration rate of e is: $p_x = 1 - (1 - e)^x$.

educational attainment was positively associated with emigration. Among working aged immigrants, other indicators of economic integration (school enrollment (age 18–34 only), homeownership, having an income above the poverty threshold (age 18–34 only) and cash welfare recipiency tended to be negatively associated with emigration. Among middle aged adults (35–64), being unemployed or not in the labor force was also associated with emigration (hence, the return of the less successful). As expected, economic integration in the labor force did not appear to be as strongly associated with emigration among older adults compared with younger adults, and worked in the opposite direction in the case of some factors. For example, home ownership was associated with emigration of elderly adults, perhaps signifying seasonal emigration among retirees (i.e., snowbirds).

The social network perspective predicts that social integration in the receiving society will deter emigration while the maintenance of ties in the sending country will increase it. Indeed, indicators of family and social integration were strongly associated with emigration in all models, although again, the relationships tended to be stronger and more often in the predicted directions among younger than older adults. For those aged 18–34 and 35–64, unmarried immigrants and married immigrants who were not living with their spouse were more likely to emigrate than married (spouse present) immigrants, but among older immigrants aged 65+, married persons were more likely to emigrate. Similarly, younger immigrants living in households with children were less likely to emigrate, but elderly persons living in a household with children (possibly their grandchildren) were more likely. Finally, duration of U.S. residence was negatively associated with emigration, but these effects were much stronger and more significant among younger than older immigrants.

The life course/health perspective predicts that aging and poor health will increase the likelihood of emigration, except perhaps in the case of younger labor migrants, among whom poor health may actually deter emigration. The results were generally consistent with these ideas. For younger immigrants, age and worse health were both negatively associated with emigration even after controlling for health insurance coverage. Among older adults aged 65+, age was positively associated with emigration, but those reporting “fair” or “poor” health were more likely to emigrate than those in “excellent” health.

The results also suggested that the probability of emigration increased between 2002 and 2006 before declining in 2008 back to pre-2000 levels. This may seem inconsistent with Riosmena’s (2004) conclusion that emigration rates declined in recent years due to increases border enforcement, but it is important to note that Riosmena (2004) examined emigration trends from 1980 to 1996, and we examined emigration from 1996 through 2005. Moreover, our results are consistent with research on current emigration trends. At least among Mexicans, return migration declined around 2007/2008 (Rendall et al. 2009). Perhaps circular migration (and thus emigration) increased despite greater border enforcement during the economic boom of the mid-2000 decade. Construction workers were in high demand during this time period before the crash of the housing market in 2007.

Which factors were most important? To answer this question, we examined the decrease in the deviance statistic, or −2 logged likelihood (−2LL), between the null model (with only the intercept and the control for attrition for reasons other than emigration) and the full model (the models shown in Table 2). We then compared this with the reduction in deviance obtained in models that exclude the blocks of economic, social, and health factors. The results suggest that economic integration and social capital indicators were important among working aged adults (18–34 and 35–64), accounting for 21–24% of reduction in −2LL in the case of the economic factors, and 31–35% in the case of the social factors. Other variables (year, gender, and Mexican origin) were also important, together accounting for 38–42%. Finally, the health/aging indicators were the least important, explaining 3–6%. For older

ages, the contributions of economic and social factors were less (<20%) but health and aging were much more important (34%).

Do the factors affecting emigration vary by gender and Mexican origin? We estimated the models separately by age, gender, and Mexican origin. Due to sample size limitations, we focus here on the emigration models for working aged adults aged 18–64 (Table 3), although models estimated on the elderly population are available from the authors for interested readers.

Three notable differences emerged. First, the associations of some indicators of economic integration with emigration were weaker or operated in the opposite direction than expected among Mexican immigrants. Education was not strongly associated with emigration for Mexicans or non-Mexican men. Also, among Mexican men, employment, the income-to-poverty ratio, and welfare receipt were unrelated to emigration. Overall, the contributions of economic factors to emigration was similar across groups, but the social and family variables tended to account for greater reductions in the −2LL among Mexicans than non-Mexicans.

Second, we found gender differences in the strength of the effects of marital status and the presence of children. Men who were unmarried or married but not living with their spouse were more likely to emigrate than married men living with their wives. These effects were similar but weaker for women. Conversely, children deterred emigration for women but not men. This suggests that men may be more likely to return to their home country if they are separated from their families or wish to find a spouse, while women are more likely to remain in the U.S. to form families here.

Third, health was associated with emigration only among non-Mexicans. For all groups, age was negatively associated with emigration, but poor health was associated with lower levels of emigration for non-Mexicans only.

The Magnitude of Estimated Emigration Bias

The results suggest that emigration does not occur randomly. But what impact might emigration have on the foreign-born population? Are the differences in emigration across outcomes like home ownership or health large enough to significantly alter the composition of the foreign-born population? And, are they large enough to change the association between duration of residence and the outcome?

To help answer these questions, we simulated the effects of emigration selectivity on the home ownership, marital status, and health composition of hypothetical immigrant cohorts. We selected these three outcomes because of their substantive significance as indicators of economic, cultural, and health incorporation. In the simulations, we gave newly arrived immigrants the composition we observed in the CPS for new arrivals. For example, about 20% of new arrivals (in the United States less than 2 years) aged 18–34 reported owning their home, so we assigned 20% of the hypothetical cohort to be home owners at the time of entry. This initial percentage remained fixed for the cohort over time and no deaths occurred. Rather, the composition of those in the cohort who remained in the host society (i.e., those who did not emigrate) shifted over time due solely to selective emigration. We subjected the cohort to emigration rates derived from the estimated models. For example, we estimated that homeowners are about 61% as likely to emigrate as renters. If the overall emigration were 1%, an odds ratio of .61 would imply that homeowners emigrate at a rate of 0.66% versus 1.08% among renters.[4] If the overall emigration rate were higher—say 7%—the emigration rates for home owners and renters would then be 4.63 and 7.59%, respectively.

The results of the simulation are shown in Table 4. One of the largest odds ratios was for home ownership: home owners were 61% as likely (39% less likely) to emigrate as renters. If emigration rates were as low as 1%, emigration would have very little impact on the percentage of home owners among immigrants remaining in the U.S. The percentage would increase from 20 to 21% over 15 years on account of differential emigration, and the overall percentage for all durations of residence combined would be 20.7—not much different from the true level of 20%. However, if emigration were as high as 7% (the rate observed among Mexican men), emigration bias becomes more apparent. Home ownership would increase from 20 to 29% over 15 years solely on account of emigration, and the overall level would reach 24%.

The association of marital status with emigration was also large, particularly for working-aged adults. For example, formerly married adults aged 18–34 were 63% more likely to emigrate as married (spouse present) adults. If emigration rates were about 3%, the percentage formerly married (primarily divorced or separated) among ever-married adults would decrease from 9.3 to 7.2% over 15 years on account of emigration. If emigration were as high as 7%, it would decline to 4.9%, and the overall percentage for all durations of residence combined would be 7.3%, roughly one-fifth less than its true value.

Most of the odds ratios were smaller than those estimated for home ownership and marital status. For example, health was associated with emigration among immigrants aged 18–34, but the effect was small. Those in fair or poor health were 91% as likely to emigrate as those in excellent health. Not surprisingly, the results of the simulation suggest that this difference is not large enough to significantly impact the health composition of the immigrant population. Even if emigration rates were as high as 7%, the percentage in fair or poor health would increase by only half a percentage point (from 5.3 to 5.8%) over 15 years. We also found evidence of salmon bias among the elderly. In this case, the impact of emigration was a bit larger but still small. If emigration were as high as 3% (probably the upper limit on emigration among the elderly), the percentage in fair or poor health would decline from 40 to 38% over 15 years, and the overall percentage across all durations would be 39%–very nearly the same as the true 40%.

As a final note, it is important to keep in mind that these results probably represent “worst-case” estimates of emigration bias since we do not take into account the degree to which emigration is balanced out by the subsequent returns of emigrants back to the United States. The simulations thus suggest that emigration is unlikely to significantly bias estimates of immigrant health, particularly among younger immigrants, but may bias estimates of other characteristics such as home ownership and marital status.

Discussion

Very little is known about the number and characteristics of emigrants from the United States. This study used the CPS Matching method for estimating emigration to provide a first-time comprehensive look at the factors affecting foreign-born emigration. The social capital perspective was most consistently supported, the neoclassical economic perspective received mixed support, and the health perspective was least important for understanding emigration.

The expectations of social capital theory were consistently supported, particularly among Mexican men. It is noteworthy that Mexican working-aged men were particularly

[4]These estimates satisfy two constraints: that 0.66/1.08 = 0.61, and that the weighted average emigration rate = 1% (0.2 * 0.66 + 0.8 * 1.08 = 1).

unresponsive to many economic indicators, such as employment status and poverty, but very responsive to the presence or absence of family in the U.S. and (probably) abroad. This is consistent with the large body of evidence that emerged from the Mexican Migration Project concerning the importance of social and family networks for maintaining migration flows. Emigration patterns among Mexican men also provide an example of how migration decisions may be shaped more by social networks than economic considerations of individual migrants themselves. But the results also suggest that the results from the Mexican Migration project may not be readily generalized to other immigrant groups or to Mexican women. Mexicans are far more likely to be unauthorized than non-Mexicans, which may help explain why Mexican emigration is more strongly related to family connections than to economic factors such as unemployment. Unauthorized immigrants must depend on family and friends to survive, and they may be reluctant to return home during an unemployment spell because of sunk costs incurred from crossing the U.S.–Mexico border. The migration process may be different for Mexican women too. Unmarried Mexican women were less likely to emigrate than married women, while unmarried Mexican men were more likely. If unmarried Mexican women immigrate to the U.S., this may mark a departure for their family from practices of circular labor migration and movement toward permanent settlement in the U.S (Massey et al. 2002).

The economic perspective received mixed support. In favor of the economic perspective, economic investments–either in human capital (school enrollment) or in real estate—discouraged return migration, particularly among working-aged immigrants. Also, those receiving cash welfare were less likely to emigrate. In contrast with the economic perspective, however, emigration was not always associated with being economically unsuccessful. For example, unemployment was unrelated to emigration among Mexicans (men and women) and non-Mexican women. This is consistent with other studies and our own results showing no increases (and even a decline) in Mexican emigration during the current recession in the late 2000s (Rendall et al. 2009). Also, emigrants sometimes were positively selected on education (in the case of non-Mexican women). These mixed effects could help explain why prior work has not found strong evidence of emigration-related bias on immigrant earnings trajectories (e.g., Duleep and Regets 1997; Duleep and Dowan 2002; Hu 2000; Constant and Massey 2003).

Emigration was less strongly associated with health and aging than with economic integration and social capital. We found that among the elderly, the unhealthy were more likely to emigrate than the healthy, but the effects tended to be small. Moreover, the opposite pattern, whereby healthy people are more likely to emigrate, emerged among younger immigrants. This is noteworthy because much of the literature on immigrant and Hispanic health and mortality has assumed that unhealthy immigrants would be more likely to emigrate. As far as we know, the opposite has not been observed before, so we are cautious in our interpretations. Nevertheless, the finding is consistent with the general fact that young (probably healthy) working-aged adults are more residentially mobile than other demographic groups (White and Lindstrom 2006).

Our overarching conclusion is that emigration is not a random process. This means that emigration has the capacity to shape the characteristics of immigrant cohorts through the force of attrition. Prior research has shown a strong relationship between duration of U.S. residence and home ownership (Myers 2006). But we show here that emigration is strongly negatively associated with home ownership, and our simulation suggests that at least some of the increase in home ownership among immigrants is due to departures by non-homeowners rather than increases in the percentage that purchase homes. In addition, selective emigration could help explain the high levels of marriage and low levels of divorce commonly observed among immigrants. However, selective emigration is unlikely to have

much impact on the health composition of the foreign-born population living in the United States. This is consistent with other research showing that health-selection on emigration—while it may exist—is not large enough to explain the immigrant health paradox (Hummer et al. 2007; Turra and Elo 2008). These results lend researchers some assurance that health patterns by duration of residence reflect changes in health for individuals (perhaps due to acculturation, “weathering”, or some combination).

One policy implication of this research relates to the production and usage of population estimates. The population estimates produced by the U.S. Census Bureau depend in part on residual estimates of foreign-born emigration (U.S. Census Bureau 2010), which are likely to miss the higher emigration levels of recent arrivals and may fail to pick up variations by age, sex, and national origin, particularly among recently arrived immigrants. This could introduce error in the population estimates series, particularly for areas with large numbers of recently arrived immigrants. We caution, however, that even if emigration rates are in fact higher than assumed, this does not necessarily mean that official statistics overestimate the foreign-born stock. Demographic research consistently shows that the foreign-born population is growing rapidly despite relatively high levels of emigration. This points to the possibility that both in- and out-migration may be underestimated, meaning that the share of recently arrived immigrants (composed primarily of temporary and circular migrants) may be higher than what appears in official statistics. More research on in-migration is necessary to assess this idea.

Another implication concerns temporary labor migration policy. The results reinforce the depiction of Mexican male migration as a highly temporary and circular migration flow that is more responsive to the presence or absence of family/ kinship ties than to short-term unemployment spells. Perhaps most remarkable is that emigration back to Mexico did not increase during the late 2000s economic downturn. Some observers would argue that this suggests the need for a legal temporary worker program that is permitted to expand and contract with the demands of the economy (Hanson 2007; Massey et al. 2002). If people were permitted to freely cross the U.S.-Mexico border, temporary migrants may be more willing to go home during periods of slow economic growth. However, other observers may interpret the high rates of emigration among Mexican immigrants as a reason to craft immigration admission and integration policies that encourage the settlement and successful integration of immigrant families and their children. In future research, it would be valuable to estimate emigration levels by occupation and education. This would help tease apart the types of labor migrants that are most likely to return home versus settle in the United States, and this information could be used to help guide the expansion of temporary labor migration programs.

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Table 1

Annual emigration of the foreign-born by Mexican origin, gender, and age group, 1996–2009

	All Ages	Age 18–34	Age 35–64	Age 65+
Annual rates of emigration				
Mexican men	4.3	7	2.1	0
Mexican women	1.6	3.8	0	4
Non-Mexican men	2.8	3.2	2.9	1.7
Non-Mexican women	2.8	3.7	2.5	2.9
All foreign-born	2.9	4.2	2.3	2.3
Annual number of emigrants (in 1,000s, based on 2000 population of 27 million)				
Mexican men	205	167	38	0
Mexican women	73	65	0	8
Non-Mexican men	269	100	150	20
Non-Mexican women	298	140	120	38
All foreign-born	846	472	308	66
Sample size for emigration rates				
Mexican men	14,644	7,520	6,516	608
Mexican women	12,811	6,081	5,972	758
Non-Mexican men	30,217	9,599	16,749	3,869
Non-Mexican women	35,180	10,073	19,362	5,745
All foreign-born	92,852	33,273	48,599	10,980

Table 2

Models of emigration by age (odds ratios)

	Age 18–34	Age 35–64	Age 65+
Educ. and econ. factors			
High school (vs. < HS)	0.96	1.13***	1
Some college (vs. < HS)	1.02	1.18***	0.93
College+ (vs. < HS)	1.07	1.20***	0.94
Enrolled in school	0.75***		
Homeowner	0.60***	0.66***	1.60***
Unemployed (vs. employed)	1.02	1.22***	0.61
Not in LF (vs. employed)	0.94	1.12***	0.88
100–150% of poverty (vs. poverty)	0.86**	0.95	1.11
150+% of poverty (vs. poverty)	0.84***	1.03	1.13
Cash welfare or SSI recipient	0.82	0.68***	0.74***
Family and social networks			
Mar., spouse absent (vs. mar, sp present)	1.35***	1.69***	0.58***
Formerly married (vs. mar, sp present)	1.63***	1.36***	0.79***
Never married (vs. mar, sp present)	1.18***	1.27***	0.81
# Adults in HH	0.99	1.01	1.02
Children ages 0–4 (vs. no children)	0.84***	1.09+	0.81
Children 5–17 (vs. no children)	0.77***	0.90***	1.34**
Children 0–4 and 5–17 (vs. no children)	0.82***	0.92*	1.46*
In U.S. 5–9 years (vs. 0–4)	0.76***	0.86**	0.91
In U.S. 10–14 years (vs. 0–4)	0.67***	0.70***	0.82
In U.S. 15+ years (vs. 0–4)	0.71***	0.71***	0.75*
Citizen of the U.S.	0.97	0.95	0.89
Health and aging			
Age	0.97***	0.99***	1.04***
“Very good/good” health (vs. “excellent”)	0.94*	0.96	1.02
“Fair/poor” health (vs. “excellent”)	0.91	0.98	1.22*
Covered by health insurance	1	0.97	1
Other			
1998/1999 (vs. 96/97)	0.94	1.03	0.99
2000/2001 (vs. 96/97)	0.94	0.98	1.62***
2002/2003 (vs. 96/97)	1.35***	1.61***	1.64***
2004/2005 (vs. 96/97)	1.44***	1.95***	1.40**
2006/2007 (vs. 96/97)	1.33***	1.60***	1.78***
2008 (vs. 96/97)	1.13+	0.99	1.64***

	Age 18–34	Age 35–64	Age 65+
Mexican female	0.88**	0.86**	1.47**
Non-Mexican male	0.82***	0.95	1.13
Non-Mexican female	0.82***	0.87	1.34
Intercept	1.90***	0.44***	0.00***
N	33,273	48,599	10,980
−2LL (null model)[a]	44,208	58,714	13,057
−2LL (full model)	42,909	57,426	12,618
Difference	−1,299	−1,288	−438
Proportion of difference accounted for by:			
Education and economic factors	0.21	0.239	0.173
Family and social network factors	0.309	0.346	0.199
Health and aging factors	0.059	0.035	0.342
Other/residual factors	0.422	0.38	0.286

[a]Null model includes only the intercept and the propensity for non-follow-up

*$p < .05$;

**$p < .01$;

***$p < .001$

Table 3

Models of emigration by Mexican origin and gender, age 18–64 (odds ratios)

	Mexican men	Mexican women	Non-Mexican men	Non-Mexican women
Educ. and econ. factors				
High school (vs. < HS)	1.13*	1.01	0.94	1.17**
Some college (vs. < HS)	0.9	1.08	1.01	1.26***
College+ (vs. < HS)	1.15	1.21	1.05	1.26***
Enrolled in school	0.59***	0.63**	0.88	0.77**
Homeowner	0.63***	0.72***	0.59***	0.67***
Unemployed (vs. employed)	0.99	1.03	1.27**	1.07
Not in LF (vs. employed)	1.01	0.84***	1.13*	1.04
100–150% of poverty (vs. poverty)	0.89	0.92	0.83**	0.91
150+% of poverty (vs. poverty)	1	0.84**	0.88*	0.90*
Cash welfare or SSI recipient	0.72	0.98	0.61***	0.70***
Family and social networks				
Mar., sp. absent (vs. mar, sp present)	1.44**	1.01	1.85***	1.28**
Formerly married (vs. mar, sp present)	1.39**	1.38***	1.53***	1.25***
Never married (vs. mar, sp present)	1.23***	1.03	1.31***	1.16**
# Adults in HH	1.02	1.05	0.98	0.98
Children ages 0–4 (vs. no children)	0.88	0.79**	1.12*	0.86**
Children 5–17 (vs. no children)	0.85**	0.68***	0.98	0.85***
Children 0–4 and 5–17 (vs. no children)	0.88	0.71***	1.01	0.87**
In U.S. 5–9 years (vs. 0–4)	0.71***	0.73***	0.83***	0.83***
In U.S. 10–14 years (vs. 0–4)	0.58***	0.67***	0.69***	0.73***
In U.S. 15+ years (vs. 0–4)	0.59***	0.65***	0.75***	0.72***
Citizen of the U.S.	0.93	0.98	0.93	0.97
Health and aging				
Age	0.99***	0.98***	0.99**	0.99***
“Very good/good” health (vs. “excellent”)	0.98	1.03	0.93*	0.93*
“Fair/poor” health (vs. “excellent”)	1.04	1.05	0.93	0.94
Covered by health insurance	0.94	1.07	0.97	1.03
Other				
1998/1999 (vs. 96/97)	10	0.88	1.05	0.99
2000/2001 (vs. 96/97)	0.98	0.93	1.02	0.89
2002/2003 (vs. 96/97)	1.20*	1.23**	1.62***	1.56***
2004/2005 (vs. 96/97)	1.43***	1.51***	1.77***	1.78***
2006/2007 (vs. 96/97)	1.44***	1.29**	1.53***	1.45***
2008 (vs. 96/97)	1.08	1.21	1.11	0.89

	Mexican men	Mexican women	Non-Mexican men	Non-Mexican women
Intercept	0.74	0.79	0.66*	0.46***
Number of observations	14,036	12,053	26,348	29,435
−2LL (null model)[a]	17,465	15,248	33,630	36,642
−2LL (full model)	17,003	14,898	32,682	35,739
Difference	−462	−350	−948	−903
Proportion of difference accounted for by:				
Education and economic factors	0.307	0.241	0.292	0.22
Family and social network factors	0.376	0.447	0.288	0.205
Health and aging factors	0.078	0.166	0.024	0.043
Other/residual factors	0.24	0.145	0.396	0.532

[a]Null model includes only the intercept and the propensity for non-follow-up

*p <.05;

**p <.01;

***p < .001

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Table 4

Simulated effects of selective emigration on immigrants' characteristics by years in the United States

	Emigration rate				
	0%	**1%**	**3%**	**5%**	**7%**
% Homeowner, age 18–34 (odds ratio = 0.61)					
Initial value (0 years in U.S.)	20	20	20	20	20
5 years in U.S.	20	20.3	21.1	21.8	22.6
10 years in U.S.	20	20.7	22.2	23.8	25.5
15 years in U.S.	20	21	23.3	25.8	28.6
Total population	20	20.7	21.8	22.8	23.6
% Formerly married among ever-married, ages 18–34 (odds ratio = 1.63)					
Initial value (0 years in U.S.)	9.3	9.3	9.3	9.3	9.3
5 years in U.S.	9.3	9.1	8.6	8.1	7.6
10 years in U.S.	9.3	8.8	7.9	7	6.1
15 years in U.S.	9.3	8.6	7.2	6	4.9
Total population	9.3	8.9	8.1	7.6	7.3
% Unhealthy, age 18–34 (odds ratio = 0.91)					
Initial value (0 years in U.S.)	5.3	5.3	5.3	5.3	5.3
5 years in U.S.	5.3	5.3	5.3	5.4	5.4
10 years in U.S.	5.3	5.3	5.4	5.5	5.6
15 years in U.S.	5.3	5.3	5.5	5.6	5.8
Total population	5.3	5.3	5.4	5.5	5.5
% Unhealthy, age 65+ (odds ratio = 1.22)					
Initial value (0 years in U.S.)	40	40	40	40	40
5 years in U.S.	40	39.8	39.3	38.7	38.2
10 years in U.S.	40	39.5	38.5	37.5	36.4
15 years in U.S.	40	39.3	37.8	36.2	34.7
Total population	40	39.5	38.7	38.1	37.6