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# Labor migration: a coping strategy used by agricultural farmworkers to mitigate poverty in Mexico

## Migración laboral: estrategia de enfrentamiento de los jornaleros agrícolas para mitigar la pobreza en México

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## Abstract

The occupation of farmworker has historically been highly precarious. Many farmworkers migrate from some of the poorest regions of the country, such as Chiapas, Oaxaca, and Guerrero, to the main agro-exporting regions, including Jalisco, Michoacán, Baja California, and Sonora. As a result, farmworker migration has become a key strategy for mitigating poverty. This article estimates the treatment effect of farmworker migration on poverty using a recursive bivariate probit model that jointly considers the probability of migrating and the probability of being poor, drawing on data from the 2024 National Household Income and Expenditure Survey (ENIGH). The results indicate a statistically significant treatment effect of migration on poverty reduction. When gender and ethnicity are taken into account, the treatment effect is much larger among Indigenous farmworkers (both women and men) than among non-Indigenous workers. The findings support the interpretation of migration as a coping strategy against poverty. A limitation is that ethnic self-identification may be endogenous; addressing this would require a substantially more complex model, although it would be unlikely to change the direction of the main results.

**Keywords**: farmworker; migration and poverty; recursive bivariate probit.

**JEL Codes:** J61; I32; C35

## Resumen

La ocupación de jornalero agrícola ha sido históricamente precarizada. Muchos de los jornaleros migran desde las regiones más pobres del país, como Chiapas, Oaxaca y Guerrero hacia las grandes regiones agroexportadoras como Jalisco, Michoacán, Baja California y Sonara. Esto ha convertido a la migración laboral jornalera en una estrategia de mitigación de la pobreza. El objetivo del artículo es estimar el efecto de tratamiento de la migración jornalera en la mitigación de la pobreza, mediante el uso de un modelo probit bivariado recursivo, el cual considera la probabilidad de migrar y la probabilidad de ser pobre, utilizando la base de datos de la Encuesta Nacional de Ingreso y Gasto de los Hogares (ENIGH) 2024. Los resultados muestran un efecto de tratamiento de la migración en la mitigación de la pobreza. Cuando consideramos diferencias de género y etnicidad, el efecto de tratamiento es mucho mayor entre personas jornaleras indígenas (hombres y mujeres) que en los no indígenas. Los resultados confirman el uso de la migración como una estrategia de enfrentamiento de la pobreza. Una limitación es que la auto-adscripción étnica podría ser endógena, y su corrección complejizaría más el modelo, pero no afectaría el sentido de los resultados.

**Palabras clave:** personas jornaleras agrícolas; migración y pobreza; probit bivariado recursivo.

**Códigos JEL**: J61; I32; C35

## 1. Introduction

The labor precariousness of farmworkers begins when they are transported from their communities of origin to the farming regions where they will be employed, such as western Mexico (Guanajuato, Jalisco, and Michoacán), the northwest (Baja California, Baja California Sur, Sinaloa, and Sonora), and, to a lesser extent, the state of Chihuahua. There, this precariousness is reflected in the absence of employment contracts, low wages, and the lack of benefits that should be non-negotiable, such as access to medical care (see Vanackere, 1988; Posadas, 2015). If we consider that—as has been widely documented—farmworkers come from the poorest states in the country (Chiapas, Oaxaca, Guerrero, and Veracruz), labor precariousness is layered on top of structural poverty. This combination makes farmwork the most precarious occupation among all salaried jobs in Mexico. Although there is evidence that the sustained increase in the minimum wage has helped reduce multidimensional poverty among farmworkers, it remains high[1]: it fell from 65.9 percent in 2018 to 53.7 percent in 2024, while among non-agricultural wage workers it is below 20 percent.

Working conditions in farmworkers' states of origin are so precarious that staying implies not only earning substantially lower wages but also facing a higher risk of unemployment, which would further depress their already rudimentary living standards and deepen their poverty. For this reason, farmworker migration has long operated as a strategy for coping with the poverty conditions that characterize labor-sending regions (Lara & de Grammont, 2000; Guzmán & López, 2005; Lara, 2008; Saldaña & Sánchez, 2012; Salgado-Viveros, 2015). This makes it possible to formally examine the link between migration and poverty by asking whether migrant farmworkers face a lower likelihood of being poor than those who do not migrate.

The objective of this article is therefore to determine whether labor migration reduces poverty among farmworkers in Mexico. We use microdata from the 2024 National Household Income and Expenditure Survey (ENIGH), which for the first time includes a migration variable based on place of birth. This information makes it possible to identify and quantify the treatment effect of farmworker migration on poverty. The treatment-effect framework (Zhang et al., 2010; Stabridis & van Gameren, 2018) allows us to estimate the impact of migration as a mitigator of multidimensional poverty among farmworkers.

Using a recursive bivariate probit model (Wooldridge, 2010), we find that migration reduces the probability of being poor by 8.3 percentage points relative to non-migrants. When we disaggregate this effect by gender and ethnicity (male/female and Indigenous/non-Indigenous), the reduction exceeds 18 percentage points among Indigenous farmworkers, while it is not statistically significant among non-Indigenous workers. This provides evidence that migration is an effective poverty-coping strategy for Indigenous farmworkers. These findings support the case for comprehensive public policies to regulate and improve working conditions for farmworkers, ensure access to training on labor rights, and establish effective coordination mechanisms with employers to guarantee decent working conditions.

## 2.Literature Review

The topic of poverty has been widely studied both in Mexico and in other countries; in the case of farmworkers, it has been addressed through the study of their working and living conditions. In the case of migration among farmworkers, research has focused mainly on the journeys and routes from their places of origin to the most important agricultural regions in the country. In general, these studies rely on qualitative research approaches (Salgado-Viveros, 2015; Velasco, Zlolniski, & Coubès, 2014; Lara, Sánchez, & Saldaña, 2014; Trujillo, 2020).

Following the enactment of the General Law of Social Development (LGDS) in 2004, the National Council for the Evaluation of Social Development Policy (CONEVAL) was created. One of its main objectives was to establish the guidelines and criteria for the definition, identification, and measurement of poverty. On this basis, in 2010 CONEVAL published the *Guidelines and general criteria for the definition, identification and measurement of poverty*[1], which gave rise to the *Methodology for the multidimensional measurement of poverty in Mexico*, since before 2008 poverty was measured using household income with three poverty lines:

1 DOF. (June 16th 2010). General Guidelines and Criteria for the Definition, Identification, and Measurement of Poverty

2010. Access to:https://www.dof.gob.mx/nota_detalle.php?codigo=5146940&fecha=16/06/2010#gsc.tab=0

asset, capabilities, and food or extreme poverty (Boltvinik, 2012).

The official *Methodology for the multidimensional measurement of poverty in Mexico* published by CONEVAL was updated in 2018[2]. According to CONEVAL, poverty is measured in the space of economic wellbeing and also in the space of social rights, with measurements at various levels: national, state, municipal, and urban and rural. Since 2018, the economic wellbeing space has been measured through the indicator called current per capita income, and the social rights space is measured through dichotomic indicators of average educational lag in the household, access to health services, access to social security, quality and spaces of the dwelling, access to basic services in the dwelling, and access to nutritious and quality food. As of 2025, following the dissolution of CONEVAL, poverty measurement has been under the responsibility of INEGI.

To identify people as multidimensionally poor, they must have at least one deprivation in any of the social rights and, in addition, an income below the economic wellbeing line. If a person has one or more social rights deprivations but an income equal to or higher than the economic wellbeing line, they are considered vulnerable due to deprivations. Conversely, if a person has no deprivations but their income is below the economic wellbeing line, they are considered vulnerable due to income. People who have no deprivations and whose income is equal to or higher than the economic wellbeing line are classified as non-poor and non-vulnerable. To identify people considered multidimensionally extreme poor, they must have at least three deprivations and a current per capita income below the minimum wellbeing line or extreme income poverty line. Although this methodology is widely accepted because it was developed by a committee of academics from Mexico and several other countries, it has not been exempt from criticism; one of the most elaborate critiques is that of Boltvinik (2024), who argues that although the methodology is robust, it does not take into account an important component, namely people's time availability, and proposes an alternative poverty measurement approach that includes time poverty and a corresponding threshold. In this approach, being classified as poor does not require that a household has at least one deprivation and an income below a basket; satisfying either one of these conditions is sufficient to be considered poor. For the purposes of this study, the official poverty measurement methodology is used because it is a widely accepted and comparable measure.

Previous studies related to the one proposed here have indirectly addressed poverty measurement among farmworkers, largely due to the lack of information. Hernández (2019), using the 2009 Survey of Farmworkers, analyzes the profile of farmworkers and their precarious working conditions. Posadas (2015, 2018), Morales and Natera (2020), and Flores-Mariscal (2021) make important contributions by showing poverty as an inherent feature of their living conditions and as one of the main reasons why they resort to seasonal migrant labor as a strategy to cope with poverty and extreme poverty. Trujillo (2020), drawing on an ethnographic study of farmworkers in the Valle del Mezquital, confirms that poverty, migration traditions, and violence are the three main drivers of farmworkers migration, and that for many rural and Indigenous families internal migration represents the only survival strategy available.

The situation of farmworkers has worsened over the last forty years due to the deterioration of the Mexican countryside, which has eroded the already scarce earnings of peasants, forcing them to supplement their income by hiring themselves out as laborers or farmwokers for larger agricultural producers or, in the case of those without land, by migrating to agro-industrial regions where they could obtain employment and wages that were not as low as in their places of origin (Vanackere, 1988).

Among the factors that led to the impoverishment of small production units in rural Mexico are the absence of public policies to address their technological needs; lack of access to agricultural credit and to subsidies such as those received by European or U.S. farmers; and the absence of protection by the Mexican State when the country opened to international trade. Within this literature, day-labor migration is conceptualized as a family strategy of economic reproduction in the face of rural poverty and the lack of local alternatives (Velasco, Zlolniski, & Coubès, 2014; Lara Flores, 2016).

Boltvinik (2019) notes that peasant poverty stems from the fact that, in rainfed agriculture,

2 DOF. (October 30 h 2018). Agreement updating the Guidelines and general criteria for the definition, identification and measurement of poverty (Guidelines,2018). Access to: https://www.dof.gob.mx/nota_detalle.php?codigo=5542421&fecha=30/10/2018

there are periods when the labor force remains underemployed, affecting the costs of its reproduction, which are borne by the peasants themselves. For this reason, agricultural prices during the marketing season do not cover the costs of labor, which undermines their wellbeing levels. In this peasant economic logic, migration becomes a central strategy to escape poverty.

At the international level, quantitative evidence on the link between migration and poverty is consistent. Adams and Page (2005), using data from 71 developing countries, find that a 10 percent increase in the share of international migrants reduces on average by 2.1 percent the incidence of extreme poverty, a result that holds even after instrumenting for the potential endogeneity of migration. In the Mexican context, Núñez (2024) estimates, using Bayesian spatial models at the municipal level, that a one-percentage-point increase in the municipal poverty index increases the net migration rate by 1.3 percentage points, showing an expulsive effect: poorer municipalities expel more population, confirming that migration responds to poverty conditions in places of origin. From a vulnerability perspective, De la Fuente (2010) shows that, in rural Mexican households, migration and associated transfers are part of the strategies through which households reduce their exposure to future poverty, although the benefits do not reach the most vulnerable households uniformly, underscoring the importance of analyzing the effect of migration in a differentiated way across groups with different levels of disadvantage.

In many cases, migrant farmworker mobility started as temporary, either pendular or "golondrina" across different agricultural regions. As export agriculture expanded, migration periods lengthened, eventually giving rise to large settlements of farmworkers, as in the case of the municipality of San Quintín, Baja California (Velasco, Zlolniski, & Coubès, 2014). Damián and Pacheco (2019) document the extreme poverty conditions prevailing in the Mexican countryside in labor-sending areas, which explains why migration is sometimes the only way out to alleviate poverty. Lara, Sánchez, and Saldaña (2014) trace the migration routes of poor farmworkers who leave the state of Guerrero and migrate first to Morelos and later to Sinaloa, following the agro-export circuit. Velasco, Zlolniski, and Coubès (2014) document the pattern of farmworker migration from the poorest regions of the country to San Quintín, Baja California, and Velasco and Campos (2018) document the settlement of Indigenous farmworkers in Baja California Sur; in both cases, these processes take place around agro-export activities, as in Sariego (2007) for the case of Sonora. In all these cases, farmworkers migrate to major export regions in search of jobs that share a common feature of labor precariousness but that, despite this, are a better option than remaining in their states of origin and enduring extreme poverty.

Despite the value of these studies, their qualitative approach does not allow for generalizable estimates of the effect of migration on poverty among farmworkers at the national level. In the quantitative domain, Stabridis (2022) estimates labor vulnerability among farmworkers in export crops and finds that the accumulation of disadvantages—being a woman, being Indigenous, or having low schooling—significantly increases the risk of receiving low wages and suffering violations of labor conditions, evidence suggesting that the effect of migration on wellbeing may be heterogeneous across different groups of farmworkers. The present study contributes to this literature by estimating, using nationally representative data from the 2024 ENIGH and a recursive bivariate probit model that accounts for the endogeneity of migration, the treatment effect of migrant status on the probability of multidimensional poverty among farmworkers, distinguishing by gender and ethnic status.

## 3. Methodology

Building on previous work, as well as Stark and Taylor (1991), migration can be understood as a strategy to cope with poverty and, therefore, as one of the determinants of farmworkers' poverty status, as shown in equation (1):

$$y_{1i}^{*} = \alpha_1 + \beta_1 M_i + X_{1i}'\gamma_1 + \varepsilon_{1i} \quad (1)$$

where $y_{1i}^{*}$ is the latent level of multidimensional poverty, which is not observed; instead, one observes $y_{1i}^{*} = 1$ if $y_{1i}^{*} > 0$, and $y_{1i}^{*} = 0$ otherwise. Equation (1) relates poverty to whether the person is a migrant, as well as to control variables such as age, schooling, and Indigenous status, among others. However, migration is not an exogenous event, since individuals decide whether to migrate or not. For this reason, we specify equation (2) for the determinants of migration:

$$M_i^* = \alpha_2 + X_{2i}'\gamma_2 + Z_i'\delta + \varepsilon_{2i} \quad (2)$$

Here, $M_i^*$ is the latent migration variable, which is not observed; instead, one *observes* $M_i = 1$ if $M_i^* > 0$, and $M_i = 0$ otherwise. Equation (2) relates migration to individual-level controls such as age, schooling, and Indigenous status. Using equation (2), we obtain the probability of migrating as a function of individuals' sociodemographic and labor characteristics. To introduce an exogenous source of variation in the determinants of migration, we include fully exogenous variables, such as municipal-level poverty data from the 2020 municipal poverty measurement and information from the 2020 Population and Housing Census. In this way, equations (1) and (2) are linked through the correlation between their error terms (Greene, 2008). Given that both dependent variables are binary, we estimate a recursive bivariate probit model (Wooldridge, 2010). The probabilities of interest are given by:

$$P(y_{1i} = m, M_i = p \mid X_i, Z_i) = \Phi_2(q_m\mu_{1i}, q_p\mu_{2i}; \rho) \quad (3)$$

where the joint probabilities of migration and poverty status follow a bivariate normal distribution, and $\rho$ is the correlation parameter between the error terms. In this expression, $q_m = 2m - 1$ and $q_p = 2p - 1$, with $m, p \in \{0,1\}$, so that when they take values 0 or 1, the four joint probabilities arise: $p_{11}$, $p_{01}$, $p_{10}$ and $p_{00}$. By maximizing the log-likelihood associated with equation (3), we estimate the bivariate probit model and obtain these predicted joint probabilities. The correlation parameter ρ captures the dependence between the migration and poverty equations.

Following Zhang et al. (2009) and Stabridis and van Gameren (2018), the treatment effect of migration on poverty can be estimated from the predicted probabilities obtained with the bivariate probit, as follows:

$$TE = P(y_1 = 1 \mid M = 1, X) - P(y_1 = 1 \mid M = 0, X) \quad (4)$$

From (4), the effect of migration on poverty is the difference between two conditional probabilities: the probability of being poor, conditional on being a migrant, minus the probability of being poor, conditional on being a non-migrant farmworker, which serves as the counterfactual. As shown in equations (5) and (6), both conditional probabilities are obtained as ratios of the four joint probabilities $p_{11}$, $p_{01}$, $p_{10}$ and $p_{00}$:

$$P(y_1 = 1 \mid M = 1, X) = \frac{p_{11}}{p_{11} + p_{01}} \quad (5)$$

$$P(y_1 = 1 \mid M = 0, X) = \frac{p_{10}}{p_{10} + p_{00}} \quad (6)$$

By controlling for the possible endogeneity of migration and modeling the determination of poverty as a function of migrant status, this approach can be interpreted as a treatment-effect framework for the impact of migration on poverty, as suggested by Wooldridge (2010: 594–599).

The sample consists of farmworkers, defined as wage workers employed in agricultural activities. The effect of migration on poverty is estimated for any farmworker in the sample, and separately for Indigenous (non-Indigenous) women (men). The objective is to obtain group-specific probabilities in order to assess whether the poverty impact of migration differs by gender and ethnicity. The advantage of this approach is that the treatment effect of migration on poverty can be identified as long as the unobserved effect of poverty on migration is time-invariant, which is a key assumption in endogeneity-control approaches in two-stage models (Maddala, 1983; Davidson & MacKinnon, 1993).

## 4. Data

To analyze the effect of migration on multidimensional poverty, we use the microdata from the 2024 wave of the National Household Income and Expenditure Survey (ENIGH). ENIGH is a complex household survey designed to provide an overview of household income and expenditure in Mexico, in terms of frequency, type, and amount. It also collects information on household characteristics (and the dwellings in which they live) and on the sociodemographic and labor characteristics of household members. Its complex sampling design ensures representativeness at the national, state, and urban–rural levels, and the survey is fielded in even-numbered years. Beyond its importance for analyzing income and expenditure patterns among Mexican households, ENIGH is the main data source used to construct Mexico's multidimensional poverty measures.

Despite the richness of ENIGH, prior to 2024 it did not collect information on individuals' place of

birth and, therefore, did not allow researchers to distinguish migrants from local residents. The 2024 round includes this variable for the first time. This innovation makes it possible not only to estimate poverty, but also to differentiate individuals by migration status, which is a central contribution of this article.

From ENIGH 2024, we select the sample using the following criteria:

Individuals who are employed, are between 18 and 64 years of age, belong to the Economically Active Population, are adults, and are wage workers.

From this group, we retain only those who report their occupation as farmworker and who work in the agricultural sector, following the definition in Salgado-Viveros and Stabridis (2024). Finally, we classify a farmworker as a migrant if they were not born in the state where they currently live.

Out of 97,682 wage workers in the original sample, the final analytical sample consists of 7,559 farmworkers, representing 4.8 percent of all wage workers. Using the ENIGH sampling weights, this sample represents approximately 1.98 million farmworkers in Mexico.

To examine regional differences in agricultural labor markets, we construct five regions: South–Southeast, Center, Center–West, Northeast, and Northwest, following Stabridis and Salgado-Viveros (2023), as detailed in Table 1. The South–Southeast region is predominantly labor-sending and includes Chiapas, Guerrero, and Oaxaca. In the Center–West, Michoacán, Jalisco, and Guanajuato stand out as major receivers of farm labor, while in the Northwest, the main receiving states are Baja California, Baja California Sur, Sinaloa, and Sonora. Table 1 reports the regional composition of the farmworker sample and the expanded number of farmworkers in each region when applying the ENIGH sampling weights. The South–Southeast is the region with the largest share of farmworkers (34.3 percent). Although the Northwest concentrates only 9 percent of all farmworkers, it is the region where wages are highest and working conditions are comparatively less precarious than in the rest of the country, although still far below the conditions observed among other wage workers.

Of the total number of farmworkers in the sample, 9.0 percent are migrants. When considering the share of migrants by region, it is observed that in the South–Southeast the proportion is 4.0 percent, whereas in the Northwest it is 33.9 percent, the highest among all regions. The sociodemographic variables used are age, a dichotomous variable indicating whether the person is a woman, years of schooling, and a variable indicating whether the person self-identifies as Indigenous; in addition, 2020 census variables are included (the share of migrants at the municipal level), as well as the municipal-level percentages of poverty and extreme poverty in 2020, which serve as an exogenous source of variation in the migration equation.

**Table 1.** Farmworkers by region , 2024

| Region[1] | Sample | Weighted Sample | %. Weighted sample |
|---|---|---|---|
| South-southeast | 1,536 | 677,673 | 34.3% |
| Center | 1,077 | 406,238 | 20.5% |
| Center-west | 2,544 | 615,097 | 31.1% |
| Northeast | 869 | 102,385 | 5.2% |
| Northwest | 1,533 | 177,340 | 9.0% |
| Total | 7,559 | 1,978,733 | |

**Source:** Self-elaboration with microdata of ENIGH 2024, collected by INEGI
**Notes:** South-southeast include: Campeche, Chiapas, Guerrero, Oaxaca, Quintana Roo, Tabasco, Veracruz and Yucatán.
Center include: Ciudad de México, Hidalgo, Estado de México, Morelos, Puebla and Tlaxcala
Center-west: Aguascalientes, Colima, Guanajuato, Jalisco, Michoacán, Nayarit, Querétaro, San Luis Potosí and Zacatecas
Northeast include: Coahuila, Chihuahua, Durango, Nuevo León and Tamaulipas
Northwest include: Baja California, Baja California Sur, Sinaloa and Sonora

Local and migrant workers are similar in age, but migrants include a higher proportion of women than local workers, whereas the latter have slightly more years of schooling (7.2 years) than migrants (6.8 years). Indigenous self-identification is very similar in both groups, who live primarily in rural areas. Regarding the 2020 municipal information (from the population census and the multidimensional poverty measurement), it is observed that, for migrants, the municipal share of migrants in 2020 was higher than for locals, while local workers lived in municipalities with higher levels of poverty and extreme poverty than migrants.

Regarding labor conditions and poverty among farmworkers, Table 3 reports separate statistics for local and migrant workers. The average monthly wage is close to 5,800 pesos, but migrant farmworkers earn almost 24 percent more than local

**Table 2.** Sociodemographic characteristics by migration status, 2024

| Variables | Non-migrant | Immigrant | Total |
|---|---|---|---|
| Age | 38.265 | 38.065 | 38.247 |
| Female=1 | 0.134 | 0.250 | 0.145 |
| Years of schooling | 7.231 | 6.817 | 7.194 |
| Self-declared indigenous=1 | 0.400 | 0.378 | 0.398 |
| Rural area=1 | 0.716 | 0.575 | 0.703 |
| Prop. Immig. By municipality (2020 census) | 0.068 | 0.286 | 0.088 |
| Perc. of poor people by municipality in 2020 | 61.254 | 45.168 | 59.811 |
| Perc. of ext. poor people by municipality in 2020 | 16.745 | 8.452 | 16.001 |

**Source:** Self-elaboration with microdata of ENIGH 2024, collected by INEGI
Multidimensional poverty by municipality in 2020
**Note:** Reports only means

workers, work about 1.5 additional hours per week, and receive hourly wages that are 17 percent higher. One defining feature of farmwork is the very limited access to employment benefits such as health care: only 8.6 percent of farmworkers have access to health benefits, with a lower rate among local farmworkers (7.5 percent) than among migrants (19.1 percent). Household size is similar in both groups. In terms of poverty and extreme poverty, local farmworkers exhibit higher poverty rates than migrants, suggesting that migrant farmworkers tend to earn higher wages, work more hours, and face a lower probability of being poor than their local counterparts, consistent with the idea that migration operates as a poverty-coping strategy.

## 5. Results

Based on the selected sample, we estimate a recursive bivariate probit model with two equations. The first models the determinants of the probability of migrating as a function of individual- and municipal-level characteristics in 2020. The second models the probability of being multidimensionally poor as a function of individual-level characteristics and the (endogenous) migration decision, with the results presented in Table 4. Because the model is estimated by maximum likelihood and is nonlinear, the coefficients cannot be interpreted as marginal effects, as in a linear regression model;

**Table 3.** Labor characteristics by migration status, 2024

| Variables | Non-migrant | Immigrant | Total |
|---|---|---|---|
| Monthly wage (pesos) | 5,656.31 | 7,044.49 | 5,780.83 |
| Hours worked per week | 42.732 | 44.198 | 42.864 |
| Hourly wage (pesos) | 32.139 | 37.588 | 32.627 |
| Health Benefit coverage=1 | 0.075 | 0.191 | 0.086 |
| Total of household members | 4.628 | 4.531 | 4.619 |
| Prop. poverty 2024 | 0.551 | 0.392 | 0.537 |
| Prop. extreme poverty 2024 | 0.149 | 0.069 | 0.142 |

**Source:** Self-elaboration with microdata of ENIGH 2024, collected by INEGI
Multidimensional poverty by municipality in 2020
**Note:** Reports only means

instead, interpretation focuses on their sign, while joint probabilities, conditional probabilities, and treatment effects are computed subsequently. The second column of Table 4 reports the migration equation and the third column reports marginal effects from the poverty equation.

The migration equation indicates that women are more likely to migrate than men, that individuals with more schooling are less likely to migrate, and that those living in larger households are more likely to migrate. Migration does not differ by household headship or by rural residence. With respect to municipal marginalization, the probability of migrating does not differ significantly across municipalities with different levels of marginalization, nor across the agricultural regions defined in this study. Farmworkers living in the Northern Border Free Zone—where minimum wages doubled in 2019—are less likely to migrate. There is a positive association between migration and access to health benefits, and municipalities with higher migration rates in 2020 tend to have a higher current share of migrant farmworkers.

The poverty equation shows that women and more educated individuals have a lower probability of being poor, whereas farmworkers who are household heads and those living in larger households are more likely to be poor. Farmworkers who self-identify as Indigenous have a higher probability of being poor. Migrant farmworkers have a lower probability of being poor, and because migration is modeled in the first equation, this can be interpreted—under the identification assumptions—as a treatment effect of migration on poverty. The estimated correlation

**Table 4.** Recursive bivariate probit model of migration and poverty, 2024

| Variables | Probability of migration | Probability of poverty | Average marginal effects on P(poverty = 1 \| migrant = 1) |
|---|---|---|---|
| Age | 0.002 | 0.055*** | 0.014*** |
| | (0.011) | (0.009) | (0.002) |
| Age squared | -0.000 | -0.001*** | -0.000*** |
| | (0.000) | (0.000) | (0.000) |
| Female=1 | 0.209*** | -0.205*** | -0.069*** |
| | (0.053) | (0.047) | (0.012) |
| Years of schooling | -0.046*** | -0.033*** | -0.004*** |
| | (0.006) | (0.005) | (0.001) |
| Total of household members | 0.024** | 0.151*** | 0.036*** |
| | (0.010) | (0.008) | (0.003) |
| Household head=1 | 0.074 | 0.357*** | 0.084*** |
| | (0.048) | (0.039) | (0.011) |
| Rural area=1 | 0.045 | -0.531*** | -0.138*** |
| | (0.055) | (0.043) | (0.013) |
| Very low marginalization index =1 | 0.346* | -0.743*** | -0.217*** |
| | (0.205) | (0.087) | (0.030) |
| Low marginalization index =1 | 0.158 | -0.528*** | -0.146*** |
| | (0.203) | (0.085) | (0.028) |
| Medium marginalization index=1 | 0.085 | -0.272*** | -0.076*** |
| | (0.193) | (0.084) | (0.026) |
| High marginalization index=1 | 0.097 | -0.105 | -0.035 |
| | (0.180) | (0.087) | (0.026) |
| R. Center=1 | 0.048 | -0.102* | -0.030** |
| | (0.091) | (0.056) | (0.015) |
| R. Center-west=1 | -0.066 | -0.492*** | -0.118*** |
| | (0.087) | (0.051) | (0.016) |
| R. Northeast=1 | 0.053 | -0.298*** | -0.079*** |
| | (0.105) | (0.063) | (0.018) |
| R. Northwest=1 | 0.094 | -0.448*** | -0.121*** |
| | (0.107) | (0.079) | (0.023) |
| Northern border municipality=1 | -0.547*** | -0.307*** | -0.031 |
| | (0.081) | (0.088) | (0.026) |
| Health Benefit coverage=1 | 0.170*** | -0.730*** | -0.198*** |
| | (0.056) | (0.063) | (0.019) |
| Hourly wage (pesos) | 0.003*** | | |
| | (0.001) | | |
| Owner-occupied housing = 1 | -0.378*** | | |
| | (0.046) | | |
| Share of migrants in the municipality (2020 census) | 3.273*** | | |
| | (0.143) | | |
| Municipal poverty 2020 | 0.002 | | |
| | (0.003) | | |
| Municipal extreme poverty 2020 | -0.016** | | |
| | (0.006) | | |
| Self-declared indigenous=1 | 0.337*** | 0.216*** | 0.026** |
| | (0.048) | (0.038) | (0.011) |
| Migrant=1 | | -0.795*** | -0.200*** |
| | | (0.118) | (0.021) |
| Constant | -1.746*** | -0.209 | |
| | (0.352) | (0.187) | |
| Observations | 7,517 | 7,517 | 7,517 |
| Tthetrachorical correlation (athrho) | 0.451*** | | |
| | -0.087 | | |

**Source:** Self-elaboration with microdata of ENIGH 2024, collected by INEGI
**Note**: Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

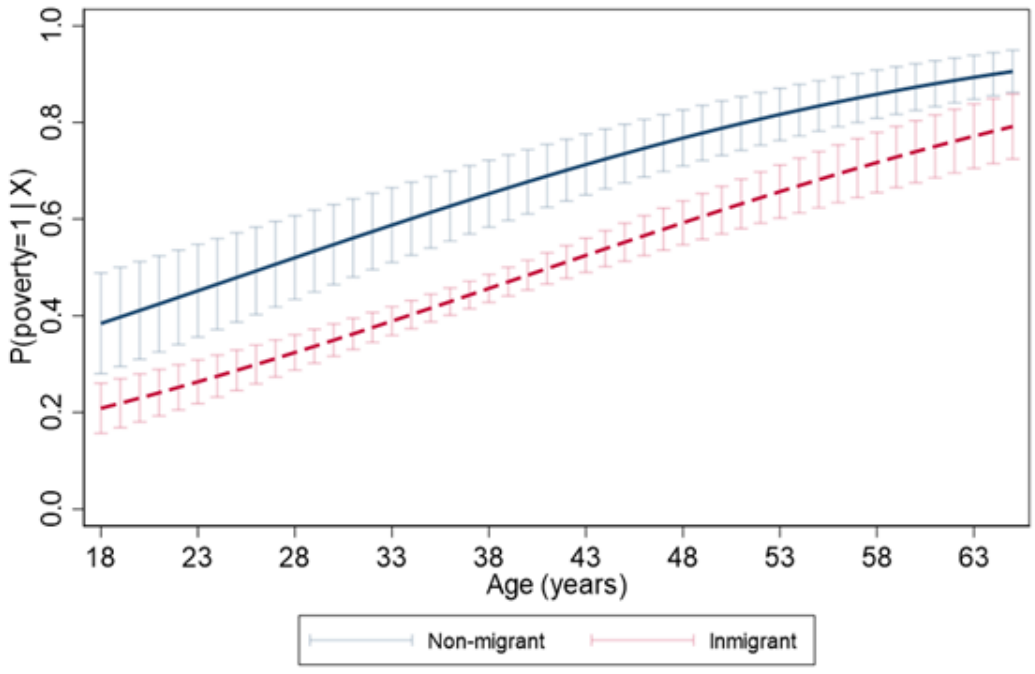
coefficient between the unobservables in the migration and poverty equations (transformed via the arctangent of $\rho$) is 0.451 and highly statistically significant, indicating that modeling both equations jointly is appropriate to address endogeneity. The positive and significant correlation implies that unobserved factors that increase the likelihood of migrating are also associated with a higher propensity to be poor; failing to control for this endogeneity would therefore underestimate the poverty-reducing effect of migration. In terms of marginal effects, women are 6.9 percentage points less likely to be poor than men; each additional year of schooling reduces the probability of poverty by 0.4 percentage points; and migrant farmworkers are 20 percentage points less likely to be poor than non-migrants.

Figure 1 plots the predicted probability of being poor as a function of age, separately for local and migrant farmworkers. For every age within the 18–65 working-age range, local farmworkers display a higher probability of poverty than migrants, and this gap is statistically significant across the entire age range, as the 95 percent confidence intervals do not overlap. The probability of poverty increases with age for both groups, consistent with the gradual deterioration of physical work capacity inherent to farmwork. This pattern visually reinforces the estimated treatment effect of migration reported in Table 5: regardless of age, migration is associated with a lower probability of poverty.

Figure 2 presents the average marginal effects of all covariates included in the poverty equation on $P(\text{poverty}) = 1$, estimated from the recursive bivariate probit. Each dot represents the average marginal effect of a given covariate, and the bars show the 95 percent confidence intervals; covariates whose intervals do not cross the zero line are statistically significant. The largest poverty-reducing effects correspond to the 2020 municipal share of migrants, access to health benefits, and migrant status, with marginal effects of approximately –0.29, –0.20, and –0.20, respectively. By contrast, being the household head has the largest positive effect on the probability of poverty (around +0.09), suggesting that, among farmworker households, headship entails a greater dependent burden. Variables such as owner-occupied housing, Indigenous status, and a high marginalization index are not statistically significant at the 95 percent level.

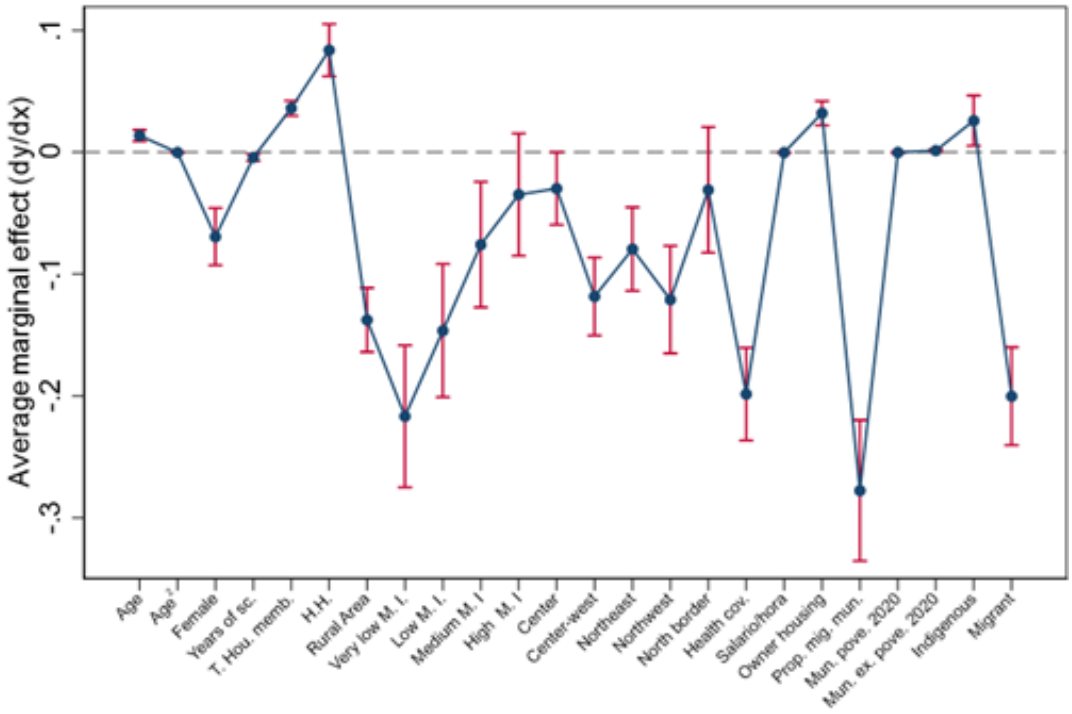
To quantify the treatment effect of migration on poverty, Table 5 reports coefficients, standard errors, and 95 percent confidence intervals. Panel A presents the joint probabilities of migration

**Figure 1.** Predicted probability of poverty as a function of age, by migration status




Source: Authors' calculations based on the recursive bivariate probit model of migration and poverty

**Figure 2.** Average marginal effects of all covariates included in the poverty equation on P(poverty = 1)




Source: Authors' calculations based on the recursive bivariate probit model of migration and poverty

status and poverty: migrant poor account for only 6.4 percent of the sample, non-migrant poor for 36.2 percent, and 57.7 percent of farmworkers are non-poor (migrants and non-migrants combined). Panel B reports the conditional probabilities of being poor among migrants and among non-migrants[3]. The difference between the probability of being poor for non-migrants and that for migrants corresponds to the treatment effect of migration on poverty, which shows that migration reduces the probability of being poor by 8.3 percentage points. This confirms that migration can be interpreted as a poverty-mitigation strategy among farmworkers.

Panels C–F of Table 5 explore treatment-effect heterogeneity by gender and ethnicity. Panel C shows that, among Indigenous women farmworkers, migration reduces the probability of being poor by 18.9 percentage points relative to Indigenous non-migrant women, an effect larger than the average for the full sample. Panel D shows that, for non-Indigenous women, the treatment effect is not statistically significant. Panel E indicates that, among Indigenous men, migration reduces the probability of being poor by 18.4 percentage points compared with Indigenous non-migrant men. Finally, Panel F shows that, for non-Indigenous men, migration has no statistically significant effect on poverty. Taken together, these results indicate that the poverty-reducing effect of migration is concentrated among Indigenous farmworkers—both women and men—consistent with the interpretation of migration as a poverty-coping strategy for the most disadvantaged groups.

M = 1 indicates migrant status (a person who was not born in the state where they currently live); M = 0 indicates a local worker. P = 1 indicates multidimensional poverty; P = 0 indicates non-poverty. Predicted probabilities are obtained from the recursive bivariate probit model estimated by maximum likelihood. Conditional probabilities $P(P=1 \mid M=1)$ and $P(P=1 \mid M=0)$ are computed as ratios of predicted joint probabilities, following Zhang, Zhao, and Harris (2009) and Stabridis and van Gameren (2018). Reported means are simple averages of individual predicted probabilities, calculated over the unweighted sample. Standard errors obtained by bootstrapping with 200 replications and a fixed random seed. *p<0.01,* p<0.05, *p<0.1

3 These probabilities are obtained as joint probabilities from the model results, following the approach used by Zhang et al. (2009) and Stabridis and van Gameren (2018).

**Table 5.** Predicted probabilities and treatment effects for types of farmworkers

| Panel A: Genereal probabilities | Average | Standard error | Confidence interval at 95% | |
|---|---|---|---|---|
| Pr(M=0, P=0) | 0.459 | 0.224 | 0.003 | 0.952 |
| Pr(M=1, P=1) | 0.064 | 0.068 | 0.000 | 0.690 |
| Pr(M=1, P=0) | 0.115 | 0.207 | 0.000 | 0.958 |
| Pr(M=0, P=1) | 0.362 | 0.275 | 0.000 | 0.980 |
| Panel B: Treatment effect of migration on poverty for the full sample | | | | |
| Prob(P=1\| M=1) | 0.357 | 0.005 | 0.347 | 0.368 |
| Prob(P=1\| M=0) | 0.441 | 0.003 | 0.435 | 0.447 |
| Treatment effect | -0.083*** | 0.005 | -0.094 | -0.073 |
| Panel C: Treatment effect of migration on poverty for Indigenous women | | | | |
| Prob(P=1\| M=1) | 0.259 | 0.015 | 0.229 | 0.288 |
| Prob(P=1\| M=0) | 0.448 | 0.013 | 0.422 | 0.474 |
| Treatment effect | -0.189*** | 0.017 | -0.222 | -0.156 |
| Panel D: Treatment effect of migration on poverty for non-Indigenous women | | | | |
| Prob(P=1\| M=1) | 0.235 | 0.009 | 0.217 | 0.253 |
| Prob(P=1\| M=0) | 0.239 | 0.008 | 0.224 | 0.254 |
| Treatment effect | -0.004 | 0.010 | -0.024 | 0.016 |
| Panel E: Treatment effect of migration on poverty for Indigenous men | | | | |
| Prob(P=1\| M=1) | 0.428 | 0.012 | 0.405 | 0.451 |
| Prob(P=1\| M=0) | 0.612 | 0.005 | 0.602 | 0.623 |
| Treatment effect | -0.184*** | 0.012 | -0.206 | -0.161 |
| Panel F: Treatment effect of migration on poverty for non-Indigenous men | | | | |
| Prob(P=1\| M=1) | 0.385 | 0.007 | 0.371 | 0.399 |
| Prob(P=1\| M=0) | 0.387 | 0.004 | 0.379 | 0.395 |
| Treatment effect | -0.002 | 0.007 | -0.016 | 0.012 |

**Source:** Authors' calculations based on the estimation of the recursive bivariate probit
**Notes:** Standard errors obtained by bootstrapping with 200 replications and a fixed random seed.
* p<0.01, p<0.05, * p<0.1

## 6. Discussion and Conclusions

Farmwork in Mexico has historically been highly precarious, characterized by low wages, informality, and the absence of contracts and benefits. Many farmworkers originate from the poorest regions of the country, with a high share of Indigenous population, which has made seasonal farm labor migration one of the main strategies for coping with poverty. The main sending states for migrant farmworkers are Chiapas, Guerrero, Oaxaca, and Veracruz, whose residents migrate temporarily or permanently to export-oriented agricultural regions such as the Center–West (Michoacán, Jalisco, and Guanajuato) and the Northwest (Baja California, Baja California Sur, Sinaloa, and Sonora). To assess whether migration has a treatment effect as a poverty-mitigation mechanism, we estimated a recursive bivariate probit model in which we first model the probability of migrating and then the probability of being poor.

The migration equation shows that women are more likely to migrate than men, while individuals with higher levels of schooling are less likely to do so, and those living in the Northern Border Free Zone have a lower probability of migrating. There is also a positive association between migration and access to health benefits. The poverty equation indicates that women and better-educated individuals have a lower probability of being poor, whereas farmworkers who are household heads and those living in larger households are more likely to be poor.

Farmworkers who self-identify as Indigenous are also more likely to be poor. Migrant farmworkers, in contrast, have a lower probability of being poor, and because migration is estimated in the first equation, this can be interpreted—under the identification assumptions—as a treatment effect of migration on poverty. The positive and statistically significant correlation between the migration and poverty equations implies that unobserved factors that increase the likelihood of being a migrant are also associated with a higher likelihood of being poor.

With respect to the treatment effect of migration on poverty, we find that migrant farmworkers are 8.3 percentage points less likely to be poor than their non-migrant counterparts. When we account for gender and ethnicity, the treatment effect is large and statistically significant among Indigenous farmworkers (both women and men), but not significant among non-Indigenous workers. The stronger treatment effect among Indigenous farmworkers is consistent with the fact that, being on average poorer, they make more intensive use of migration as a strategy to cope with poverty. A limitation of the analysis is that we treat Indigenous self-identification as exogenous; if this assumption did not hold, a more complex modeling strategy would be required to properly address this issue.

These results suggest that, although migration is an effective strategy for coping with poverty among farmworkers—especially those who self-identify as Indigenous—there is an urgent need to implement public policies targeted to this labor group. An effective regulatory framework is needed to improve farmworkers' labor conditions, incorporate cultural relevance, promote awareness of labor rights, and foster greater sensitivity within the agricultural business sector. Ensuring effective access to health care for farmworkers would reduce their labor vulnerability and, consequently, their probability of falling into poverty. Strengthening local economies in farmworkers' communities of origin would also help reduce both their poverty risk and their need to migrate.

Future research should consider a dynamic framework with at least two periods of information on poverty and migration, which would make it possible to estimate the impact of migration on vulnerability to poverty. In addition, a multivariate probit specification that explicitly models the decision to self-report Indigenous ethnicity could be used to examine whether this choice affects the role of migration as a poverty-mitigation mechanism.

An additional limitation of this study is that ENIGH 2024 does not contain individual-level information on remittance sending or receipt, which prevents us from modeling the link between farmworker migration, remittances, and poverty within the present framework. Likewise, because the sample is restricted to farmworkers, there is no variation in employment status that would allow us to estimate a model with that variable as the dependent outcome. Future work could address these extensions using data sources that include direct information on remittances and occupational trajectories, enabling a more comprehensive analysis of farmworker migration as a poverty-mitigation strategy over the life course.

Finally, the definition of migrant used in this study—a farmworker who was not born in the state where they currently live—captures lifetime interstate migration and not necessarily recent labor migration or temporary mobility, which is the most common pattern among farmworkers. This restriction is imposed by the information available in ENIGH 2024 and may include cases of childhood migration or fail to capture pendular mobility across agricultural regions. Future research using data that record the timing and frequency of migration episodes will allow for a more precise measurement of this phenomenon.

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