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Abstract

We study the effects of female immigration on the labor supply of Dominican-born women of different educational levels and family structures. Using individual-level data for 2003–2016 and exploiting geographic variation in early immigrant settlements together with time variation in female immigration inflows, we find that female immigration is associated with a decrease in the labor supply of low educated women at the intensive margin (hours worked) and with an increase in the labor supply at the intensive margin of highly educated women with family dependents (relative to equally educated women without dependents). We fill a gap in the literature on the links between migration and female labor supply, which has focused on developed countries.

Keywords: Immigration, Female labor supply, Developing countries
JEL classification: F22, J16, J22, J61, R23

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1 Introduction

Despite significant improvements during the twentieth century, almost one in two women of working-age in Latin American and the Caribbean (LAC) remains out of the labor market, rendering the participation gap between men and women in the region—at 30 percent—one of the largest in the world (Mateo Díaz and Rodríguez-Chamussy, 2016). Increasing the participation of women in the labor market is not only a goal in itself but it can also have positive spillovers on GDP growth, economic development, and poverty reduction. Understanding the reasons that lie behind changes in female labor supply is highly relevant for policy-making. Several factors have played a role in the increase of female labor supply in LAC ranging from technological change and changes in economic activity to increases in educational attainment and declines in fertility rates (Busso and Fonseca, 2015). However, migration has received relatively less attention as a factor that can affect female labor supply, especially in developing regions.

Some of the same forces that have pushed women to participate in the domestic labor market most likely have also empowered them to migrate internationally and to participate in foreign labor markets. This phenomenon, known as the *feminization* of migration can have sizeable impacts on the female labor supply of destination countries.[1] The literature has identified two channels through which immigration can affect native women's labor supply: (1) complementarity or substitutability in the labor market and (2) substitution for domestic labor. The first one is the traditional channel studied at least since Card (1990) and operates through complementarity or substitutability in production tasks.[2] The second one operates through the availability of domestic services, such as child and elderly care and cleaning. An increase in female (low-skilled) immigration increases the supply of domestic services and reduces their price, allowing native women who were constrained to work because of their domestic responsibilities, to increase their labor supply.

The existing evidence on the link between migration and female labor supply only applies to South–North migration flows to a handful of developed countries such as the US (Cortés and Tessada, 2011; Forlani et al., 2015), Spain (Farré et al., 2011), Italy (Barone and Mocetti, 2011; Peri et al., 2015), Hong Kong (Cortés and Pan, 2013), and Australia,

[1] The feminization of migration is not a global phenomenon, but it can be relevant in some countries or regions. In LAC, the ratio of female to male immigrants is close to 1, however, the stock of female immigrants aged 25–64 grew faster in the past two decades than the stock of males, pushing the ratio of female to male immigrants in that age bracket from 0.88 in 1990 to 0.96 in 2019 (UN International Migrant Stock Database).

[2] If female immigrants are substitutes for native women, an increase in immigration will decrease labor demand for native women, lowering their wages. If on the other hand, female immigrants complement native women in production, an increase in immigration will increase labor demand for native women, increasing their wages.

Germany, Switzerland, and the UK (Forlani et al., 2015). However, it is not obvious *a priori* if the results obtained for developed countries can be generalized to developing countries, where immigrants are likely more similar to the native population than in the former and labor markets' structure and regulations—in addition to social norms—are in general different. Understanding the effects of migration on female labor market outcomes in the developing world is of utmost policy relevance, given that these countries lag behind in terms of female labor force participation, a key factor determining women's economic and social empowerment.

In this paper, we investigate whether and how the influx of female immigrants—most of whom are from Haiti and have low educational levels—affects the labor supply of Dominican-born women of different skills and different family structures. We use individual-level data from the Dominican Republic Labor Force Survey annually for the period between 2003 and 2016 to obtain labor market outcomes for Dominican-born women aged 25–64 such as labor force participation and hours worked, hourly wages, and monthly earnings for those employed. Our empirical strategy combines regional variation in early settlements of immigrants with time variation in female immigrant inflows to construct a measure of *exposure* to female immigrant inflows. This measure allows us to overcome two concerns. The first one is that actual immigration inflows to a region can be related to current economic opportunities, biasing OLS estimates. Our measure of exposure therefore exploits the tendency of immigrants to settle in places where they have a historic network, and is unlikely to be affected by current economic conditions in the regions. The second one is that the Dominican Republic, like several other developing countries, does not have annual sub-national records of immigrant stocks.[3]

We focus on the Dominican Republic (DR) for several reasons. First, the Haiti-DR corridor is one of the most important migration corridors in Latin America and the Caribbean, placing the DR as one of the countries with the largest foreign-born share in the region.[4] Second, although migration from Haiti is still predominantly male, in the past two decades—coinciding with our sample period—there has been an increase in the feminization of these flows; which gives us time variation to estimate their effect. Third, the DR has conducted two immigrant surveys that allow to measure the immigrant population—a large part of which is under irregular immigrant status—more accurately than population censuses. In addition, the DR has a biannual Labor Force Survey (LFS)

[3]This measure of exposure is generally used in the literature as an instrument for *observed* regional migration flows. Since we use this measure directly as the independent variable, our estimates should be interpreted as reduced form.

[4]The foreign-born share in the DR is above 5 percent and approximately nine in ten foreign-born are Haitian. See section 2 for detailed immigration figures and background information.

that is representative at the regional level and that gathers information on a wide array of variables. These factors are not trivial, since one of the reasons South–South migration remains so elusive to study is the lack of high-quality data. Finally, this migration corridor has remained largely understudied in the economics literature.

We find that exposure to female immigration is associated with a decline in the intensive margin of female labor supply (hours worked) and in earnings and that these effects are concentrated on women with low education (less than completed high school). The magnitudes of the effects are not trivial. Our estimates imply that when a province moves from the 25th to the 75th percentile of exposure, low-skilled women decrease weekly hours worked by 0.38 (around 23 minutes) and their earnings decline by 1.95 percent, on average. This finding is contrary to the evidence from developed countries where immigration is associated with an increase in the labor supply of high-skilled women and no negative effects for low-skilled women. This suggests that female immigration to the DR increased competition in the labor market for low-skilled women. We find that this decreased working hours of Dominican domestic workers as well as those of women in other sectors.

When we estimate heterogeneous effects across women with and without family dependents (children and the elderly), we find that exposure to immigration is associated with an increase in the labor supply of high-skilled (completed high school or more) women with dependents relative to equally skilled women without dependents. This differential effect for women with dependents does not take place among low-skilled women. This is consistent with the channel that the literature identified in developed countries: migration reduces the price of household services and allows women who have a higher opportunity cost of time to buy these services and increase their labor supply.

Our paper belongs to a literature that studies the effects of international migration on native women's labor supply decisions. This literature has focused exclusively on high-income migration-recipient countries. Therefore, we fill a gap in this literature by studying an important South–South migration corridor for which the findings for developed countries may not hold.

The literature has found a range of effects of migration on the labor supply of native women depending on their education, family structure, and age. Cortés and Tessada (2011) investigate the effects of low-skilled immigration to the US on the labor supply decisions of high-skilled women—who have the highest opportunity cost of time—and find that migration flows increase labor supply at the intensive margin (hours worked) of women at the top of the female wage distribution while reducing the time they devote to household work. However, they do not find that mothers of young children react differently to

the increase in immigration.[5] This last finding contrasts with Farré et al. (2011) and Cortés and Pan (2013) who, for Spain and Hong Kong, respectively, find a role for family structure. Farré et al. (2011) estimate a positive effect of immigration at the extensive margin (probability of being employed) for women with family responsibilities—caring for young children and elderly dependents—relative to equally skilled women without these responsibilities, they attribute this effect to the fact that immigration increased the local availability of domestic services and reduced their price. Cortés and Pan (2013) compare mothers of young versus mothers of older children and find that the help provided by foreign domestic workers increased the labor force participation of the former group relative to the latter across all education groups, but with the largest effects concentrated on highly educated women. Lastly, Peri et al. (2015) provide evidence that in Italy recent immigrants caused an increase in the supply of workers performing domestic services allowing older native women—who face the increasing burden of taking care of the elderly in their families due to population ageing—to stay longer in the labor force and to retire later relative to men.

Our paper contributes more generally to the small literature on South–South economic migration. Data limitations have prevented the empirical study of this phenomenon even though South–South migration is as important in magnitude as South–North migration (UN, 2019). An early study of the impacts of economic migration on destination labor markets is Gindling (2009) for Costa Rica which was followed more recently by Del Carpio et al. (2015) on Malaysia and Biavaschi et al. (2018) on South-Africa.[6]

The paper by Gindling (2009) deserves a separate mention. This paper studies the effect of Nicaraguan migration in Costa Rica—a large migration corridor in LAC—using a skill-cell approach and estimates separate effects on the earnings of female workers. The author finds supportive evidence of immigrants being substitutes to low-skilled women and complements to high-skilled women and hypothesizes that these effects can be explained by the fact that Nicaraguan women disproportionately work in domestic services. We add to this paper by estimating the effects on labor supply (at the extensive and intensive margins) and by estimating the differential effect of migration on those

[5]Barone and Mocetti (2011) find results mostly aligned with Cortés and Tessada (2011) for Italy. They estimate a positive effect of immigrants who provide household services on female labor supply at the intensive margin but not at the extensive margin and this effect is explained by high-skilled women. Forlani et al. (2015) provide complementary evidence for Australia, Germany, Switzerland, the UK and the US. The authors find that an increase in migrants working in household services increases female labor supply at the intensive margin for high-skilled women and at the extensive margin for low-skilled women. These effects are larger for countries with less supportive family policies.

[6]An even more recent branch of research studies the impacts of *forced* migration in the South, such as those of the Syrian refugee flows to Turkey (Altındağ et al., 2020; Del Carpio and Wagner, 2015) or the Venezuelan exodus to Colombia (Rozo and Vargas, 2019 and Ibañez and Rozo, 2020).

women with family dependents.

The only study to our knowledge that analyzes the labor market effects of Haitian migration in the DR is Sousa et al. (2017), which uses a *mixed* approach that combines geographic variation in the concentration of immigrants with gender-skill variation in their composition and does not find a correlation between immigration and natives' wages.[7] One limitation of this study is that it uses only a cross-section of data for the year 2010, therefore its estimates can only be interpreted as a cross-regional correlation. We add to this study in two ways. First, we use time variation, which allows us to control for (time–invariant) unobserved regional characteristics and second, besides estimating effects on wages we estimate effects on the intensive and extensive margins of labor supply.

The evidence we present differs from that of developed countries. We find negative effects of immigration on hours worked by low-skilled native women and small and insignificant effects on the labor supply of high-skilled women, suggesting that immigration to the DR increased competition in the labor market for low-skilled women. This contrasts with the evidence for the US (Cortés and Tessada, 2011), Italy (Barone and Mocetti, 2011), and Hong-Kong (Cortés and Pan, 2013), where low-skilled immigration had positive effects on the labor-supply of high-skilled women (either at the extensive or intensive margins) and no negative effects on low-skilled women.[8] However, the differential positive effect on hours we find for high-skilled women with dependents—relative to equally skilled women without them—is aligned with the evidence from Spain (Farré et al., 2011).

Our results underline the importance of implementing policies to ameliorate the negative impacts of migration on native women of low education and other vulnerable groups. Such policies could take the form of assisting native women whose hours are reduced in obtaining new qualifications or finding a new occupation. They could also take the form of facilitating and promoting the integration of migrants in the labor market. A great part of South-South migrants, including those in the DR, have irregular immigration status. Granting work permits to migrants could allow them to reach better employment opportunities, reducing the pressure on the most vulnerable native workers. Even when they have a work permit, immigrants may also lack the necessary human capital (including language skills) or can face information frictions in accessing high-quality jobs. Providing immigrants with skills training and information about employment opportunities could also work in that same direction.

[7] In another chapter of the same book, Kone and Özden (2017) simulate the effects of immigration in the Dominican Republic using a structural model.

[8] The results from Spain in Farré et al. (2011) differ somewhat from the rest of developed countries since they find evidence of migration exerting downward pressure on the employment rates of women of lower education.

2 Background: Female Migration to the Dominican Republic

The Dominican Republic (DR) is a unitary state with 31 provinces and a federal district (Santo Domingo). It occupies the eastern two-thirds of Hispaniola island, which it shares with the Republic of Haiti, located on the western third. While the DR is the largest economy in the Central America and the Caribbean region, Haiti is the poorest country in LAC and one of the poorest nations in the world.[9]

These stark differences in standards of living between both countries paired with an unusual geography of two countries sharing one island—for Haitians who want to emigrate, the only other country accessible by land is the Dominican Republic—and with the high frequency of natural disasters in Haiti, all contribute to large migration flows originating in Haiti. According to the most recent immigrant survey in 2017, 5.6 percent of the DR's more than 10 million inhabitants are foreign-born and 87 percent of the foreign-born are from Haiti. These figures place the DR as one of the countries with the largest foreign-born share in LAC.[10]

Women represent, according to the 2017 Immigrant Survey, 39 percent of the foreign-born in the Dominican Republic, a share that increased from 36 percent in 2012. Exploiting information on the year of arrival reported in the immigrant surveys, we can see that this trend in the *feminization* of migration has been taking place since 2005 (see Figure 1). The ratio of female to male immigrants, increased from .44 women per one man in 2005 to .61 women per one man in 2017.[11]

Foreign-born women in the Dominican Republic, 84 percent of which are born in Haiti, are on average younger and have lower educational attainment than Dominican-born women. Table 1 shows these and other characteristics—obtained with the 2010 population census—for Dominican and foreign-born women aged 25–64, according to their level of skill. Foreign-born women are on average four years younger (36 versus 41) than Dominican women, a difference that is larger when we consider the low-skilled sample (i.e. those who have not completed secondary education), where natives are on average

[9] In 2017, the DR had a *per capita* GDP of $7,875 (current) U.S. dollars, ten times that of Haiti, which was $766 (WEO database, IMF).

[10] According to the UN's Population Division estimates for 2019, the foreign-born share in the Dominican Republic is 5.3 percent, a figure only surpassed in the continent by Belize (15.4), Costa Rica (8.3), and Suriname (7.9). There are some outliers in the Caribbean (like Antigua and Barbuda, Bahamas, Barbados, Dominica, and Grenada), which have unusually large foreign-born shares due to their small population size.

[11] We use the terms "Dominican-born" and "natives" interchangeably as we do with the terms "foreign-born" and "immigrants". Although this may seem imprecise, in our context it makes sense given that we do not analyze second-generation immigrants separately due to data limitations (the Labor Force Survey does not ask about the country of birth of parents).

42 years old while immigrants are on average 34 years old. Educational achievement for Dominican-born women is substantially higher. Literacy rates are 89 percent among Dominican-born women, while they are at 61 percent for the foreign-born. The share of women with completed secondary education is 42 percent among natives and 37 percent among all immigrants (while it is only 20 percent among Haitians).[12]

Regarding family decisions, immigrant women have a marriage rate of 22 percent, similar to that of native women, at 23 percent. The marriage rate increases with skill levels, but more so for immigrants. The number of births per woman is similar across groups: 2.9 for native and 2.7 for immigrant women (it is slightly higher for Haitian women, at 3).

In terms of labor market outcomes, employment rates are similar for both groups, however, this masks heterogeneity across educational achievement. For low-skilled women, employment rates are larger for immigrants (at 40 percent) than for natives (34 percent). The opposite happens for high-skilled women, for which employment rates are 54 percent for natives versus 46 percent for immigrants (and 36 for Haitians).[13] The 2010 population census does not include questions on hours worked and on earnings. Therefore, to compare wages across natives and immigrants, we resort to the 2010 Labor Force Survey (LFS).[14] According to information in the LFS, which we do not report in Table 1, hourly wages of native high-skilled women are 53 percent larger than those of high-skilled foreign-born women. This gap reduces to 13 percent for those that are low-skilled. Haitian immigrants, in particular, may be at a disadvantage in the labor market given that their native language is not Spanish but French and Haitian Creole. Research has found that language skills are an important determinant of immigrants' labor market performance (Bleakley and Chin, 2004; Yao and van Ours, 2015). In this sense, it is worrisome that only 27 percent of female Haitian immigrants report in the 2017 Immigrant Survey (ENI) that they speak Spanish well or very well, whereas this share increases to 87 percent among the non-Haitian.

A large majority of working women in the DR are employed in the service sector and this is true for both natives and immigrants. Among native women, 84 percent of those employed work in services compared to 77 among immigrants. However, immigrant women are much more likely than natives to work in agriculture than their native counter-

[12]The information for Haitians in this section should be interpreted with caution given that 42 percent of them have a missing value in the education question.

[13]We have limited information on time-use decisions by women. Results from a special module of the 2016 Household Survey, reveal that women are almost 50 percent more likely than men to perform domestic tasks (93 percent of women perform domestic tasks versus 63 percent of men) and that women dedicate 13.6 hours a week to domestic work whereas men dedicate 5.7 hours. Unfortunately, we cannot assess the impact of immigration on these decisions, since the survey module was only conducted once in 2016.

[14]These figures should be interpreted with caution given that the LFS may not be representative for the foreign-born population.

parts (12 percent versus 1 percent). This is compensated by a larger proportion of native women working in manufacturing. These differences are explained by low-skilled women, since among the high-skilled, the sectoral composition of employment is much more homogeneous. Contrary to what happens in high-income countries, where immigrants disproportionately specialize in domestic services, in the Dominican Republic the share of female workers in those services is similar across native and foreign-born women.[15] In 2010, the share of the employed female population working in domestic services was 16 percent for Dominican women and 17 percent for all foreign-born women (while it was 22 percent for Haitian women).

In conclusion, female immigrants are more likely to be low-skilled and more likely to work than native women. Therefore, we expect female immigrants as a whole to compete in the labor market with low-education native women more than with high-skilled women. However, given that there is a high degree of similarity between high-skilled immigrant and high-skilled native women—both in demographic characteristics and in labor market outcomes—, we do not expect positive effects of immigration on high-skilled native women driven by complementarities with high-skilled immigrant women. In addition, given that female immigrants are not more likely to work in domestic services than natives, we expect modest positive effects on the labor supply of highly educated women—through a decline in the price of domestic services—and negative effects on native domestic workers. We confirm these hypotheses in section 5.

3 Empirical Strategy

Our empirical strategy exploits both regional variation in the historic concentration of the foreign-born population and time variation in the inflow of female immigrants to identify the effects of female migration on the labor market outcomes of native Dominican women.[16] More precisely, we want to estimate the following specification for the time period from 2003 to 2016

$$y_{ipt} = \beta \text{Exposure}_{pt} + \phi_p + \phi_t + \mathbf{X}'_{it}\beta_X + \mathbf{R}'_{pt}\beta_R + \varepsilon_{ipt}. \quad (1)$$

The dependent variable y_{ipt} is a labor market outcome of native woman i, in province p,

[15] For example, in the US, low-skilled immigrant women represent less than 2 percent of the labor force but more than 25 percent of the workers in private households (Cortés and Tessada, 2011).

[16] An alternative to our approach is the so called *skill-cell approach*, which uses variation in immigrants' educational achievement. We cannot use this approach since data availability on educational achievement of immigrants is limited; importantly, we cannot retrieve the educational level of immigrants at the time of arrival.

and year t.[17] These outcomes can be: a dummy for being employed, weekly hours worked, wages per hour, and monthly earnings. The variable Exposure_{pt} is the exposure to female immigration of native women in province p in year t (see below for details). The term ϕ_p is a set of province fixed-effects that capture time-invariant province characteristics. The term ϕ_t is a set of year fixed-effects that capture aggregate shocks that affect all provinces in a given year. The vector $\mathbf{X}'_{it}$ controls for individual characteristics such as age, education, and marital status and the vector $\mathbf{R}'_{pt}$ controls for province characteristics in year t by interacting predetermined values of population, literacy rates, and night light density with year dummies. In this way, we avoid controlling for contemporaneous values of province-level variables that can be endogenous to migration exposure. To analyze differential effects across women of different skills, we also estimate (1) splitting the sample into high-skilled (completed secondary education) and low-skilled women (less than secondary education).

A straightforward approach would be to use immigrant inflows as a share of the population in province p as the independent variable. This approach faces two concerns. The first one is that the location of recent immigrants is non-random because they are probably attracted to regions that have better economic opportunities. Therefore, immigrant settlements can be related to contemporaneous regional demand shocks, which in turn determine labor market outcomes. This would render the concentration of immigrants endogenous, biasing the OLS estimate of β upwards. The second challenge, which is specific to our context, is that there are no systematic yearly data in the Dominican Republic on immigrant stocks at the sub-national level. In conclusion, even a *naïve* OLS regression that relates current immigrant concentration with current labor market outcomes is not possible to estimate in our context.

We use *predicted* rather than actual female immigrant inflows as our measure of exposure to immigration to address these concerns. Exposure is given by

$$\text{Exposure}_{pt} = \frac{1}{\text{FemPop}_{p,93}} \left(\delta_{p,93} \times \text{FemImmFlow}_{C,t} \right) \times 100, \tag{2}$$

where $\text{FemImmFlow}_{C,t}$ is the country-level net inflow of immigrant women to the Dominican Republic in year t, which is equal to the difference between the stock in t

[17] In 1993 the DR had 30 provinces: the current provinces of Distrito Nacional and Santo Domingo, on one side, and Peravia and San José de Ocoa, on the other, were merged. In our estimations, exposure is defined at the level of the provinces existing in 1993 (N=30). Results (not presented here) are qualitatively similar when we define exposure at the level of regions of the labor force survey (N=10). Provinces have an average area of 1,489 squared kilometers, half the average area of a US county (2,911 sq. km). Regions are much larger, with an average area of 4,765 sq. km. See Table B2 in the Appendix for a list of regions and provinces.

and the stock in $t-1$; $\delta_{p,93} = \text{ImmStock}_{p,93}/\text{ImmStock}_{C,93}$ is the share of the national stock of foreign-born (including both women and men) in 1993 living in province p; and $\text{FemPop}_{p,93}$ is the Dominican-born female population living in province p in 1993. [18,19]

This measure distributes contemporaneous inflows of female immigrants at the country-level $\text{FemImmFlow}_{C,t}$, to each province p according to the pattern of immigrants' early settlements.[20] This pattern is captured by $\delta_{p,93}$, the share of immigrants that settled in province p in 1993, a decade prior to the start of our sample period. The predicted inflow of female migrants is then normalized by the province's female native population in 1993. We avoid using contemporaneous population values since they can be correlated with immigration inflows and therefore we fix the population at an initial value. This measure of exposure is inspired by the shift-share instruments proposed by Altonji and Card (1991) and Card (2001).[21] Most of the literature that followed uses predicted inflows as an instrument for actual inflows. Since we do not have data on actual inflows, our estimates should be interpreted as reduced form.

This measure of exposure addresses the first concern by allocating contemporaneous female flows according to early immigrant settlements, exploiting the tendency of immigrants to locate in regions where there are previous enclaves of immigrants from their country of origin. Immigrant networks facilitate the process of job search and improve labor market outcomes (Munshi, 2003). In the Dominican Republic, social networks seem to be an important determinant of migration: 78 percent of the foreign-born in the DR declared in the 2017 National Survey of Immigrants (ENI) that they had family or friends in the DR prior to immigration, this share reaches 80 percent among women and 82 percent among Haitian women. We use all immigrants—rather than just the female—to capture this network effect since female migration can be largely driven by family unification (especially in a country were most of the migrant stock is male). Our measure of exposure captures the exogenous *supply-push* component of immigration flows (Card, 2001) if two identifying assumptions hold. First, that the unobserved factors that determined the

[18] See Figure 2 for a map of the province shares $\delta_{p,93}$, Figure 3 for the evolution of the country-level stock of female immigrants, and Figure 4 for an histogram of exposure at the province-year level.

[19] The literature has used variation in countries of origin of immigrants when constructing measures like 2. We do not use this type of variation since we do not have the data to cosntruct $\delta_{p,93}$ by country of origin. Also, more than 80 percent of the female foreign-born population in the DR come from Haiti. Therefore, $\text{FemImmFlow}_{C,t}$, includes immigrants from all origin countries. Since Haitian immigrants have different characteristics, in Table C1 we show estimates of the baseline model but using the flow of Haitian women only to measure exposure.

[20] An alternative to our measure would use stocks of immigrants instead of flows. We chose flows because our sample period is fourteen years long (2003–2016), which implies that using stocks or cumulative inflows would have bundled together different waves of immigrants. Older waves are more likely to have assimilated as time goes by and can have different effects on the labor market than newer waves.

[21] See Lewis and Peri (2015) for a recent survey of this approach among other approaches.

location of immigrants in a province in 1993, are uncorrelated with future demand shocks in that province. Second, that the contemporaneous aggregate flow of women to the DR is driven by domestic conditions in Haiti (or by aggregate conditions in the DR) and not by demand shocks in a specific province.

It also addresses the second concern—data availability—by combining available census data on initial immigrant shares with data on yearly inflows from immigrant surveys (see section 4.2 for details). Note that in many developing countries there are no systematic records at the sub-national level of immigrant population as the ones that exist, for example, in Europe.[22,23] Therefore, our measure has the advantage that it can be easily applied in contexts where data availability acts as a constraint.[24]

A possible threat to our identification strategy is that provinces that were outperforming others in 1993 continue on a differential path in economic outcomes during the following decades and therefore the shares $\delta_{p,93}$ are confounding the effect of improved economic opportunities with the effect of immigration—a violation of the first identifying assumption. These factors are province-time specific and therefore, are not controlled by province and year fixed effects, ϕ_p and ϕ_t. To ameliorate this problem, we include province characteristics in 1993 that can be associated with future demand shocks, interacted with year dummies. Given the unavailability of sub-national GDP estimates, we use province-level night light density in 1993 as a proxy for the level of economic activity (Henderson et al., 2012; Donaldson and Storeygard, 2016). We also include the provinces' population size and literacy rates in 1993 interacted with year dummies. In robustness exercises we control, in addition, for unemployment rates and the share of skilled women in female population at the regional level in 2003 interacted with year dummies.[25] This strategy also ameliorates the concerns in Goldsmith-Pinkham et al. (2018), who argue that exogeneity of shift-share instruments depends crucially on the exogeneity of the shares.

A second threat, laid out in Jaeger et al. (2018), is that estimates obtained with the shift-share approach could be conflating the short-run effect of immigration on wages with the long-run dynamic adjustment of the economy to a new equilibrium, since settlement

[22]Such records exist for the years in which population censuses are conducted. However, there are two problems with using census data for immigrant concentration in our setting. First, the microdata is available for only two censuses: 2002 and 2010, which does not provide enough time variation in outcomes. Second, the 2010 census questionnaire does not include a question on earnings.

[23]See for example, the datasets used in Farré et al. (2011) for Spain and Peri et al. (2015) for Italy.

[24]Recent research on the effects of the Venezuelan exodus in Colombia also uses predicted inflows due to the unavailability of detailed sub-national data, see for example Rozo and Vargas (2019) and Ibañez and Rozo (2020).

[25]The microdata for the 1993 census is not public, so we can only construct a limited number of variables using data from summary statistics. We use the labour force survey in 2003 to calculate the additional controls, which can only be done at the more aggregate regional level.

patterns of immigrants tend to be persistent. This causes, according to the authors, the coefficient on wages to be biased towards zero. The bias is proportional to the degree of serial correlation of the shift-share instrument. Several comments are due here. First, our estimates on the effects of exposure to migration on earnings and wages are negative, so they could be interpreted as a lower bound for the absolute value of the effect. Second, one of the main takeways from our baseline results is that the low-skilled are relatively more affected than the high-skilled. Under the assumption that the dynamic adjustment to the new equilibrium does not vary with skill levels, then the bias will be of similar magnitude in both groups, rendering our relative result still valid. A similar argument can be applied to the case of our results for women with family dependents. In that case we also care about the differential effect for women with dependents relative to women with no dependents, therefore under the assumption—which seems plausible—that both groups have a similar dynamic adjustment process, the estimate on this differential effect (β_D) should not be biased. Finally, in the DR the high correlation of the δ shares across years is a concern. However, a great part of this correlation is driven by the capital, Santo Domingo. The province of Distrito Nacional, which includes today's capital district, was home in 1993 to 32.6 percent of the foreign-born population. The correlation of the shares in 1993 with the shares in 2010 is 0.97, this correlation drops to 0.79 when we exclude Santo Domingo. To deal with this concern, we show in section 6 that our point estimates (and their corresponding confidence intervals) are virtually unaltered when we exclude this province.

Since exposure varies at the province × year level, we cluster standard errors at that level in our baseline regressions. This choice of clustering accounts for the potential correlation of labor outcomes within a province in a given year. Clustering at the province level, would account for correlation in outcomes within a province across years. We choose the province × year level in the baseline rather than the province level to avoid the problem of having a small number of clusters. In section 6 we show our point estimates are even more precise when clustering at the province level.

Finally, to estimate heterogeneous impacts of immigration according to native women's family structure we follow Farré et al. (2011) and estimate a version of (1) where we interact the shock with a dummy indicating the presence of family dependents

$$\begin{aligned} y_{ipt} = \beta_E \text{Exposure}_{pt} + \beta_D \text{Exposure}_{pt} \times \text{Dep}_{it} \\ + \phi_p + \phi_t + \mathbf{X}'_{it}\beta_X + \mathbf{R}'_{pt}\beta_R + \varepsilon_{ipt}, \end{aligned} \tag{3}$$

where Dep_{it} is a dummy variable equal to one if woman i has family dependents living

in her same household and zero otherwise. We define a family dependent as a child 8 years old or younger or an adult 65 years old or older.[26] Therefore, β_D measures the differential effect of exposure on women with family responsibilities. We expect this effect to be positive at least in the high-skill group, which is the group with the highest opportunity cost of time and that could benefit more from an increase in the supply of domestic services.[27]

4 Data and Measurement

We combine data from Labor Force Surveys, Population Censuses, and specialized Immigrant Surveys to construct our dataset. Our estimation sample consists of repeated annual cross-sections for the period 2003–2016 of Dominican-born women aged 25–64 . Depending on the outcome, we further restrict the sample to women who work. Summary statistics of our estimation sample are in Table 2.[28]

4.1 Labor Market Outcomes

Labor market outcomes come from the Dominican Republic National Labor Force Survey (*Encuesta Nacional de Fuerza de Trabajo*, ENFT), which is representative at the region level.[29] The ENFT is conducted every year in April and October by the Central Bank surveying on average 8,000 households and close to 29,000 individuals in each round. The survey questionnaire includes a household module and an individual module, which provides information on all household members. The survey covers a wide range of topics related not only to the labor market but also to living conditions.

We use the April and October rounds of the ENFT during the period 2003–2016 to obtain outcome variables for native women at the individual level, including: a dummy for being employed, weekly hours worked, hourly wages, and monthly earnings. We also use the data from this survey to control for certain individual characteristics in the regressions such as age, years of education, marital status, and having family dependents in the household.

[26] The term for family dependents without interaction is included in the vector $\mathbf{X}'_{it}$ of individual characteristics.

[27] Although for this purpose, it would be interesting to construct a measure of female immigration specialized in domestic services (as in Barone and Mocetti, 2011), data limitations prevent us from doing so because we do not know the sector of employment of migrants in each year, see section 4.2.

[28] See Table B1 for a complete list of the variables used in the analysis, their definitions, and their sources.

[29] The survey divides the country in 10 regions, which are listed in Table B2.

4.2 Immigration Inflows

We use data from the two existing rounds of the National Immigrant Survey (*Encuesta Nacional de Inmigrantes*, ENI), conducted for the first time in 2012 and repeated in 2017, to obtain annual immigration flows at the country-level. ENI is a nationally-representative survey conducted by the National Office of Statistics (*Oficina Nacional de Estadística*, ONE). In 2017 it surveyed 223,528 individuals in 73,286 households, of which 17,397 were foreign born. ENI asks respondents about their country of birth and their year of arrival at the Dominican Republic.[30]

We use information from the 2017 round to calculate the stock of female immigrants in the country in each year for the period 2002–2016 and denote the measure *raw* because it does not include those immigrants that arrived prior to 2017 but could not be surveyed because they were no longer in the country either because they passed away or because they left the country (either by their own will or by deportation), therefore underestimating the true stock in each year. To correct for this, we adjust the raw measure using an adjustment factor which we calculate by exploiting the information from the first round of ENI, in 2012. More precisely, we look for an adjustment factor, $(1+\gamma)$, that allows to transform raw values of the stock as measured in 2017, $M_t^{\mathrm{R},17}$, to *adjusted* values, by inflating the observed 2017 values to correct for deaths and emigration. The adjusted stock $\widetilde{M}_t$ in any year t is given by

$$\widetilde{M}_t = (1+\gamma)^s M_t^{\mathrm{R},17}. \tag{4}$$

We adjust the immigrant stock in each year by using two different approaches, that we denote *arithmetic* and *geometric*, to calculate the rate γ (which we assume is constant). Details of the calculation of γ are in Appendix A.

A comparison between the different measures of the female immigrant stock is displayed in Figure 3. Both measures, arithmetic and geometric, share the property that they match the value of ENI 2012. The migrant stock adjusted with the arithmetic assumption is reasonably close to the values obtained with the 2010 population census. It is, however, larger than the value from the 2002 census, but this is consistent with reports that the 2002 census considerably underestimated the immigrant population in the Dominican Republic (see Sousa et al., 2017). Therefore, we use the arithmetic measure throughout our analysis and we show in section 6 that baseline results are robust to using the geometric measure.

[30] Since there is a high level of circular migration between Haiti and the DR, the ENI asks respondents with multiple entries to the DR what was the year of their first and of their last entry. We construct our baseline measures using the first year of entry.

In January 2010 Haiti was hit by a magnitude-7.0 earthquake with epicenter just 16 miles from Port-au-Prince, Haiti's capital. The event had devastating consequences in terms of material and human loss and reportedly caused massive displacements of people within the country and to the DR. The exact number of Haitians that crossed the border to the DR in the aftermath of the earthquake remains unknown. The different measures of migration inflows we calculate yield a net entry of 12,000–15,000 Haitian women to the DR in 2010, much larger than the average net entry during 2003–2016 (of 4,000–9,700) but probably lower than reported by the media. The reason our measure yields a lower entry than what most likely took place in that year is that we use self-reported entries in the 2017 and 2012 immigrant surveys. It is possible that a large part of displaced Haitians entered temporarily after the earthquake and were no longer in the DR by the times of the surveys. We think of our measures, therefore, as capturing a more permanent net inflow of migrants.

4.3 Foreign-born Shares

The share of each province in the country's foreign-born population come from the 1993 Population Census of the Dominican Republic. Figure 2 shows a map of these shares. The provinces with the largest shares are Distrito Nacional and Santiago with shares of 32.6 percent and 8.6 percent, respectively. These two provinces are home to the country's largest urban centers.

4.4 Regional Characteristics

Early characteristics of provinces come from different data sources. Night light density in 1993 comes from the Defense Meteorological Satellite Program (DMSP).[31] Population and literacy rates come from the 1993 Census. In robustness checks we control for initial unemployment rates for females and the share of high-skilled women in 2003, which we construct with the ENFT.

5 Results

5.1 Baseline Estimates: Female Migration and Women's Labor Supply

We begin the empirical analysis by estimating (1) via OLS. Table 3, reports a series of estimates of the coefficient β for four labor market outcomes: the probability of working,

[31] We use the average visible, stable lights, and cloud-free coverage composite from DMSP - OLS data.

weekly hours worked, hourly wages, and monthly earnings. For each outcome, the first column includes province and year fixed effects and the second one includes, in addition, province characteristics in 1993 interacted with year dummies. Panel A displays these estimates for the whole sample of women aged 25–64. The effects on the extensive margin (columns 1 and 2) of labor supply are close to zero in magnitude, not statistically significant, and their sign is not robust to changes in specification. The effects on the intensive margin (columns 3 and 4) are negative, although somewhat imprecise since they are only significant at the 10 percent level. Finally, the effects on wages and earnings are also negative, but only the effects on earnings are statistically significant (5 percent level).

However, these effects vary in magnitude and significance when we split the sample according to education levels (panels B and C). For the high-skilled (those with complete secondary or more education), the effect on the extensive margin is positive, and larger than for the whole sample (although not statistically significant) while for the low-skilled (those with incomplete secondary or less education), it is negative and small. On the contrary, the effect on hours—the intensive margin—is negative for both groups but substantially larger in magnitude (and statistically significant) for the low-skilled. The effect on wages is still negative and non-significant for both samples but the point estimate is an order of magnitude larger for the low-skilled. Finally, the negative effect on earnings is also more pronounced for the low-skilled, while it is also estimated more precisely for this group, being highly statistically significant.

In summary, the negative effects of exposure to female immigration on hours worked and on earnings seem to be concentrated in the low-skilled group. High-skilled women experience slightly positive effects on the extensive margin and small negative effects in the rest of the variables but none of these effects are statistically significant. The effects on low-skilled women are not trivial in magnitude. A one percentage point increase in exposure to female immigration is associated—for low-skilled Dominican women—with a decrease of around 2 hours worked per week—equivalent to 5.4 percent—and a 10.3 percent decrease in monthly earnings.[32] Put another way, moving a province from the 25th to the 75th percentile of exposure—which implies a change in exposure of 0.19 percentage points—, decreases hours by 0.38 (around 23 minutes) and earnings by 1.95 percent.

Given the relevance of the Haiti-DR migration corridor (female Haitian migration constitutes around 90 percent of all female migration in the DR), it is likely that the results we find are driven by Haitian immigrants. Unfortunately, we do not have data on the

[32]Note that exposure is not measured in shares but rather multiplied by 100 (equation 2). We calculate the percentage-equivalent of the decrease in hours by dividing 2 by the average number of weekly hours worked by low-skilled women in our sample, which is 37.3. Since earnings are measured in natural logarithms, we apply an exponential transformation to interpret the coefficient.

1993 shares, $\delta_{p,93}$, broke-down by nationality so we cannot construct an exposure measure based entirely on Haitian migration. But we do have data on yearly flows by nationality of origin from the ENI. In table C1 we replace the measure of female immigrant inflows, $\text{FemImmFlow}_{C,t}$, in equation 2 with the annual inflows of Haitian women. We find that results are qualitatively similar to the baseline but the coefficients on hours and earnings are more negative for the low-skilled sample. This results make sense since Haitian-born women in the DR tend to have lower educational achievement than the rest of immigrants (see section 2) which implies they are more likely to compete with native low-skilled women in the labor market.

Why do low-skilled women reduce their labor supply? One explanation is that female immigrants are a close substitute to low-skilled native women and the arrival of immigrants increases the availability of low-skilled labor and therefore, reduces low-skilled wages. The point estimate for wages is imprecise but it implies a decrease of 3.4 percent in wages associated with an increase in exposure of 1 percent. This decrease in wages, increases the opportunity cost of time spent at work—and increases the relative value of time at home, either spent on leisure or on domestic tasks—making some women at the margin reduce their working hours and causing an average decrease in hours for the whole sample. The combined effects of the reduction in wages and the reduction in hours causes a sizeable decrease in earnings.

Before concluding this section, we would like to emphasize that we are not estimating the effect of an increase in regional immigrant inflows but rather the effect of an increase in regional *exposure* to immigration. Exposure measures like ours are generally used in the literature as an instrument for observed regional migration inflows, but—due to lack of data—we use it directly as the independent variable. Although actual flows and exposure are likely to be highly correlated, our estimates should not be interpreted as structural but rather as reduced form.

5.2 The Role of Family Dependents and Domestic Labor

In the previous section, we established that the effects of migration vary substantially across women of different education levels. In this section, we further explore if migration has heterogeneous effects depending on the family structure of women. The idea is—as explained in section 1—that an increase in immigration increases the supply of domestic services, reducing their price (see Cortés and Tessada, 2011 and Farré et al., 2011 for theoretical models). Services such as cleaning or child care become more affordable, allowing women to increase their labor supply. It is reasonable to expect that this effect

will prevail in groups that have higher wages and therefore a higher opportunity cost of time, typically women of higher education. For women of lower skill who tend to earn lower wages, the additional income made by working more hours may not compensate the cost of childcare and other domestic services.

We explore these mechanisms by estimating equation 3. Table 4 displays the results for the high-skilled and the low-skilled samples. We are interested in the coefficient β_D which captures the additional effect on labor supply of having family dependents, which we define as a dummy equal to one if the woman has children 8 years old or younger or adults 65 and older living in her same household. As panels A and B show, having a dependent substantially reduces labor supply and earnings for both high-skilled and low-skilled women. However, the effect of the interaction of exposure to female migration and the dependents dummy on hours worked is positive and significant for high-skilled women while it is negative and non-significant for low-skilled women. The total effect of exposure on hours for women with dependents ($\beta_E + \beta_D$) is positive and equal to 0.29 but is not statistically significant.[33] Finally, the coefficients for the extensive margin of labor supply are small and positive for both groups but not significant.

These results imply that high-skilled women with dependents increase their hours worked as a result of increased exposure to migration, compared to equally skilled women with no dependents. Although the total effect of exposure on hours worked by high-skilled women with dependents is not statistically different from zero, the *differential* effect of exposure on hours worked by high-skilled women with dependents compared to equally skilled women with no dependents is positive. A one percentage point increase in exposure to female immigration is associated with an increase of 1.48 weekly hours for high-skilled women with a dependent relative to equally skilled women with no dependents. In other words, moving a province from the 25th to the 75th percentile of exposure, increases hours worked by a high-skilled woman with a dependent by 0.28 (17 minutes) relative to one equally skilled without dependents. This effect is absent in women of lower education. In fact, for them the effect goes in the opposite direction. As we discussed, for this group migration probably increases the opportunity cost of working and being away from their families, and they respond by decreasing the intensive margin of labor supply.

The results we find for high-skilled women could be explained by female immigration increasing the supply (and lowering the price) of household services or by complementarities in production. Although we cannot rule out the existence of complementarities, we show suggestive evidence of a household channel: this is that the increase in female

[33]The p-value of the F-test for joint-significance is equal to 0.7.

immigration led to an increase in the supply of domestic services.[34] Ideally, one would use data on the price of these services. Since that data is not available for the DR, we resort to the LFS to construct proxies for their availability and price. Unfortunately, this survey is only representative at the region level, so we can only construct these variables for 10 regions in each survey round. As a measure of the relative supply of domestic workers we use the region's share of female domestic workers in total female employment.[35] As proxies for prices of domestic services we use the region's average wage and average earnings of female domestic workers. As Columns 1 and 2 of Table show, regions that were more exposed to female immigration experienced an increase in the share of domestic workers in total female employment. A one percentage point increase in exposure leads to a 2 p.p. decrease in the share of domestic employment. The regressions where the dependent variable is the log hourly wage of domestic workers yield a large negative point estimate (a one p.p. increase in exposure to female immigration leads to an 11 percent decline in domestic wages) but that is not precisely estimated. This could be explained by lack of power to reject the null hypothesis due to the low number of observations in the regressions (N=280).[36] It could also be due to measurement error in hours (we calculate hourly wages by dividing earnings by hours worked). Therefore, we also analyze regional average earnings of domestic workers, which are less prone to measurement error. As column 3 shows, earnings for domestic workers drop significantly: 21 percent per one p.p. increase in regional exposure.

As we did in the previous subsection, we analyze if our results are driven by Haitian female migration. Table C2 shows results for equation 3 but using the inflow of Haitian women in the exposure measure. We find results qualitatively similar to those of Table 4 but point estimates for β_D for hours worked in the high-skilled sample are larger. This is reasonable since Haitian women, as we show in section 2 are more likely to work in domestic services (22 percent) than other immigrants (17 percent).

Finally, we can explore if the fall in hours worked by low-skilled native women that

[34]The existence of a household channel has been argued in the literature by comparing the labor supply decisions of women to those of men. This is problematic since, as noted in Cortés and Tessada (2011), men's time use decisions can also be affected by the increased availability of domestic services (either directly or through their spouses' decisions). Taking this caveat into account, when we use men as a control group we find that high-skilled women with dependents compared to similar men, still supply more hours as a consequence of increased migration but the effect is not precisely estimated. We also find that high skilled women have not received higher wages as a consequence of increased migration compared to similar men, therefore, it is unlikely that our results are explained by immigrants being better complements to high-skill women than to high-skilled men.

[35]We define domestic workers as those employed in ISIC's Rev. 3 activity 9500 "Activities of private households as employers of domestic staff".

[36]One of the survey rounds in April 2004 has no reported earnings for women in domestic services in the El valle region, therefore the number of observations in columns 2 and 3 is 279 instead of 280.

we document in section 5.1 is explained by increased competition in non-domestic occupations, in domestic occupations, or in both. As we showed in section 2, a large share of native Dominican women (16 percent) work in domestic services. Therefore, we can expect some competition between immigrants and low-skilled natives happening in the domestic service sector. This is in contrast to findings in developed countries, where the effects of immigrants specialized in domestic services are not negative for low-skilled native women, probably because immigrants are more likely to work in the domestic service sector than natives (see Barone and Mocetti, 2011). Table 6 shows an estimation of our baseline model for the sample of low-skilled women but interacting the exposure measure with an indicator of domestic worker. The point estimate for the effect of *Exposure* on hours worked is negative which means that non-domestic workers (the control group) experience a decline of 1.6 working hours per one p.p. increase in exposure. The interaction with domestic is negative and insignificant, however the total effect for domestic workers is negative (a decline of 2.7 hours per one p.p increase in exposure) and significant at 5 percent.[37] We interpret these results as suggestive evidence that female immigration increased competition with low-skilled natives working both in domestic and non-domestic occupations.

6 Robustness

We address potential concerns with our empirical strategy by conducting a series of robustness checks on the baseline results.[38]

Clustering We check that our baseline results are robust to clustering standard errors at the province level. As we mentioned before, we have 30 clusters in this case which could be considered a small number. As Panels A1 and B1 in Table 7 show, clustering at this level makes all the estimates more precise and therefore does not affect the statistical significance of our main results.

Additional regional shocks Our baseline results control for population, night-light density, and literacy rates at the province-level in 1993 interacted with year dummies. We show that our results are robust to including in Panels A2 and B2 of Table 7 female unemployment rates and the share of women with secondary education at the region level

[37] The p-value of the F-test for joint-significance of the coefficients for *Exposure* and *Exposure* × *Domestic* is equal to 0.03.

[38] All the results for the specification with family dependents (equation 3) are also robust to the checks conducted in this section and are available upon request.

in 2003 interacted with year dummies. Data availability prevents us from calculating these variables at the province level in 1993 (see section 3).

Adjustment of migration inflows As we explained in section 4.2, we calculate the stock of female immigrants using the year of arrival in the 2017 immigrant survey (ENI 2017) but adjusting for deaths and re-emigration exploiting information from ENI 2012 on the actual stock of that year. Our baseline analysis is done using an arithmetic adjustment in equation (A2). In Panels A1 and B1 of Table 8, we show that baseline results are robust to using the geometric adjustment instead.

Measure of past immigrant settlements Our measure of exposure to female immigration in equation (2) uses the past settlement pattern of all immigrants—female and male—to distribute the inflows of female immigrants. An alternative explored in Barone and Mocetti (2011) is to use the settlement pattern of men immigrants, which according to the authors, could be more exogenous since the supply-push factors for men and women are different. We re-estimate our baseline model using the settlements of male immigrants to construct the shares, $\delta_{p,93}$. Panels A2 and B2 in Table 8 show that baseline results are qualitatively similar when using this definition of exposure, however the effects on hours and earnings are more negative than in the baseline. To remain conservative, we prefer as a baseline the results in Table 3.

Excluding metropolitan areas All our estimates are virtually unaltered when we exclude from the sample the provinces of Distrito Nacional and Santiago. Distrito Nacional, which in 1993 included the capital, concentrates 32.6 percent of the foreign-born in the country and Santiago concentrates 8.6 percent of the foreign born. This alleviates the concerns in Jaeger et al. (2018), as explained in section 3. Results are shown in Panels A1 and B1 of Table 9.

Excluding certain years Our sample period (2003–2016) spans 14 years and there are other events that took place during that period that could have an influence on labor market outcomes of natives. One such case, is the banking crisis of 2003–2004 which prompted a notable fall in real earnings. Another, is the massive regularization of the migratory situation of thousands of Haitians without documentation that took place in 2014 through the National Plan of Regularization (and the passage of Law 169–14). Although these events should be controlled for by the year fixed effects, we check that our results are robust to excluding the periods 2003-2004 and 2014–2016. Panels A2 and B2

in Table 9 show that results are similar to the baseline.

7 Concluding Remarks

We study the effects of female immigration to the Dominican Republic, most of which is from Haiti and of low education, on native women's labor supply. The Haiti-Dominican Republic migration corridor is an important corridor in the developing world that has remained largely understudied by the economics literature.

Using individual-level data from the Labor Force Survey for 2003–2016 and combining variation in early immigrant settlements and in yearly inflows of female migrants, we find that female immigration has led to disparate outcomes across women of different education levels and family structures. While it led to a decrease in hours worked and earnings of low-skilled native women, it lead to an increase in hours worked of high-skilled women with dependents, compared to equally skilled women without dependents.

The results we find for women of different skills differ importantly from those found in developed countries, where low-skilled immigration is associated with an increase in the labor supply of high-skilled women and no negative effects on low-skilled women. This suggests that simply generalizing the evidence on the labor market effects of migration in the developed world to developing countries may not be the correct approach.

Finally, our results point to the importance of policies aimed at alleviating the disparate effects of migration on vulnerable groups in developing countries. Such policies could take the form of assistance (or re-training) to native workers who are displaced or see their hours or wages reduced. They could also take the form of promoting the integration of migrants in labor markets. A large part of migrants in Latin America and the Caribbean are undocumented or do not have the right to work. Even when they are allowed to work, they may face information frictions or lack the necessary skills to obtain high-quality jobs. Policies like granting work permits and providing job-search or skills training can promote a better economic integration of migrants, alleviating the pressures on native vulnerable groups, such as low-skilled women. To guide policy-making, more research is needed to understand the barriers that immigrant women face to access high-quality jobs in developing countries.

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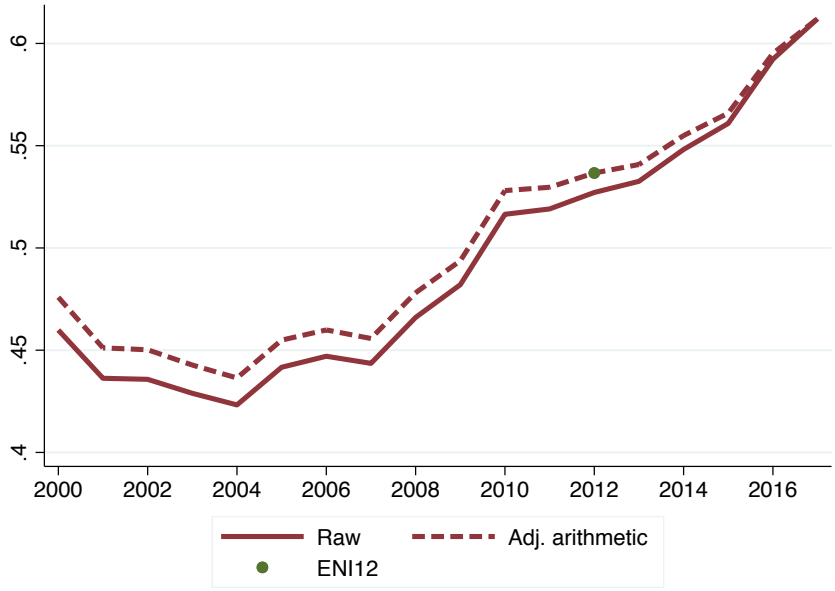
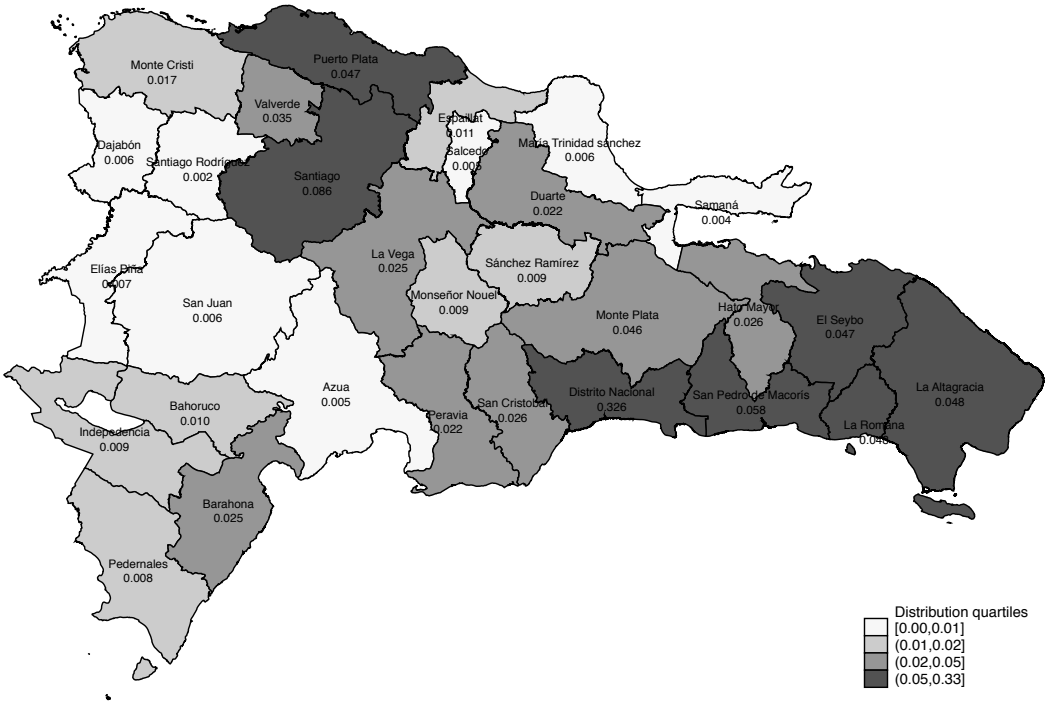
Figures

Figure 1
Female/Male Foreign-born Population, 2000–2017



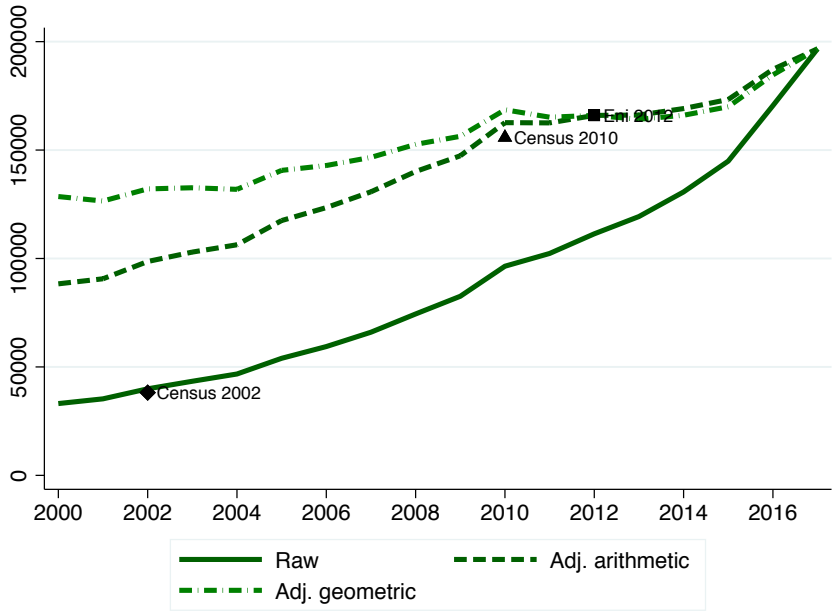
Note.–Ratio of the female foreign-born stock to the male foreign-born stock. Raw figures are calculated using the reported year of arrival in the 2017 National Immigrant Survey (ENI). Adjusted figures are calculated exploiting information in the 2012 ENI and adjusting the past-stock for deaths and re-emigration, see section 4.2 for details.

Figure 2
Province Shares of the National Stock of Foreign-born Population, 1993



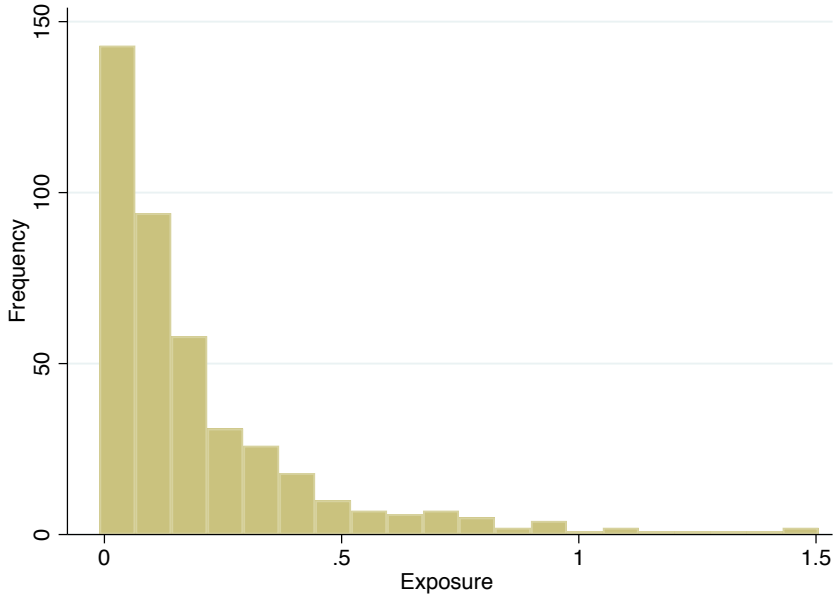
Note.–Figures for the 30 provinces existing at the time come from the 1993 Population Census. The province of Distrito Nacional includes today's capital district, Santo Domingo, and the province of Peravia includes today's province of San José de Ocoa.

Figure 3
Stock of Female Foreign-born Population, 2000–2017



Note.–Raw figures are calculated using the reported year of arrival in the 2017 National Immigrant Survey (ENI). Adjusted figures are calculated exploiting information in the 2012 ENI and adjusting the past-stock for deaths and re-emigration, see section 4.2 for details. The stocks obtained from the 2002 and 2010 population censuses are marked with symbols.

Figure 4
Distribution of Exposure to Female Immigration across Provinces



Note.– Absolute frequency of exposure to female immigration at the province-year level (N=420). $\text{Exposure}_{pt} = \delta_{p,93} \times \text{FemImmFlow}_{C,t} \times 100/\text{FemPop}_{p,93}$.

Tables

Table 1
Characteristics of Native and Immigrant Women in 2010

	Total		*Low-skilled*		*High-skilled*	
	(1) Natives	(2) Immigrants	(3) Natives	(4) Immigrants	(5) Natives	(6) Immigrants
Panel A. Demographics						
Age	40.9	36.0	42.5	34.0	37.4	36.9
Literacy (%)	89.0	61.1	93.9	92.0	100.0	100.0
High Skilled (%)	41.8	36.5	0.0	0.0	100.0	100.0
Married (%)	23.3	21.5	17.9	18.9	33.4	44.8
Births	2.9	2.7	3.4	2.7	2.0	1.5
Panel B. Labor market						
Worker (%)	41.1	38.7	34.5	40.0	53.5	46.1
Primary sector (%)	1.4	11.8	1.7	7.8	0.5	0.8
Manufacturing sector (%)	9.2	4.8	11.0	4.1	7.9	7.0
Services sector (%)	84.4	76.5	81.2	80.3	87.7	87.1
Domestic services (%)	16.1	17.4	27.5	24.6	3.9	5.6
Observations	2,008,786	80,947	1,078,797	33,074	775,484	19,013

Note.– Figures are based on the 2010 Population Census of the Dominican Republic. Women aged 25–64. *High-skilled* are women with completed secondary or more education. *Low-skilled* are women with incomplete secondary or less education. Response rates for each row variable may vary across education groups. Due to missing values in the education variable (an issue more prevalent in immigrants), totals do not correspond to the weighted averages across education groups.

Table 2
Summary Statistics–Estimation Sample

			Sub-sample | $Hours > 0$	
	(1)	(2)	(3)	(4)
	Low-Skilled	High-Skilled	Low-Skilled	High-Skilled
Panel A. Exposure				
Exposure	0.18	0.18	0.19	0.18
Panel B. Demographics				
Age	43.4	37.2	42.2	37.5
Years of education	5.4	14.0	5.9	14.3
Married (%)	18.5	34.1	15.7	34.8
Dependent (%)	55.5	58.8	51.0	56.3
Panel C. Labor market				
Worker (%)	39.4	62.2	100.0	100.0
Hours	37.3	39.0	37.3	39.0
Real wage	58.1	106.3	58.1	106.3
Real earnings	7569.7	16032.6	7569.7	16032.6
Observations	112,542	57,390	44,371	35,703

Note.– Figures correspond to simple averages and are based on the April and October rounds of the Labor Force Surveys (ENFT) for the period 2003–2016. The sample consists of Dominican-born women aged 25–64. *Dependent* is a dummy equal to one if the woman lives in a household with a child 8 years old or younger or with an adult 65 years old or older. Columns 3 and 4 correspond to the sample of women who work a positive number of hours. *Hours* are measured weekly. *Hourly wages* and *monthly earnings* are expressed in Dominican pesos of 2016.

Table 3
Effects of Exposure to Female Immigration on Native Women

	Sub-sample | $Hours > 0$							
	Employed		Hours		Wage		Earnings	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A. Whole Sample								
Exposure	-0.0034	0.0043	-1.1810*	-1.3172*	-0.0232	-0.0232	-0.0715**	-0.0747**
	(0.0150)	(0.0143)	(0.6622)	(0.6814)	(0.0237)	(0.0258)	(0.0307)	(0.0323)
Observations	169,932	169,932	80,074	80,074	76,780	76,780	76,780	76,780
R2	0.09	0.09	0.03	0.03	0.25	0.25	0.29	0.30
Panel B. High-skilled Sample								
Exposure	0.0026	0.0185	-0.2366	-0.3851	-0.0095	-0.0040	-0.0269	-0.0255
	(0.0190)	(0.0183)	(0.7713)	(0.7677)	(0.0368)	(0.0397)	(0.0384)	(0.0403)
Observations	57,390	57,390	35,703	35,703	34,642	34,642	34,642	34,642
R2	0.08	0.08	0.02	0.02	0.31	0.31	0.30	0.30
Panel C. Low-skilled Sample								
Exposure	-0.0052	-0.0017	-1.8345**	-1.9968**	-0.0346	-0.0346	-0.1049***	-0.1086***
	(0.0168)	(0.0164)	(0.8519)	(0.8938)	(0.0293)	(0.0307)	(0.0376)	(0.0388)
Observations	112,542	112,542	44,371	44,371	42,138	42,138	42,138	42,138
R2	0.05	0.05	0.03	0.03	0.05	0.05	0.08	0.09
1993 vars × year	✗	✓	✗	✓	✗	✓	✗	✓

Note.– Estimation sample of Dominican-born women aged 25–64. All estimations include year, survey round (April, October), and province fixed effects and individual-level covariates (age, age-squared, years of education, and a dummy for being married). Even-numbered columns include additional controls in 1993 (log of population, average night-light density, and literacy rates) at the province-level interacted with year dummies. $\text{Exposure}_{pt} = \delta_{p,93} \times \text{FemImmFlow}_{C,t} \times 100/\text{FemPop}_{p,93}$. *Employed* is a dummy for being employed. *Hours* are weekly hours worked. *Wage* is the log hourly wage. *Earnings* is the log of monthly earnings. Robust standard errors in brackets are clustered at the province × year level. *** p<0.01, ** p<0.05, * p<0.1.

Table 4
Effects of Exposure to Female Immigration on Native Women: Interaction with Family Dependents

	Sub-sample | $Hours > 0$							
	Employed		Hours		Wage		Earnings	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A. High-skilled Sample								
Exposure	-0.0016	0.0145	-1.0396	-1.2023	-0.0055	-0.0033	-0.0449	-0.0473
	(0.0214)	(0.0210)	(0.9184)	(0.9071)	(0.0423)	(0.0449)	(0.0412)	(0.0430)
Dependent=1	-0.0391***	-0.0389***	-1.0443***	-1.0441***	-0.0132	-0.0140	-0.0477***	-0.0485***
	(0.0054)	(0.0054)	(0.1959)	(0.1964)	(0.0104)	(0.0104)	(0.0115)	(0.0115)
Dependent=1 × Exposure	0.0062	0.0056	1.4728**	1.4880**	-0.0077	-0.0019	0.0324	0.0388
	(0.0188)	(0.0188)	(0.6180)	(0.6147)	(0.0331)	(0.0328)	(0.0323)	(0.0323)
Observations	57,390	57,390	35,703	35,703	34,642	34,642	34,642	34,642
R2	0.08	0.08	0.02	0.02	0.31	0.31	0.30	0.30
Panel B. Low-skilled Sample								
Exposure	-0.0125	-0.0095	-1.4662	-1.6301	-0.0439	-0.0431	-0.0957**	-0.0995**
	(0.0198)	(0.0196)	(0.9593)	(0.9951)	(0.0323)	(0.0335)	(0.0453)	(0.0458)
Dependent=1	-0.0645***	-0.0645***	-1.1442***	-1.1357***	0.0139	0.0143	-0.0226***	-0.0221**
	(0.0040)	(0.0040)	(0.2035)	(0.2036)	(0.0087)	(0.0087)	(0.0087)	(0.0087)
Dependent=1 × Exposure	0.0098	0.0105	-0.7504	-0.7574	0.0183	0.0169	-0.0182	-0.0184
	(0.0141)	(0.0142)	(0.7385)	(0.7356)	(0.0262)	(0.0260)	(0.0351)	(0.0351)
Observations	112,542	112,542	44,371	44,371	42,138	42,138	42,138	42,138
R2	0.05	0.05	0.03	0.03	0.05	0.05	0.08	0.09
1993 vars × year	✗	✓	✗	✓	✗	✓	✗	✓

Note.– Estimation sample of Dominican-born women aged 25–64. All estimations include year, survey round (April, October), and province fixed effects and individual-level covariates (age, age-squared, years of education, and a dummy for being married). Even-numbered columns include additional controls in 1993 (log of population, average night-light density, and literacy rates) at the province-level interacted with year dummies. $\text{Exposure}_{pt} = \delta_{p,93} \times \text{FemImmFlow}_{C,t} \times 100/\text{FemPop}_{p,93}$. *Dependent* is adummy equal to one if the woman lives with a child 8 years old or younger or an adult 65 years old or older. *Employed* is a dummy for being employed. *Hours* are weekly hours worked. *Wage* is the log hourly wage. *Earnings* is the log of monthly earnings. Robust standard errors in brackets are clustered at the province × year level. *** p<0.01, ** p<0.05, * p<0.1.

Table 5
Effects of Exposure to Female Immigration on Regional Domestic Employment, Wages, and Earnings

	(1) % Domestic	(2) Avg. Wage	(3) Avg. Earnings
Exposure	0.0249**	-0.1196	-0.2079***
	(0.0078)	(0.1134)	(0.0569)
Observations	280	279	279
R2	0.70	0.37	0.69

Note.– All estimations include year, survey round (April, October), and region fixed effects. *% Domestic* is the share of domestic female workers in total female employment in region-year-semester. *Avg. Wage* is the log average hourly wage of female workers in domestic services in the region-year-semester. *Avg. earnings* are the log average monthly earnings of female workers in domestic services in the region-year-semester. $\text{Exposure}_{rt} = \delta_{r,93} \times \text{FemImmFlow}_{C,t} \times 100/\text{FemPop}_{r,93}$. Robust standard errors in brackets are clustered at the region level. *** p<0.01, ** p<0.05, * p<0.1.

Table 6
Effects of Exposure to Female Immigration on Low-skilled Native Women: Interaction with Domestic Worker Dummy

	Sub-sample | $Hours > 0$					
	(1)	(2)	(3)	(4)	(5)	(6)
	Hours	Hours	Wage	Wage	Earnings	Earnings
Exposure	-1.4638*	-1.6099*	-0.0494	-0.0509	-0.1057***	-0.1097***
	(0.8626)	(0.9057)	(0.0327)	(0.0343)	(0.0404)	(0.0418)
Domestic=1	0.0406	0.0415	-0.1626***	-0.1627***	-0.1851***	-0.1849***
	(0.2534)	(0.2529)	(0.0115)	(0.0115)	(0.0132)	(0.0131)
Domestic=1 × Exposure	-1.0902	-1.0982	0.0386	0.0401	-0.0024	-0.0026
	(0.8623)	(0.8630)	(0.0379)	(0.0383)	(0.0406)	(0.0409)
Observations	44,371	44,371	42,138	42,138	42,138	42,138
R2	0.03	0.03	0.06	0.07	0.10	0.10

Note.– Estimation sample of low-skilled Dominican-born women aged 25–64. All estimations include year, survey round (April, October), and province fixed effects, individual-level covariates (age, age-squared, years of education, and a dummy for being married), and province-level controls in 1993 (log of population, average night-light density, and literacy rates) interacted with year dummies. $\text{Exposure}_{pt} = \delta_{p,93} \times \text{FemImmFlow}_{C,t} \times 100/\text{FemPop}_{p,93}$. *Hours* are weekly hours worked. *Wage* is the log hourly wage. *Earnings* is the log of monthly earnings. Robust standard errors in brackets are clustered at the province × year level. *** p<0.01, ** p<0.05, * p<0.1.

Table 7
Robustness–Effects of Exposure to Female Immigration on Native Women–Clustering of s.e at Province Level And Additional controls in 2003

		Sub-sample | $Hours > 0$		
	(1)	(2)	(3)	(4)
	Employed	Hours	Wage	Earnings
Panel A. High-skilled Sample				
A.1 Clustering s.e. at province level				
Exposure	0.0185	-0.3851	-0.0040	-0.0255
	(0.0200)	(0.7631)	(0.0428)	(0.0429)
Observations	57,390	35,703	34,642	34,642
R2	0.08	0.02	0.31	0.30
A.2 Additional controls in 2003				
Exposure	0.0191	-0.1072	-0.0001	-0.0126
	(0.0197)	(0.7419)	(0.0378)	(0.0364)
Observations	57,390	35,703	34,642	34,642
R2	0.08	0.02	0.31	0.30
Panel B. Low-skilled Sample				
B.1 Clustering s.e. at province level				
Exposure	-0.0017	-1.9968***	-0.0346	-0.1086***
	(0.0135)	(0.6573)	(0.0253)	(0.0308)
Observations	112,542	44,371	42,138	42,138
R2	0.05	0.03	0.05	0.09
B.2 Additional controls in 2003				
Exposure	-0.0002	-2.0762**	-0.0518	-0.1337***
	(0.0187)	(0.9286)	(0.0364)	(0.0403)
Observations	112,542	44,371	42,138	42,138
R2	0.05	0.03	0.06	0.09

Note.– Estimation sample of Dominican-born women aged 25–64. All estimations include year, survey round (April, October), and province fixed effects, individual-level covariates (age, age-squared, years of education, and a dummy for being married), and province-level controls in 1993 (log of population, average night-light density, and literacy rates) interacted with year dummies. Panels A1 and B1 cluster standard errors at the province level. Panels A2 and B2 include additional controls in 2003 at the region-level (unemployment rate of women and the share of women that are high-skilled) interacted with year dummies. $\text{Exposure}_{pt} = \delta_{p,93} \times \text{FemImmFlow}_{C,t} \times 100/\text{FemPop}_{p,93}$. *Employed* is a dummy for being employed. *Hours* are weekly hours worked. *Wage* is the log hourly wage. *Earnings* are the log of monthly earnings. Robust standard errors in brackets are clustered at the province × year level (Panels A2 and B2). *** p<0.01, ** p<0.05, * p<0.1.

Table 8
Robustness–Effects of Exposure to Female Immigration on Native Women–Geometric Adjustment and Male Immigrant Settlements

		Sub-sample | $Hours > 0$		
	(1)	(2)	(3)	(4)
	Employed	Hours	Wage	Earnings
Panel A. High-skilled Sample				
A.1 Geometric adjustment				
Exposure	0.0163	-0.3810	-0.0298	-0.0520
	(0.0167)	(0.6736)	(0.0374)	(0.0361)
Observations	57,390	35,703	34,642	34,642
A.2 Using male immigrant settlements				
Exposure	0.0339	-0.5908	-0.0079	-0.0443
	(0.0326)	(1.3809)	(0.0704)	(0.0718)
Observations	57,390	35,703	34,642	34,642
Panel B. Low-skilled Sample				
B.1 Geometric adjustment				
Exposure	-0.0058	-1.9510**	-0.0375	-0.1113***
	(0.0157)	(0.8452)	(0.0290)	(0.0390)
Observations	112,542	44,371	42,138	42,138
B.2 Using male immigrant settlements				
Exposure	-0.0058	-3.5294**	-0.0620	-0.1933***
	(0.0298)	(1.6080)	(0.0546)	(0.0699)
Observations	112,542	44,371	42,138	42,138

Note.– Estimation sample of Dominican-born women aged 25–64. All estimations include year, survey round (April, October), and province fixed effects, individual-level covariates (age, age-squared, years of education, and a dummy for being married), and province-level controls in 1993 (log of population, average night-light density, and literacy rates) interacted with year dummies. $\text{Exposure}_{pt} = \delta_{p,93} \times \text{FemImmFlow}_{C,t} \times 100/\text{FemPop}_{p,93}$. For panels A1 and B1 $\text{FemImmFlow}_{C,t}$ calculated with geometric adjustment. For panels A2 and B2 $\delta_{p,93}$ are constructed using the settlement of male immigrants. *Employed* is a dummy for being employed. *Hours* are weekly hours worked. *Wage* is the log hourly wage. *Earnings* are the log of monthly earnings. Robust standard errors in brackets are clustered at the province × year level. *** p<0.01, ** p<0.05, * p<0.1.

Table 9
Robustness–Effects of Exposure to Female Immigration on Native Women–Excluding Metropolitan Areas and Crisis (2003-2004) and Regularization (2014-2016) Years

		Sub-sample | $Hours > 0$		
	(1)	(2)	(3)	(4)
	Employed	Hours	Wage	Earnings
Panel A. High-skilled Sample				
A.1 Excluding Metropolitan Areas				
Exposure	0.0192	-0.3204	-0.0027	-0.0233
	(0.0181)	(0.7644)	(0.0397)	(0.0407)
Observations	36,816	22,440	21,789	21,789
A.2 Excluding Crisis & Regularization Years				
Exposure	0.0423*	-1.0973	0.0342	-0.0156
	(0.0216)	(0.9789)	(0.0406)	(0.0455)
Observations	36,603	22,541	21,823	21,823
Panel B. Low-skilled Sample				
B.1 Excluding Metropolitan Areas				
Exposure	-0.0041	-1.9894**	-0.0340	-0.1064***
	(0.0166)	(0.8664)	(0.0311)	(0.0386)
Observations	80,707	30,108	28,566	28,566
B.2 Excluding Crisis & Regularization Years				
Exposure	-0.0001	-2.3045**	-0.0315	-0.1074***
	(0.0166)	(0.9972)	(0.0334)	(0.0341)
Observations	75,014	29,203	27,722	27,722

Note.– Estimation sample of Dominican-born women aged 25–64. All estimations include year, survey round (April, October), and province fixed effects, individual-level covariates (age, age-squared, years of education, and a dummy for being married), and province-level controls in 1993 (log of population, average night-light density, and literacy rates) interacted with year dummies. Panels A1 and B1 exclude the (1993) provinces of Distrito Nacional and Santiago. Panels A2 and B2 exclude years 2003–2004 and 2014–2016. $\text{Exposure}_{pt} = \delta_{p,93} \times \text{FemImmFlow}_{C,t} \times 100/\text{FemPop}_{p,93}$. *Employed* is a dummy for being employed. *Hours* are weekly hours worked. *Wage* is the log hourly wage. *Earnings* are the log of monthly earnings. Robust standard errors in brackets are clustered at the province × year level. *** p<0.01, ** p<0.05, * p<0.1.

Appendix

A Adjustment of Foreign-born Stocks

We adjust the *raw* measure of the immigrant stock as reported in the ENI 2017 using an adjustment factor which we calculate by exploiting the information from the first round of ENI, in 2012. More precisely, we look for an adjustment factor, $(1+\gamma)$, that allows to transform raw values of the stock as measured in 2017, $M_t^{\mathrm{R},17}$, to *adjusted* values, by inflating the observed raw 2017 values to correct for deaths and emigration. The adjusted stock $\widetilde{M}_t$ in any year t is given by

$$\widetilde{M}_t = (1+\gamma)^s M_t^{\mathrm{R},17}. \tag{A1}$$

We adjust the immigrant stock in each year by using two different approaches, that we denote *arithmetic* and *geometric*, to calculate the rate γ (which we assume is constant).

We start by defining $M_{2012}^{\mathrm{T},12}$, the stock of immigrants in 2012 obtained with ENI's first round (2012)—i.e. the true stock in 2012—and $M_{2012}^{\mathrm{R},17}$ the raw stock in 2012 calculated with ENI's second round (2017)—i.e. the stock of immigrants in 2012 that were still living in the country in 2017.

The arithmetic adjustment of the migrant stock, uses $\gamma^A = \frac{1}{5}\left(M_{2012}^{\mathrm{T},12}/M_{2012}^{\mathrm{R},17} - 1\right)$, the simple per-year average of the growth rate necessary to inflate the raw $M_t^{R,17}$ value to get the true $M_t^{T,12}$ value. Under this formulation, the arithmetic adjustment of the migrant stock, $\widetilde{M}_t^A$, is given by

$$\widetilde{M}_t^A = \left[1 + \gamma^A (2017 - t)\right] M_t^{\mathrm{R},17}. \tag{A2}$$

The geometric adjustment of the migrant stock, uses $\gamma^G = \left(M_{2012}^{\mathrm{T},12}/M_{2012}^{\mathrm{R},17}\right)^{1/5} - 1$, the compound average growth rate necessary to inflate the raw $M_t^{R,17}$ value to get the true $M_t^{T,12}$ value. Under this formulation, $\widetilde{M}_t^G$, is given in each year $t < 2017$ by

$$\widetilde{M}_t^G = \left(1 + \gamma^G\right)^{2017-t} M_t^{\mathrm{R},17}. \tag{A3}$$

B Data Appendix

Table B1
Variable Definitions and Sources

Variable	Description	Source
Exposure		
Migrant stock	Number of foreing-born that report having arrived to the Dominican Republic in a given year.	National Immigrant Survey (ENI) 2012, 2017
$FemmImmFlow_{C,t}$	Difference between the national stock of foreign-born women in *t* and *t-1*.	ENI 2012, 2017
$\delta_{p,93}$	Province *p* share in the national stock of foreign-born in 1993	1993 Census
$FemPop_{p,93}$	Province *p* female population in 1993.	1993 Census
Outcome variables		
Employed	Dummy variable = 1 if individual reports to be employed in the week prior to the survey.	National Labor Force Survey (ENFT) 2003–2016
Hours	Reported weekly hours worked by employed individuals.	ENFT 2003–2016
Hourly wage	Ratio of reported weekly earnings to reported weekly hours. We exclude the 1st and 99th percentile.	ENFT 2003–2016
Earnings	Monthly earnings. Monthly hours worked times hourly wage.	ENFT 2003–2016
Demographics– Native sample		
Age	Age reported by subjects at the time of survey.	ENFT 2003–2016
Years of education	Calculated based on reported maximum education level achieved and on the number of years necessary to reach that level according to the Dominican Republic's educational system.	ENFT 2003-2016
High-Skilled	Dummy variable = 1 if the individual completed high school. It is equivalent to having 12 years or more of education according to the Dominican Republic's system.	ENFT 2003-2016
Low-Skilled	Dummy variable = 1 if the individual has less than high school education. It is equivalent to having less than 12 years of education according to the Dominican Republic's system.	ENFT 2003-2016
Married	Dummy variable = 1 if the individual reports being married at the time of survey.	ENFT 2003-2016
Dependent	Dummy variable = 1 if in the household of the individual surveyed lives at least one kid under the age of 9 or at least one adult 65 years old or older.	ENFT 2003-2016

Continues in the next page

Table B1 – Variable description (continued from last page)

Variable	Description	Source
Domestic	Dummy variable = 1 if the individual reports working in ISIC's Rev. 3 Activity 9500 "Activities of private households as employers of domestic staff".	ENFT 2003-2016

Table B2
Regions and provinces in Dominican Republic

Region	**Province**
Cibao Norte	Espillat Puerto plata Santiago
Cibao Sur	La vega Monsenor Nuel Sanchez ramirez
Cibao Nordeste	Duarte Hermanas Mirabal Maria Trinidad Sanches Samana
Cibao Noroeste	Dajabon Monte Cristi Santiago Rodriguez Valverde
Valdesia	Azua Peravia San Jose de Ocoa San Cristobal
El Valle	Elias Pina San Juan
Enriquillo	Barahona Bahoruco Independencia Pedernales
Yuma	El seibo La Romana La Altagracia
Higuamo	Hato mayor Monte Plata San pedro de macoris
Ozama	Distrito Nacional Santo Domingo

C Supplementary Results

Table C1
Effects of Exposure to Female Haitian Immigration on Native Women

			Sub-sample | $Hours > 0$					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Employed	Employed	Hours	Hours	Wage	Wage	Earnings	Earnings
Panel A. Whole Sample								
Exposure	-0.0069	0.0025	-1.3251	-1.4787*	-0.0296	-0.0276	-0.0853**	-0.0871**
	(0.0180)	(0.0171)	(0.8238)	(0.8464)	(0.0283)	(0.0309)	(0.0369)	(0.0390)
Observations	169,932	169,932	80,074	80,074	76,780	76,780	76,780	76,780
R2	0.09	0.09	0.03	0.03	0.25	0.25	0.29	0.30
Panel B. High-skilled Sample								
Exposure	-0.0023	0.0164	-0.1798	-0.3662	-0.0108	-0.0009	-0.0287	-0.0241
	(0.0233)	(0.0225)	(0.9545)	(0.9523)	(0.0442)	(0.0476)	(0.0468)	(0.0491)
Observations	57,390	57,390	35,703	35,703	34,642	34,642	34,642	34,642
R2	0.08	0.08	0.02	0.02	0.31	0.31	0.30	0.30
Panel C. Low-skilled Sample								
Exposure	-0.0080	-0.0033	-2.1108**	-2.2830**	-0.0435	-0.0427	-0.1260***	-0.1292***
	(0.0199)	(0.0195)	(1.0386)	(1.0905)	(0.0356)	(0.0372)	(0.0456)	(0.0466)
Observations	112,542	112,542	44,371	44,371	42,138	42,138	42,138	42,138
R2	0.05	0.05	0.03	0.03	0.05	0.05	0.08	0.09
1993 vars × year	✗	✓	✗	✓	✗	✓	✗	✓

Note.– Estimation sample of Dominican-born women aged 25–64. All estimations include year, survey round (April, October), and province fixed effects and individual-level covariates (age, age-squared, years of education, and a dummy for being married). Even-numbered columns include additional controls in 1993 (log of population, average night-light density, and literacy rates) at the province-level interacted with year dummies. $\text{Exposure}_{pt} = \delta_{p,93} \times \text{FemImmFlow}_{C,t} \times 100/\text{FemPop}_{p,93}$, where $\text{FemImmFlow}_{C,t}$ corresponds to the inflow of Haitian women only. *Employed* is a dummy for being employed. *Hours* are weekly hours worked. *Wage* is the log hourly wage. *Earnings* is the log of monthly earnings. Robust standard errors in brackets are clustered at the province × year level. *** p<0.01, ** p<0.05, * p<0.1.

Table C2
Effects of Exposure to Female Haitian Immigration on Native Women: Interaction with Family Dependents

			Sub-sample | Hours > 0					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Employed	Employed	Hours	Hours	Wage	Wage	Earnings	Earnings
Panel A. High-skilled Sample								
Exposure	-0.0073	0.0113	-1.1686	-1.3695	-0.0026	0.0032	-0.0474	-0.0473
	(0.0260)	(0.0256)	(1.1249)	(1.1154)	(0.0508)	(0.0538)	(0.0498)	(0.0522)
Dependent=1	-0.0392***	-0.0391***	-1.0827***	-1.0819***	-0.0120	-0.0130	-0.0475***	-0.0484***
	(0.0056)	(0.0056)	(0.2016)	(0.2021)	(0.0108)	(0.0108)	(0.0118)	(0.0118)
Dependent=1 × Exposure	0.0076	0.0070	1.8303**	1.8425**	-0.0156	-0.0083	0.0338	0.0415
	(0.0218)	(0.0219)	(0.7273)	(0.7239)	(0.0394)	(0.0391)	(0.0380)	(0.0379)
Observations	57390	57390	35703	35703	34642	34642	34642	34642
R2	0.08	0.08	0.02	0.02	0.31	0.31	0.30	0.30
Panel B. Low-skilled Sample								
Exposure	-0.0171	-0.0132	-1.6507	-1.8242	-0.0556	-0.0537	-0.1164**	-0.1196**
	(0.0231)	(0.0228)	(1.1453)	(1.1909)	(0.0388)	(0.0400)	(0.0536)	(0.0539)
Dependent=1	-0.0647***	-0.0647***	-1.1222***	-1.1130***	0.0132	0.0137	-0.0227**	-0.0221**
	(0.0041)	(0.0041)	(0.2078)	(0.2080)	(0.0089)	(0.0090)	(0.0089)	(0.0090)
Dependent=1 × Exposure	0.0119	0.0127	-0.9422	-0.9539	0.0234	0.0220	-0.0195	-0.0198
	(0.0160)	(0.0160)	(0.8340)	(0.8310)	(0.0298)	(0.0296)	(0.0397)	(0.0397)
Observations	112542	112542	44371	44371	42138	42138	42138	42138
R2	0.05	0.05	0.03	0.03	0.05	0.05	0.08	0.09
1993 vars × year	✗	✓	✗	✓	✗	✓	✗	✓

Note.– Estimation sample of Dominican-born women aged 25–64. All estimations include year, survey round (April, October), and province fixed effects and individual-level covariates (age, age-squared, years of education, and a dummy for being married). Even-numbered columns include additional controls in 1993 (log of population, average night-light density, and literacy rates) at the province-level interacted with year dummies. $\text{Exposure}_{rt} = \delta_{r,93} \times \text{FemImmFlow}_{C,t} \times 100/\text{FemPop}_{r,93}$. *Dependent* is adummy equal to one if the woman lives with a child 8 years old or younger or an adult 65 years old or older. *Employed* is a dummy for being employed. *Hours* are weekly hours worked. *Wage* is the log hourly wage. *Earnings* is the log of monthly earnings. Robust standard errors in brackets are clustered at the province × year level. *** p<0.01, ** p<0.05, * p<0.1.