FOREIGN-BORN EMIGRATION: NEW APPROACHES AND ESTIMATES BASED ON MATCHED CURRENT POPULATION SURVEY (CPS) FILES

by

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Abstract

Postcensal population estimates depend on the accuracy with which the components of demographic change are measured. Of the components, emigration, especially of the foreign-born, has proved to be extremely difficult to gauge. Without "direct" methods (i.e., those identifying individuals who emigrate), demographers have come to rely on indirect methods, such as the residual method. The residual estimates are extremely sensitive to underlying assumptions and are particularly ill-suited for measuring emigration of recent arrivals.

Here, we introduce a new method for estimating emigration that takes advantage of the unique sample design of the Current Population Survey (CPS)—which conducts a series of interviews in the same housing units over a period of 16 months. Individuals appearing in one March Supplement to the CPS but not the next include those who died in the intervening year, those who moved within the country, and those who emigrated. We use statistical methods to estimate the proportion of emigrants among those not followed up. Our method produces emigration estimates that are comparable to those from residual methods in the case of longer-term residents (immigrants who arrived more than 10 years ago), but yields higher—and what appear to be more accurate—estimates for recent arrivals. Although limited by sample size, we produce emigration estimates by age, sex, region of birth, duration of residence, and legal status.

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INTRODUCTION

Postcensal population estimates depend on the accuracy with which the components of demographic change are measured. Of the components, emigration, especially of the foreign-born, has proved to be extremely difficult to gauge. In general, population estimates built up through the cohort component method will be too low if emigration is overestimated and too high if emigration is underestimated. In the case of residual estimates of unauthorized migrants, the accuracy of emigration rates among the legal foreign-born is critical and the estimates of unauthorized migrants vary directly with inaccuracies in measurement of emigration. If emigration estimates of the legal foreign-born are too high, the estimated legal population will be too low and the unauthorized migrant population will be overestimated.

Official statistics on emigration from the United States are virtually non-existent. The Immigration and Naturalization Service (INS) counted departing foreign-born emigrants from 1908 to 1957 (Woodrow-Lafield 1998), but eventually discontinued this practice due to concerns about the quality of the resulting data (Kraly 1998). In contrast, the number of births, deaths, and arrivals of legal immigrants are known with virtual certainty because the United States vital registration system and the former INS (now "U.S. Citizenship and Immigration Services") are required by law to count these events and collect other data on them. Other direct methods for measuring emigration, such as multiplicity surveys attempting to identify emigrants by interviewing their relatives in the United States (Passel and Woodrow 1989) and use of administrative records (Duleep 1994), have met with, at best, limited success.

Out of necessity then, emigration has been estimated with a variety of indirect demographic methods, the most prominent of which is the residual method. We develop below an alternative method for estimating emigration that takes advantage of the longitudinal nature of the Current Population Survey. By producing new foreign-born emigration estimates with which residual estimates can be compared, we offer a new way to evaluate and update the emigration estimates last used by the Census Bureau and others for the production of population estimates.

PREVIOUS RESEARCH

Warren and Peck (1980) first developed indirect methods for estimating emigration that occurred during the 1960s. Their technique—referred to here as the residual method—has since served as the major approach used by the U.S. Census Bureau for developing and updating foreign-born emigration estimates. Warren and Peck (1980) initially estimated annual emigration of 114,000 for the 1960–70 decade. This figure was later increased to 133,000 on the basis of Warren and Passel's (1987) analysis of INS Alien Address Registration data for 1965-80. This figure was used as an official annual "point estimate" by the U.S. Census Bureau until the mid-1990s, when the number was increased to 195,000 based on residual estimates for the 1980s developed by Ahmed and Robinson (1994). The only estimates for the 1990s were developed by Mulder (Mulder et al. 2002; Mulder 2003), which showed an estimated annual foreign-born emigration of 225,000. Mulder's estimates were never used as official Census Bureau estimates, however; the Census Bureau continues to rely on Ahmed and Robinson's (1994) work.

The residual method for estimating emigration in the decade between two censuses involves the comparison of two population figures: (1) the "expected" population if no emigration had occurred during the decade, and (2) the resident population at the end of the decade. For example, Mulder's (2003) residual estimates of emigration during the 1990s were constructed by surviving immigrants who arrived prior to 1990, based on the 1990 Census, forward to 2000 (i.e., by aging all cohorts by ten years and subtracting the estimated numbers of deaths) and then comparing the survived population with pre-1990 arrivals in 2000 based on Census 2000. When the former estimate is larger than the latter, the difference is attributed to emigration. The number of emigrants among those arriving between 1990 and 2000 is estimated by applying emigration rates derived from the analysis of earlier arrivals.

Previous emigration estimates based on the residual method as well as other methods have been reviewed in detail elsewhere (Kraly 1998; Woodrow-Lafield 1998; Mulder 2003). To help summarize this literature, we compare the results of various studies in Table 1 showing estimates of the annual number of emigrants together with emigration rates and the rate at which immigration is offset by emigration as would be implied by the various estimates (Table 1). Even though the residual estimates show increases in the annual number of emigrants over time—from 114,000 to 195,000 and then to 225,000, the associated *rates* of emigration appear to have declined from 1.18 percent during the 1960s to 1.15 in the 1980s and then to 0.88 in the 1990s.[1] Similarly, the ratio of emigrants during the decade to immigrants during the period appears to have declined.

The remaining estimates in Table 1 show less consistency across studies, especially for the recently-arrived immigrants and for those from Mexico. Mulder's (2003) estimate of 21,000 emigrants per year among 1990s arrivals implies a much lower emigration rate than prior studies. Emigration rates for Mexican immigrants show even more variability, although all of the residual estimates show much lower levels of emigration than the estimates made by Massey and his colleagues (Massey and Singer 1995; Massey et al. 2002). These much higher measures of emigration 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. The discrepancy between the residual estimates and those made by Massey demonstrates the importance of distinguishing between net outflows (which is measured by the residual estimates) and gross flows (measured by the estimates of the number of trips out of the United States in Massey's work). We discuss this issue in further detail below.

[1] Rates are calculated as emigrants divided by the average population exposed to the risk of emigrating, or the mid-period foreign-born population.

LIMITATIONS OF THE RESIDUAL APPROACH

One major drawback of the residual method is that the estimates are sensitive to differences in census coverage between the two censuses. For example, net undercounts were higher in the 1990 Census than in Census 2000 (Robinson et al. 1993; Hogan 1993; U.S. Census Bureau 2001), so in many cases the “expected” populations in 2000 turned out to be smaller than the enumerated populations in 2000 and thus implied a negative emigration rate (an impossibility). This problem is especially evident for country-of-origin groups that contain large proportions of unauthorized migrants, such as Mexicans. For Mexicans, the “expected” population is significantly lower than the enumerated population for both 1980–1990 and 1990–2000 entrants (Mulder 2003). Ahmed and Robinson (1994) found similar problems in their analysis of the 1980 and 1990 Censuses. In spite of an overall increase in net undercount between 1980 and 1990, they suggest that census coverage of migrants from Mexico and Central America may have improved substantially by the 1990 Census because many former unauthorized migrants had acquired legal status under the provisions of the Immigration Reform and Control Act of 1986 (IRCA). Ahmed and Robinson handle the problem of differential undercount and negative emigration estimates by computing emigration rates for race/ethnic groups (not country-of-origin groups) while excluding those country groups with negative rates. They then use the race/ethnic-specific rates as proxies to estimate emigration rates for countries that initially had negative rates; they match countries with the races of immigrants, for example, matching Hispanic rates to countries sending high proportions of Hispanic immigrants. Mulder (2003) handles the problem by adjusting the 1990 and 2000 census figures for undercount. This approach results in a set of final emigration estimates that are highly sensitive to the coverage estimates on which the adjustments are based.

A second limitation of the residual method is its dependence on the accurate comparison of age- and year-of-entry cohorts across two censuses. The consistency of reporting on year of entry is of particular concern. In the 1990 Census, one-third of immigrants who reported having come to the U.S. between 1985 and 1990 on the year-of-entry question probably lived in the United States prior to 1985 based on their responses to the residence-five-years-ago question (Ellis and Wright 1998). If a significant number of recently-arrived immigrants understate the length of time they have lived in the United States in the first census, estimates of recent arrivals would be overstated, as would estimates of the “expected” population ten years later. If reporting on year of entry were more accurate in the later census, it would appear that more recent arrivals emigrated than was in fact the case. Changes in reporting of year of entry across the two censuses may also contribute to the problem of negative emigration estimates caused by change in census coverage.

A third weakness of the residual method is its inability to estimate emigration for recently-arrived immigrants—i.e., those who arrived since the previous census. For this group, the earlier census is not available, so emigration rates are not derived 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. The alternative of comparing the recent cohorts in the second census to estimates of survivors of legal immigrants who arrived between the two censuses is not a viable solution because such estimates are used to measure unauthorized migrations (e.g., Passel et al. 2004a). One consequence of the lack of residual measures for the recent arrivals is evident from Table 1—the inconsistent, in some cases, unrealistically low residual-based estimates of emigration for recently-arrived immigrants.

Finally, another limitation of the residual method is that it does not permit the estimation of emigration by migration status because migration status is not identified in the census.[2] Massey and Singer (1995) make a significant contribution by using life histories from the Mexican Migration Project (MMP) to estimate in-, out-, and net unauthorized migration flows from Mexico to the United States. Their work suggests that the unauthorized flow from Mexico consists of large proportions of temporary and circular migrants (those making numerous yet relatively short trips to the United States) whose emigration rates are likely to be much higher than those of lawful permanent residents. Their work is not directly transferable to other populations because their universe consists principally of persons living in Mexico and so, may overstate levels of return migration to Mexico. Further, their data are for Mexicans only so that one cannot prepare comparable estimates for legal immigrants or unauthorized migrants from countries other than Mexico. Such comparisons of emigration patterns of unauthorized and legal immigrants remain important. For example, residual methods for estimating unauthorized migration uses the emigration patterns of *legal* migrants, not all foreign-born. (The residual estimate of unauthorized migration requires an accurate estimate of the number of legal immigrants living in the United States, and the legal foreign born estimate, in turn, is estimated as the number of legal admissions minus the estimated numbers of deaths and emigrants *among the legal foreign born*.)

A NEW APPROACH

In this report, we supplement prior residual estimates with emigration estimates that do not depend on assumptions about differential census coverage or consistent reporting on year-of-entry. We refer to this new approach as the "CPS Matching Method." This method takes advantage of the unique sample design of the Current Population Survey (CPS), which conducts 8 interviews in the same housing units—but not necessarily with the same individuals—over a 16-month period. For our estimates, we use the Annual Social and Economic Supplement (ASEC) to the CPS conducted every March. 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 houses in the U.S.), emigrants who moved out of the country, and a residual group who cannot be matched for other reasons. Based on prior knowledge and some assumptions about factors affecting the rates of internal migration, mortality, and non-follow-up, we use statistical methods to estimate the probability that non-matched individuals died, moved internally, emigrated, and were not followed for other reasons. We do not explicitly assign individuals categorically as an emigrant or not an emigrant. Rather, each individual is assigned a *probability* that they emigrated. We average the probability of emigration across all foreign-born who first appeared in the March CPS in year t to estimate the proportion of emigrants among the foreign-born.

One advantage of the CPS Matching Method is that it does not depend critically on the consistency of year-of-entry or coverage. The residual method compares the sizes of foreign-born cohorts between two data sources collected ten years apart, and assumes that the foreign-born cohorts in each data source consist of the same people with the sole exceptions of those who died or emigrated. Differences in undercount or misreporting between the two data

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[2] The Census Bureau and others do produce estimates of the number and characteristics of unauthorized migrants enumerated in the Census. However, these estimates rest, in part, on assumptions about foreign-born emigration. Their subsequent usage in the production of emigration estimates by status would involve a circularity in logic.

sources—unknown factors that are difficult to quantify—are not typically factored in. However, the CPS Matching Method does not compare the size of *groups* between data sources. Rather, it follows *individuals* over time. There is no need to assume consistency in coverage or reporting between the two surveys because all social and demographic information (age, period-of-entry, place of birth, sampling weight) is obtained from a single data source: the CPS in year t.

Another advantage is that the Matching Method estimates emigration rates for recently-arrived immigrants in the same manner as earlier arrivals. The new method is therefore more likely to produce comparable estimates across different period-of-entry groups than the residual method. Finally, the CPS Matching Method permits us to estimate emigration by legal status. We use the migration status assignments developed by Passel and Clark (1998) and refined by this project team in Passel et al. (2004b). The status assignment procedures assign individual foreign-born CPS respondents in year t when they first appear in the CPS to legal status groups (some deterministically and others on a probabilistic basis) using place of birth, occupation, sex, gender, welfare recipiency, state of residence, and other characteristics. Once individuals are assigned as legal or unauthorized, we analyze one-year follow-up probabilities separately by legal status in order to produce status-specific emigration estimates.

As elaborated below, the CPS Matching estimates depend critically on the accuracy of other assumptions. Two of the most significant are (1) that emigration rates among second-generation natives are negligible and (2) that immigrants and the second generation have similar patterns of non-follow-up due to unmeasured causes (while controlling for a number of socioeconomic factors).

Emigration and Return Immigration

Emigration estimates from the CPS Matching Method are likely to be larger than those based on residual methods because the residual method does not count as "emigrants" those who leave the United States but later return within the decade (i.e., so-called "return immigrants"[3]). Specifically, the residual method does not measure the *annual* number of emigrants directly. Rather, it typically involves estimating *net* emigration over a ten-year period. Average annual emigration is then the total number of emigrants over the decade divided by ten.[4] However, some immigrants may have been living in the United States both at the beginning and end of the decade while having made several trips back and forth during the decade. This phenomenon seems particularly important in the case of Mexican migration to the United States. Massey and his colleagues estimate that the average duration of a Mexican labor migrant's first trip to the United States is only 21 months and that one-third of these migrants return to the United States in a second trip within ten years of the first trip (Massey et al. 2002). This type of circular migration is not captured by the residual method because the residual method estimates the number of *net* rather than *total* emigrants.

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[3] We use the term "return immigration" to denote immigration to the United States by former immigrants who have left the United States to live abroad, but have returned to the United States. We use this term to distinguish the phenomenon from "return migration" which is usually used to mean emigration from the United States or return by immigrants to their home country.

[4] Average annual emigration rates are sometimes computed by dividing the average annual emigration by the mid-period foreign-born population. A better method is to derive the annual rate as one minus the 10th root of the 10-year probability that an immigrant will *not* emigrate. In neither case, however, is annual emigration measured directly.

In general, the total amount of emigration measured over a time interval will be larger if sub intervals over which immigrant cohorts are followed become shorter. For example, the number of *net* emigrants over 1990–2000 (***E***) is equal to the sum of the number of emigrants each year ($\boldsymbol{E_y}$) minus the number of return immigration trips to the U.S. in each year ($\boldsymbol{R_y}$) among those who emigrated during 1990–2000, or:

$$E = \sum_{y=1990}^{y=2000} E_y - \sum_{y=1990}^{y=2000} R_y$$

Taking the annual average:

$$\frac{E}{10} = \overline{E}_y - \overline{R}_y .$$

So, by simple algebra,

$$E = 10 \cdot \left(\overline{E}_y - \overline{R}_y\right), \text{ and}$$

$$\overline{E}_y = \frac{E}{10} + \overline{R}_y .$$

By this relationship, the difference between estimates using a ten-year interval, such as most residual methods, and a one-year interval, as measured by the CPS Matching Method introduced here, is equivalent to the average annual number of returns to the U.S. by former immigrants. We introduce a method to “correct” the CPS Matching estimates for return immigration and derive *net* emigration measures comparable to those produced by residual methods and to those required for population estimates.

METHODOLOGY

Basic Approach

Our emigration estimates are based primarily on analysis of the Current Population Survey. A key feature of the CPS sample design—one that is critical for our purposes—is that it follows housing units over time. The CPS interviews occupants of the housing units for 4 consecutive months; these units do not appear in the next 8 monthly CPSs. Then, the occupants of the same housing units are interviewed for 4 additional consecutive months. Each interview, numbered 1 through 8, is referred to with the number of “months-in-sample.” With this design, those in months-in-sample 1 through 4 in year *t* appear in months-in-sample 5 through 8 in year *t+1*. It is important that the sample is of *addresses*, not *individuals*. Thus, if a CPS respondent moves to a new address, he/she is not followed. Rather, the new occupants of the original housing unit are interviewed and the original respondent is dropped from the sample. This feature of the CPS sample design permits us to use follow-up rates—i.e., the proportion of persons in months-in-sample 1–4 in one year who are successfully interviewed as members of months-in-sample 5–8 in the following year—as a basis for estimating emigration. (See U.S. Census Bureau (2002a) for a detailed description of the CPS design.)

Madrian and Lefgren (1999) estimate that 29 percent of those eligible for follow-up in the March CPS 1980-1998 surveys were not successfully followed up. There are many reasons a person may not be interviewed during a second year in the CPS sample. Based on known rates of internal migration and mortality (derived from the CPS and NCHS statistics), Madrian and

Lefgren (1999) estimate that 16.3 percent moved to another address in the United States and 0.9 percent died, leaving 11.8 percent who were not followed up for other reasons. Of the residual 11.8 percent, some may have moved to another country while others may not have been followed due to non-response, refusals, coding error, or the inability to contact a person in the housing unit. Thus, 11.8 percent is the maximum percent of emigration, and this figure is almost certainly far too high because it does not take into account other reasons for non-follow-up. The basic dilemma is how to parse out the residual into emigration and residual-non-follow-up components. Our strategy is to assume that—in the case of natives—*virtually all* of the residual is due to residual non-follow-up. We then analyze with multivariate models the factors affecting residual non-follow-up for natives, and apply these models to immigrants in order to estimate their residual non-follow-up component.

More formally, we begin by representing the proportion of eligible persons not followed up (**u**) as the sum of the proportion who migrated within the United States (**m**), the proportion who died (**d**), the proportion who emigrated (**e**), and the proportion who were not followed up for other reasons (**r**). These components can be estimated for subgroups of the population. Thus, for immigrants (**I**), we represent the relationship as:

$$u^{I} = m^{I} + d^{I} + r^{I} + e^{I} \qquad \text{(1a—immigrants)}$$

Most of these terms may be estimated from existing data. The non-follow-up probability ($\mathbf{u^I}$) may be estimated as the number of persons followed up in the March CPS in year t+1 divided by the number eligible to be matched in the March CPS in year t. The proportion of internal migrants ($\mathbf{m^I}$) may be estimated from the place-of-residence-one-year-ago question in the CPS. The probability of death ($\mathbf{d^I}$), a small component except in the older ages, may be estimated for immigrants using the National Health Interview Survey or NHIS (Palloni and Aries 2004). We are left with the proportion of emigrants ($\mathbf{e^I}$) and residual non-follow-up probability ($\mathbf{r^I}$) for immigrants. To estimate these components, we introduce the parallel relationship for the non-follow-up probability for second-generation natives:

$$u^{N} = m^{N} + d^{N} + r^{N} + e^{N} \qquad \text{(1b—natives)}$$

We choose second-generation natives rather than all natives as a comparison group for reasons explained further below. Each of the components in (1b) can be estimated in the same manner as in (1a). However, a simplifying assumption about the value of $\mathbf{e^N}$ is available. Fernandez (1995) estimates that during the 1980s, roughly 48,000 U.S. born emigrated per year. This level of emigration amounts to an annual rate of about .02 percent. Work with more recent data suggests that even this small level of native-born emigration is too high, perhaps by a factor of 3 (Gibbs et al. 2003)[5]. Thus, we make the assumption that the emigration probability of second-generation natives ($\mathbf{e^N}$) is negligible or essentially zero. Thus equation 1b reduces to:

$$u^{N} = m^{N} + d^{N} + r^{N} \qquad \text{(1c—natives)}$$

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[5] Fernandez (1994) and Gibbs et al. (2003) present the only empirically-based estimates of U.S. born emigration. Although this work has its own limitations, we have no alternative at this point other than to use it. Further work on U.S. born emigration is necessary to provide more support for the assumption that U.S. born emigration is so low.

With this assumption, subtracting the non-follow-up probability for second-generation natives (1b) from the probability for immigrants (1a) and solving for the probability of emigration for immigrants ($\mathbf{e^I}$) yields:

$$e^I = u^I - u^N + m^N - m^I + d^N - d^I + r^N - r^I \quad (1d)$$

Given equation (1d), if we were to assume that immigrants and natives had the same non-follow-up rates (i.e., $\mathbf{r^N}$- $\mathbf{r^N}$ = 0), we could calculate $\mathbf{e^I}$. But nativity differences in age, sex, income, education, and homeownership composition may lead to nativity differences in non-follow-up. Our strategy is to model non-follow-up ($\mathbf{r^N}$) among second generation natives as a function of socioeconomic and demographic factors (e.g., age, sex, education, marital status, presence of children, type of residence, and home ownership) while taking into account native probabilities of internal migration and mortality. Assuming that the same factors operating in the same way govern the residual non-follow-up rates for immigrants, we plug immigrants' characteristics into the model for second-generation natives to generate predicted probabilities of residual non-follow-up for immigrants ($\mathbf{r^I}$). Once ($\mathbf{r^N}$- $\mathbf{r^I}$) is estimated, we solve equation (1d) to yield individual-level probabilities of emigration for the foreign born ($\mathbf{e^I}$), which we average across the foreign-born and various foreign-born subgroups to produce emigration rates. Our estimation technique is explained in greater detail in the section below.

The selection of a native comparison group is an important issue. The underlying assumptions of the matching method are that (1) the native comparison groups have very low rates of emigration, and (2) the native comparison groups behave similarly to immigrants with respect to the factors affecting non-response. Satisfying both assumptions simultaneously may be difficult. On the one hand, third-or-higher generation natives may serve as a good comparison group because they may be less likely to emigrate than second-generation natives (because they tend to have fewer family connections overseas (Passel and Woodrow 1989)). On the other hand, second generation natives may serve as a good comparison group because they may behave more similarly to immigrants vis-à-vis non-response than third-or-higher generation natives (based on standard ideas about assimilation). Realistically, we may be able to relax the first assumption if we obtained up-to-date estimates of native emigration that could be factored into the final estimates. We believe it would be more difficult to relax the second assumption due to the difficulty in modeling nativity differences in attrition in the CPS. Therefore, to increase the likelihood that the second assumption about non-follow-up holds, we opt to use the second generation rather than all natives as the native comparison group.

While the choice of the second generation as a comparison group offers some advantages for adults, there remain some issues regarding the use of the second generation as a comparison group for children. Both immigrant and second-generation children are the children of immigrants; they often share the same households and are, therefore, likely to have similar emigration rates. Therefore, in the case of children, the estimated emigration rates of the second generation are not likely to be negligible, as required in the assumptions used to derive emigration rates for immigrants. Thus, if we were to use the methodology outlined above for estimating emigration rates of children, we would almost certainly underestimate emigration rates for immigrant children. For this reason, we treat children ages 0 to 14 differently from adults. Specifically, we effectively assume that immigrant children emigrate at the same rate as their parents.

Estimation Strategy

The probability of non-follow-up for each person can be estimated with logistic regression as shown below, where $\mathbf{X_i}'\boldsymbol{\beta}^{\mathbf{I}}$ is the product of a vector of socioeconomic and demographic characteristics of individual i (inversed) and a vector of corresponding logistic regression coefficients[6]:

$$\mathrm{u}_i^{\mathrm{I}} = \frac{\exp(\mathbf{X_i}'\boldsymbol{\beta}^{\mathbf{I}})}{1+\exp(\mathbf{X_i}'\boldsymbol{\beta}^{\mathbf{I}})} = F(\mathbf{X_i}'\boldsymbol{\beta}^{\mathbf{I}}) \qquad \text{(2a—immigrants)}$$

$$\mathrm{u}_n^{\mathrm{N}} = \frac{\exp(\mathbf{X_n}'\boldsymbol{\beta}^{\mathrm{N}})}{1+\exp(\mathbf{X_n}'\boldsymbol{\beta}^{\mathrm{N}})} = F(\mathbf{X_n}'\boldsymbol{\beta}^{\mathrm{N}}) \qquad \text{(2b—natives)}$$

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and each component of non-follow-up can be similarly modeled:

Immigrants:	Natives:
$\mathrm{m}_i^{\mathrm{I}} = F(\mathbf{X_i}'\boldsymbol{\mu}^{\mathbf{I}})$	$\mathrm{m}_n^{\mathrm{N}} = F(\mathbf{X_n}'\boldsymbol{\mu}^{\mathrm{N}})$
$\mathrm{d}_i^{\mathrm{I}} = F(\mathbf{X_i}'\boldsymbol{\delta}^{\mathbf{I}})$	$\mathrm{d}_n^{\mathrm{N}} = F(\mathbf{X_n}'\boldsymbol{\delta}^{\mathrm{N}})$
$\mathrm{r}_i^{\mathrm{I}} = F(\mathbf{X_i}'\boldsymbol{\rho}^{\mathbf{I}})$	$\mathrm{r}_n^{\mathrm{N}} = F(\mathbf{X_n}'\boldsymbol{\rho}^{\mathrm{N}})$
$\mathrm{e}_i^{\mathrm{I}} = F(\mathbf{X_i}'\boldsymbol{\lambda}^{\mathbf{I}})$	$\mathrm{e}_n^{\mathrm{N}} = F(\mathbf{X_n}'\lambda^{\mathrm{N}})$

(2c)

The likelihood of non-follow-up can therefore be expressed as the sum of predicted values generated from each component model. For example, for immigrants:

$$\mathrm{u}_i^{\mathrm{I}} = F(\mathbf{X_i}'\boldsymbol{\mu}^{\mathbf{I}}) + F(\mathbf{X_i}'\boldsymbol{\delta}^{\mathbf{I}}) + F(\mathbf{X_i}'\boldsymbol{\rho}^{\mathbf{I}}) + F(\mathbf{X_i}'\boldsymbol{\lambda}^{\mathbf{I}}) \qquad \text{(3a—immigrants)}$$

Averaging (3a) across all immigrants yields the non-follow-up probability among immigrants:

$$\mathrm{u}^{\mathrm{I}} = \frac{\sum_{i=1}^{i=I} F(\mathbf{X_i}'\boldsymbol{\beta}^{\mathbf{I}})}{I} = \frac{\sum_{i=1}^{i=I} F(\mathbf{X_i}'\boldsymbol{\mu}^{\mathbf{I}}) + F(\mathbf{X_i}'\boldsymbol{\delta}^{\mathbf{I}}) + F(\mathbf{X_i}'\boldsymbol{\rho}^{\mathbf{I}}) + F(\mathbf{X_i}'\boldsymbol{\lambda}^{\mathbf{I}})}{I} \qquad \text{(4a—immigrants)}$$

where ***I*** is the number of immigrants.

We reason that the difference in non-follow-up between immigrants and natives—after removing the influence of compositional differences and differences in mortality and internal

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[6] $\mathbf{X_i}'\boldsymbol{\beta}^{\mathbf{I}}$ is matrix notation for the sum of the products of X values of person i and corresponding coefficients, such as, in the case of a model with three variables, $X_{1,i}*B_1 + X_{2,i}*B_2 + X_{3,i}*B_3$.

migration—may be attributed to emigration. To measure this difference, we compare predicted values for immigrants from (4a) with predicted values for natives *assuming both groups had immigrants' characteristics* (obtained by generating predicted values for immigrants using native coefficients). The non-follow-up probability of the second generation *assuming they had the same composition as immigrants* ($\boldsymbol{u}^{N*}$) is obtained by replacing the coefficients in (4a) with natives' coefficients:

$$\mathrm{u}^{\mathrm{N}*} = \frac{\sum_{i=1}^{i=I} F(\mathbf{X_i}'\beta^{\mathrm{N}})}{I} = \frac{\sum_{i=1}^{i=I} F(\mathbf{X_i}'\boldsymbol{\mu}^{\mathrm{N}}) + F(\mathbf{X_i}'\boldsymbol{\delta}^{\mathrm{N}}) + F(\mathbf{X_i}'\boldsymbol{\rho}^{\mathrm{N}}) + F(\mathbf{X_i}'\boldsymbol{\lambda}^{\mathrm{N}})}{I} \qquad \text{(4b—natives*)}$$

By subtracting (4b) from (4a) and solving for the emigration probability of immigrants, $\sum_{i=1}^{i=I} F(\mathbf{X_i}'\boldsymbol{\lambda}^{\mathbf{I}})/\mathrm{I}$, we estimate the foreign-born probability of emigration as:

$$\mathrm{e}^{\mathrm{I}} = \sum_{i=1}^{i=I} F(\mathbf{X_i}'\boldsymbol{\lambda}^{\mathbf{I}})/\mathrm{I} \qquad \text{(5a—immigrants)}$$

$$= \sum_{i=1}^{i=I} \frac{1}{I}[F(\mathbf{X_i}'\boldsymbol{\beta}^{\mathbf{I}}) - F(\mathbf{X_i}'\boldsymbol{\beta}^{\mathrm{N}}) + F(\mathbf{X_i}'\boldsymbol{\mu}^{\mathrm{N}}) - F(\mathbf{X_i}'\boldsymbol{\mu}^{\mathbf{I}})$$

$$+ F(\mathbf{X_i}'\boldsymbol{\delta}^{\mathrm{N}}) - F(\mathbf{X_i}'\boldsymbol{\delta}^{\mathbf{I}}) + F(\mathbf{X_i}'\boldsymbol{\rho}^{\mathrm{N}}) - F(\mathbf{X_i}'\boldsymbol{\rho}^{\mathbf{I}}) + F(\mathbf{X_i}'\boldsymbol{\lambda}^{\mathrm{N}})]$$

At this point, we introduce some simplifying assumptions, including those described above. First, we assume that—if given the same socioeconomic, health, and demographic characteristics—immigrants would exhibit the same residual non-follow-up rates as second-generation natives; that is, $\sum_{i=1}^{i=I} \frac{1}{I}\left[F(\mathbf{X_i}'\boldsymbol{\rho}^{\mathrm{N}}) - F(\mathbf{X_i}'\boldsymbol{\rho}^{\mathbf{I}})\right] = 0$. Also, we assume that emigration among the second generation is small enough to ignore for the reasons noted above; in other words, $\sum_{i=1}^{i=I} \frac{1}{I}\left[F(\mathbf{X_i}'\lambda^{\mathrm{N}})\right] = 0$. Therefore,

$$\mathrm{e}^{\mathrm{I}} = \sum_{i=1}^{i=I} \frac{1}{I}\left[F(\mathbf{X_i}'\beta^{\mathrm{I}}) - F(\mathbf{X_i}'\beta^{\mathrm{N}}) + F(\mathbf{X_i}'\delta^{\mathrm{N}}) - F(\mathbf{X_i}'\delta^{\mathrm{I}}) + F(\mathbf{X_i}'\mu^{\mathrm{N}}) - F(\mathbf{X_i}'\mu^{\mathrm{I}})\right] \qquad \text{(5b)}$$

where the first two terms give the predicted nativity difference in non-follow-up, and the second two terms are the predicted difference in internal migration, and the last two are the predicted difference in mortality. Equation (5b) is particularly useful because all components can be estimated with CPS and NHIS data.

Data

To estimate foreign-born emigration rates for the late 1990s and early 2000s, we use the Annual Demographic Supplements to the March CPS, designated as the Annual Social and

Economic Supplements beginning with March 2003. The supplements from this month offer several advantages over other months. The March Supplements contain a substantial range of socioeconomic and demographic information not in other months. The information needed to identify immigrants and native generations (i.e., country of birth, citizenship, year of entry to the U.S., and country of birth of parents) appear in every monthly CPS since 1994, but only the March supplement contains the question on residence 1 year ago that we use to identify internal migrants. In addition, we have developed methodology and SAS code for assigning legal status to the foreign-born in the March CPS (Passel et al. 2004b). Thus, the groundwork has already been laid that makes it possible to develop foreign-born emigration estimates separately for legal and unauthorized migrants.

A further advantage of March supplements is that the samples are larger than in other months; since emigration is a relatively rare event, the larger samples provide more precise estimates. Since the mid-1970s, the March supplement has contained an oversample of Hispanics. Under this design, all CPS households containing one or more Hispanics in the previous November are reinterviewed as part of the March supplement.[7] Those reinterviewed households that still contain one or more Hispanics in March are included as part of the CPS supplement sample. This sampling scheme effectively doubles the number of Hispanic households in the March Supplement. The households in the Hispanic oversample carry their month-in-sample values from November and can be matched from year to year using the same procedures as for the regular CPS sample.[8]

Beginning with the March 2002 CPS, the supplement has been expanded further by adding additional households from non-overlapping rotation groups in adjacent months. The additional increase includes:

(a) Hispanic households from the April CPS in rotation groups 1 and 5;

(b) Non-Hispanic, non-white households from the previous November CPS in rotation groups 1 and 5-8 and from the April CPS in rotation groups 1 and 5;

(c) Non-Hispanic, white households containing children 18 or younger from the previous November CPS in rotation groups 1 and 5-8 and from the April CPS in rotation groups 1 and 5.

The Census Bureau also released a research version of the March 2001 Supplement with the additional oversample, designated as the State Children's Health Insurance Program (SCHIP) file.

To estimate foreign-born emigration rates for the late 1990s and early 2000s, we use data drawn from the 1998, 1999, 2000, 2001, 2002, 2003, and 2004 March CPS Supplements. We choose these years because of the availability of estimates of legal status for all years.[9] During

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[7] The previous November is the most recent month for which the sample does not overlap with March.

[8] In some years, the month-in-sample variable for the Hispanic oversample has been miscoded on the CPS public-use files. The Census Bureau provided instructions that permitted us to carry out the matching required by our method.

[9] We do not include the March 1994 CPS because it is not possible to match the 1994 with the 1995 files (due to revisions in geographic identifiers, household identifiers were altered to protect the confidentiality of survey respondents; as a result, it is impossible to match the 1994 with the 1995 file). We do not include the March 1996 and 1997 CPS because we have not completed the legal status assignments for the 1997 file. As a result, we cannot

the 1998–2004 period, the basis for CPS weights changed from the 1990 Census to Census 2000. The official change-over occurred with the March 2002 CPS which was the first to use weights based on Census 2000. However, the March 2001 SCHIP file and a research version of the March 2000 Supplement also used Census 2000-based weights developed by the Census Bureau. Where possible, we use the 2000-based weights.

The sample used for modeling non-follow-up and for producing estimates of emigration is limited to foreign-born persons in the 1998 through 2003 CPS March samples who were eligible to be followed up in the following year. This means that for most years, the sample is restricted to those in months-in-sample 1–4[10]. For comparative purposes, we also examine the patterns of non-follow-up among second generation natives in the same rotation groups. The final analytic sample includes 44,973 foreign born and 47,094 second generation natives.

The sample used to measure internal migration is limited to first- and second-generation persons in the 1999 through 2004 CPS March samples who had lived in the United States in the prior year. Although we use different years of data for examining internal migration and non-follow-up, both refer to the same time period. We use years 1999 through 2004 (years t+1) rather than 1998 through 2003 (years t) because internal migration is a retrospective question and refers to moves made by those in the t+1 data that occurred during the period between year t and t+1. Non-follow-up, on the other hand, is measure prospectively and pertains to behavior of those in the CPS in year t for the period between year t and t+1. Unlike non-follow-up, we use all months-in-sample rather than just those in months 1-4. The sample for the internal migration models is more than twice as large as for the non-follow-up models (94,995 foreign born and 92,934 second generation natives).

We also use the National Health Interview Survey-National Death Index (NHIS-NDI) data to model the probability of death for immigrants and natives. Conducted each year since 1957, the NHIS is an annual survey of individuals age 18 and older about health status, health care, and insurance coverage. Beginning with the 1986 sample, NHIS respondents were linked to the National Death Index (NDI) files (a data base of all deaths in the United States) in order to ascertain vital status and age at death. NHIS respondents are matched on a number of identifiers, including social security number, first and last name, father's surname, and month and year of birth. Details about the methodology and quality of matches are discussed in the NHIS documentation (NCHS 2000). As of the time we conducted our analysis, NHIS respondents had been linked to the 1987 through 1997 NDI files. The NHIS did not include a question on place of birth until 1989, so we use the 1989 through 1994 NHIS files, which are linked to the 1989-1997 NDI files. We organize the NHIS-NDI data in person-year records, including a record for each year of life lived by NHIS respondents from the time of the survey and the time of their death or censorship in 1997, whichever comes first. The analytic data file includes 344,536

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estimate internal migration by legal status for the 1996-97 period and we cannot estimate match probabilities for the 1997-98 period. The former makes it impossible to produce estimates for the 1996-97 period and the latter makes it impossible to produce estimates for the 1997-98 period. When assignments are completed for March 1997, we will be able to add two additional years to the estimates. With the additional years, we may be able to provide estimates for two time periods, the late 1990s and early 2000s.

[10] We encountered some problems in matching the expanded CPS samples with the March 2001 SCHIP public-use file and the 2002 file. Accordingly, for the 2001–02 match we used the regular March 2001 CPS Supplement; for the 2002–03 matches we used only rotation groups 2 and 3. We are pursuing discussions with Census Bureau personnel to try to resolve the matching problems and expand the cases available for analysis.

person-year records for immigrants (2,480 deaths) and 2,767,340 person-year records for natives (27,652 deaths).

Although we use different data for modeling non-follow-up, internal migration, and migration, this does not present a serious problem. Our method only requires that we obtain a vector of coefficients that can be used to predict the probability of non-follow-up, internal migration, and migration. Once we estimate coefficients from a given sample, we apply the coefficients to the foreign-born in months-in-sample 1-4 in the 1998-2003 CPSs to calculate predicted probabilities of non-follow-up, internal migration, and mortality.

Models

To obtain values for the components of (5b), we first estimate three sets of weighted logistic regression models: the first set predicts non-follow-up among those eligible to be followed up, and the second set predicts internal migration (i.e., living at a different address from the year before) among those who were living in the United States the year before, and the third set predicts a one-year probability of dying. The models for non-follow-up and internal migration are estimated separately by sex and race/ethnicity for immigrants and for the second generation and include as independent variables identical sets of socio-demographic variables: age, education, housing tenure (rent vs. own), type of residence (e.g., mobile home, house/apartment, dormitory, rooming house), school enrollment status, and general health status. We estimated a total of 16 models of internal migration and 16 models of non-follow-up. We present in Appendix Table 1 the coefficients of the models for Mexican first and second generation men only. The estimates for the other models are available from the authors upon request.

For adults, we generate predicted values of the likelihood of non-follow-up and of internal migration from the appropriate model. Four predicted values are calculated for each immigrant:

(1) probability of non-follow-up using immigrant coefficients;

(2) probability of non-follow-up using second- generation coefficients;

(3) probability of internal migration using immigrant coefficients; and

(4) probability of internal migration using second- generation coefficients.

Predicted values for each sex and race/ethnic group are derived from each groups' corresponding models. For example, the predicted values for Mexican males come from the "Mexican male" models. In the case of Asians, there were not enough cases in the second (or third) generation to obtain stable estimates, so we obtained their predicted models estimated for all race/ethnic groups combined.

Predicted one-year probabilities of death for adults ages 18 and over are obtained from the National Health Interview Survey-National Death Index for 1989–97. Using a person-year file, we estimate separate logistic regression models for immigrants and natives predicting whether a person died during the year. Because of unavailability of questions on parents' place of birth in the NHIS, we do not estimate separate models for second and third-and-higher generation natives. The independent variables in our models include sex, age, race/ethnicity, and general health. We use coefficients from the mortality models to generate both the "immigrant" and "native" (second generation) predicted probabilities for immigrants in the CPS, as shown in the last two terms of (5b). The model estimates are presented in Appendix Table 2.

We subtract the immigrant predicted values from the second-generation predicted values (natives, in the case of mortality) as shown in the brackets in equation (5b). We then average the results to obtain an estimate of $\mathbf{e}^{\mathbf{I}}$, the emigration probability, for all immigrants and for immigrant subgroups by age, sex, country-of-origin, year-of-entry, and migration status. That is, we average the bracketed portion of (5b) for the different groups. As noted above, children are assigned the predicted probability of emigration of their parents on the assumption that most children emigrate with their families.

COMPONENTS OF ESTIMATES

Non-follow-up

To determine whether a respondent in the March Supplement to the CPS in one year (**t**) is successfully followed up the following year (**t+1**), we match those eligible for follow-up[11] in the 1998–2003 March CPSs with respondents in the following years' CPSs, 1999–2004. Households from rotation groups 1–4 in each year t are matched to rotation groups 5–8 in the following year t+1. Then, matched individuals in these households are identified. The matched and unmatched individuals in year t are used to measure follow-up rates. We use the methodology and STATA code 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.[12] 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.

Internal Migration

The CPS asks respondents whether he/she lived in a different residence one year before. We define the internal migration probability for the time period from year t to t+1 as the proportion of movers among those who reported having lived in the United States one year before. Since the internal migration question is retrospective, we estimate internal migration rates for year **t** (1998–2003) from respondents in the CPS in year **t+1** (1999–2004). This procedure is different from the way follow-up rates are computed, which is the proportion of CPS respondents in year ***t*** who are followed up in year ***t+1***. Although the procedures are different, the reference periods for the rates are the same. The full CPS sample in year **t+1** is used for estimating internal migration rather than just the new rotation groups. For example, the proportion of internal migrants in 1998 is estimated from all rotation groups in the March 1999 CPS.

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[11] For most cases, those with month-in-sample codes 1–4 are eligible for follow-up. For those in the Hispanic oversample in most years, month-in-sample is erroneously reverse-coded (personal communication with Census Bureau). For these cases, we select months-in-sample 5-8 as eligible for follow-up. As noted earlier, we have encountered some problems in matching the expanded CPS samples from the 2002 supplement. We are attempting to resolve these issues through conversations with appropriate Census Bureau personnel.

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

Return Immigration

We define "return immigrants" as the foreign-born population who reported in year **t+1** living abroad one year earlier, but also reported having come to live in the United States more than 2 years before. "Net emigration" is then defined as the number of emigrants derived with the emigration probability as estimated with (5b) minus the number of return immigrants. Return immigration is estimated using the same analytical sample as used to estimate internal migration.

Migration Status

We produce estimates of emigration rates for the foreign-born population classified by legal status as: naturalized citizens, lawful permanent resident (LPR) aliens, refugee admissions, legal temporary migrants (or legal nonimmigrants), and unauthorized migrants. The status-specific estimates are derived from the matched March CPS Supplements in which individual foreign-born respondents have been assigned to legal status groups. Status assignments are based on responses to the CPS questions on citizenship, nativity, and year of immigration as well as additional information on characteristics of the migrants and their families including occupation, income, welfare use, school enrollment, educational attainment, and place of residence. The assignments are done only with data from the respondent'' first appearance in the March supplement. Detailed descriptions of the assignment methods can be found in Passel and Clark (1998) and Passel et al. (2004b); the definitions and assignment methods are summarized briefly here:

(1) *Naturalized Citizens*. These are legal immigrants who have acquired U.S. citizenship. The basic information used to identify naturalized citizenship comes from a question on citizenship in the CPS. However, the naturalized citizen category tends to be over-reported for people who have not been in the U.S. long enough to qualify for naturalization and for longer-term immigrants from Mexico and Central America (assuming the year-of-entry responses are correct). The former group is edited to noncitizen status. For the latter group, some individuals have citizenship changed during the assignment of unauthorized status. (See below.)

(2) *Refugees and Asylees*. These are individuals legally admitted to the United States for humanitarian purposes outside the regular family- and employment-based immigration system. Refugees tend to come from a relatively small number of countries that tend not to send many other migrants. These assignments are based on country of birth and year of immigration. The total number of refugees is adjusted to agree with a demographic estimate of the population group. In our estimates, only refugees who have come to the United States since 1980 are identified.

(3) *Lawful Temporary Migrants.* These are individuals admitted to the United States for specific periods (usually 3 years or less, but longer than 1 year) for specific purposes related to employment or study. Major groups that we identified in the CPS include foreign students, H-1B ("High-tech") guest workers, intra-company transfers, and exchange visitors. These assignments are based principally on school enrollment and occupation but draw on other CPS variables. Our assignment procedures attempt to match individual characteristics with the requirements for specific visa categories.

(4) *Lawful Permanent Residents (LPR)*. LPRs are individuals admitted as legal immigrants; they are also known as "green card" holders. In terms of status assignments, LPRs are individuals not classified in any of the other categories. Over time, most individuals admitted as refugees or asylees "adjust their status" and become LPRs. Virtually all naturalized citizens were, at one time, LPRs.

(5) *Unauthorized Migrants*. These are individuals who entered the United States without passing through ports of entry, entered with fraudulent or invalid documents, or entered with valid temporary visas but either stayed beyond the term of the visas or otherwise violated their terms of admission. The number of unauthorized migrants appearing in the CPS and the total number in the country are estimated by comparing a demographically-derived estimate of legal foreign-born residents with the foreign-born population in the CPS. (See Passel et al. 2004a for a more detailed description of the estimation methods.) Unauthorized migrants in the CPS are assigned using information on occupation, school enrollment, industry, and family relationships. The total number assigned is adjusted to agree with the demographic targets by age, sex, and country of birth for 6 large states and the balance of the country.

Standard Errors

For the estimates of return immigration, we calculate standard errors using the methodology provided in the CPS documentation by applying the "b" factors associated with Hispanics (U.S. Census Bureau 2002b)[13]. Because our methodology for producing emigration estimates involves so many computational steps, the calculation of standard errors of the final emigration estimates through the direct application of statistical formulas would be very difficult—if not virtually impossible. We instead use a bootstrap routine (the "bootstrap" command in Stata) to approximate standard errors of the emigration estimates[14].

RESULTS

The CPS Matching Method yields an estimate for the annual foreign-born emigration rate of 2.39% (with a standard error of 0.06% and 95% confidence interval of ±0.12%) (Table 2). For a population of 29,988 thousand foreign-born (as in the March 2000 CPS), this rate translates into roughly 716 thousand emigrants per year (±38.4 thousand). At the same time, we estimate a return immigration rate of 0.87% (±0.12%), translating into 260 thousand (±19.1 thousand) return immigrants annually. Subtracting return immigration from total emigration yields annual net emigration of 1.52 percent or 456 thousand net emigration per year (±0.18% or ±54.2 thousand).

Males are nearly twice as likely to emigrate as females—3.13 percent versus 1.65 percent—and are significantly more likely to be return immigrants, but not enough to offset their

[13] We checked the "Source and Accuracy" statements for all years of the CPS (1998-2004) and found that the adjustment factor "b" for Hispanics is the same across all years.

[14] The bootstrap method repeatedly draws random samples with replacement of size N from the original sample of size N and calculates the emigration estimates using the methodology outlined above. In our application, we draw 100 separate samples. The standard deviation of the estimates across the samples is interpreted as the standard error (Stata Corporation 2003).

significantly higher emigration rates. Net male emigration (2.03%) remains about twice as high as net female migration (0.94%), a difference that is statistically significant. Emigration and return immigration tend to be relatively higher for younger immigrants, decline with age, and in the case of emigration but not return immigration, increase for elderly immigrants (aged 65 years and over). Taking emigration and return immigration together, net emigration appears to be highest for children and working-age adults (25–44 years), but dips for teenagers and young adults (ages 15–24) and middle-aged adults (ages 45–64). Net emigration rates for those aged 25 to 44 are statistically significantly higher than those aged 15 to 24 and ages 45 to 64. The net figure significantly increases from a low of 0.11% for persons age 45-64 to 1.06% for elderly immigrants (aged 65 and over), perhaps as a result of retirees returning to their countries of origin.

Emigration rates vary considerably with migration/legal status, but in expected ways. Generally, the greater the attachment of the group to the United States, the less likely emigration is. Thus, refugees—many of whom cannot return to their countries of origin—and immigrants who have naturalized—who have made the greatest legal/political commitment to living in the United States—exhibit the lowest emigration rates, both for total emigration and net emigration. Legal non-citizens (LPRs) have the next highest emigration rates. Then, unauthorized migrants, many of whom engage in temporary, circular migration, exhibit relatively high emigration and return immigration rates. Finally, legal temporary migrants, who are primarily on student and temporary work visas, exhibit by far the highest emigration rates. Their net emigration rate (3.56%) is significantly higher than any other migration status group.

When we examine the emigration rates by duration of residence in the United States, we find that, in general, emigration rates are highest for recent arrivals and decline significantly with time in the United States. Return immigration rates are relatively low for both recent and earlier arrivals (0–4 and 10 or more years in the country), but higher for those who have lived in the U.S. for an intermediate length of time (5-9 years).

There is also considerable variability in emigration by country or region of birth. Much of the variation we find appears to be associated with the composition of the immigrant population by country. Thus, for example, countries that have high proportions of recent migrants and/or high proportions of unauthorized and legal temporary migrants tend to have higher emigration rates than countries where the immigrant population contains higher proportions of naturalized citizens and refugees. Immigrants from Mexico, India, Africa, Canada, and Europe appear to have the highest emigration rates and immigrants from China, the lowest. Of the high-emigration groups, the relatively high return immigration rates for Mexicans stand out as they are 25% higher than any other country. This distinction reflects the circular migration patterns that are commonly observed for Mexican migrants, especially the unauthorized group. On balance, *net* emigration rates appear lowest for groups from southeast Asia (China, Philippines, and “Other” Asia) and Central or South America; they are highest for immigrants from south Asia (India), Africa, Europe, and Canada. Because of their high return immigration rates, Mexican immigrants have net emigration levels that fall between these two groups.

How do the CPS Matching estimates of emigration compare with residual-based estimates? We compare our net emigration rates and estimates with the residual estimates prepared by Ahmed and Robinson (1994) and Mulder (2003) in Table 3. Ahmed and Robinson’s estimates pertain to the 1980-1990 period and Mulder’s estimates pertain to the 1990–2000 decade. To make a fair comparison, we derive net emigration rates from their results and then apply these rates to the 2000 foreign-born population to obtain estimates of emigration for 2000 to compare with our results.

In the case of earlier arrivals (i.e., those immigrants who have lived in the U.S. for 10 or more years), we find that our estimates are very similar to Mulder's (2003) estimates (Ahmed and Robinson (1994) did not present estimates by time in the United States). Our net emigration rate of 1.27 percent for this group is about 9 percent below Mulder's estimate of 1.39 percent. However, we estimate many more net emigrants among recent arrivals—1.81 percent versus only 0.32 percent in Mulder's estimates, or a difference of 210,000 annual emigrants. Estimating emigration among recent arrivals is particularly problematic for residual-based methodologies; in particular, they tend to be made by "borrowing" rates from the earlier arrivals. The large differences we find suggest that borrowing rates may not be appropriate and that the CPS Matching method offers an improved method for assessing emigration among recent arrivals.

Our results also show higher emigration rates for countries or regions that send large numbers of unauthorized, temporary, and recently-arrived immigrants. The CPS Matching estimates are significantly higher than Mulder's residual-based estimates for Mexico[15] (a source of unauthorized and circular migrants), and both Ahmed/Robinson's and Mulder's estimates for India (a source of many temporary migrants on H-1B visas) and Africa (a relatively new source of immigration and a source of temporary migrants). In terms of *numbers* of emigrants, the CPS-based results are higher from the residual-based estimates by the largest amounts for Mexicans, Europeans, and Indians. They are significantly lower in the cases of Ahmed and Robinson's estimates of the number of emigrants from China and the Philippines, and from Mulder's estimate for "Other" Asians.

CONCLUSIONS AND FURTHER WORK

In this report, we have developed a new method for estimating foreign-born emigration using a hitherto untapped data resource, mainly year-to-year matched data files from the March Supplements to the Current Population Survey. We have demonstrated the feasibility of this approach for producing emigration estimates for a multi-year time period from several pooled CPS samples from adjacent years, and produced initial estimates using data from the 1998–2004 period. Emigration estimates developed with the CPS Matching method could be incorporated into other applications including national (and subnational) population estimates and residual-based methods for estimating unauthorized migration (e.g., Passel et al. 2004a). We plan to continue development of the CPS Matching method, explore further the properties of the estimates, and introduce further improvements and refinements for our on-going research.

Refinements. One key parameter that permits us to solve the multi-equation system of relationships (e.g., equations 2a and 2b) is the assumption that native-born emigration is negligible. We plan to explore the sensitivity of the foreign-born emigration estimates to various assumptions about the level of U.S.-born emigration. If emigration for the second generation were not negligible, then the foreign-born emigration estimates are too low. We will re-estimate a range of foreign-born emigration estimates under varying assumptions about native emigration rates. A plausible range of assumptions about U.S.-born emigration will be drawn from the

[15] The residual methodology failed to produce "direct" estimates for Mexico so that the figures show in Table 3 apply residual-based estimates of rates for Hispanics from other areas to the Mexican-born population.

research literature (e.g., Gibbs et al. 2003 and Fernandez 1995), or be developed through further analysis of foreign data sources (i.e., by estimating the number of U.S. born present in foreign census or survey data).

Our initial estimates based on the CPS Matching method pooled data covering the broad period from 1998 through 2004. Although a single pair of matched CPSs (e.g., 1998–1999 or 2003–2004) does not, in general, provide enough cases to produce stable estimates of emigration, we plan to explore alternative collapsing schemes so that the method might be used to produce data for narrower time intervals. To achieve this end, we plan to extend the estimates in two ways. First, we plan to add data from the March 1996 and 1997 CPS Supplements and extend the time period back two years. The additional data set may permit us to produce emigration estimates for just the 1996–2000 period (if not with the full detail shown in Table 2, then for a smaller set of characteristics). In addition, we were unable to make use of the full, expanded CPS Supplements beginning with the 2001 SCHIP data set and extending through March 2004. We hope to resolve the matching difficulties that required us to restrict the estimates to 2 rotation groups per year. Fixing these problems would greatly expand the number of cases available for 2001–2004 and would probably permit development of stand-alone estimates for this period. With both extensions, we hope to be able to assess trends in emigration from 1996-2000 to 2000–2004.

Finally, we plan to produce more detailed emigration estimates than those shown in Table 2. The current estimates are for countries of birth, period of entry, and legal status separately. We plan to produce estimates of emigration that simultaneously take into account two (or more) of the dimensions shown. We also plan to produce more detailed estimates by region/state of residence and marital status. Finally, we plan to produce estimates for minor children of the foreign-born, both U.S-born and foreign-born.

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Table 1. Emigration Estimates in Prior Research

Table 1
Emigration Estimates in Prior Research

Population and Source of Estimate	Time Period of Estimate	Average Annual Emigrants	Emigration Rate[1] (percent)	Emigrants per 100 Arrivals[2]
All Foreign Born				
Warren and Peck (1980)*	1960-1970	114,000	1.18	34.3
Ahmed and Robinson (1994)*	1980-1990	195,000	1.15	27.8
Mulder (2003)*	1990-2000	225,000	0.88	24.7
Recent Arrivals: In U.S. Less Than 10 Years				
Warren and Peck (1980)*	1960-1970	62,000	4.39	18.6
Borjas and Bratsberg (1996)	1970-1980	89,000	3.20	19.7
Mulder (2003)*	1990-2000	21,000	0.32	2.3
In U.S. 10-20 Years				
Mulder (2003)*	1990-2000	142,000	1.66	13.9
Mexican Foreign Born				
Ahmed and Robinson (1994)*	1980-1990	20,000	0.62	12.1
Mulder (2003)*	1990-2000	27,000	0.40	16.1
Massey and Singer (1995)**	1965-1989	1,306,000	51.53	85.9
Recent Arrivals: In U.S. Less Than 10 Years				
Mulder (2003)*	1990-2000	3,000	0.12	1.7
Massey, et al. (2002)[a]**	1965-1985	629,000	28.75	57.5
Massey, et al. (2002)[b]**	1965-1985	182,000	12.50	25.0
In U.S. 10-20 Years				
Mulder (2003)*	1990-2000	15,000	0.68	5.1

* Residual estimate
** Life histories of all international moves
a Unauthorized Mexican migrants
b Legal Mexican migrants
1 Average annual emigrants per average person years of exposure (per 1,000)
2 Emigrants during the period per 100 immigrants during the period

Table 2. Foreign-Born Emigration Based on the CPS Matching Method, by Age, Sex, Country of Birth, Legal Status, and Time in U.S.: Circa 2000

Table 2

Foreign-Born Emigration Based on the CPS Matching Method, by Age, Sex, Country of Birth, Legal Status, and Time in U.S., Circa 2000

Group or Characteristic	*Annual Rates* (expressed as percents)			*Numbers of Persons* (in 000s)			
	Emigration	Return Immig.	Net Emigration	Emigrants	Return Immig.	Net Emigrants	2000 Population
All Foreign-Born	2.39	0.87	1.52	716.4	260.0	456.4	29,988
	(0.06)	(0.06)	(0.09)	(19.2)	(19.1)	(27.1)	
Sex							
Male	3.13	1.10	2.03	474.1	166.8	307.3	15,160
	(0.09)	(0.10)	(0.13)	(13.2)	(15.4)	(20.2)	
Female	1.65	0.71	0.94	244.1	104.7	139.4	14,828
	(0.09)	(0.08)	(0.12)	(13.0)	(12.1)	(17.7)	
Age Group							
0-14 years	3.66	1.25	2.41	74.5	25.5	49.0	2,034
	(0.33)	(0.30)	(0.45)	(6.8)	(6.1)	(9.1)	
15-24 years	2.42	1.24	1.18	105.4	54.0	51.4	4,353
	(0.09)	(0.21)	(0.23)	(4.0)	(8.9)	(9.8)	
25-34 years	3.82	1.08	2.73	259.9	73.8	186.1	6,810
	(0.13)	(0.15)	(0.20)	(8.5)	(10.2)	(13.3)	
35-44 years	3.70	1.07	2.63	238.6	69.1	169.5	6,455
	(0.18)	(0.15)	(0.24)	(11.8)	(9.8)	(15.4)	
45-64 years	0.64	0.53	0.11	45.8	37.8	8.0	7,171
	(0.09)	(0.10)	(0.13)	(6.3)	(7.2)	(9.6)	
65+ years	1.45	0.39	1.06	46.0	12.3	33.7	3,164
	(0.08)	(0.13)	(0.15)	(2.5)	(4.1)	(4.8)	
Migration Status							
Refugee	1.34	0.76	0.58	30.3	17.2	13.0	2,259
	(0.20)	(0.21)	(0.30)	(4.6)	(4.8)	(6.7)	
Naturalized	1.61	0.34	1.26	156.7	33.6	123.1	9,743
	(0.09)	(0.07)	(0.12)	(9.1)	(6.8)	(11.4)	
Legal Non-citizens	2.58	1.11	1.47	246.1	106.0	140.1	9,535
	(0.12)	(0.13)	(0.17)	(11.2)	(12.1)	(16.4)	
Unauthorized	3.33	1.26	2.06	249.7	94.8	154.9	7,503
	(0.15)	(0.16)	(0.22)	(11.3)	(11.9)	(16.4)	
Legal Temporary	4.46	0.90	3.56	42.2	8.5	33.7	948
	(0.47)	(0.37)	(0.60)	(4.4)	(3.5)	(5.7)	

Notes: Rates are based on analysis of pairs of matched CPS files from 1998-99 through 2003-04. Estimates of emigrants and return migrants derived by applying rates to March 2000 CPS population based on Census 2000 weights. Standard errors are shown in parentheses (see text for details).

Table 2 (continued)

Foreign-Born Emigration Based on the CPS Matching Method, by Age, Sex, Country of Birth, Legal Status, and Time in U.S., Circa 2000

Group or Characteristic	*Annual Rates* (expressed as percents)			*Numbers of Persons* (in 000s)			
	Emigration	Return Immig.	Net Emigration	Emigrants	Return Immig.	Net Emigrants	2000 Population
Years in United States							
0-4 years	3.61	0.98	2.63	166.9	45.2	121.7	4,618
	(0.18)	(0.15)	(0.23)	(8.2)	(6.8)	(10.6)	
5-9 years	3.01	1.60	1.41	285.2	151.7	133.4	9,479
	(0.15)	(0.21)	(0.25)	(13.9)	(19.4)	(23.9)	
10+ years	1.79	0.67	1.12	285.1	106.6	178.5	15,891
	(0.05)	(0.07)	(0.09)	(7.6)	(11.4)	(13.7)	
Country or Region of Birth							
Mexico	3.02	1.24	1.78	257.9	105.7	152.2	8,535
	(0.33)	(0.14)	(0.36)	(27.9)	(12.1)	(30.4)	
Central & S.Amer.	1.76	0.88	0.88	77.7	38.7	39.0	4,411
	(0.26)	(0.17)	(0.31)	(11.3)	(7.4)	(13.5)	
Caribbean	1.87	0.68	1.20	90.9	32.8	58.0	4,849
	(0.20)	(0.14)	(0.24)	(9.7)	(6.9)	(11.9)	
Canada (N.Amer.)	3.36	0.69	2.67	23.7	4.9	18.8	705
	(0.27)	(0.35)	(0.44)	(1.9)	(2.5)	(3.1)	
Europe	2.91	0.44	2.47	105.5	16.0	89.4	3,628
	(0.37)	(0.13)	(0.39)	(13.5)	(4.7)	(14.3)	
China	1.13	0.39	0.74	17.5	6.1	11.4	1,548
	(0.09)	(0.20)	(0.22)	(1.3)	(3.1)	(3.3)	
Philippines	1.41	0.67	0.74	19.0	9.0	10.0	1,352
	(0.24)	(0.26)	(0.36)	(3.3)	(3.5)	(4.8)	
India	3.85	0.72	3.13	43.2	8.1	35.1	1,121
	(0.18)	(0.31)	(0.36)	(2.0)	(3.5)	(4.0)	
Other Asia	1.73	0.99	0.75	54.2	30.8	23.4	3,128
	(0.13)	(0.21)	(0.25)	(4.1)	(6.6)	(7.8)	
Africa	3.45	0.71	2.73	24.5	5.1	19.4	711
	(0.13)	(0.37)	(0.39)	(0.9)	(2.6)	(2.8)	

Notes: Rates are based on analysis of pairs of matched CPS files from 1998-99 through 2003-04. Estimates of emigrants and return migrants derived by applying rates to March 2000 CPS population based on Census 2000 weights. Standard errors are shown in parentheses (see text for details).

Table 3. Alternative Emigration Estimates Based on the CPS-Matching Method and the Residual Method

Table 3

Alternative Emigration Estimates Based on the CPS-Matching Method and the Residual Method

	Net Emigration Rate (expressed as percent)					*Annual Net Emigration* (in 000s)				
		Residual Method		*Difference (CPS-Residual)*			*Residual Method*		*Difference (CPS-Residual)*	
Group or Characteristic	CPS Matching Method	Ahmed & Robinson (1994)	Mulder (2003)	Ahmed & Robinson (1994)	Mulder (2003)	CPS Matching Method	Ahmed & Robinson (1994)	Mulder (2003)	Ahmed & Robinson (1994)	Mulder (2003)
All Foreign-Born	1.52	1.15	0.88	0.37	0.64	456	346	265	111	191
	(0.09)					(27)				
Years in United States										
Less than 10 years	1.81	---	0.32	---	1.49	255	---	45	---	210
	(0.13)					(17)				
10+ years	1.27	---	1.39	---	-0.12	201	---	220	---	-19
	(0.05)					(14)				
Country or Region of Birth										
Mexico	1.78	0.62	0.42	1.17	1.37	152	53	35	99	117
	(0.33)					(30)				
Central & S.Amer.	0.88	0.93	0.98	-0.05	-0.10	39	41	43	-2	-4
	(0.26)					(14)				
Caribbean	1.20	1.17	0.96	0.03	0.23	58	57	47	1	11
	(0.20)					(12)				
Canada (N.Amer.)	2.67	1.42	1.15	1.25	1.52	19	10	8	9	11
	(0.27)					(3)				
Europe	2.47	1.34	1.05	1.12	1.42	89	49	38	41	51
	(0.37)					(14)				
China	0.74	1.79	0.87	-1.05	-0.14	11	28	14	-16	-2
	(0.09)					(3)				
Philippines	0.74	1.59	0.64	-0.85	0.11	10	21	9	-11	1
	(0.24)					(5)				
India	3.13	1.61	0.40	1.53	2.73	35	18	4	17	31
	(0.18)					(4)				
Other Asia	0.75	0.73	1.83	0.01	-1.08	23	23	57	0	-34
	(0.13)					(8)				
Africa	2.73	2.41	1.13	0.32	1.60	19	17	8	2	11
	(0.13)					(3)				

Sources: CPS Matching Method -- Table 2; Residual Method -- Ahmed and Robinson (1994); Mulder (2003).
Standard errors are shown in parentheses (see text for details). Mulder (2003) and Ahmed and Robinson (1993) did not report standard errors for their estimates.

Appendix Table 1. Logistic Regression Models of Internal Migration and Non-Follow-Up Among First and Second Generation Mexican Origin Males

Appendix Table 1

Logistic Regression Models of Internal Migration and Non-Follow-Up Among First and Second Generation Mexican Origin Males

	Internal Migration		Non-Follow-up	
	1st Generation	2nd Generation	1st Generation	2nd Generation
CPS Oversample	-1.044 ***	-0.956 ***	-0.032	-0.102
Homeowner	-0.726 ***	-1.175 ***	-1.000 ***	-0.934 ***
Age				
(0-4)				
5-9	-0.453 *	-0.416 ***	-0.127	-0.408 ***
10-14	-0.462 *	-0.757 ***	-0.480 *	-0.476 ***
15-19	-0.189	-0.297 *	0.549 *	-0.187
20-24	0.042	0.167	0.958 ***	0.801 ***
25-29	-0.049	0.067	0.383	0.419 *
30-34	-0.440 *	-0.265	0.082	0.355
35-39	-0.732 ***	0.000	-0.139	-0.214
40-44	-0.776 ***	-0.449	-0.181	-0.098
45-49	-0.861 ***	-0.522 *	-0.388	0.041
50-54	-1.158 ***	-1.374 ***	-0.222	-0.459
55-59	-0.959 ***	-1.071 **	-0.508	0.066
60-64	-1.246 ***	-1.107 **	-0.631 *	-0.936 *
65-69	-1.447 ***	0.000	-0.567	-0.074
70-74	-2.678 ***	-0.670	0.002	-0.400
75+	-1.771 ***	-1.231 **	0.262	0.556 *
CPS Year				
1998/1999	0.096	0.045	-0.119 *	-0.113
2000/2001	-0.169 **	-0.229 **	-0.049	-0.236 **
(2002-2004)				
School Enrollment				
High School	-0.415 **	-0.635 **	-0.569 ***	-0.014
College	-1.191 ***	-0.576 **	-1.139 ***	-0.181
General Health Status				
Excellent	-0.035	-0.274	0.119	-0.058
Very Good	-0.023	-0.188	0.011	-0.027
Good	0.029	-0.177	0.007	0.009
Fair	-0.221	0.038	-0.043	0.413
(Poor)				
Educational Attainment				
Less than HS	0.047	-0.318	-0.240 *	0.094
HS	-0.016	-0.388 *	-0.314 **	-0.123
Some College	-0.009	-0.269	-0.204	-0.027
(College+)				
Intercept	-0.310	0.104	0.193	-0.124
N	19,106	14,107	8,917	7,261
-2LL Chi-Square	1,519	937	967	429
Pseudo R-square	0.093	0.107	0.083	0.060

Source: 1998-2004 March Current Population Survey (see text for description of sample)
*** p<.001 ** p<.01 *p<.05
Reference categories are noted in parentheses.

Appendix Table 2. Discrete-time Event History Models of Mortality by Nativity (Logistic Regression Coefficients)

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Appendix Table 2

Discrete-time Event History Models of Mortality by Nativity (Logistic Regression Coefficients)

	Immigrants		Natives	
Intercept	-2.017	***	-1.711	***
Male	0.517	***	0.531	***
Race/Ethnicity				
Mexican	0.336	***	0.044	
Other Hispanic	0.067		0.204	***
NH White	0.230	***	0.205	***
Black	0.176		0.297	***
(Other)				
Age				
18-24	-3.984	***	-4.042	***
25-34	-3.722	***	-3.782	***
35-44	-3.324	***	-3.149	***
45-54	-2.756	***	-2.330	***
55-64	-1.818	***	-1.535	***
65-74	-0.994	***	-0.819	***
(75+)				
General Health Status				
Excellent	-1.181	***	-1.711	***
Very Good	-1.135	***	-1.501	***
Good	-0.937	***	-1.127	***
Fair	-0.586	***	-0.653	***
(Poor)				
Number of person-years	344,536		2,767,340	

Source: 1989-1994 National Health Interview Surveys
linked to the 1989-2004 National Death Index
*** p<.001 ** p<.01 *p<.05
Reference categories are noted in parentheses.