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The Effect of Family Member Migration on Education and Work among Nonmigrant Youth in Mexico

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Abstract

While academic and policy circles have given much attention to the assimilatory experiences of Mexican immigrants in the United States, less is known about those who stay behind—an especially unfortunate oversight given the increasing number of Mexican youth with migrant family members. Of the studies that do exist, most have sought to identify the effect migration has on youths' migratory and educational aspirations, often using qualitative methods in single sending communities. The present article supplements this research in two ways: (1) in addition to assessing educational outcomes, the scope of the analysis is expanded to include nonmigrants' interaction with another homeland institution of upward mobility—the labor market; and (2) using a large demographic data set, statistical techniques are employed to adjust for unobserved selectivity into the migrant family-member population, thus accounting for a potentially serious source of bias. The results suggest that youth in migrant-sending families are less likely to complete the educational transitions leading up to post-secondary school, and have a lower probability of participating in the local economy. The results also indicate that unobserved factors play a "nonignorable" role in sorting youth into migrant and nonmigrant families.

The magnitude of Mexican migration into the United States has increased steadily over the past 30 years (Office of Immigration Statistics 2005). While the assimilatory experiences and driving forces behind this new-wave of immigration have generated a large body of scholarship in the family and migration literatures, until recently academics had paid little attention to those for whom migration has a more indirect effect: the family members who stay behind. Analyses of "the other half" have focused predominantly on the macro-level implications of out-migration in terms of economic development, weighing the reduction of productive labor capacity against the aggregate financial gains from remittances (Taylor 1999). Relatively little is known, however, about the social and community-level effects of migration in sending areas, and less still about whether and how efforts abroad shape the individual behaviors and ambitions of family members "back home."

In this article, I begin to explore these questions. My principal objective is to determine whether and to what extent immediate exposure to international migration influences the way remaining family members, particularly youth, interact with the educational system and local labor markets. This is an especially timely question given the number of people crossing the border from Mexico—an outflow that totaled nearly 3 million persons between 1995 and 2000 (Passel, Van Hook and Bean 2004)—and the amount of resources that these individuals channel back to their places of origin (Sana 2008). It is also a question of considerable scholarly relevance. Although the specific effects of migration are complex and

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sometimes contentious, there is little doubt that the movement of people and goods reorganizes social, cultural, and economic structures in sending areas; a transformation that may complicate the meanings and value nonmigrants attach to homeland institutions of upward mobility.

The implications of this stance are elaborated most prominently by culture-of-migration theorists (Kandel and Massey 2002; Massey 1986, 1987; Mines and Massey 1985), who connect foreign wage labor with the cross-national transmission of capital, ideas, and culture that together promote future migratory behavior. These analyses anticipate a negative correlation between family involvement in migration and the willingness of children to invest in resources associated with upward mobility, particularly schooling. As migratory behavior becomes normalized the prospect of participating in homeland institutions is eclipsed by more profitable opportunities in foreign markets. Evidence supporting this position has come primarily from fieldworkers in individual communities (Alarcón 1992; Cohen 2004; Levitt 2001; Mines 1981; Reichert 1982; Rouse 1992; Smith 1998; Wiest 1973). To date only a few quantitative studies have yielded insight into the micro-level implications of the culture-of-migration thesis (Kandel and Kao 2001; Kandel and Massey 2002).

From another point of view, it is plausible to suppose that family member migration might have the opposite effect. Youth in transnational families may approach school and labor as a way to “do their part,” effectively acknowledging the sacrifices of migrant family members through diligence and hard work. This hypothesis has some sway among feminist scholars of migration and the family (Hondagneu-Sotelo and Avila 1997; Parreñas 2001, 2005a). Parreñas (2005a), for example, detected a tendency among youth in transnational families to achieve highly in school as a way of reciprocating the efforts and contributions of remitting relatives. Benefiting from repatriated earnings, there is also reason to imagine that youth in migrant-sending families might fare comparatively better in school due to their ability to offset or otherwise defray the opportunity costs associated with continued enrollment. Edwards and Ureta (2003), López-Córdova (2005), and Taylor (1999) have speculated that this may in fact be the case.

In investigating these hypotheses, this paper makes two primary contributions. First, rather than focusing exclusively on the education of nonmigrant children, I extend the analysis to include their labor market activity. In taking a wider view of migration and its implications for nonmigrant children, I am able to evaluate the extent to which the theoretical frameworks described above can be generalized to stratifying institutions other than the school system. Second, and perhaps more importantly, I rely on a series of endogenous switching regression models and Heckman’s index sufficient method to measure and adjust for potential selection bias. The inability of prior studies to address selectivity has, with few exceptions (Hanson and Woodruff 2003; McKenzie and Rapoport 2007; Miranda 2007), led to results that potentially conflate the effect of migration with the effect of unobserved factors—such as economic need and/or a shared sense of ambition and responsibility within the family—that promote migratory behavior in the first place.

The remainder of the paper is organized into five sections. After briefly reviewing the literature on individual- and community-level effects of out-migration in the first section, I introduce and describe my data source. Next, I develop two multivariate models to estimate the impact of family migratory behavior on youths’ educational attainment and labor force activity. In the fourth section, I summarize results from these models, focusing in particular on questions related to self-selection. Finally, in closing, I recapitulate my main findings and suggest various avenues for further research.

THREE APPROACHES TO STUDYING THE EFFECTS OF MIGRATION IN SENDING AREAS

As issues surrounding international migration have come to the foreground in recent decades scholarly orientations have become increasingly diverse. While the bulk of the literature has sought to predict migratory behavior using competing theoretical models (Fussell 2004; Fussell and Massey 2004; Massey 1990a, 1990b; Piore 1979; Portes and Walton 1981; Stark 1991; Todaro 1976), or to situate the experiences of contemporary immigrants within debates over assimilation (Alba and Nee 2003; Gans 1992; Hirschman 2001; Portes, Fernández-Kelly and Haller 2005; Portes and Rumbaut 1996; South, Crowder and Chavez 2005; Zhou 1997), an emerging line of research has begun to consider the meaning and impact of migration from the perspective of the homeland.

The existing research on sending areas can be divided roughly into three groups. The first, and most familiar, understands *migration as an avenue for capital flow* (Durand, Parrado and Massey 1996; Massey and Parrado 1994; Orozco 2002; Taylor et al. 1996a, 1996b; Wise and Marquez 2007). Conceptually, this approach seeks to isolate and draw conclusions about the mechanisms through which economic changes in migrant-sending areas occur. For instance, Durand and colleagues (1996) examined the effects of foreign earnings, or "migradollars," on development in a set of Mexican communities. Using a case study approach, the authors concluded that the financial resources that migration affords have a positive effect on households' budgets, leading to increased consumer spending and, consequently, additional demand in the labor market for new workers.

Others working in this tradition have taken a decidedly more pessimistic stance, characterizing migration as a sort of self-perpetuating "syndrome" in which sending areas serve as "nurseries and nursing homes" for their mostly migrant labor force (Reichert 1981; Russell 1992; Stuart and Kearney 1981). Rather than triggering economic growth and productive investment, these scholars suggest that a large share of remitted earnings are used for recurrent expenses and consumer goods (Canales 2007), a tendency that increases import demand and fosters an overall climate of economic dependency (Kapur 2005). However, as Taylor (1999) and others point out, this understanding rests on a rather restrictive (and arbitrary) definition of productive spending, which largely ignores the importance of human capital investment. Although it may be true that migrant-sending families rely on remittances to augment consumption, households may also use these monies to finance the continued schooling of younger family members. Recent work by Edwards and Ureta (2003), Taylor and Mora (2006), and López-Córdova (2005) gives some support to this conjecture, showing that remittances relax credit constraints among receiving families, enabling parents to invest more freely and more heavily in their children's education.

Owing to a web of cross-national social relations, remittance-based economies also have important sociocultural implications for nonmigrants. This observation, which identifies *migration as a mode of cultural diffusion* within transnational fields, orients the second approach (Guarnizo 1994; Guarnizo and Díaz 1999; Levitt 1998, 2001; Portes and Rumbaut 1996; Roberts, Frank and Lozano-Ascencio 1999). Rather than concentrating on the net economic effect of migration, research rooted in transnationalism calls attention to a more complex interplay between emigrants and their places of origin. For example, in a multi-sited qualitative study of the transnational Dominican community, Levitt (1998) traced streams of "social remittances" from emigrants to their homelands, contending that the micro-level flow of culture—moving in tandem with material goods, money, and people—makes up an important and often overlooked component of global transfers. Social remittances, according to her argument, include normative belief structures, various systems of practice and conduct, and social capital. Together with repatriated earnings, these resources are thought to alter the incentive structure in sending communities as nonmigrants grow increasingly reliant on support from abroad (Levitt 2001).

Massey, Goldring, and Durand (1994) reached a similar conclusion when examining Mexican sending communities. As migration gains prevalence, these authors argue, places of origin experience extensive cultural and socioeconomic transformations. With time, the inflow of ideas and customs creates a "culture of migration" (Massey et al. 1998), in which aspirations and conceptions of community, familial relations, and economic responsibility take on new "transnationalized" meanings. These ideas and customs, in turn, lead nonmigrants to disinvest in traditional institutions of upward mobility with the intention of capitalizing on more attractive alternatives in other countries. Kandel and Massey (2002:1002) recently corroborated this hypothesis using survey data from the Mexican state of Zacatecas, concluding that intimate exposure to migratory behavior prompts nonmigrants "to look northward rather than locally for opportunities and social mobility."

Finally, a third framework recasts *migration as a behavior that transforms the practices and meanings of family life*. Proponents of this approach point out that young people's actions are often responsive to or constrained by the changing needs, interpersonal dynamics, and composition of migrant-sending families (Hondagneu-Sotelo and Avila 1997; Meza and Pederzini 2008; Parreñas 2001, 2005a). For example, in her case study of the Philippines, Parreñas (2005a) examined the effect of the global "care deficit" on nonmigrant dependents. Although emotional stress and hardship were commonplace in transnational families, children in the study received emotional support from abroad, as well as from parental surrogates within the community. Even more interestingly, at least for the purposes of this paper, Parreñas (2005a) observed a tendency among youth in migrant-sending families to achieve highly in school as a way of reciprocating the hard work of remitters.

Of course, reciprocating the efforts of family members living abroad need not only (or even primarily) entail improved performance in the classroom; young people may also signal their support by taking on new adult roles and responsibilities. If, for instance, the emigration of an adult earner is followed by a short-term decrease in financial resources, youth may choose to invest more heavily in income-generating activities in an effort to stabilize the household economy (Duryea, Lam and Levison 2007; Latapí and González de la Rocha 1995; Meza and Pederzini 2008; Skoufias and Parker 2006). Likewise, if the departure of a family member creates an unmet need within the household for chores or other domestic duties, children may shift their time into home production, resulting in interruptions in their schooling and potential declines in educational attainment (Meza and Pederzini 2008; Parreñas 2005b).

Taken together these three perspectives articulate a general account of international migration and its effect on sending areas. Cross-border movements of capital, people, ideas, norms, and emotional support clearly reorganize the social, cultural, and economic playing field in places of origin. These changes have concrete consequences for individuals—and especially for children. Nevertheless, these literatures contain a number of inconsistencies and opportunities for elaboration. For example, is there a desire to repay efforts abroad by succeeding in school, as Parreñas and other scholars suggest? Or, following the culture-of-migration argument, might young adults opt out of the education system in order to pursue more profitable opportunities in foreign markets? From a financial perspective, does the migration of a family member—and remittances that ensue—serve to offset the opportunity costs associated with nonmigrants' continued enrollment? Finally, moving beyond education, to what extent does the likelihood of joining migrant family members in foreign markets alter nonmigrants' interactions with other homeland institutions, particularly local labor markets?

Although prior research has spoken to some of these questions, the empirical base for doing so has been largely restricted to qualitative studies in specific sending areas. This is problematic for at least two reasons. First, focusing on a limited number of communities in a subset of Mexican states makes it difficult to validate more generic theoretical arguments about migration and its implications for nonmigrants. Although the present research makes no claim to resolving this issue in its entirety, it does represent a deliberate move in that direction. Second, existing studies generally do not distinguish the effect of the treatment (e.g., family member migration) from the effect of unobserved factors that jointly determine migratory behavior and children's educational attainment and economic activity. Instead, most researchers assume that unmeasured variables—including shared family traits, aspirations, and expectations—do not influence both the outcome of substantive interest and family-level migratory decisions. If this position proves to be unjustified, as a handful of recent studies suggest that it might (Hanson and Woodruff 2003; McKenzie and Rapoport 2007; Miranda 2007),[1] then inferences concerning the effect of family member migration may be biased. In the next two sections, I introduce a data set and analytic framework that will help to better address these concerns.

DESCRIPTION OF THE DATA

Data for this analysis were drawn from the Integrated Public Use Microdata Series-International (IPUMS-International) archive, a collaborative project that houses microdata from Mexico's XII General Population and Housing Census, 2000 (Minnesota Population Center 2008). The census, administered in February 2000 by the Instituto Nacional de Estadística, Geografía e Informática (INEGI), was the first to ask respondents to indicate the number of family members who left to go live in another country during the five years preceding enumeration. Respondents were to include persons who might have migrated only for a short time or who might have already returned by the time of the census.

The IPUMS-International sample used a stratified cluster design based on enumeration areas and localities. Constructed from 100% of the long-form questionnaires, the data set represents a 10.6% anonymized sample of the total non-institutionalized population, which amounts to approximately 10 million person records and just over 2.3 million households. The sample was designed to yield representative statistics for all localities with 50,000 inhabitants or more, including all 2,443 of Mexico's municipalities. As is the case with other IPUMS data products (Ruggles et al. 2008), logical edits and probabilistic hot-deck imputation methods were applied in rare instances where values were either inconsistent or missing entirely (Esteve and Sobek 2003). Weights were provided to inflate the sample to the total population.

The Mexican census data, like all data sets, is not without limitations. Relying on a large-scale, cross-sectional sample makes it impossible to fully characterize the dynamic nature of demographic processes like international migration, and it makes it equally difficult to articulate precise statements with respect to underlying mechanisms. Furthermore, variable availability and the wording of particular census items occasionally serve to complicate interpretation. The indicator used to identify migrant member households, for example, lacks the specificity and detail to ascertain which family member migrated, their familial relationship to the focal child, and the exact timing of their departure and, where applicable, return. Although, as I describe in more detail below, steps are taken to mitigate against some

[1]Hanson and Woodruff (2003), for example, found that family member migration is influenced by some of the same hard-to-observe factors that contribute to the educational attainments of nonmigrant children. It is unclear whether the same is true for youths' decisions concerning labor force participation.

of these limitations, the nature and content of these data nevertheless impose restrictions on the scope of analyses that can be conducted.

These imperfections notwithstanding, the Mexican census data promise at least two attractive advantages. First, the census provides a sample of transnational families that is sufficiently large to allow for complex statistical analyses, and, in particular, empirical tests for unobserved selectivity. In all, there are over 1.1 million households with at least one international migrant and close to 300,000 households with more than one. Second, the demographic, social, and economic characteristics that were collected reflect the age, sex, place of residence, family composition and background, migration history, educational attainment, labor force participation, and various types of earnings (including those derived from foreign remittances) for all members of each household. No other data set, at present, has included all of these variables while also using such a large and geographically representative sampling frame.

METHODS

Sample and measures

My analytic strategy draws on a series of multivariate models to estimate the role of family member migration in determining (1) youths' educational attainment, measured in terms of their probability of moving from one educational level to the next; and (2) the likelihood that a youth participates in the labor force. To minimize any complications introduced by youth who themselves migrate or who otherwise leave their family of orientation prior to enumeration,[2] I restricted the sample to adolescents ages 15–18 who were not heads of household or their spouses and who were living with one or more of their nuclear parents.[3] The sample was further limited to include only non-institutionalized and native-born persons, a technique that yielded a weighted sample of 6,433,326 youth from 4,117,484 families. For each youth in the sample, I matched the child's record with variables relating to the child's household, the child's siblings, the head of the household, and the head of household's spouse.

To operationalize the first outcome, I separated the Mexican school system into a sequence of successive educational transitions: from primary (elementary school) to lower secondary (middle or junior high school); and from lower secondary to secondary education (high school).[4] I defined the dependent variable by a successful transition (1 = those who matriculated; 0 = those who elected to stop), given that the child completed the previous level. Individuals who did not complete a given level were deemed "ineligible" for future transitions, and were coded to missing for all subsequent transitions. For instance, lower secondary was coded one for respondents who had completed primary and were enrolled in lower secondary, or any higher level of education; zero for respondents who had completed primary but never enrolled in or completed a year of lower secondary; and missing for respondents who did not complete primary. These transitions represent milestones in the schooling process that signify movement across institutional divisions in Mexico's

[2]In theory, for a youth to count as a migrant *and* enter the analysis sample they would have to (1) temporarily out-migrate; (2) return to Mexico prior to enumeration; and (3) do so before turning age 19. Although the Mexican census data offer no way to formally assess how often this situation occurs, other analyses using different data sources suggest that it is infrequent. Durand et al. (2001), for example, found that over 90% of temporary Mexican migrants are over the age of 18 upon their initial departure. Thus, if 10% of temporary Mexican migrants are 18 and under (thereby meeting the first criterion), one can presume that an even smaller fraction of temporary Mexican migrants satisfy criteria number two and three.

[3]Auxiliary analyses indicated that only a small proportion of Mexican youth start their own household by age 18.

[4]Again, because individuals are increasingly likely to themselves engage in migratory behavior as they approach adulthood, these analyses are not able to assess the final transition in Mexico's education system (e.g., the move from secondary to university-level studies). This is a regrettable but necessary restriction in scope.

educational system. As Mare (1980) and others have shown, one of the advantages of this approach over other strategies, which commonly regress years of schooling on social origins, is that background variables may exert different amounts of influence at various points in a student's scholastic career.

To measure labor force activity, I relied on a binary variable to indicate the child's employment status during the week preceding the census. Because youth labor encompasses a wide range of activities, which vary considerably in terms of time commitment and intensity, I restricted the definition of employment to remunerated labor of greater than or equal to 20 hours per week. Given the necessarily arbitrary nature of any cutoff, I wanted to be as inclusive as possible in recognizing the economic activities of youth and young adults. In examining the distribution of weekly hours spent working, I found that moving the threshold down to 10 or 15 hours would add only a very small fraction of youth to the universe of economically active. The specific variable of interest, a detailed measure of employment, reports the labor force activity of all respondents, even those whose primary activity was something other than work. The measure is thus more responsive to younger workers, particularly school-age adolescents, who only participated in secondary economic activities.

The main "treatment" variable, a binary indicator of family member migration, was constructed from a question asking respondents to report whether one or more family members had engaged in international migration over the previous five years. In order to get a sense for how this largely arbitrary time horizon might influence my results, I ran a parallel set of analyses (not shown) in which migrant-member families were identified as those who acknowledged at least one incidence of family member migration within the past five years *and/or* reported receiving at least one dollar in remittances from family members living abroad. The inclusion of the latter group—which effectively extends the allotted five-year time span outwards—added a relatively trivial number of cases to the migrant-member population, and did not alter my findings in a substantively meaningful sense.

To control for determinants of educational attainment and labor force activity, I included measures of children's demographic characteristics (gender and indigenous group membership), social background, dwelling characteristics, geographic and community-level characteristics, and family composition. Because Mexican youth with more highly educated parents, or with parents who have higher earnings, are likely to spend more time in school and less time working (Levison, Moe and Knaul 2001), the models controlled for social background using variables for total family income, and the maximum of the father's and mother's years of completed schooling.[5] Dwelling characteristics comprise a variety of proxies for the wealth and standard of living in the household that income might not otherwise capture, including dummy variables for the existence of piped sewage disposal, a private telephone line, and the dwelling's flooring material (0 = dirt; 1 = concrete).

The community-level covariates consist of dummy variables indicating the type of locale (urban or rural) and the state of residence. Including these regressors is particularly important given that geographic differences in labor market conditions and educational opportunities are likely to influence the enrollment and employment decisions of youth, as well as family-level migratory behaviors. Unfortunately, due to data limitations, it was impossible to determine whether the family lived in their enumerated area for the entire five-year period in question; although preliminary analysis revealed that only 6.1% of families claimed residence in a different state in 1995. Following Blake (1989) and the extensive

[5]The four levels of total family income—a variable that does not include earnings that children themselves contribute, nor does it include income derived from foreign remittances—correspond to quartiles in the income distribution.

literature on family disruption and schooling outcomes (see, e.g., Astone and McLanahan 1991), the family composition variables contain measures for the number of dependents in the family and their age structure, as well as an indicator for households with only one parent present. Accounting for these determinants gives a rough control for the interaction between, on the one hand, children's employment and enrollment status, and on the other, variability in the care and material needs of the family.

Selection on Unobservables into the Migrant Pool

Finally, it is necessary to address the issue of selectivity on unobserved factors that influence a family's migratory behavior and the outcomes of substantive interest. Since out-migration is a nonrandom process (Feliciano 2005, 2006), and because some unmeasured factor(s) may determine both family's migratory behavior and youths' school and labor market outcomes, models that treat the migratory decisions of family members as exogenous are likely to produce biased and inconsistent estimates. Suppose, for example, that the propensity to matriculate y^* and migrate z^* can be expressed as

$$y_i^* = \sum_k \beta_k x_{ki} + \delta z_i + \varepsilon_i \quad (1)$$

$$z_i^* = \sum_k \beta_k x_{ki} + \sum_k \gamma_k w_{ki} + \mu_i, \quad (2)$$

where, for the ith individual ($i = 1, \ldots, N$), x_k and w_k represent the kth characteristics ($k = 1, \ldots, K$) thought to be associated with matriculation and family member migration, respectively; z is an observed binary indicator of family member migration with coefficient δ; and β_k and γ_k are coefficients corresponding to x_k and w_k. The problem with this system of equations is that the unobservable error terms μ and ε may be correlated, such that Cov(ε, μ | x, z, w) ≠ 0. For instance, if an unobservable emphasis on work ethic within the household were to affect migration strategies and youths' educational outcomes net of observed covariates, Eq. (1) would violate the assumption that explanatory variables be independent and uncorrelated with the error term. Under these conditions the association between migration and educational outcomes would be spurious owing to unobserved attitudes, characteristics, and/or ambitions.

One non-experimental approach to resolving this inconsistency makes use of the endogenous switching regression model. The switching model is designed to demonstrate the effects of a categorical variable, such as family member migration, on an outcome when the outcome and observed variable are thought to be jointly determined by some unmeasured factor or set of factors (Winship and Mare 1992). Social scientists have invoked this procedure in a number of settings, including analyses of curriculum tracking and students' achievement (Gamoran and Mare 1989; Mare and Winship 1988), marital dissolution and women's economic outcomes (Smock, Manning and Gupta 1999), labor market sector and wages (Sakamoto and Chen 1991a, 1991b), interracial contact and racial attitudes (Powers and Ellison 1995), and earnings among Hispanic migrants (Tienda and Wilson 1992).

The switching regression framework can be summarized as follows. Keeping with the matriculation example, let y_{0i}^* and y_{1i}^* represent two latent continuous educational transition outcomes for the ith individual (as indicated by the observed dichotomous variables y_{0i} and y_{1i}), where y_{0i}^* denotes the outcome that would have occurred had the individual lived in a family without migrant members ($z = 0$), and y_{1i}^* denotes the outcome that would have been realized had they resided in a migrant-member family ($z = 1$). As in Eq. (2), let z_i^* be equal to a latent score indexing the *i*th individual's probability of belonging to a migrant-member family, so that $z = 1$ if $z^* > 0$ and $z = 0$ if $z^* \leq 0$. Following conventional notation (see, e.g., Mare and Winship 1988), the resulting structural system may be expressed as

$$y_{0i}^{*}=\sum_{k}\beta_{0k}x_{ki}+\varepsilon_{0i}\text{if}z=0, \quad (3)$$

$$y_{1i}^{*}=\sum_{k}\beta_{1k}x_{ki}+\varepsilon_{1i}\text{if}z=1, \quad (4)$$

$$z_{i}^{*}=\sum_{k}\gamma_{k}w_{ki}+\eta_{0}y_{0i}+\eta_{1}y_{1i}+\zeta_{i}, \quad (5)$$

where β_{jk} ($j = 0,1$) reflects the effect of x_{ki} on the likelihood of matriculating; γ_k denotes the effect of w_{ki} on the likelihood of belonging to a migrant-member family; η_0 and η_1 are parameters representing the effect of (expected) educational outcomes on family migratory decisions; and ε_{1i}, ε_{2i}, and ζ_i are the stochastic components of the model.

Since only a single educational outcome is observed per youth, Eq. (5) must be estimated indirectly through its reduced form:

$$z_{i}^{*}=\sum_{k}\pi_{k}w_{ki}+\varepsilon_{2i}, \quad (6)$$

where $\pi_k = \eta_0\beta_{0k} + \eta_1\beta_{1k} + \gamma_k$ and $\varepsilon_{2i} = \eta_0\varepsilon_{0i} + \eta_1\varepsilon_{1i} + \zeta_i$. The disturbance term in the structural version of the migration selection equation (e.g., ζ) is generally uncorrelated with the errors in the matriculation equations (e.g., ε_0 and ε_1). The disturbance terms in the reduced-form of the model (e.g., Eqs. (3), (4), and (6)), however, are free to correlate with one another. More specifically, ε_0, ε_1, and ε_2 are assumed to follow a trivariate normal distribution, with mean zero and $\text{Var}(\varepsilon_0)=\sigma_0^2$, $\text{Var}(\varepsilon_1)=\sigma_1^2$, $\text{Var}(\varepsilon_2)=\sigma_2^2$, $\text{Cov}(\varepsilon_0, \varepsilon_1) = \sigma_{01}$, $\text{Cov}(\varepsilon_0, \varepsilon_2) = \sigma_{02}$, and $\text{Cov}(\varepsilon_1, \varepsilon_2) = \sigma_{12}$.

The latter two disturbance covariances, σ_{02} and σ_{12}, are indicative of a common but unmeasured factor that influences both migratory behavior and remaining family members' educational outcomes. When $\sigma_{12} > 0$ there is evidence for positive and nonignorable selection on unobservables into the migrant family member pool. Such a finding would suggest that youth in migrant-member families are more likely to matriculate (net of observed social, demographic, and economic variables) than randomly selected individuals from the larger population. When $\sigma_{02} < 0$ there is evidence for positive unobserved selection into the *no*-migrant family member pool, since the covariance implies a negative correlation between successful transitions and belonging to a *migrant*-member family (Mare and Winship 1988). More generally, if either σ_{02} or σ_{12} are nonzero, then estimates obtained using conventional approaches to modeling the impact of family member migration would be biased due to unobserved selectivity.

Provided that there are theoretically grounded reasons for doing so, excluding some of the elements in w from the outcome equations can improve the precision with which estimates of σ_{02}, σ_{12}, β_{0k}, β_{1k}, and π_k are identified (Mare and Winship 1988).[6] For my purposes, findings from prior research suggest at least two plausible restrictions. First, on the basis of studies that stress the salience of social networks and contextual factors to the migratory enterprise (Massey 1986, 1987; Stecklov et al. 2005), let w (but not x) contain municipaland state-level measures of the percent of households with at least one migrant family member and the percent of household receiving remittances. Second, following Durand and Massey (1992) and Root and De Jong (1991), both of whom identify the life-cycle stage of the household as a strong predictor of family migratory behavior, let w (but not x) include indicators of the head of household's age. Excluding these variables from the outcome equations implies that their effect on youths' educational transitions and labor force participation is either weak (conditional on the control variables described earlier and contained in x) or indirect, operating via family-member migration. The data support this assumption, yielding extremely slim empirical associations between the additional "pre-treatment" variables in w and the outcomes of substantive interest.[7]

Eqs. (3), (4), and (6) can be modeled using a multistage estimation procedure (Mare and Winship 1988). In the first stage, a probit model of Eq. (6) provides consistent estimates of π_k. The resulting estimates are used to construct the inverse Mills' ratios shown below

$$\lambda_{0i}=\phi\{\widehat{z_i}\}/\Phi\{\widehat{z_i}\};\text{and}\lambda_{1i}=-\phi\{\widehat{z_i}\}/[1-\Phi\{\widehat{z_i}\}], \quad (7)$$

where $\phi\{\cdot\}$ and $\Phi\{\cdot\}$ signify the normal probability density function and the cumulative normal probability function, respectively.[8] The ratios, which conceptually represent the probability or hazard of assignment to each of the family member migration statuses, are subsequently entered into the matriculation equations (Eqs. (3) and (4), above) as additional regressors:[9]

$$y^*_{0i}=\sum_k \beta_{0k} x_{ki}+\sigma_{02}\lambda_{0i}+\varepsilon_{0i}, \quad (8)$$

$$y^*_{1i}=\sum_k \beta_{1k} x_{ki}+\sigma_{12}\lambda_{1i}+\varepsilon_{1i}. \quad (9)$$

To complete the second stage, Eqs. (8) and (9) are estimated over the subsamples for which y_0 and y_1 are observed.[10] In the next section, I draw on results from these models and from the more naïve approach described in Eq. (1) to (1) demonstrate the extent to which unobserved variables influence family member migration and Mexican youths' interactions with institutions of upward mobility; and (2) quantify the effect of migration net of measured and unmeasured factors.

FINDINGS

Descriptive Results

Table 1 summarizes the descriptive statistics for each of the variables in my analysis, disaggregated according to family migration history. The first three sections provide gross

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[6]It is not necessary that the elements in x and w be entirely disjoint. Thus, in addition to specifying exclusion restrictions, the right-hand side of the selection equation also contains indicators of parental education, age structure and number of siblings, urban-rural status, dwelling characteristics and amenities, and indigenous group membership. See Table A1 for more details.

[7]In addition to examining empirical associations between the outcomes of interest and the pre-treatment variables at the bivariate level, I regressed each of my dependent variables onto the socioeconomic, demographic, geographic, and family compositional variables included in x, as well as the additional predictors in w. In all three models, the resulting parameter estimates for the pre-treatment variables were substantively trivial in magnitude. These results are available upon request.

[8]By construction, z and λ in Eq. (7) are inversely related, such that a larger estimated value for z implies a smaller λ.

[9]A Huber-White sandwich estimator was used in the second stage to correct for heteroscedasticity (White 1980).

[10]The standard errors generated by the switching regressions may be deflated due to the presence of an estimated quantity (the inverse Mills' ratio) in the second stage probit models. Because of the unusually large sample size, however, it is unlikely that any such bias would meaningfully alter my substantive conclusions.

differences in the control variables. Relative to those in no-migrant families, these tabulations suggest that youth in migrant-sending families have parents with lower levels of educational attainment and income, tend to reside in rural homes with fewer amenities, are more likely to be female, and tend to live in households with only one parent present. With respect to the additional pre-treatment variables in the selection equation, which are shown in the fourth panel, it is not surprising that these youth are also likely to reside in municipalities and states with disproportionately high rates of out-migration and remittance receipts.

Of more importance for the present analysis are the terms for education and labor force activity. As the fifth section of the table demonstrates, the transition rate among youth in families with migrants is consistently lower than for those with stationary family members. For instance, roughly three-quarters of eligible youth in no-migrant families made the transition from lower secondary to secondary school, compared to only 59.6% of eligible children in the migrant family member group. Such gross differences are virtually nonexistent, however, when considering patterns of youth labor force participation. In fact, rather than decreasing for adolescents in the migrant family member subgroup, as culture of migration theorists might anticipate, rates of labor force activity are actually higher among youth in non-stationary families. Whether these gross differences hold in a multivariate context, and to what extent unmeasured selectivity shapes the results, will occupy my focus in the remainder of this section.

Results from a Conventional Regression Approach

Table 2 reports uncorrected probit estimates of the family member migration effect on educational transitions and participation in the labor force. Here and throughout I evaluate statistical significance according to the more conservative Bayesian Information Criterion (BIC). The Bayesian approach, which extracts a penalty proportionate to the sample size (Raftery 1995), is well-suited for the present analysis given the unusually large number of cases involved. To ease interpretation, marginal effects are shown in the third column of each panel, indicating the estimated change in the probability of an outcome for a unit change in one of the predictors, holding all others to the sample mean.

The results are suggestive in two respects. First, the coefficient on family member migration, albeit somewhat small in magnitude, is statistically significant and negative in both of the models predicting educational transitions. The probability of matriculating from lower-secondary to secondary school among youth in families with at least one migrant member, for example, is 4.5 percentage points less than it is for their counterparts in nonmigrant families, net of individual, familial, and geographic characteristics Second, if the analyses were to end here, one might reasonably conclude that family member migration has little practical effect on Mexican youths' labor force participation, as evidenced by the statistically significant but substantively modest coefficient in the rightmost column. Although this finding would be generally consistent with the negligible gross differences observed in Table 1, it says nothing about the role unobserved factors play in "matching" youth to migrant and no-migrant member families.

Testing for Nonrandom Selection on Unobservables

To assess whether unobserved factors bias the inferences drawn from the uncorrected models reported above, Table 3 provides results from endogenous switching regressions predicting the probability of successfully completing educational transitions. As discussed earlier, these estimates—unlike those in Table 2—are adjusted for nonrandom assignment to migrant and nonmigrant families on both measured and unmeasured variables. The slope coefficients for the former generally behave similarly regardless of migration status and are consistent with expectations. The only indicator that does not fit this description is indigenous group membership, a measure whose sign and significance varies depending on family member migration status.

The coefficients associated with σ_{02} and σ_{12}, given near the bottom of the table, contain information pertaining to unmeasured factors that influence migration and matriculation. Recall that if unobserved determinants of family member migration are unimportant in explaining differences in children's educational outcomes, these parameters will equal zero. This does not appear to be the case in either of the transition models. Instead, the estimated error covariances are each positive and statistically significant according to the BIC. Substantively, this implies that youth who actually belong to migrant-sending families are *more likely* to matriculate from one educational level to the next than would a random sample of children with identical measured characteristics, whereas young people who live in no-migrant families are *less likely*. Thus, efforts that fail or are otherwise unable to adjust for selection on unmeasured factors—such as the regression analyses reported in Table 2—run the risk of *understating* the magnitude of the negative migration effect.

A similar set of conclusions emerge from the switching regressions predicting labor force participation, summarized in Table 4. Here, again, the parameter estimates on the control variables for migrants and nonmigrants tend to parallel one another quite closely, suggesting that the net effects of familial, demographic, geographic, and socioeconomic characteristics on youths' labor force participation are largely homogenous with respect to family member migration. The exceptions include the effect of piped sewage and the presence of young children in the household, both of which obtain significance for one group (youth in nonmigrant families) but not the other.

Turning to the disturbance covariances at the bottom of Table 4, the fact that σ_{02} and σ_{12} are statistically significant and show a positive sign implies that there is positive selection into migrant families and negative selection into no-migrant families, net of measured variables. Put differently, if a random sample of Mexican youth with equal individual-level and background characteristics were placed in migrant member families, these youth would be *less likely* to work than those who actually reside in such families. On the other hand, if the same sample of youth were matched to no-migrant families, they would be *more likely* to participate in the labor force than the children who in fact live in such households. Either way, these coefficients cast doubt onto the accuracy of the results obtained from the uncorrected probit model and the descriptive analysis presented earlier. Although the exact nature of the selection mechanism cannot be ascertained from these data, it is clear that nonrandom selection on unobserved factors attenuates the negative association that would otherwise obtain between family member migration and youths' labor force participation.

Quantifying the Family-Member Migration Effect

What, then, is the "treatment" effect of family member migration on youths' educational outcomes and labor force participation? To better answer this question, Figures 1–3 plot index sufficient (or control function) estimates of youths' school outcomes and labor market participation (Heckman et al. 1998; Heckman, LaLonde and Smith 1999). The estimator uses parameters from the switching regression models as an apparatus to recover the average treatment effect among those in migrant member families, or what can be thought of as the average treatment effect for the treated. Staying with the variables and notation defined earlier, the formula used to derive the estimates is

$$E\{y_1 - y_0 | x, w, z{=}1\} = x_{ki}(\widehat{\beta}_{1k} - \widehat{\beta}_{0k}) + \frac{\phi\{-w_{ki}\widehat{\pi}_k\}}{\Phi\{w_{ki}\widehat{\pi}_k\}}[\widehat{\sigma}_{02} - \widehat{\sigma}_{12}], \quad (10)$$

where $\hat{\beta}_{1k}$, $\hat{\beta}_{0k}$, $\hat{\pi}_k$, $\hat{\sigma}_{02}$, and $\hat{\sigma}_{12}$ are quantities estimated in Eqs. (3)–(5), and where $\phi\{\cdot\}$ is the normal density function and $\Phi\{\cdot\}$ is the cumulative normal distribution (Heckman et al. 1999:86–91).[11] Because the second stage of the switching regressions used a probit model, Eq. (10) yields estimates that can be interpreted either as effects on *z*-scores, or, alternatively, as effects of family member migration on the cumulative normal probability of the dependent variable net of measured and unmeasured factors. The results graphed below reflect the latter interpretation.

Figures 1 and 2 display the effect of family member migration on youths' likelihood of making educational transitions. On each graph, the horizontal axis gives baseline probabilities of matriculating from one educational level to the next (or various counterfactual scenarios in which there are no migrant family members) and the vertical axis gives the change in probability associated with living in a migrant-sending family. The average treatment effect for the treated is plotted in black; the lower and upper bounds on this effect, which were derived using the 95% confidence interval on the parameter estimates obtained from the switching regression models, are shown in gray.

Two patterns in particular stand out. First, both figures show that family member migration is negatively and strongly related to youths' probability of matriculation, with effect sizes generally ranging from −.10 to −.20 for the transition to lower secondary and −.15 to −.30 for the transition to secondary school. Second, the parabolic shape of the plots implies that the migration effect diminishes gradually as children's baseline probability of matriculating tends toward 0 or 1. In other words, if a young person's eventual school continuation decision is a foregone conclusion—that is, if they are positioned at either end of the horizontal axis—then the experience of family member migration will have less of a bearing on whether they transition than would otherwise be the case.[12]

Figure 3 presents the analogous graph for nonmigrant youths' economic activity. The graph shows an inverse relationship between family member migration and youths' probability of paid employment, a story that stands in contrast to the descriptive analyses presented earlier. Furthermore, unlike the estimate obtained from the uncorrected probit model, the change in probability associated with family member migration appears to be substantial, falling between −.15 and −.35 for the majority of "treated" youth. To better indicate the magnitude of this effect, consider a hypothetical child whose baseline probability places them right on the margin between working and not working. Inspection of the graph reveals that, net of measured and unmeasured characteristics, exposure to family member migration more than halves their probability of labor force participation, from .50 to just over .20.[13]

DISCUSSION

This article has presented analyses of the effects of international migration on Mexicans who remain in their place of origin. Two main results emerged. First, using nationally

[11] I evaluated Eq. (10) using the mean values of the *k* characteristics in *x* and *w*.

[12] In supplementary analyses, I estimated average treatment effects on the treated separately for boys and girls. For both educational transitions, the resulting estimates did not differ significantly by gender. Although not completely unsurprising, this result is consistent with at least some previous work (Kandel 2003; Kandel and Kao 2001), and may reflect, at least in part, Mexico's closing gender gap in compulsory education.

[13] Given what is known about the gender composition of the Mexican labor force, one might reasonably surmise that this pattern reflects heterogeneity in the relationship between family member migration and youths' tendency to engage in paid labor, most notably between nonmigrant boys and girls. That is, females would seem more likely to be situated near the origin of the horizontal axis, and thus be relatively less susceptible to the influence of family member migration. An auxiliary by-gender analysis (not shown) bore this speculation out. Although the family member migration effect was negative for both boys and girls, the index sufficient estimate obtained for the female subsample was smaller in magnitude, resulting in a flatter and less pronounced curve. As mentioned, this finding is not entirely unexpected, particularly given the low rates at which females participate in the Mexican economy and the large differences between boys and girls in terms of their domestic roles and responsibilities (Levison et al. 2001).

representative data from the 2000 Mexican census, I found that the emigration of family members has a large, negative, and statistically significant effect on educational attainment and labor force participation among nonmigrant youth. Second, I found that estimates of these effects—and, in some instances, the inferences that follow—are sensitive to how one deals with selection on unobservables. More specifically, unmeasured heterogeneity works to dampen the estimated effect of family member migration on youths' educational outcomes and labor force participation.

To what underlying mechanism may we attribute these findings? Overall, the empirical evidence just described would seem consistent with a culture of migration argument, in which the intergenerational transmission of migratory expectations reorients nonmigrants away from homeland institutions of upward mobility and toward opportunities in foreign markets. Rather than prolonging their schooling as a result of relaxed financial constraints, the data give support to the notion that immediate exposure to migratory behavior prompts nonmigrant youth to take on a transnationalized view of the opportunity structure. Like Levitt (2001) and others have argued, the path to social mobility for these individuals is not tethered to their ancestral homes and communities, a reality that appears to have an immediate and substantial influence on the educational and economic strategies they enact.

This is not to deny the possibility that other mechanisms are also operating. That family member migration decreases the likelihood that youth participate in local labor markets could, in part, be symptomatic of the ways in which roles and responsibilities are distributed within migrant-sending families. A reduction in paid labor among nonmigrant youth, for example, might reflect an increased demand for housework, childcare, or other (unpaid) domestic duties, and, at the same time, a decreased demand for adolescent earnings. Unfortunately, data constraints prevent me from exploring this possibility further. This is a limitation, of course, and an obvious next step would be to replicate my analyses using data that feature a more detailed indicator of youth labor, preferably one that is sensitive to variation in the intensity and type of work activities (both paid and unpaid) that young people engage in.

Several other aspects of this study also suggest opportunities for further research. First, the cross-sectional nature of the census data used in the foregoing analysis makes it difficult to establish a clear temporal ordering, and thus precludes an outright claim to causality. Nevertheless, insofar as it is unlikely that the outcome variables under consideration have a direct effect on whether family members migrate, the findings are at the very least suggestive that family member migration causally influences—in the classic counterfactual sense of the word—nonmigrant youths' educational outcomes and economic activity. Subsequent work would do well to investigate this issue further, presumably using a longitudinal research design.

Future research also stands to benefit from incorporating more refined measures of municipal-level characteristics, particularly those pertaining to labor market conditions and the availability of educational opportunities. This would in part be useful as an additional robustness check on my results. But it would also help to illuminate whether, to what extent, and in what ways contextual factors moderate the association between family-member migration and non-migrants' work behaviors and schooling decisions. Does the promise of gainful employment in the local labor market, for instance, attenuate the negative migrant family-member effect on youths' economic activity? Likewise, how might the quality, quantity, and accessibility of local schools and educational services change young people's calculus concerning migration, formal education, and socioeconomic mobility?

Finally, researchers could usefully explore the types of questions posed here from a comparative-historical perspective, using similar census microdata sets that are currently available for other countries like Colombia, Ecuador, Panama, and South Africa (Minnesota Population Center 2008). Doing so may help to shed light on whether the migration effects documented above are site specific or amenable to more generic theoretical applications across space and time, as well as contribute further insights into how and under what conditions the cross-border movement of people influences the behaviors and future orientations of individuals who remain in their places of origin. Either way, I hope that the results and analytic techniques described herein will help to move this important line of work forward.

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Appendix Table A1. Parameter Estimates for the Selection Equation Predicting Family-Member Migration

Independent variables	**π**	**|π/SE(π)|**
Demographic characteristics		
Member of an indegenous group	−.156 ***	11.501
Youth's socioeconomic background		
Max of father's and mother's years of schooling	−.044 ***	61.103
Dwelling has piped sewage	.002	.327
Dwelling has private phone	.243 ***	38.165
Dwelling has dirt flooring	−.252 ***	31.117
Geographic characteristics		
Reside in a rural area	.239 ***	34.555
Family composition		
Number of children ages 5 or younger	.040 ***	12.366
Number of children ages 6 to 18	−.008 **	4.532
"Pre-treatment" variables		
Households in state with 1+ migrant family member (%)	.011 ***	13.665
Households in municipality with 1+ migrant family member (%)	.002 ***	16.616
Households in state receiving remittances (%)	.034 ***	31.838
Households in municipality receiving remittances (%)	.055 ***	105.892
Head of household's age	.004 ***	14.770
Constant	−2.052 ***	118.474
χ^2 likelihood ratio (*df* = 13)	44949.527 ***	
n	654,069	

Notes: All parameter estimates represent probit coefficients. Reference categories for the polytomous variables are ethnicity: not a member of an indigenous group; sewage: no piped sewage; phone: no access to private phone line; flooring: concrete; urban-rural status: urban. See text for details.

† Weak evidence;

* positive evidence;

** strong evidence;

*** very strong evidence using the BIC

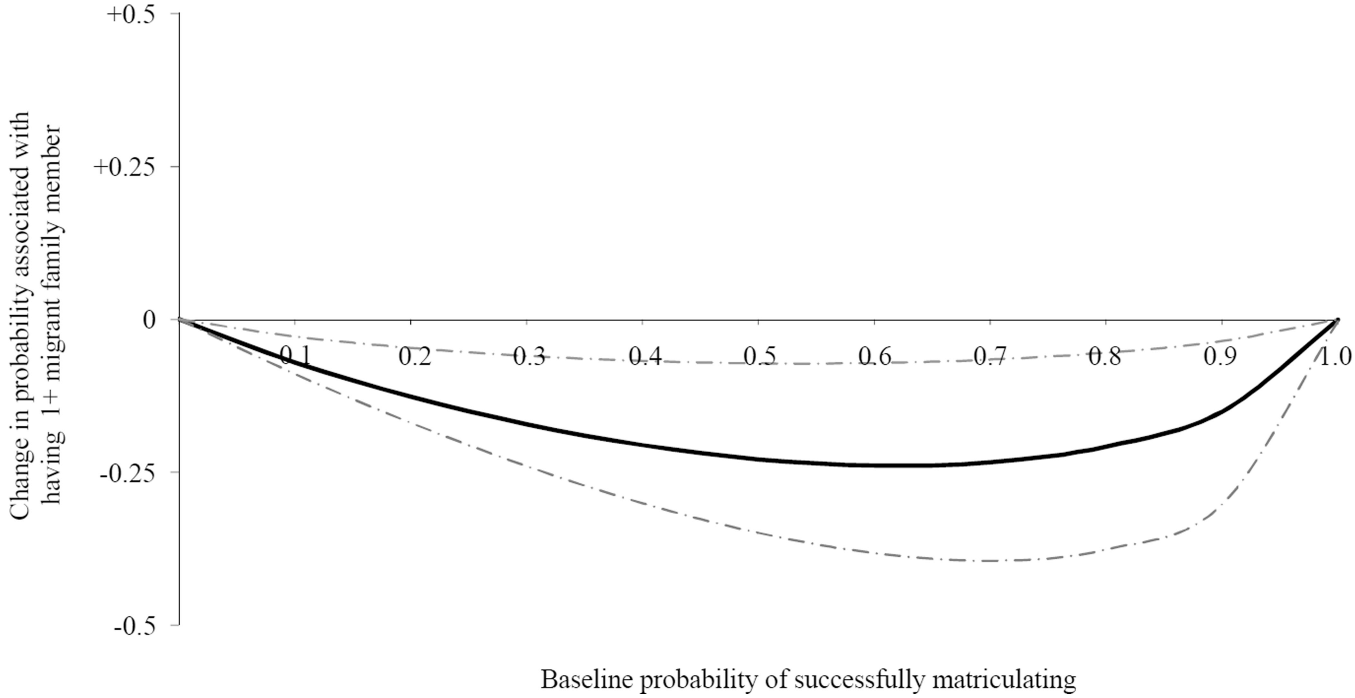

Figure 1. Estimate of the migrant family member effect on the the transition from primary to lower secondary school, Mexican youth ages 15–18 (2000)

Notes: The dotted gray lines provide a measure of uncertainty on the average treatment effect for the treated, which is given by the solid black line. These bands, which do not represent confidence limits in the traditional sense, were derived from the 95% confidence interval on the parameter estimates given in Table 3. See text for further details.

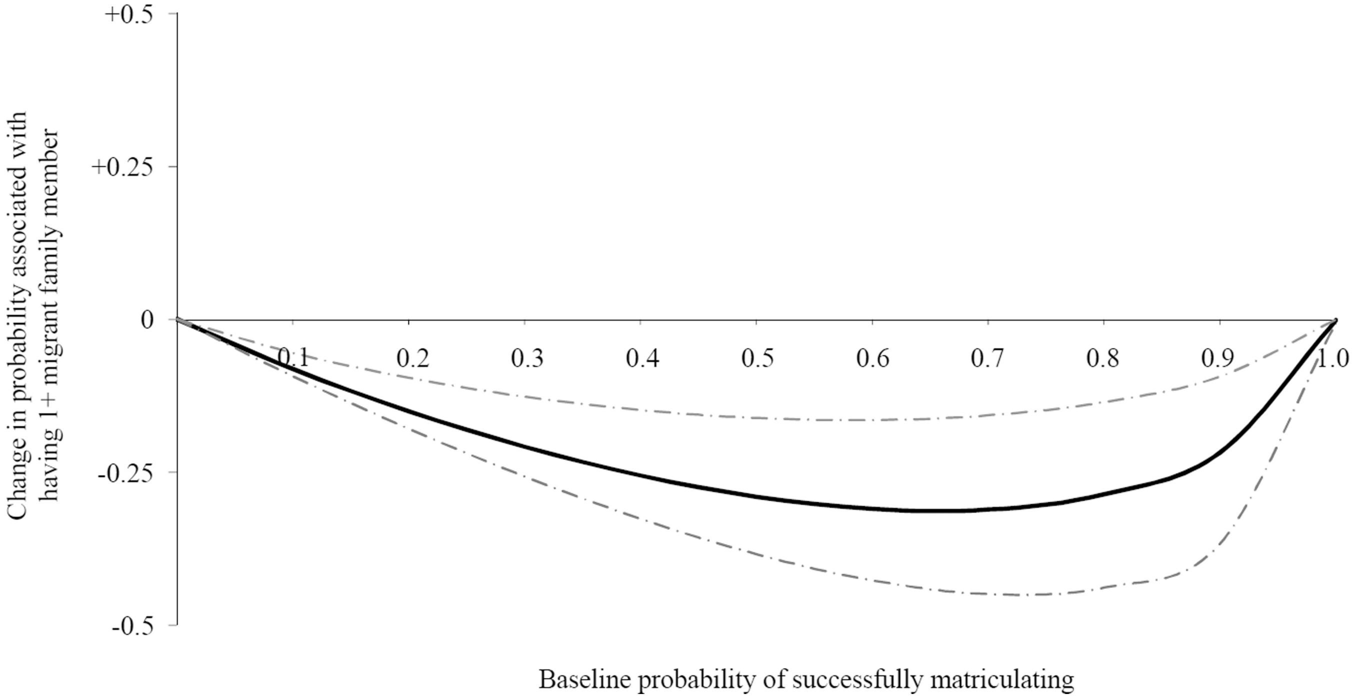

Figure 2. Estimate of the migrant family member effect on the the transition from lower secondary to secondary school, Mexican youth ages 15–18 (2000)

Notes: The dotted gray lines provide a measure of uncertainty on the average treatment effect for the treated, which is given by the solid black line. These bands, which do not represent confidence limits in the traditional sense, were derived from the 95% confidence interval on the parameter estimates given in Table 3. See text for further details.

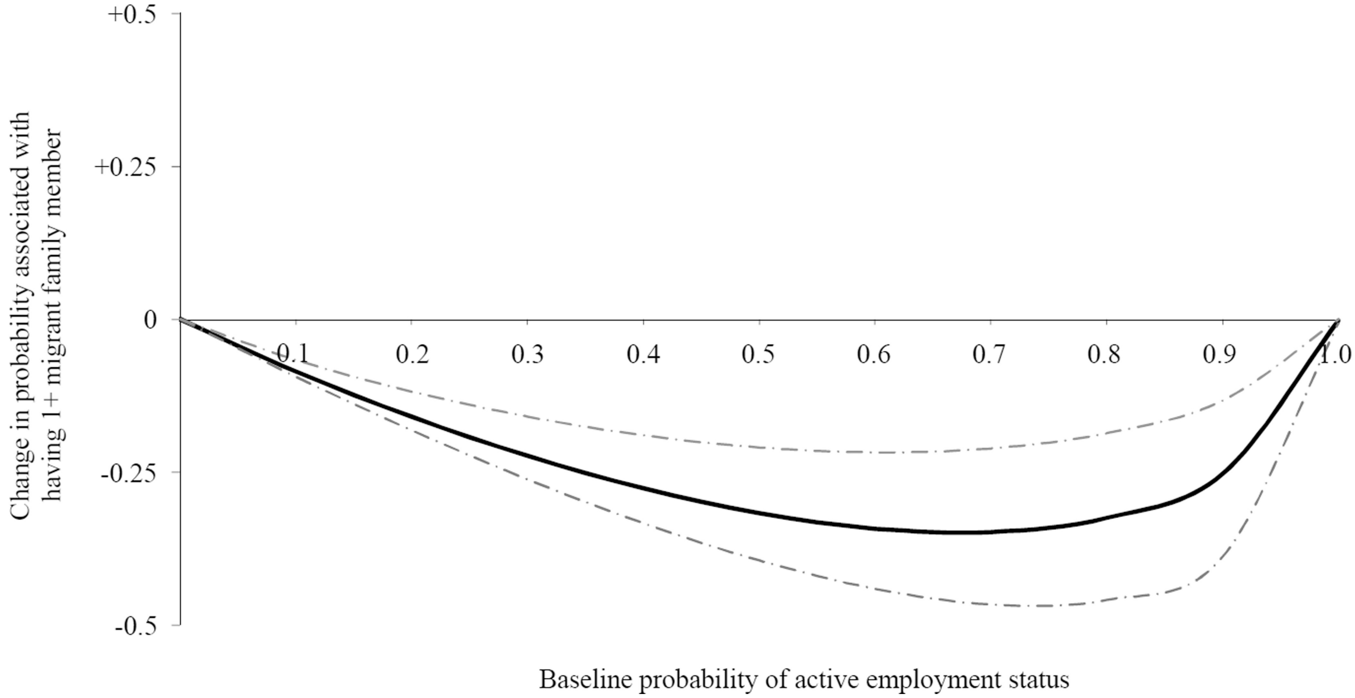

Figure 3. Estimate of the migrant family member effect on the likelihood of labor force participation, Mexican youth ages 15–18 (2000)

Notes: The dotted gray lines provide a measure of uncertainty on the average treatment effect for the treated, which is given by the solid black line. These bands, which do not represent confidence limits in the traditional sense, were derived from the 95% confidence interval on the parameter estimates given in Table 4. See text for further details.

Table 1

Descriptive Statistics for Educational Outcomes, Labor Force Activity, Social Background, Demographic Characteristics, Family Composition, and Selection Variables among Mexican Youth Ages 15–18

Independent and dependent variables	Family member migration	
	No migrants	1 or more migrant
Youth’s socioeconomic background		
Mean parental education in years	7.0	5.0
Mean family income (per month)	4072.6	3477.5
Have piped sewage	72.9%	64.8%
Have private phone	35.0%	31.8%
Have concrete flooring	84.9%	86.7%
Demographic characteristics		
Female	48.6%	52.5%
Member of an indigenous group	6.1%	3.1%
Live in a rural area	25.1%	42.7%
Family composition		
Mean number of children ages 5 or younger	0.4	0.5
Mean number of children ages 6–18	2.8	3.0
Live in single-parent household	15.9%	27.2%
“Pre-treatment” variables		
Households in state with at least one migrant family member	4.8%	6.1%
Households in municipality with at least one migrant family member	5.1%	12.1%
Households in state receiving remittances	3.9%	6.0%
Households in municipality receiving remittances	3.8%	8.8%
Head of household's age	44.4	47.2
Youth’s educational outcomes		
Successfully transitioned from primary to lower secondary school	85.6%	77.0%
Successfully transitioned from lower secondary to secondary school	74.0%	59.6%
Youth’s labor force activity		
Employed 20 or more hours per week	31.8%	35.8%
Weighted *N*	5,964,848	468,478
Unweighted *n*	596,006	58,063

Source: IPUMS-International (Mexico 2000 sample)

Table 2

Uncorrected Probit Models Predicting the Completion of Educational Transitions and Labor Force Participation

Independent variables	Dependent variable: Successful transition from primary to lower secondary			Successful transition from lower secondary to secondary			Active labor force participation		
	β	|β/SE(β)|	∂*P*/∂*x*	β	|β/SE(β)|	∂*P*/∂*x*	β	|β/SE(β)|	∂*P*/∂*x*
Family member migration status									
1+ migrant member	−049 ***	19.144	−.008	−.140 ***	45.850	−.044	−.014 ***	6.762	−.005
Demographic characteristics									
Female	−.018 ***	11.986	−.003	102 ***	66.928	.030	−.616 ***	560.273	−.211
Member of an indegenous group	.057 ***	17.169	.009	.186 ***	39.158	.051	−.014 ***	5.209	−.005
Youth's socioeconomic background									
Max of father's and mother's years of schooling	.105 ***	409.375	.017	.095 ***	412.609	.028	−.061 ***	373.620	−.021
Income of parent(s): $443–1,671 (monthly)	.082 ***	39.660	.015	.018 ***	6.844	.007	.104 ***	60.936	.037
Income of parent(s): $1,672–3,400	.136 ***	62.765	.024	.064 ***	25.020	.023	.129 ***	73.429	.045
Income of parent(s): $3,401+	.235 ***	91.641	.038	.206 ***	74.800	.071	.088 ***	44.796	.031
Dwelling has piped sewage	.213 ***	111.518	.036	.153 ***	65.234	.047	−.029 ***	18.228	−.010
Dwelling has private phone	.399 ***	189.147	.059	.371 ***	204.807	.109	−.205 ***	145.532	−.070
Dwelling has concrete flooring	.173 ***	83.317	.030	.146 ***	47.895	.045	−.036 ***	20.578	−.012
Geographic characteristics									
Reside in a rural area	−.143 ***	74.660	−.024	−.221 ***	94.274	−.069	−.106 ***	67.342	−.036
Family composition									
Number of children ages 5 or younger	−.120 ***	133.518	−.019	−.168 ***	143.248	−.050	.045 ***	62.552	.016
Number of children ages 6 to 18	−.082 ***	161.933	−.013	−.072 ***	117.599	−.021	.057 ***	143.609	.020
Single-parent household	−.083 ***	40.891	−.013	−.039 ***	18.084	−.011	−.090 ***	60.671	−.032
Constant	.386 ***	51.589		−.411 ***	51.284		.320 ***	55.784	
n	571,216			343,248			654,069		
χ^2 likelihood ratio (*df* = 45)	1025829.740 ***			780085.094 ***			801772.499 ***		

Notes: Reference categories for the polytomous variables are gender: male; ethnicity: not a member of an indigenous group; family income: $0–442; sewage: no piped sewage; phone: no access to private phone line; flooring: dirt; urban-rural status: urban; and single-parent household: two-parent household. Indicators of state of residence were included in the model but are not reported here. Results for these estimates are available upon request. Marginal effects are given in the columns headed $\partial P/\partial x$. The derivative is evaluated at the sample mean and, for dummy variables, is for a discrete change from 0 to 1.

† Weak evidence;

* positive evidence;

** strong evidence;

*** very strong evidence using the Bayesian Information Criterion (BIC)

Table 3

Parameter Estimates for Endogenous Switching Regression Models of Educational Transitions

Independent variables	Dependent variable: (1) Successful transition from primary to lower secondary — Youth in migrant-member families: β	|β/SE(β)|	(1) Youth in no migrant-member families: β	|β/SE(β)|	(2) Successful transition lower secondary to secondary — Youth in migrant-member families: β	|β/SE(β)|	(2) Youth in no migrant-member families: β	|β/SE(β)|
Demographic characteristics								
Female	−.010	.759	−.017 ***	11.169	.106 ***	6.373	.099 ***	19.201
Member of an indegenous group	−.017	.415	.045 ***	13.064	−.069	1.391	.151 ***	11.984
Youth's socioeconomic background								
Max of father's and mother's years of schooling	.086 ***	31.778	.103 ***	359.649	.075 ***	25.169	.092 ***	108.865
Income of parent(s): $443–1,671 (monthly)	.042	2.263	.083 ***	37.982	.007	.290	.031 *	3.920
Income of parent(s): $1,672–3,400	.065 †	3.503	.138 ***	59.784	.005	.203	.090 ***	11.152
Income of parent(s): $3,401+	.186 ***	9.019	.232 ***	84.473	.140 ***	5.661	.238 ***	26.350
Dwelling has piped sewage	.254 ***	15.482	.209 ***	103.103	.174 ***	8.498	.163 ***	23.439
Dwelling has private phone	.325 ***	17.835	.430 ***	186.957	.300 ***	14.213	.380 ***	56.607
Dwelling has concrete flooring	.176 ***	8.397	.194 ***	86.018	.153 ***	4.987	.137 ***	15.296
Geographic characteristics								
Reside in a rural area	−.022	1.194	−.129 ***	59.861	−.272 ***	12.206	−.170 ***	24.125
Family composition								
Number of children ages 5 or younger	−.073 ***	9.172	−.123 ***	129.005	−.123 ***	10.670	−.148 ***	38.411
Number of children ages 6 to 18	−.077 ***	17.347	−.084 ***	156.506	−.047 ***	7.926	−.063 ***	31.818
Single-parent household	−.173 ***	11.253	−.066 ***	30.229	−.085 ***	4.540	−.065 ***	8.604
Constant	.134	1.874	.420 ***	50.396	−.715 ***	8.347	−.390 ***	13.622
σ_{02}	—		.261 ***	30.503	—		.327 ***	11.805
σ_{12}	.198 ***	7.303	—		.142 ***	4.522	—	

Independent variables	Dependent variable							
	(1) Successful transition from primary to lower secondary				(2) Successful transition lower secondary to secondary			
	Youth in migrant-member families		Youth in no migrant-member families		Youth in migrant-member families		Youth in no migrant-member families	
	β	|β/SE(β)|	β	|β/SE(β)|	β	|β/SE(β)|	β	|β/SE(β)|
n	50,942		520,274		26,266		316,982	
χ^2 likelihood ratio (*df* = 45)	8129.780 ***		943044.965 ***		4654.439 ***		69484.149 ***	

Notes: All parameter estimates represent probit coefficients. Reference categories for the polytomous variables are gender: male; ethnicity: not a member of an indigenous group; family income: $0–442; sewage: no piped sewage; phone: no access to private phone line; flooring: dirt; urban-rural status: urban; and single-parent household: two-parent household. Indicators of state of residence were included in the model but are not reported here. Results for these estimates are available upon request.

† Weak evidence;

* positive evidence;

** strong evidence;

*** very strong evidence using the Bayesian Information Criterion (BIC)

Table 4

Parameter Estimates for Endogenous Switching Regression Models of Youths' Labor Force Participation

Independent variables	Youth in migrant-member families		Youth in no migrant-member families	
	β	|β/SE(β)|	β	|β/SE(β)|
Demographic characteristics				
Female	−.759 ***	67.759	−.609 ***	529.478
Member of an indegenous group	−.044	1.514	−.048 ***	17.445
Youth's socioeconomic background				
Max of father's and mother's years of schooling	−.046 ***	22.057	−.066 ***	354.839
Income of parent(s): $443–1671 (monthly)	.131 ***	8.513	.098 ***	54.222
Income of parent(s): $1,672–3,400	.170 ***	10.865	.122 ***	66.162
Income of parent(s): $3,401+	.124 ***	7.139	.083 ***	40.145
Dwelling has piped sewage	.039	2.838	−.037 ***	22.470
Dwelling has private phone	−.081 ***	5.095	−.190 ***	123.987
Dwelling has concrete flooring	.024	1.455	−.005	2.856
Geographic characteristics				
Reside in a rural area	−.083 ***	5.510	−.061 ***	34.633
Family composition				
Number of children ages 5 or younger	.022	3.274	.049 ***	64.427
Number of children ages 6 to 18	.033 ***	8.844	.058 ***	138.810
Single-parent household	.051 *	3.953	−.107 ***	67.421
Constant	−.127	2.164	.411 ***	64.693
σ_{02}	—		.426 ***	56.275
σ_{12}	.198 ***	9.478	—	
n	58,063		596,006	
χ^2 likelihood ratio (*df* = 45)	6600.970 ***		760001.763 ***	

Notes: All parameter estimates represent probit coefficients. Reference categories for the polytomous variables are gender: male; ethnicity: not a member of an indigenous group; family income: $0–442; sewage: no piped sewage; phone: no access to private phone line; flooring: dirt; urban-rural status: urban; and single-parent household: two-parent household. Indicators of state of residence were included in the model but are not reported here. Results for these estimates are available upon request.

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