

Immigrating into a Recession: Evidence from Family Migrants to the United States

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We analyze the impact of economic conditions at arrival on the economic integration of family-sponsored migrants in the United States. A 1 percentage point higher unemployment rate at arrival decreases annual wage income by 4% in the short run and 2% in the longer run. The loss in wage income results primarily from lower hourly wages due to occupational downgrading. Migrant and family networks help mitigate the negative labor market effects. Migrants who arrive during a recession take up occupations with higher concentrations of fellow countrypeople and are more likely to reside with family members, potentially reducing their geographical mobility.

I. Introduction

The economic integration of immigrants is key to reaping the benefits of international migration for immigrants and citizens of destination countries

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alike. Yet immigrant integration is often imperfect and varies substantially across and within immigrant cohorts. A large literature has examined the economic assimilation of immigrants over time, paying particular attention to changes in immigrant characteristics and selective return migration.

This paper takes a different perspective and analyzes how economic conditions at the time of arrival shape the economic integration of immigrants over time. We introduce a new identification strategy that exploits the inability of family migrants to the United States to synchronize their arrival with labor market conditions in the United States. Our focus is on key labor market outcomes, above all employment and wage income. We also provide evidence on the underlying mechanisms and coping strategies. In particular, we show that adverse economic conditions at the time of arrival can explain occupational downgrading and migrants' reliance on family and migrant networks.

Family migrants typically arrive without a job, as their visas are not sponsored by an employer. Their economic integration might hence be susceptible to labor market conditions at arrival. Economic theory suggests that entering the labor market in a recession increases the odds of unemployment and job mismatch. Immigrants might also be pushed to enter occupations with larger ethnic networks. These jobs may provide fewer opportunities to accumulate destination-specific human capital and experience, thus limiting upward mobility. In addition, future employers may perceive past labor market outcomes as a signal of productivity. Unemployed or mismatched immigrants may then face persistently lower wages. We therefore hypothesize that immigrating into a recession is associated with worse labor market outcomes, even in the longer run.

Our identification strategy exploits long waiting times for family-sponsored visas to the United States that effectively decouple the migration decision from economic conditions at the time of immigration. Because of caps on the number of available visas and excess demand, it usually takes several years until a family-sponsored visa is granted. Once a visa becomes available, family members have a limited time window to move to the United States. In addition, family migrants typically move to the sponsor's location in the United States. Hence, family migrants do not choose the US state based on economic

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conditions either. Thus, the macroeconomic conditions family migrants face at arrival in the United States are exogenous. Some family migrants happen to immigrate into a boom, others into a recession.

Family migrants constitute the largest group of permanent immigrants in the United States. In 2015, 65% of all persons obtaining lawful permanent resident status in the United States were family migrants, and only 14% were labor migrants (i.e., their visa was sponsored by their employer). Humanitarian migrants (refugees and asylees) and winners of the diversity lottery in the US account for most of the remaining share. The picture looks similar for the Organization for Economic Cooperation and Development (OECD) more generally. In 2015, family migrants accounted for 48% of all permanent immigrant arrivals in the OECD. Labor migrants accounted for only 16%.¹ Despite their relevance, family migrants have received little attention in economics.²

We base our analysis on data from the American Community Survey (ACS) and the US census for the period 2000–2019, which provide information on the year of immigration. We analyze the effect of state-level unemployment rates in the year of immigration on immigrants' labor market outcomes in the year of observation. Our econometric specification includes state, year-of-observation, year-of-immigration, and years-in-the-US fixed effects. This specification allows us to control for persistent differences in economic conditions and immigrant characteristics across states, nationwide economic conditions at the time of observation and the year of immigration, changes in the characteristics of immigrant cohorts, and the general path of economic integration over time.

The key challenge for identification is that the migration decision is endogenous to economic conditions. If the characteristics of immigrants to a specific state differ between good and bad economic times, observed differences in economic integration may be due to differences in immigrant characteristics, not differences in initial economic conditions. As argued above, family migrants cannot choose their date of immigration based on economic conditions. And they do not choose a location within the United States based on economic conditions but join their sponsor's household in the United States. As a result, the local economic conditions family migrants face at arrival are exogenous. We offer support for our identifying assumption by showing that initial unemployment rates (IURs) cannot predict the

¹ Figures for the United States come from the 2015 Yearbook of Immigration Statistics. Figures for the OECD come from the 2017 International Migration Outlook and exclude migration movements within areas of free circulation (mainly within the European Union).

² A small related literature studies migration decisions of couples and households and the resulting selection of immigrants. See, e.g., Mincer (1978), Borjas and Bronars (1991), Cobb-Clark (1993), Foged (2016), and Munk, Nikolka, and Poutvaara (2022).

size and composition of inflows of family migrants to US states. We also show that selective return migration is unlikely to bias our results. In general, rates of return migration are very low for family migrants. In addition, survival rates of migrants in our sample do not systematically differ by year of arrival and hence different initial economic conditions.

Currently available datasets do not provide information on visa type. We are thus not able to precisely identify family migrants. In our main analysis, we therefore restrict the sample to immigrants from countries for which family migration is the dominant mode of migration to the United States. We also offer alternative strategies that identify (i) family migrants from the Philippines, based on administrative data from the Philippine government, and (ii) subgroups of immigrants whose individual characteristics even more likely identify them as family migrants. In addition, we show that the inclusion of migrants who do not face waiting times for their visa biases our estimates toward zero. We therefore consider our results to be conservative.

We have three main findings. First, immigrating into a recession substantially worsens labor market outcomes, even in the longer run. A 1 percentage point higher IUR has a small and relatively short-lasting effect on employment rates. However, it decreases real annual wage income in the first 3 years by about 4%, with slow convergence to a persistent negative effect of about 2% afterward. A 5 percentage point rise in the IUR, a typical increase in a large recession (von Wachter 2020), would reduce the net present value wage income of a family migrant over a period of 10 years by USD 35,500. This is a large effect. For comparison, the equivalently estimated average earnings loss of graduating in an equally large recession is USD 23,000. Second, the loss in wage income is the result of a combination of substantial occupational downgrading, leading to lower hourly wages, and a reduction in working hours. Third, family migrants rely on migrant and family networks to cope with adverse economic conditions. In regions with larger migrant networks, family migrants who arrive at times of high unemployment experience a smaller negative effect on initial employment but take up occupations with substantially higher concentrations of fellow countrypeople. They are also more likely to reside with family members, likely the sponsor of their visa. The immobile support received from the family, however, may reduce the geographic mobility of migrants. Indeed, migrants who arrive at times of high unemployment do not increase their geographical mobility. The resulting job search frictions provide a potential explanation for the observed persistence of the effects.

Our paper contributes to two strands of literature. The first strand examines how immigrant earnings evolve after their arrival in the destination country and whether immigrant earnings assimilate to native earnings over time (starting with the seminal papers by Chiswick [1978] and Borjas [1985]). This large literature typically finds that immigrants have low initial earnings relative to natives but enjoy high earnings growth over time, with the degree of

catch-up depending on immigrant cohorts and destination countries.³ We contribute to this literature by showing that initial economic conditions generate substantial heterogeneity in the evolution of earnings across immigrant cohorts. In addition, initial economic conditions constitute a key mechanism for explaining widely documented immigrant-specific phenomena, such as downgrading (e.g., Dustmann, Frattini, and Preston 2013; Dustmann, Schönberg, and Stuhler 2016) and the use of migrant and family networks (e.g., Munshi 2003; Patel and Vella 2013; Battisti, Peri, and Romiti 2022). We also provide first evidence on the economic integration of family migrants, who represent the most relevant group of permanent immigrants in the United States and the OECD more generally.

The second strand of the literature studies how entering the labor market in a recession affects labor market outcomes. Most studies focus on the labor market effects of graduating in a recession (e.g., Raaum and Røed 2006; Kahn 2010; Kwon, Milgrom, and Hwang 2010; Oreopoulos, von Wachter, and Heisz 2012; Brunner and Kuhn 2014; Altonji, Kahn, and Speer 2016; Schwandt and von Wachter 2019; Stuart 2022; Rothstein 2023).⁴ They find that college and school students who graduate in a recession are less likely to be employed and earn lower-wage incomes in the first years of their careers, with mixed evidence on the persistence of these effects. A few studies document similar effects for immigrants and refugees who arrive at the destination country in a recession (Chiswick, Cohen, and Zach 1997; Chiswick and Miller 2002; Åslund and Rooth 2007; Godøy 2017; Mask 2020; Fasani, Frattini, and Minale 2021b; Aksoy, Poutvaara, and Schikora 2023).⁵ We contribute to this literature by introducing a novel identification strategy that exploits the inability of family migrants to the United States to synchronize their arrival with labor market conditions in the United States. We thus broaden the evidence that labor market outcomes are path dependent and not only depend on current economic conditions (e.g., see Beaudry and DiNardo 1991). It is not clear a priori that previous findings for native labor market entrants and refugees also apply to family migrants. Family migrants may enter different segments of the labor market and can draw on established networks that may cushion the effects of initial economic conditions.

II. Theoretical Considerations

Recessions affect the number and types of initial job opportunities available on the labor market. Entering the labor market in a recession hence

³ For a review of the literature, see Duleep (2015).

⁴ For a review of the literature, see von Wachter (2020).

⁵ Going beyond economic conditions at arrival, Marbach, Hainmueller, and Hangartner (2018) and Fasani, Frattini, and Minale (2021a) provide evidence that the policy environment at immigration has a long-term effect on the labor market integration of refugees. Aksoy, Poutvaara, and Schikora (2023) and Jaschke, Sardoschau, and Tabellini (2021) show that attitudes of natives also affect refugees' integration.

likely increases the odds of being unemployed or experiencing a job mismatch with the associated lower wages (Bowlus 1995; McLaughlin and Bils 2001; Devereux 2002). This may be particularly true for immigrants, as their labor market outcomes are generally more responsive to business cycle fluctuations than those of natives (Dustmann, Glitz, and Vogel 2010; Orrenius and Zavodny 2010). If future employers perceive past spells of unemployment or low wages as a signal of low productivity, initially unemployed or mismatched individuals may then face persistently lower wages (Jacobson, LaLonde, and Sullivan 1993; Arulampalam 2001; Gregg 2001; Gregory and Jukes 2001; Couch and Placzek 2010; Kroft, Lange, and Notowidigdo 2013).

The scarring effect is likely more pronounced for immigrants than for natives. Initial unemployment and job mismatching make them less likely to accumulate destination country-specific human capital and experience. Given the limited transferability of human capital and experience from the origin to the destination country (e.g., Friedberg 2000), it may thus be more difficult for immigrants than natives to signal their idiosyncratic productivity to future employers. Similarly, if immigrants initially enter the ethnic economy and accumulate skills that are specific to the ethnic economy, they may be more likely to remain in the ethnic economy and potentially face limited upward mobility (similar to the arguments made by Borjas [1992] and Battisti, Peri, and Romiti [2022]).

However, adverse economic conditions at arrival do not need to have persistent labor market effects. Immigrants may simply take more time to complete the job-matching process. With diminishing marginal returns to experience, immigrants may be able to recover from initial shocks over time. The ability to do so depends on the presence of search frictions and immigrants' occupational, sectoral, and geographical mobility (on the importance of mobility, see Topel and Ward 1992).

On the one hand, because of the scarring effect and discrimination on the labor market, immigrants may face more search frictions than natives. On the other hand, because of lower degrees of geographical attachment, immigrants may have lower moving costs and could be more mobile than natives (Green 1999; Braun and Kvasnicka 2014; Cadena and Kovak 2016). The mobility of family migrants, however, might be lower, as they usually join their sponsor's household on arrival and might depend on the immobile support of family members. From a theoretical perspective, it is thus not clear how persistent the negative labor market effects of immigrating into a recession are.

III. The System of Family-Sponsored Immigration to the United States

The Immigration and Nationality Act of 1965 established the current system of lawful permanent immigration to the United States. The latest major amendments occurred through the Immigration Act of 1990. The system

defines four main pathways of permanent immigration: family-sponsored immigration, employer-sponsored immigration, the Diversity Immigrant Visa Program, and admission on humanitarian grounds through refugee and asylum programs.⁶ US immigration law has traditionally prioritized family-sponsored immigration, which remains the most important avenue for permanent immigration to the United States.

Family members residing in the United States as US citizens or lawful permanent residents (LPRs; aka, green-card holders) can act as sponsors for family members abroad and provide legal entitlement to a visa. Sponsors need to prove that they can support their own family and the sponsored family member at an income level at or above 125% of the federal poverty level. They also need to sign an affidavit of financial support, which obliges them to support sponsored family members who cannot support themselves for 10 years or until they become US citizens.

The admission categories depend on the relationship between the sponsor and the family member, the age and marital status of the family member, and whether the sponsor is a US citizen or LPR (see table 1). Importantly, there are caps on the number of visas available in each category per year. The only exceptions are immediate relatives of US citizens (i.e., spouses, parents, and unmarried children under the age of 21 years). They are not subject to numerical limitations (Kandel 2014).

The Immigration Act of 1990 caps the total number of LPR visas at 675,000 per year. This number includes 480,000 family-sponsored immigrants, 140,000 employer-sponsored immigrants, and 55,000 diversity immigrants. It excludes refugees and asylees. To account for the fact that immediate relatives of US citizens are not subject to numerical limitations, the annual cap of 480,000 family-sponsored immigrants is adjusted in the following way:

$$\begin{aligned}
 &480,000 \text{ (annual total cap on family-sponsored immigrants)} \\
 &- \text{ number of immediate relatives granted LPR status in the prior year} \\
 &- \text{ number of aliens paroled into the US for at least a year in the prior year} \\
 &+ \text{ number of unused employer-sponsored visas from the prior year.}
 \end{aligned}$$

The minimum adjusted number of family-sponsored immigrants, however, is fixed at 226,000 per year. As the number of immediate relatives granted LPR status has exceeded 254,000 every year, the annual cap for other family-sponsored immigrants has effectively remained at the minimum of 226,000 for the past two decades (Kandel 2018). Table 1 shows the resulting number of family visas available for the different categories.

⁶ There are a few other pathways to lawful permanent resident status, but they are quantitatively not important.

Table 1
Family-Sponsored Admission Categories

Category	Sponsor	Eligible Family Members	Number of Visas
IR	US citizen	Spouses, parents, and unmarried children under 21	Unlimited
F1	US citizen	Unmarried sons and daughters and their minor children	23,400
F2A	LPR	Spouses and minor children	87,900
F2B	LPR	Unmarried sons and daughters	26,300
F3	US citizen	Married sons and daughters	23,400
F4	US citizen	Siblings and their minor children	65,000

SOURCE.—Congressional Research Service, summary of Immigration and Nationality Act §203(a) and §204; 8 U.S.C. §1153.

NOTE.—Unused visas from the previous year can increase the number of visas in individual categories. For more details, see Kandel (2014).

In addition, per-country ceilings regulate that citizens from a single country cannot account for more than 7% of all available visas.⁷

Demand for family migration far exceeds the number of available visas per year. As a result, the caps have created a large backlog of individuals whose LPR visa petitions have been approved by the US Citizenship and Immigration Services but for whom visas are not available. As of November 1, 2017, 3.95 million approved LPR visa petitions in the family track were pending and waiting for visa processing (Kandel 2018). Within admission categories, the Department of State processes visa applications in the order in which petitions were filed. Visa applicants typically wait for several years before they receive their visa.

The Department of State regularly publishes the waiting times for applicants who are currently invited for visa processing in its monthly Visa Bulletin. Figure 1 shows the waiting times, defined as the time between filing a petition and being invited for visa processing, for different categories of family migrants for the period 1992–2019. For example, in 2005 family migrants receiving F1 visas (unmarried sons and daughters of US citizens and their minor children) or F2A visas (spouses and minor children of LPRs) had been in the queue for about 4 years, those receiving F2B visas (unmarried sons and daughters of LPRs) for more than 8 years, those receiving F3 visas (married sons and daughters of US citizens) for about 7 years, and those receiving F4 visas (siblings of US citizens and their minor children) for about 12 years. Because of the binding per-country ceilings, applicants from the most important countries of origin—Mexico, China, India, and the Philippines—have different and often considerably longer waiting times (see fig. A.1; figs. A.1–A.16 are available online).

Prospective family migrants face not only long waiting times but also uncertainty about the length of waiting times. At the time of filing their petition,

⁷ Exceptions are possible for category F2A (spouses and children of LPRs) and some employment-based visas. For more details, see Kandel (2018).

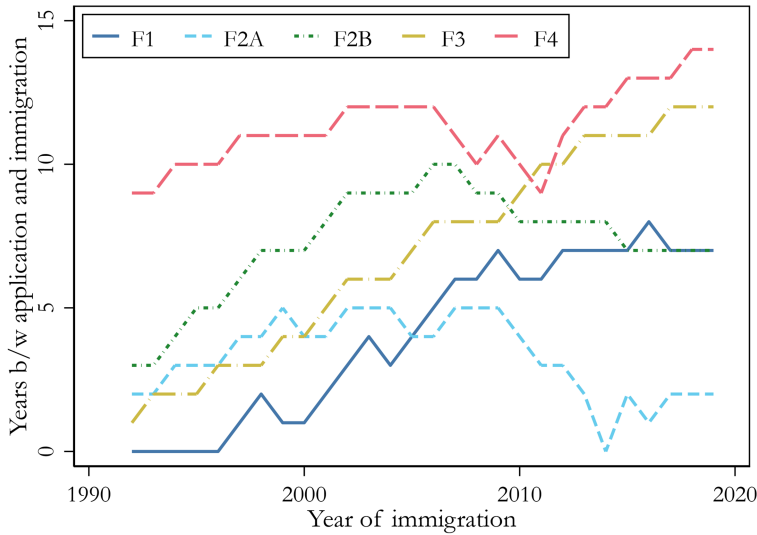


FIG. 1.—Waiting times for family migrants by admission category. The figure shows waiting times for unmarried sons and daughters of US citizens and their minor children (F1), spouses and minor children of LPRs (F2A), unmarried sons and daughters of LPRs (F2B), married sons and daughters of US citizens (F3), and brothers and sisters of US citizens (F4). Data source: US Department of State Visa Bulletins (January), own calculations. Figure A.1 shows waiting times for applicants from Mexico, China, India, and the Philippines, where the per-country ceiling is binding.

they know only the waiting times of applicants who have just been invited for visa processing in each admission category. Their effective waiting times, however, may differ depending on how many other prospective migrants from the same and other countries have filed petitions in the same admission category. Indeed, as figure 1 shows, waiting times have not been constant over time. They have considerably increased for most admission categories.

Once a visa becomes available, the National Visa Center informs the applicant who then has to apply for a visa within a period of 12 months. Failure to apply for the visa results in termination of the petition (Immigration and Nationality Act §203(g)). Visas are valid only for up to 6 months, so migrants have to enter the United States within that period. Migrants receive LPR status only after arriving in the United States.⁸ Thus, family migrants have limited opportunities to time their move to the United States.

⁸ Some migrants also adjust their visa status while being in the United States. However, relative to other types of immigrants, adjustment of status is not common among family migrants. According to the 2015 Yearbook of Immigration Statistics, only 16,783 family migrants did so in 2015. Status adjustment is particularly common for spouses of US citizens and LPRs.

The combination of long waiting times, uncertainty about the length of waiting times, and the short validity of their visa forces family migrants to enter the United States within a narrow time window that is beyond their control and not known *a priori*. Family migrants are thus not able to synchronize their arrival with labor market conditions in the United States. Our identification strategy exploits this particular feature of the US immigration system to overcome the problem of endogenous migration decisions.

IV. Data

A. Data Sources and Sample Restrictions

We base our analysis on data from the 2000 US census and the ACS for the period 2000–2019, obtained via IPUMS. These data have two advantages over other potential data sources. First, the sample is large, even if we restrict it to immigrants with specific characteristics. Second, the data provide information on the year of immigration and country of origin as well as a wide range of labor market outcomes.⁹

Like other currently available datasets, however, our data do not provide information on the visa type. The US census and the ACS sample all types of permanent residents as well as aliens who reside in the United States on temporary visas or irregularly. We are hence not able to directly identify family migrants. In our main analysis, we therefore restrict the sample to immigrants from countries for which family migration is the dominant mode of migration to the United States. In section VI.C, we explore alternative strategies, which yield similar results.¹⁰

To understand the relative importance of different modes of migration to the United States, we use data from the Yearbook of Immigration Statistics for the period 1992–2019.¹¹ Published every year by the US Department of Homeland Security, the yearbooks provide information on the number of immigrants and aliens who were admitted to the United States by admission category and country of origin in that year. We only consider admission categories that are likely to be sampled by the census and the ACS. They include LPRs, students and exchange visitors on temporary visas, and temporary workers. They exclude tourists, business travelers, and diplomats. For

⁹ We cannot use the 2010 US census, as it was a short-form-only census and does not provide the relevant information.

¹⁰ Borjas and Bronars (1991) use family relationships and the timing of immigration within a household to identify family migrants in census data. We abstain from pursuing this strategy, as our results in sec. VI.B show that household formation is a function of economic conditions at immigration. Focusing on family migrants who live in the same household as the sponsor would lead to a sample selection problem.

¹¹ The Yearbook of Immigration Statistics was not published in 1990, 1991, and 1997. We approximate 1997 values as the average of 1996 and 1998 values.

each country of origin, we then calculate the share of family migrants, labor migrants (permanent and temporary), and all other migrants.¹²

We define the dominant mode of migration to account for more than 50% of yearly admissions in a given year. We classify countries as other when neither family nor labor migration alone accounts for more than 50% of yearly admissions. For our estimation sample, we consider only immigrants from an origin country that was classified as “family dominated” in the year of immigration. Table 2 lists the main countries of origin with predominantly family migration to the United States. Individuals from the Philippines (36% among all family migrants in our sample), Vietnam (16%), the Dominican Republic (16%), and Haiti (10%) make up the large majority of family migrants in our sample. For these countries, family migration remains the dominant mode throughout most of the observation period. Individuals from Guyana (4%), Bangladesh (3%), Iran (3%), Ecuador (3%), Peru (2%), Cambodia (1%), Ghana (1%), and Syria (1%) play a minor role. Table A.1 (tables A.1–A.6 are available online) lists the main countries of origin with predominantly labor migration to the United States. Most labor migrants in the sample come from India (48%), the United Kingdom (12%), and Japan (9%). Less important countries of origin include Germany (5%), France (4%), Argentina (3%), Australia (3%), Venezuela (2%), Italy (2%), Israel (2%), South Africa (1%), and the Netherlands (1%). We classify Mexico and countries in Central America as other countries. They have high rates of irregular migration to the United States (Cohn, Passel, and Gonzales-Barrera 2017), which are not captured by the Yearbook of Immigration Statistics.

Table A.2 shows that our definition of country types indeed captures meaningful differences in the composition of migrants. Family migrants account for 65% of migrants from countries with predominantly family migration. With 5% and 15%, the share of family migrants is much lower for countries with predominantly labor migration and other countries. Similarly, labor migrants account for 61% of migrants from countries with predominantly labor migration but only for 12% and 4% of migrants from countries with predominantly family migration and other countries.

In the following analysis, we hence focus on migrants from countries with predominantly family migration to the United States. We limit the sample to individuals who were between 22 and 60 years old at the time of immigration and the time of observation. This age range captures individuals who are likely to become active in the labor market at arrival. In addition, it excludes

¹² The Yearbook of Immigration Statistics does not provide information on the number of immigrants admitted to the United States in a given year but does provide information on the number of immigrant and nonimmigrant visas granted. A single individual might sequentially hold multiple nonimmigrant visas, resulting in an overcount of this individual. Since multiple visas are a potential concern for non-family migrants only, our estimated share of family migrants can be considered a lower bound.

Table 2
Main Countries of Origin with Predominantly Family Migration

	Number of Observations in Estimation Sample	% of Years Family Dominated	% of Migrants Family Based
Philippines	41,860	81.5	56.8
Vietnam	18,676	77.8	61.1
Dominican Republic	18,198	100	76.2
Haiti	11,687	100	73.9
Guyana/British Guiana	4,743	100	85.7
Bangladesh	3,437	48.1	60.3
Iran	3,428	33.3	58.7
Ecuador	3,176	25.9	57.4
Peru	2,453	22.2	53.9
Cambodia	1,722	77.8	73.0
Ghana	1,130	33.3	55.4
Syria	1,114	51.9	54.4
Other	5,351	100	62.8
Total	116,975	73.5	65.7

NOTE.—Share of years family dominated is the share of years between 1992 and 2018 in which a country’s dominant mode of migration to the United States was family migration. Number of observations in estimation sample is the number of observations in our sample, i.e., individuals aged 22–60 years at the time of immigration and observation who were born outside the United States as non-US citizens, did not get naturalized within the first 3 years of immigration, immigrated between 1992 and 2018, and had spent at most 10 years in the United States at the time of observation. We only consider individuals who immigrated in a year in which the country was classified as family dominated. Countries are ordered by number of observations in the estimation sample. Table A.1 presents corresponding numbers for labor migration.

minor children of US citizens who can be sponsored without waiting times (see table 1). It also excludes young adults who immigrate as students or exchange visitors as well as older parents of US citizens. These groups do not face waiting times either. We also restrict the sample to individuals who were born outside the United States as non-US citizens, did not get naturalized within the first 3 years of immigration,¹³ immigrated between 1992 and 2018, and had spent at most 10 years in the United States at the time of observation. Despite these restrictions, our sample likely includes considerable numbers of nonfamily migrants, and some family migrants could be immediate relatives who do not face long waiting times. We revisit this issue in section VI.C and show that the inclusion of migrants who do not face waiting times for their visa biases our estimates toward zero. For family migrants who face waiting times, the effects of immigrating into a recession are hence likely more severe than our conservative main results suggest.

B. Descriptive Statistics

Table 3 provides summary statistics of demographic characteristics and main outcome variables. Column 1 shows statistics for family migrants. For

¹³ Information on the year of naturalization is available in the ACS only starting in 2008. Therefore, we can apply this restriction only to part of the sample.

Table 3
Summary Statistics

	Family Migrants		Labor	Refugees	Natives	Graduates
	All	Filipinos	Migrants			
Personal characteristics:						
Female (0/1)	.62 (.49)	.66 (.47)	.49 (.50)	.50 (.50)	.51 (.50)	.50 (.50)
Age at immigration	34.9 (9.03)	35.2 (8.95)	31.8 (8.00)	33.8 (8.67)		
Age at observation	40.2 (9.17)	40.6 (9.07)	36.4 (8.28)	39.1 (8.93)	41.9 (11.2)	24.7 (3.62)
Years of schooling	12.8 (4.15)	14.7 (2.78)	15.9 (2.92)	12.0 (4.98)	13.7 (2.65)	13.7 (2.16)
At least 4 years of college (0/1)	.36 (.48)	.61 (.49)	.73 (.44)	.29 (.45)	.31 (.46)	.30 (.46)
Economic conditions:						
Unemployment rate at immigration	6.21 (2.13)	6.22 (2.11)	5.87 (2.03)	6.26 (2.27)		5.98 (2.04)
Labor market outcomes:						
Employed (0/1)	.69 (.46)	.74 (.44)	.70 (.46)	.68 (.47)	.75 (.43)	.73 (.45)
Real annual wage income (in 1,000 USD)	36.0 (36.7)	43.7 (38.5)	81.0 (74.5)	33.9 (35.5)	54.4 (54.1)	31.9 (32.2)
Real hourly wage	21.6 (22.5)	25.6 (24.0)	40.7 (34.1)	20.3 (21.0)	28.0 (25.9)	19.1 (17.7)
Occupational income score (USD)	25.0 (11.5)	28.0 (12.5)	37.7 (14.7)	23.8 (11.0)	27.4 (12.0)	25.0 (10.8)
Occupational education score	13.1 (1.81)	13.6 (1.86)	15.0 (1.90)	12.9 (1.81)	13.8 (1.82)	13.5 (1.78)
Self employed (0/1)	.061 (.24)	.038 (.19)	.065 (.25)	.063 (.24)	.088 (.28)	.035 (.18)
Public welfare assistance (0/1)	.016 (.13)	.0070 (.083)	.0032 (.056)	.060 (.24)	.017 (.13)	.017 (.13)
Social outcomes:						
Previous migrant in household (0/1)	.49 (.50)	.45 (.50)	.21 (.41)	.27 (.44)		
Being household head (0/1)	.34 (.48)	.33 (.47)	.47 (.50)	.48 (.50)	.54 (.50)	.39 (.49)

Table 3 (Continued)

	Family Migrants		Labor		Natives	Graduates
	All	Filipinos	Migrants	Refugees		
Moved between states last year (0/1)	.025 (.16)	.027 (.16)	.058 (.23)	.033 (.18)	.025 (.16)	.052 (.22)
Observations	116,975	50,520	185,761	26,686	28,023,040	6,972,754

NOTE.—This table shows summary statistics for six different samples: family migrants, Filipino migrants (the single largest group of family migrants), labor migrants, refugees, natives (US-born individuals), and US-born graduates (all education levels). Immigrant samples are restricted to individuals who were born outside the United States as non-US citizens, did not get naturalized within the first 3 years of immigration, and immigrated between 1992 and 2018. In addition, the samples are restricted to individuals who were between 22 and 60 years old at the time of immigration/observation and observed at most 10 years after immigration/graduation. For graduates, the sample is restricted to individuals who were between 19 and 33 years old at the time of observation (Schwandt and von Wachter 2019). The analysis is based on data from the 2000 US census and the 2001–19 ACS. Observations are weighted by the average annual sample weights to account for the fact that the census and ACS have very different sample sizes and that the sample size of the ACS is not constant over time. All wages are in 2019 USD. The occupational income score is the log average wage income of a worker in the same occupation, state, and year of observation. The occupational education score is the average number of years of schooling of a worker in the same occupation and year of observation. See sec. IV for more details.

comparison, the table also shows the same information for migrants from the Philippines, who constitute the largest group of family migrants, in column 2; labor migrants in column 3; refugees in column 4; natives (i.e., US-born individuals) in column 5; and US-born graduates (all education levels) in column 6. We restrict the sample to individuals who were between 22 and 60 years old at the time of immigration/observation and observed at most 10 years after immigration/graduation. For graduates, we follow Schwandt and von Wachter (2019) and restrict the sample to individuals who were between 19 and 33 years old at the time of observation.

On average, family migrants immigrated at the age of 35. They are disproportionately female (62%) and have on average 12.8 years of schooling. While this is about a year less than the average native, family migrants are more likely than natives to hold a college degree (36% vs. 33%). By comparison, labor migrants immigrated at the age of 32. Their sex ratio is more balanced (49% female), and they have considerably higher levels of schooling (15.9 years). Family migrants are slightly younger than natives at the time of observation (40.2 vs. 41.6 years). Labor migrants are considerably younger (36.4 years).

Family migrants faced an IUR of 6.21%.¹⁴ This is higher than the IUR faced by labor migrants (5.87%) and in line with our argument that family migrants cannot synchronize their arrival with labor market conditions in the United States.

In terms of labor market outcomes during the first 10 years in the United States, family migrants do substantially worse than natives. Their employment

¹⁴ State-level unemployment rates come from the Local Area Unemployment Statistics, published by the US Bureau of Labor Statistics.

levels (69% vs. 84%), annual wage income conditional on being employed (USD 36,000 vs. 54,000 in 2019 USD), real hourly wages conditional on being employed (USD 22 vs. 28), and occupational income scores (USD 25 vs. 27.5; i.e., the average hourly wage of a worker in the same occupation,¹⁵ state, and year of observation) are lower. Not surprisingly, labor migrants have a strong labor market performance. They have substantially higher wage incomes (USD 81,000), hourly wages (USD 41), and occupational income scores (USD 38) than family migrants or natives.

Family migrants are about as likely to receive public welfare assistance as natives (1.6% vs. 1.3%). This can be explained by the fact that family migrants have no access to welfare benefits in their first years after arrival. Finally, family migrants are much less likely to be household heads than natives (34% vs. 54%). They are not more likely to have moved across states within the last year (2.5% vs. 2.5%). About half of family migrants also live with a family member with longer tenure in the United States, potentially their sponsors. By contrast, labor migrants are generally more independent and mobile.

The demographic characteristics of migrants from the Philippines are comparable to those of all family migrants. However, they have significantly higher levels of schooling (14.7 vs. 12.8 years) and college graduation rates (61% vs. 36%). They also do better in the labor market than all family migrants but still worse than natives.

The table also shows characteristics of refugees and US-born graduates. Like family migrants, these population groups have little control over when they enter the US labor market and therefore provide an interesting comparison. Consistent with our approach for family and labor migrants, we identify refugees as coming from countries for which refugee migration accounts for more than 50% of yearly admissions. Individuals from Iraq (14.5%), Myanmar (11.4%), Ethiopia (10.5%), Ukraine (8.8%), and Somalia (6.5%) make up most refugees in our sample. By and large, the demographic and labor market characteristics of refugees are similar to those of family migrants. Graduates are much younger (24 vs. 40 years), as we consider graduates from all levels of education and restrict the sample to individuals who were between 19 and 33 years old at the time of observation. Graduates are close to family migrants regarding average years of schooling (13.3 vs. 12.8 years). And despite being considerably younger, they perform almost as well as family migrants on the labor market.

¹⁵ We classify occupations following the IPUMS approach occ1990 and using the third hierarchical level with 101 distinct occupations. We deviate from this approach only in minor instances to combine small occupations with very few observations. Specifically, we (i) combine all mining occupations; (ii) combine legislators and top executives; (iii) add funeral directors, postmasters, and mail superintendents to managers of service organizations, not elsewhere classified; and (iv) combine managers of horticultural speciality farms and horticultural speciality farmers. Overall, we distinguish between 90 occupations in our sample.

Before explaining our empirical approach, we explore the bivariate relationship between initial economic conditions and labor market outcomes in our sample of family migrants. Figure 2 plots average employment rates (fig. 2A) and real annual wage incomes (fig. 2B) for different IURs. As expected, higher IURs are associated with lower employment rates and lower annual wage incomes conditional on being employed in subsequent years.

V. Empirical Approach

A. Econometric Specification

We use two specifications to estimate how state-level IURs shape the economic integration of family migrants over time. The first specification takes the following form:

$$y_{i,s,t,m} = \alpha + \sum_{t=m+1}^{m+10} \beta_{(t-m)} \text{IUR}_{s,m} + \theta_s + \lambda_t + \chi_m + \gamma_{(t-m)} + X_{it} \Delta + \epsilon_{it}, \quad (1)$$

where $y_{i,s,t,m}$ measures the outcome of interest for family migrant i who was observed in state s in year t and immigrated in year m . Our focus is on four key labor market outcomes: employment, log real annual wage income, log real hourly wages, and occupational quality as measured by the log occupational income score (the log average wage income of a worker in the same occupation, state, and year of observation).

The explanatory variable of interest is the state-level $\text{IUR}_{s,m}$. We use a fully flexible model and allow the effect of $\text{IUR}_{s,m}$ to vary with the number

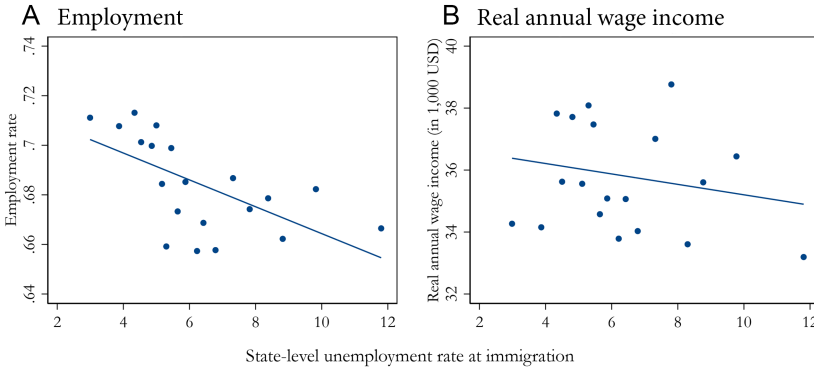


FIG. 2.—Correlation between the IUR and average labor market outcomes. The graphs show the correlation between the state-level unemployment rate in the year of immigration and the employment rate (A) and the real annual wage income (in 2019 USD) conditional on being employed (B) for our entire sample of family migrants. Binned scatterplot with 20 equally sized bins.

of years in the United States ($t - m$). We estimate the effect for the first 10 years in the United States. The coefficients $\beta_{(t-m)}$ capture the effect of the IUR plus the weighted sum of the effects of unemployment rates in subsequent years. These parameters are of interest because they capture the overall effect of initial economic conditions for a typical evolution of state-level unemployment rates afterward (Oreopoulos, von Wachter, and Heisz 2012; Schwandt and von Wachter 2019; von Wachter 2020). In other words, the parameters capture the full difference in labor market outcomes of family migrants who arrive at different stages of the business cycle. This qualification matters because state-level unemployment rates exhibit serial correlation, which only disappears after about 5 years.¹⁶ For simplicity, we will follow the literature and only refer to the IUR in the following. However, the reader should keep in mind that the coefficient also reflects subsequent labor market conditions that are correlated with the IUR.

The terms θ_s , λ_t , χ_m and $\gamma_{(t-m)}$ represent a full set of state, year-of-observation, year-of-immigration, and years-in-the-US fixed effects.¹⁷ They control for persistent differences in economic conditions and immigrant characteristics across states (θ_s), nationwide economic conditions at the time of observation (λ_t), nationwide economic conditions at the time of immigration and changes in the characteristics of immigrant cohorts (χ_m), and the general path of economic integration over time ($\gamma_{(t-m)}$). The term X_{it} is a vector of migrant-level controls. It includes age, age squared, gender, education levels, and a full set of country-of-origin dummies. The term ϵ_{it} is an independent and identically distributed error term.

The second specification models the dynamic effects of the IUR with a fifth-order polynomial:

$$y_{i,s,t,m} = \alpha + \sum_{j=1}^5 \beta_j \text{IUR}_{s,m}(t-m)^j + \theta_s + \lambda_t + \chi_m + \gamma_{(t-m)} + X_{it}\Delta + \epsilon_{it}. \quad (2)$$

The polynomial specification is less flexible and introduces more restrictions on the functional form. However, it generates more precise estimates, as there are fewer parameters to be estimated. For the main results, we present both specifications side by side. Both specifications yield very similar results. For additional results, we therefore use the polynomial specification.

¹⁶ The first-order autoregressive (AR(1)) coefficient is 0.81 (including state and year fixed effects). Serial correlation weakens considerably in the following years. The AR(2), AR(3), and AR(4) coefficients are 0.58, 0.36, and 0.15, respectively. Serial correlation disappears after 5 years. The AR(5) coefficient is close to zero (−0.01) and no longer statistically significant.

¹⁷ As year of immigration, years in the United States, and year of observation are not separately identified, we drop two dummies from the year of observation instead of just one, as for the other fixed effects.

We estimate all models using ordinary least squares (OLS) and cluster standard errors at the state-year-of-immigration level. We also weight observations by the average annual sample weights. Doing so accounts for the fact that the census and ACS have very different sample sizes and that the sample size of the ACS is not constant over the period 2000–2019.

B. Identification

Our key identifying assumption is that the immigration of family migrants is independent of state-level IURs. As we have argued above, family migrants cannot choose their date of immigration based on economic conditions. The combination of long waiting times, uncertainty about the length of waiting times, and the short validity of their visa does not allow them to synchronize their arrival with labor market condition in the United States. In addition, family migrants typically join their sponsor's household in the United States. Our own calculations using administrative data from the Philippines on the universe of family migrants from the Philippines to the United States reveal that this is the case for 98.5% of all family migrants. Hence, family migrants do not choose the US state on the basis of economic conditions either. We therefore argue that the local economic conditions family migrants face at arrival in the United States are exogenous.

While migrants cannot influence waiting times, they could decide to forgo their visa when it becomes available. If this decision is correlated with migrants' characteristics, this would create a selection problem. We argue that it is unlikely that prospective migrants decide to forgo their visa to avoid temporary unfavorable conditions in the United States. First, forgoing a visa is an extremely costly decision for family migrants. They would need to file a new application and be placed at the end of the visa queue. Second, income gains from migrating to the United States for migrants from lower-income countries are large (McKenzie, Stillman, and Gibson 2010; Clemens, Montenegro, and Pritchett 2019), which makes migration attractive even if conditions in the United States are not optimal.

We can conduct two indirect tests of this identifying assumption. First, if family migration is indeed exogenous to local economic conditions, the IUR should not be correlated with the composition and size of family migrant inflows to US states. We test this prediction by regressing (i) the state-level IUR on individual-level migrant characteristics (including the full set of fixed effects outlined above) and (ii) the log number of family migrants arriving in a given state and year on the state-level unemployment rate in that year (including state and year-of-immigration fixed effects). Table 4 summarizes the results. The upper panel shows that individual characteristics of migrants (age, gender, years of schooling) are not correlated with the IUR. All coefficients are close to zero. This observation holds for our entire sample (col. 2) as well as when we restrict the sample to migrants who were interviewed within the first 3 years of arrival and thus can be considered new arrivals

Table 4
Correlation between the IUR and Migrant Characteristics and the Number of Migrants

	Family Migrants		Labor Migrants	
	Three Years	Ten Years	Three Years	Ten Years
Migrant Characteristics				
Age at immigration (/100)	.0237 (.0442)	.0438 (.0295)	-.0138 (.0404)	-.0182 (.0299)
Female (0/1)	.0045 (.0081)	-.0033 (.0052)	-.0057 (.0044)	-.0043 (.0029)
Years of schooling	-.0011 (.0012)	-.0022*** (.0008)	-.0029** (.0011)	-.0019** (.0008)
<i>N</i>	37,163	116,975	82,109	185,761
<i>F</i> -test (<i>p</i> -value)	.677	.0111	.0680	.100
Number of Migrants				
Number of migrants (log)	-.0086 (.0186)	-.0208 (.0228)	.0121 (.0363)	-.0090 (.0336)
<i>N</i>	1,122	1,377	1,122	1,377
<i>F</i> -test (<i>p</i> -value)	.642	.360	.739	.788

NOTE.—The upper panel shows the coefficients from OLS regressions of the state-level unemployment rate in the year of immigration on migrant characteristics as well as a full set of state, year-of-observation, year-of-immigration, and years-in-the-US fixed effects. The lower panel shows the coefficients from OLS regressions with the data collapsed by state and year of immigration. We regress the log number of family migrants observed in each state and year-of-immigration cell on the respective unemployment rate in that cell, controlling for state and year-of-immigration fixed effects. We replace missing values in empty cells with no observed migrants with zeros. In cols. 1 and 3, the sample is restricted to migrants who immigrated in the 3 years prior to observation in order to capture recent inflows. Standard errors are clustered at the state-year-of-immigration level.

** Statistically significant at the 5% level.

*** Statistically significant at the 1% level.

(col. 1). For the entire sample, we detect that an additional year of schooling is associated with a 0.0022 percentage point lower IUR. This is a very weak association and not economically significant. The lower panel shows that the size of family migrant inflows is unrelated to the IUR, too. The last two columns show the same associations for labor migrants, for whom we observe similar results.

Second, we test whether the unemployment rates before immigration are related to labor market outcomes of family migrants. If our identification strategy succeeds in decoupling the migration decision from economic conditions at the time of arrival, unemployment rates before arrival should not be correlated with labor market outcomes of family migrants. We test this prediction by using the sample specified above and regressing our four labor market outcomes on state-level unemployment rates at different points before immigration. Table A.5 summarizes the results. It shows that unemployment rates in the years prior to immigration are not correlated with labor market outcomes. All coefficients are close to zero. Only the unemployment rate in the year of immigration has a significant and negative effect on labor

market outcomes. Overall, these results support our argument that migration decisions of family migrants are exogenous to economic conditions at the time of immigration.

C. Internal and Return Migration

Measuring the unemployment rate at the state rather than the national level provides us with a more accurate measure of economic conditions at arrival. Figure A.2 shows that there is considerable heterogeneity in unemployment rates between states over time, even after accounting for state and year fixed effects.

We do not observe the state of arrival, only the current state of residence and the state of residence in the previous year. We use the earliest state of residence to proxy for the state of arrival. To the extent that migrants move between states, we measure IURs with an error. Figure A.3 shows that family migrants are less likely to move between states than labor migrants. This is likely due to family ties. Still, in the first 3 years in the United States, more than 3% of family migrants report having moved between states in the previous year. However, as we show later, family migrants do not seem to move in response to initial economic conditions. If interstate migration is independent of initial economic conditions, measurement error will increase over time. This classical measurement error would lead to attenuation bias, so we would overestimate the speed of convergence. The magnitude of our estimated effects should therefore be conservative.

Given the cross-sectional nature of our data, we are not able to track individuals over time. Selective return migration could thus potentially bias our results. For instance, we may wrongly conclude that immigrating into a recession is not associated with poor labor market outcomes if immigrants with poor labor market outcomes are more likely to leave the United States and disappear from our sample. However, Dustmann and Görlach (2015) document that family migrants, and those with an Asian background in particular, are least likely to return. Borjas and Bratsberg (1996) analyze the patterns of return migration and arrive at the same conclusion.

We provide additional evidence that selective return migration is unlikely to bias our estimates. Figure A.4 plots the number of migrants observed in year t relative to the number of migrants in the year of immigration m for different years of immigration over time. There is no evidence of return migration for family migrants. If anything, the number of family migrants observed in the sample increases over time, especially for the early 2000s, potentially indicating that newly arriving family migrants were undersampled. The picture looks different for labor migrants. Their number in the sample substantially decreases over time, suggesting that many eventually return. This result is consistent with recent evidence from Akee and Jones (2019). Figure A.5 shows that these results hold when we aggregate over all years of immigration and look at changes in cohort size of family and labor migrants over time.

VI. Results

A. Economic Conditions at Arrival and Labor Market Outcomes

1. Main Results

Figure 3 summarizes our main regression results. It shows by how much a 1 percentage point increase in the IUR affects four key labor market outcomes: employment, log real annual wage income, log real hourly wages, and occupational quality as measured by the log occupational income score (the log average wage income of a worker in the same occupation, state, and year of observation). It plots the year-specific coefficient β for the first 10 years in the United States. Diamonds refer to the flexible specification (eq. [1]),

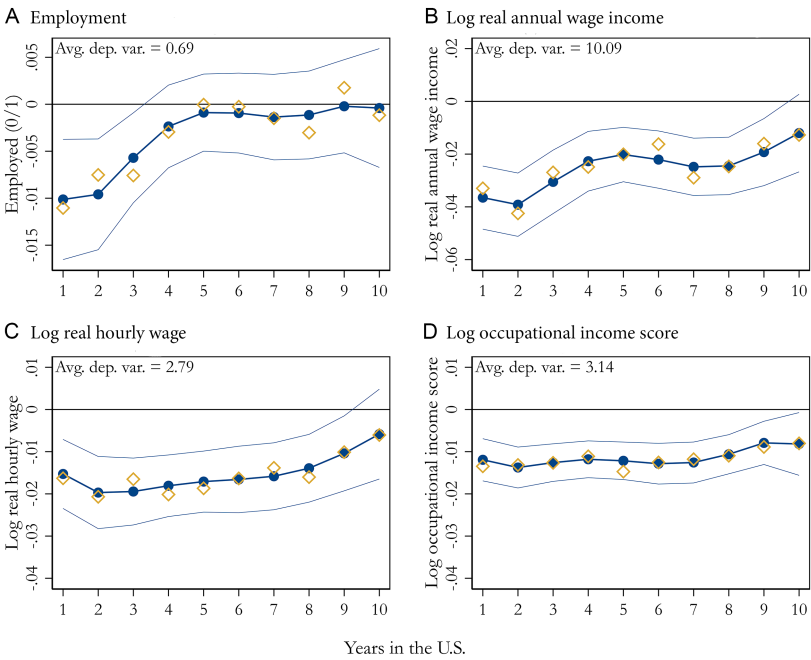


FIG. 3.—Effect of the IUR on labor market outcomes of family migrants. The x -axis shows the years since immigration. The y -axis shows the effect of a 1 percentage point higher state-level unemployment rate in the year of immigration on the outcome variable in the respective year. Diamonds refer to the flexible specification (eq. [1]), and circles (with 95% confidence intervals) refer to the less flexible polynomial specification (eq. [2]). All regressions include a full set of state, year-of-observation, year-of-immigration, and years-in-the-US fixed effects. They also include age, age squared, gender, education, and country-of-origin dummies. Standard errors are clustered at the state-year-of-immigration level. The corresponding regression tables are in table A.3.

circles to the less flexible polynomial specification including the 95% confidence interval (eq. [2]). The results for both specifications are very similar. In the following, we therefore focus on the polynomial specification. The corresponding regression tables are in table A.3.

A 1 percentage point increase in the IUR has only a small and short-lasting effect on the employment status of family migrants (fig. 3A). In the first year after arrival, family migrants are 1.0 percentage points less likely to be employed. Compared with a scenario where family migrants have the same labor force participation rate (LFPR) and propensity to be employed as the overall working-age population, this effect is relatively small. In that case, a 1 percentage point increase in the IUR should lower the initial employment rate by $1/\text{LFPR}$. With a LFPR of about 0.75 in this age group, this effect amounts to about 1.3 percentage points. Family migrants hence do relatively well in terms of finding a job. Consistent with this result, employment rates converge quickly and are no longer affected after 5 years in the United States.

The picture looks different for real annual wage income (fig. 3B). Conditional on being employed, a 1 percentage point increase in the IUR decreases annual wage income in the first 3 years by about 4%. There is only slow convergence afterward to a persistent negative effect of about 2%.

The negative effect on wage income is largely due to lower wage rates (fig. 3C). A 1 percentage point increase in the IUR decreases hourly wages by about 2%. The effect is relatively constant over time, with some convergence at the end of the 10-year period. The gap between the effect on annual wage income and the effect on hourly wages indicates that family migrants who arrive at higher unemployment rates work fewer hours.

We also find evidence for occupational downgrading (fig. 3D). A 1 percentage point increase in the IUR persistently decreases occupational income scores by about 1%. Family migrants who immigrate into a recession are hence pushed into lower-paid occupations.¹⁸

To better understand the nature of occupational downgrading, we consider two additional outcomes: (i) an occupational education score and (ii) task measures that characterize the occupations that family migrants take up. The occupational education score is defined analogously to the occupational income score and measures the average number of years of schooling of a worker in the same occupation and year of observation. The task measures come from Acemoglu and Autor (2011) and are composite measures of

¹⁸ Occupational income scores are expressed in hourly wage rates and can thus be compared to actual hourly wage rates. Interestingly, occupational income scores are substantially higher than actual hourly wage rates (the respective log values are 3.14 vs. 2.79). This is an indication that immigrants earn less than natives within occupations. It also implies that occupational downgrading accounts for almost the entire reduction in hourly wages in absolute terms (see fig. 5).

O*NET (Occupational Information Network) work activities. We distinguish between five different measures: (i) nonroutine cognitive—analytical, (ii) nonroutine cognitive—interpersonal, (iii) routine cognitive, (iv) routine manual, and (v) nonroutine manual. Each measure is standardized to have a cross-occupation mean of 0 and a standard deviation of 1. Nonroutine cognitive tasks are typical in managerial, professional, and technical occupations. Routine cognitive tasks are typical in clerical, administrative, and sales occupations. Routine manual tasks are typical in production and operative occupations. Nonroutine manual tasks that require flexibility and adaptability are typical in service occupations. We provide a more detailed description of the task measures at the end of the appendix (available online).

A 1 percentage point increase in the IUR decreases the occupational education score by about 0.3 years of schooling (fig. 4A). Family migrants who arrive at times of high unemployment also end up in occupations with different task profiles (fig. 4B). A higher IUR increases the likelihood of having an occupation with manual tasks (both routine tasks, such as controlling machines and processes, and nonroutine tasks, such as operating vehicles or

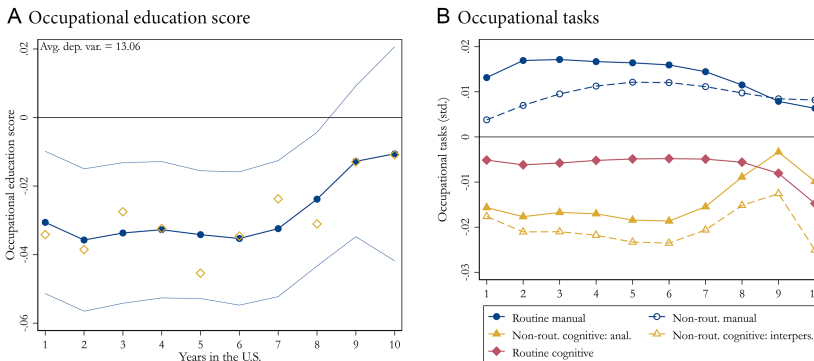


FIG. 4.—Effects of the IUR on additional occupational characteristics. The x -axis shows the years since immigration. The y -axis shows the effect of a 1 percentage point higher state-level unemployment rate in the year of immigration on the outcome variable in the respective year using the polynomial specification (eq. [2]). The outcome in A is an occupational education score, which is the average number of years of schooling of a worker in the same occupation and year of observation. The outcomes in B are occupational task measures defined by Acemoglu and Autor (2011). They are standardized to have a cross-occupation mean of 0 and a standard deviation of 1. We describe the task measures at the end of the appendix. All regressions include a full set of state, year-of-observation, year-of-immigration, and years-since-immigration fixed effects. They also include age, age squared, gender, education, and country-of-origin dummies. Standard errors are clustered at the state-year-of-immigration level.

mechanized devices). At the same time, a higher IUR decreases the likelihood of having an occupation with nonroutine cognitive tasks (both analytical tasks, such as analyzing data or thinking creatively, and interpersonal tasks, such as guiding and directing subordinates). The IUR has no major effect on the likelihood of having an occupation with routine cognitive tasks. Family migrants who immigrate into a recession hence experience occupational downgrading along a number of dimensions. They are more likely to enter occupations with lower average wages, lower average educational requirements, and an emphasis on manual rather than nonroutine cognitive tasks. This downgrading might also affect the type of human capital they accumulate on the job and explain why the effects are relatively persistent.

To more systematically understand the income loss of immigrating into a recession, we decompose the effect into four components: income loss due to occupational downgrading as measured by the occupational income score, due to a reduction in residual hourly wages (i.e., beyond the income loss due to occupational downgrading), due to a reduction in working hours (intensive margin of labor supply), and due to a lower probability of being employed (extensive margin of labor supply).

To do so, we first calculate the average employment rate, annual working hours, annual wage income, hourly wage, and occupational income score of family migrants by years in the United States. We can now obtain the reduction in annual wage income due to occupational downgrading (a) by multiplying the effect of a 1 percentage point increase in the IUR on log occupational income scores times the average occupational income score and average hours worked. Multiplying the effect of a 1 percentage point increase in the IUR on log hourly wages times the average hourly wage rate and average working hours gives us the total effect on annual wages due to lower hourly wages (b). To obtain the effect due to a reduction of the labor supply at the intensive margin (c), we multiply the effect of a 1 percentage point increase in the IUR on annual hours worked (shown in fig. A.14*d*) times the average hourly wage rate. The sum of (b) and (c) gives us the total effect on annual wage income conditional on being employed. To obtain the effect due to a reduction of the labor supply at the extensive margin (d), we multiply the effect of a 1 percentage point increase in the IUR on being employed times the average annual wage income.

Figure 5 shows the results of the decomposition for a 1 percentage point increase in the IUR. All four components play a considerable role. Over the first 10 years in the United States, the estimated cumulative loss due to occupational downgrading is about USD 4,000, which also explains the reduction in income due to lower hourly wages that amounts to USD 4,591. The cumulative income loss due to a reduction in working hours is about USD 1,600, and that due to lower employment rates is about USD 900 (all in 2019 USD discounted with 5%). Overall, a 1 percentage point increase in the IUR decreases annual wage income by about USD 1,100 in the first years and is still

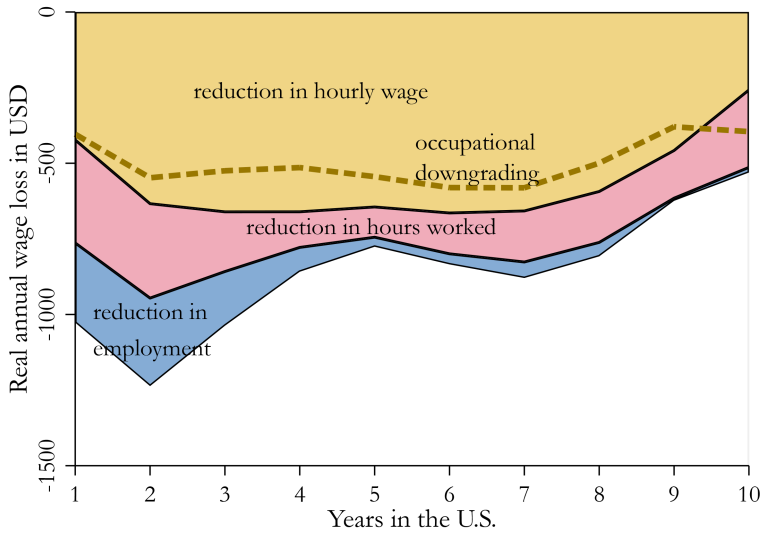


FIG. 5.—Decomposition of the effect of a 1 percentage point increase in the IUR on annual wage income. The figure shows the absolute loss in real annual wage income (in 2019 USD) as a result of a 1 percentage point higher unemployment rate in the year of immigration. It decomposes the income loss into four components: (i) income loss due to a reduction in hourly wages (top area), (ii) reduction in hourly wages that is due to occupational downgrading (dashed line), (iii) income loss due to a reduction in working hours conditional on being employed (middle area), and (iv) income loss due to a lower probability of being employed (bottom area).

more than USD 500 after 10 years in the United States.¹⁹ The estimated loss in the net present value wage income over the entire 10-year period amounts to about USD 7,100. A 5 percentage point rise in unemployment rates, a typical increase in a large recession (von Wachter 2020), would hence reduce the 10-year wage income of an average family migrant by USD 35,500 (assuming linearity of the effect). This is a large effect. We can compare this earnings loss to the earnings loss of graduating in an equally large recession. We follow Schwandt and von Wachter (2019) and consider all education levels, but we restrict the sample to US-born graduates. We use the ACS data and our econometric specification, replacing the year of immigration with the year of graduation (also see fig. 6). A 5 percentage point rise in unemployment rates reduces the 10-year wage income of an average graduate by USD 23,000. The effects are hence more severe for family migrants.

¹⁹ The effect on annual wage income in absolute amounts appear to be more persistent than the relative effect presented in fig. 3. This is because the relative effect translates into a fairly constant absolute effect, as overall income levels rise quickly with time spent in the United States.

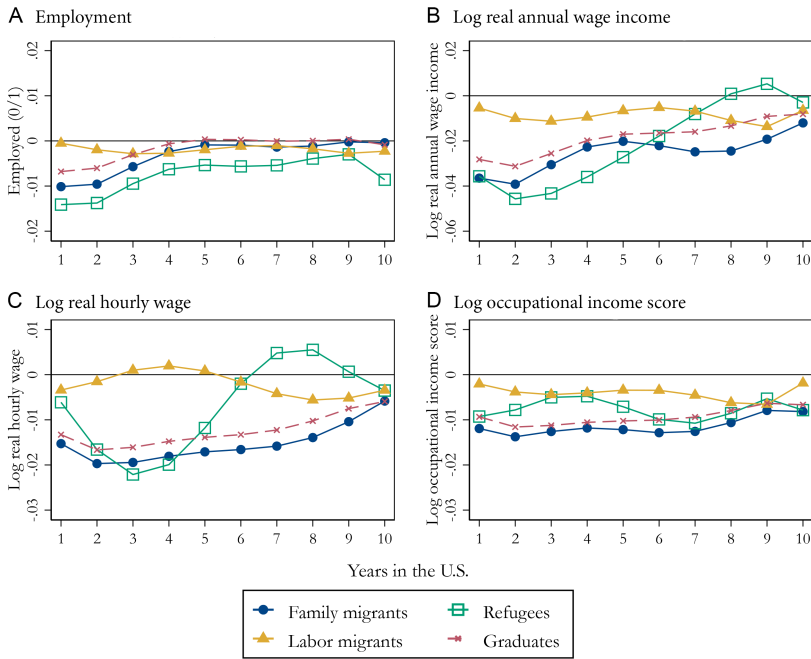


FIG. 6.—Effect of the IUR on labor market outcomes of family migrants, labor migrants, refugees, and US-born graduates. The figure reruns the main analysis for four different samples: family migrants, labor migrants, refugees, and US-born graduates (all educational levels). Section IV outlines the construction of these samples. The x -axis shows the years since immigration/graduation. The y -axis shows the effect of a 1 percentage point higher state-level unemployment rate in the year of immigration/graduation on the outcome variable in the respective year using the polynomial specification (eq. [2]). All regressions include a full set of state, year-of-observation, year-of-immigration/graduation, and years-since-immigration/years-since-graduation fixed effects. They also include age, age squared, gender, education, and country-of-origin dummies.

By how much does immigrating into a recession slow down the economic assimilation of family migrants? To study the convergence toward natives, we first calculate the average employment rate and annual wage income of family migrants by years in the United States. We then use the estimated marginal effects to derive the employment rate and annual wage income for a 5 percentage point increase in the IUR. Figure A.6*a* shows that the IUR has only a relatively small and short-lasting effect on assimilation in terms of employment. The employment rate of family migrants also converges to the native employment rate after about 9 years. A higher IUR, however, slows down assimilation in terms of annual wage income (fig. A.6*b*). According

to our estimates above, a 1 percentage point increase in the IUR decreases annual wage incomes by 2%–4%. This relative effect translates into a fairly constant absolute effect, as overall income levels rise quickly with time spent in the United States. The figure suggests that a 5 percentage point increase in the IUR delays assimilation in terms of annual wage income by about 2 years. The figure also shows that the income levels of family migrants do not fully catch up with those of natives. After 10 years, the average immigrant-native gap in annual wage income is about USD 12,100. A 5 percentage point increase in the IUR widens the gap by an additional USD 2,500.

How does entering the labor market in a recession compare across different population groups? Figure 6 compares the effects across (i) family migrants, (ii) labor migrants, (iii) refugees, and (iv) US-born graduates (all education levels). Our results are qualitatively similar for family migrants and graduates. Both groups have little control over when they enter the US labor market. However, all effects are more pronounced for family migrants than for graduates. Refugees experience slightly larger and more persistent effects on employment. However, they also see more convergence in annual wage income and hourly wages. By contrast, we hardly observe any effects on the labor market outcomes of labor migrants, especially in the first years after arrival. By definition, these migrants arrive with a job. Their economic integration should hence be less susceptible to economic conditions at arrival.

We also test whether the IUR has a nonlinear effect on our main outcomes by using a quadratic specification. Table A.4 shows that the IUR has a nonlinear effect on employment, which decreases as the IUR increases (col. 1). Still, the effect of the IUR on employment remains small. We do not find such evidence for other outcomes (cols. 2–4). The coefficients on the quadratic terms are small and far from being statistically significant.

In addition, we test for effect heterogeneity by gender and level of education. Overall, we find limited effect heterogeneity for both dimensions. The employment effect is slightly more negative for female than for male family migrants (fig. A.7). However, the opposite is true for the effect on hourly wages. There are no major differences in annual wage income, suggesting that male family migrants work longer hours to make up for lower hourly wages. The initial labor market effects are relatively similar for family migrants with and without a college degree (fig. A.8). The effects on log occupational income scores are a bit more pronounced for higher-skilled family migrants, though. These results also hold when we restrict the sample to individuals who were older than 30 years at the time of immigration to minimize the possibility that the level of education is potentially affected by the IUR (fig. A.9).

2. Robustness Checks

We check the robustness of our main results to address several potential concerns. First, our analysis of annual wage income is conditional on being

employed and might therefore suffer from sample selection bias. So far, we have implicitly assumed that sorting into employment is unrelated to the IUR and annual wage income. This assumption may not be unreasonable in our case. The IUR has only a small and short-lasting effect on employment. It hence does not considerably affect the pool of individuals for whom we observe annual wage income. In addition, the analysis of effect heterogeneity does not suggest that the IUR systematically affects employment of particular groups of family migrants more strongly. Nevertheless, we address potential sample selection bias and conduct an unconditional analysis of annual wage income. We first assign zero earnings to individuals who do not report any wage income and then increase all earnings by a small amount to maintain the log specification. The effect on annual earnings depends on the amount we add (fig. A.10). When adding USD 1, the effect becomes much more pronounced. A 1 percentage point increase in the IUR decreases annual earnings in the first 4 years by about 10%. There is quick convergence in the years that follow, and the negative effect is reduced to about 5% after 8 years in the United States. When adding USD 500 and 1,000, however, the effect on annual earnings is much more similar to our original estimate. The effect of a 1 percentage point increase in the IUR is slightly larger in the first years but converges to a similar negative level after 8 years in the United States. To avoid adding an arbitrary amount, we also run a similar analysis in absolute levels with the log transformation. We look at three outcomes: annual wage income conditional on being employed, annual wage income with incomes for nonemployed individuals set to zero, and total earnings (labor income and all other forms of income, including transfers). Figure A.11 shows that the effects are extremely similar across the three outcomes. It is also similar to the results of our decomposition exercise that is based on the conditional log specification but shows the wage income loss in levels (fig. 5). These exercises suggest that potential sample selection bias has no major implications for our results.

Second, economic conditions at the time of observation could also affect labor market outcomes. To the extent that current and initial economic conditions are correlated, our results reflect the effect of both initial and subsequent conditions. We use state-year-of-observation fixed effects to address this issue. The effects are qualitatively similar but smaller (fig. A.12). The smaller effect sizes could be due to state-year-of-observation fixed effects capturing some of the typical evolution of state-level unemployment rates after arrival (which were previously captured by the IUR).

Third, local economic conditions in the United States and in the country of origin might not be fully independent. For instance, migrant networks could transmit economic shocks from different US states to countries of origin (via remittances, trade, or foreign direct investment links). Because of long waiting times for a visa, such links would not change the composition of new family migrants in the short run. But family migrants could potentially change their

labor market behavior in the United States to support family members left behind through remittances. We use country-of-origin-year-of-immigration fixed effects to capture all shocks that are specific to a country of origin in the year of immigration. Our results remain unchanged (fig. A.13).

Fourth, we also investigate the effects on self-employment, labor income (i.e., wage, business, and farm income), and total earnings (i.e., labor income and all other forms of income, including transfers). Our results are robust to these more comprehensive definitions of income (fig. A.14). The effect on self-employment is positive but close to zero and not statistically significant.

B. Coping Strategies

1. *Migrant Networks*

A large literature has shown that migrant networks substantially help migrants to find a job (e.g., Munshi 2003; Patel and Vella 2013; Battisti, Peri, and Romiti 2022). The support provided by migrant networks might be particularly beneficial in a recession when the scarcity of jobs makes it more difficult for recent migrants to compete in the labor market. With relatively low levels of destination-specific human capital and work experience, it is difficult for migrants to signal their ability to potential employers. Migrant networks can help migrants to overcome at least part of this information asymmetry and thus increase migrants' attractiveness to employers. Network support should matter less in a boom when more jobs are available and job search is less competitive.

To test this prediction, we include an interaction in our specification and analyze whether the employment and wage effects of immigrating into a recession differ by the size of the local migrant network. We define local network size as the share of fellow countrypeople among the working-age population (defined as 22–60 years). We calculate this variable at the state level and for three points in time: for 1990 and 2000, based on census data, and for 2010, based on data from the ACS waves conducted in 2009, 2010, and 2011.²⁰ The initial network size of each migrant in our sample is then given by the value of the network variable in the year that predates the year of immigration. For instance, a migrant who arrived in 2003 is assigned the network as measured in 2000.

We restrict the sample to migrants who had not spent more than 3 years in the United States to capture the initial integration into the labor market. We look at average effects across these years to increase the precision of our estimates. We also focus on migrants from the Philippines, Vietnam, the Dominican Republic, and Haiti. The samples of migrants from other countries with predominantly family migration are too small, especially for the analysis of ethnic occupations introduced below. For the same reason, we also

²⁰ The 2010 US census was a short-form-only census and does not provide the relevant information.

exclude migrants when there are fewer than 100 observations of migrants from the same country of origin in the same state to calculate the network measure and occupational distribution.

Table 5 shows the results. Consistent with our hypothesis, larger migrant networks mitigate the negative employment effect of adverse economic conditions at arrival (col. 1). A 1 percentage point higher population share of fellow countrypeople weakens the employment effect of a 1 percentage point higher IUR by 0.26 percentage points (from -1.32 to -1.06 percentage points). However, conditional on being employed, networks are not able to cushion the effect on annual wage incomes (col. 2), hourly wages (col. 3), or log occupational income scores (col. 4). Here, the interaction terms are not statistically or economically significant. In a recession, networks can thus help family migrants to find a job but do not improve the quality of the job found.

If networks indeed facilitate job search in a recession, they should push migrants into occupations with larger networks of fellow countrypeople. We investigate this hypothesis by following Patel and Vella (2013) and constructing a measure of the concentration of fellow countrypeople in different occupations:

$$concentration_{c,o,s,t} = \frac{100 \times mig_{c,o,s,t}}{mig_{c,s,t}},$$

where $concentration_{c,o,s,t}$ measures the percentage of workers from country of origin c in state s in year t who are employed in occupation o . High values of $concentration_{c,o,s,t}$ indicate that workers from a given country of origin concentrate in these occupations. We distinguish between the same occupations as for calculating the occupational income score and again calculate the measure for the years 1990, 2000, and 2010.

Column 5 of table 5 shows that in a recession networks indeed make family migrants more likely to enter occupations with higher concentrations of fellow countrypeople. In response to a 1 percentage point higher IUR, a 1 percentage point higher population share of fellow countrypeople increases the share of fellow countrypeople in the same occupation by 3.11 percentage points. Compared with the average share of fellow countrypeople (4.98%), this is a large effect, amounting to a 62% increase. We check the robustness of this result in table A.6. It holds when we base the concentration measure on industries instead of occupations or change the required minimum number of observations of migrants from the same country of origin in the same state from 100 to zero or 200. The interaction between IUR and the share of fellow countrypeople only becomes weaker when we no longer restrict the sample to recent arrivals and look at the entire 10-year period. As some migrants will have changed their jobs over that period, the initial network effects could be attenuated.

Table 5
Effect of the IUR on Labor Market Outcomes of Family Migrants by Network Size

	Employment and Wages				% of ... in Occupation			
	Employment (1)	Log Wage Income (2)	Log Hourly Wage (3)	Log Occupational Income Score (4)	Country People (5)	Natives (6)	Mexicans (7)	Chinese (8)
Network size $\times$ unemployment rate at immigration (/100)	.26** (.11)	.02 (.21)	.14 (.16)	.01 (.07)	3.11* (1.88)	.27 (.42)	-1.27 (1.78)	1.14 (1.46)
Unemployment rate at immigration (/100)	-1.32** (.54)	-2.30* (1.28)	-1.80** (.91)	-.72* (.41)	-3.91 (8.27)	-2.84 (2.08)	-2.49 (8.70)	4.44 (6.85)
Network size	.01* (.01)	.00 (.02)	-.04*** (.01)	-.01** (.01)	-.27** (.13)	-.04 (.03)	.08 (.13)	-.01 (.11)
Observations	27,334	17,864	17,864	17,864	17,864	17,864	17,864	17,864
Mean outcome	.61	9.82	2.68	3.11	4.98	2.34	3.66	3.02

NOTE.—This table reports OLS estimates. Column heads show outcome variables. The outcome variables in the last four columns are the state-level shares of workers from the same country of origin who work in the same occupation (in percent). Unemployment rate at immigration is the state-level unemployment rate in the year of immigration (divided by 100 to improve readability). Network size is the share of migrants from the same country of origin among all working-age adults in the state at the time of immigration. The sample is restricted to migrants who immigrated in the 3 years prior to observation in order to capture recent inflows. We include only observations for which we observe at least 100 migrants from the same country of origin in the same state for calculating the occupational distribution. All regressions include a full set of state, year-of-observation, year-of-immigration, and years-in-the-US fixed effects. They also include age, age squared, gender, education, and country-of-origin dummies. Standard errors are clustered at the state-year-of-immigration level.

* Statistically significant at the 10% level.

** Statistically significant at the 5% level.

*** Statistically significant at the 1% level.

We also run a placebo regression and check whether family migrants are more likely to enter occupations in which many US-born workers (i.e., generally large occupations) or other migrant groups are employed. We calculate the concentration measure for US-born workers and for the two most important countries of origin of nonfamily migrants, Mexico and China, and use them as outcomes. The last three columns of table 5 summarize the results. In line with the hypothesis that migrant networks should not make family migrants more likely to enter occupations with higher concentration of other population groups, all three interaction terms turn out to be insignificant. Thus, migrants do not appear to be pushed into large occupations or occupations with a high concentration of migrants in general. Taken together, these results suggest that migrant networks help family migrants to find employment during periods of higher unemployment and channel them into ethnic occupations.

2. *Family Support*

Family migrants are generally not entitled to welfare benefits for at least 5 years after arrival. Instead, sponsors are obliged to support sponsored family members for 10 years or until they become US citizens (Kolker 2016). The ACS provides information on welfare receipt and the composition of household members, which we can use as a proxy for receiving support from the family.

Figure 7A confirms that adverse economic conditions at arrival have no effect on receiving welfare assistance. Family migrants seem to rely on support from the family instead. A 1 percentage point increase in the IUR increases the likelihood of living with a family member with longer tenure in the United States by about 1 percentage point over the entire 10-year period (fig. 7B). At the same time, a 1 percentage point increase in the IUR decreases the likelihood of being a household head by 0.5 percentage points in the first years, with slow convergence afterward (fig. 7C). These results are consistent with the idea that adverse economic conditions make it more difficult for recently arrived family migrants to move out of their sponsors' households because they depend on support from the family.

The immobile support received from the family could also reduce the geographic mobility of family migrants. In general, the literature has found that migrants are more mobile than natives (Green 1999; Braun and Kvasnicka 2014; Cadena and Kovak 2016). Figure 7D, however, shows that family migrants do not increase their geographical mobility in response to adverse local economic conditions on arrival. The increased dependence on family members likely increases the job search frictions of family migrants who immigrate into a recession.

Having no access to welfare benefits may also explain the small effects of a higher IUR on employment and the large effects on wages and occupational

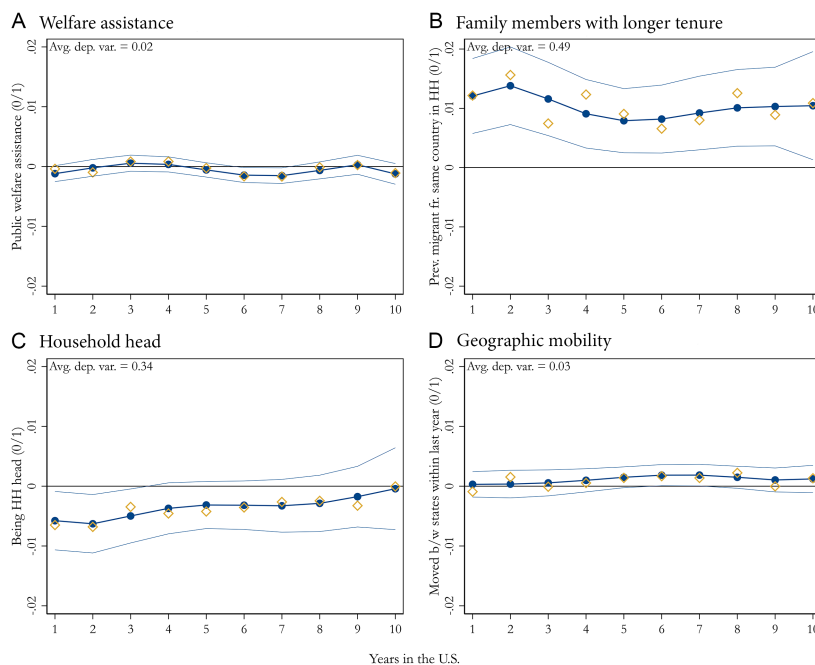


FIG. 7.—Effect of the IUR on receiving welfare assistance and support from the family. The x -axis shows the years since immigration. The y -axis shows the effect of a 1 percentage point higher state-level unemployment rate in the year of immigration on the outcome variable in the respective year. Diamonds refer to the flexible specification (eq. [1]), and circles (with 95% confidence intervals) refer to the less flexible polynomial specification (eq. [2]). All regressions include a full set of state, year-of-observation, year-of-immigration, and years-in-the-US fixed effects. They also include age, age squared, gender, education, and country-of-origin dummies. Standard errors are clustered at the state-year-of-immigration level.

quality. Family migrants may lower their reservation wage and accept lower-paying jobs to earn own income already shortly after arrival to reduce the burden on their family.

3. Investment in Human Capital

Younger family migrants might also invest in additional human capital in response to adverse economic conditions on arrival.²¹ Indeed, the opportunity costs of acquiring human capital through additional formal education might be lower in times of high unemployment. We test this hypothesis by

²¹ Battisti, Peri, and Romiti (2022) provide some evidence that seemingly unfavorable initial condition in the form of smaller coethnic networks can increase investment in human capital.

checking whether the IUR affects enrollment in tertiary education. We restrict the sample to individuals who were at most 30 years old at the time of immigration. Arguably, younger individuals should have the largest incentives to invest in human capital. We also restrict the sample to individuals who had spent at most 3 years in the United States at the time of observation to capture their initial response to the IUR. We look at average effects across these years to increase the precision of our estimates.

Table 6 shows the results. We do not find an effect of the IUR on enrollment in tertiary education (col. 1), even if we further restrict the sample to individuals who were at most 25 years old at the time of immigration (col. 2). As expected, we do not find an effect if we restrict the sample to individuals who were older than 30 years at the time of immigration (col. 3). Older individuals should be less likely to invest in additional human capital.

C. Remaining Biases and Alternative Strategies of Identifying Family Migrants

We cannot precisely identify family migrants who have to queue for their visa. We have therefore restricted the sample to immigrants from countries for which family migration is the dominant mode of migration to the United States. Yet the sample likely includes considerable numbers of nonfamily migrants and immediate relatives who do not face long waiting times for their visa. Their migration decisions may well be endogenous to initial economic conditions. In the case of endogenous migration decisions, however, our estimates are likely to be biased toward zero.

Labor migrants are less susceptible to labor market conditions at the time of arrival, as they arrive with a job. Indeed, as shown in figure 6, IURs are hardly related to labor market outcomes of employment migrants. The effects of immigrating into a recession are hence potentially more severe for family migrants than our main results suggest.

Table 6
Effect of the IUR on Family Migrants' Enrollment in Tertiary Education

	Years		
	22–30	22–25	31–60
Unemployment rate at immigration	–.002 (.004)	–.004 (.007)	.003 (.002)
Observations	13,711	5,944	23,452
Mean outcome	.176	.227	.065

NOTE.—This table reports OLS estimates. Column heads show the age restriction applied to the sample. The outcome variable is a binary indicator of enrollment in tertiary education at the time of the interview. Unemployment rate at immigration is the state-level unemployment rate in the year of immigration. The sample is restricted to migrants who immigrated in the 3 years prior to observation in order to capture recent inflows. All regressions include a full set of state, year-of-observation, year-of-immigration, and years-in-the-US fixed effects. They also include age, age squared, gender, education, and country-of-origin dummies. Standard errors are clustered at the state-year-of-immigration level.

When it comes to immediate relatives who can be sponsored without waiting times, sponsors might be more likely to select more productive family members in times of high unemployment. Otherwise, they might find themselves obliged to offer financial support. A positive selection of immediate relatives would again bias our estimates toward zero. Unfortunately, our data do not allow us to test this argument. In any case, as we have shown in table 4, IURs are not correlated with observable migrant characteristics. The potential for donors to endogenously sponsor family members might thus be limited.

To gauge the magnitude of such biases toward zero, we include an interaction between the IUR and the share of capped family migrants among all migrants moving to the United States from the same country and in the same year. If the above arguments are correct, the effect of the IUR should become more negative the higher the share of capped family migrants (as they face waiting times). Indeed, as panel B of table 7 shows, the interaction term is

Table 7
Effect of the IUR on Labor Market Outcomes of Family Migrants by Share of Capped Family Migrants from Same Country of Origin

	Employed (0/1)	Log Real Annual Wage Income	Log Real Hourly Wage	Log Occupational Income Score
A. Average Effect of Unemployment Rate at Immigration for Family Migrants				
Unemployment rate at immigration (/100)	-.344* (.205)	-2.451*** (.452)	-1.573*** (.329)	-1.183*** (.207)
Observations	116,975	84,864	84,864	84,769
B. Interaction between Unemployment Rate at Immigration and Share of Capped Family Migrants				
Share capped family migrants × unemployment rate at immigration (/100)	-.004 (.005)	-.083*** (.014)	-.066*** (.011)	-.032*** (.006)
Unemployment rate at immigration (/100)	-.623*** (.091)	-.827** (.332)	-.166 (.261)	-.150 (.131)
Share capped family migrants	.002*** (.000)	-.002** (.001)	-.002*** (.001)	-.001*** (.000)
Observations	661,431	465,844	465,844	465,419

NOTE.—This table reports OLS estimates. Column titles show outcome variables. Unemployment rate at immigration is the state-level unemployment rate in the year of immigration. The share of capped family migrants is the share of capped family migrants among all migrants moving to the United States from the same country of origin and in the same year, as measured by the Yearbooks of Immigration Statistics. All regressions include a full set of state, year-of-observation, year-of-immigration, and years-in-the-US fixed effects. They also include age, age squared, gender, education, and country-of-origin dummies. Standard errors are clustered at the state-year-of-immigration level.

* Statistically significant at the 10% level.

** Statistically significant at the 5% level.

*** Statistically significant at the 1% level.

negative for all wage-related labor market outcomes (but not for being employed). Increasing the share of capped family migrants from 0% to 100% decreases the negative effect of a 1 percentage point higher IUR on annual wage incomes by an additional 8.3%, on hourly wages by an additional 6.6%, and on occupational income scores by an additional 3.2% over the first 10 years after arrival. These effects are about three times as large as our conservative main results averaged over the same period (table 7, panel A).

We also offer two alternative strategies of identifying family migrants with waiting times. First, we use data from the Philippine government that provide information on the visa type for permanent migrants from the Philippines to the United States since 1988. The administrative data capture the universe of migrants, as every permanent migrant from the Philippines needs to register with the Commission on Filipinos Overseas (CFO) for departure clearance. In addition to the exact admission category, the data include migrant-level information on age, gender, education, date of migration, and the destination state in the United States. We use all of this information to reweight our sample of Filipino migrants in the ACS, so they resemble the characteristics of family migrants in visa categories with waiting times. Compared with Filipino migrants in the ACS sample, CFO family migrants with waiting times are on average older at the time of immigration (39 vs. 35 years), less likely to be female (53% vs. 66%), and less likely to have a college degree (55% vs. 61%). Nevertheless, as figure A.15 shows, the reweighted results are similar to the baseline results without such weights. Consistent with a cleaner identification of family migrants who faced waiting times, the wage effects of immigrating into a recession are even larger. The figure also shows that immigrating into a recession has generally somewhat more pronounced effects for Filipino migrants than other (non-Filipino) family migrants.

Second, we try to exclude immediate relatives who can be sponsored without waiting times from the sample. Such relatives include spouses, parents, and unmarried children under 21 years of US citizens (see table 1). Our main sample already excludes minor children and older parents, as it is limited to individuals aged 22–60 at the time of immigration. To also exclude younger parents, we further reduce the age range to individuals aged 22–40. A person sponsoring parents as immediate relatives needs to be a US citizen and at least 21 years old. The reduced age range hence excludes all parents who were at least 20 years old at birth of their child. To exclude spouses of US citizens, we further restrict the sample to individuals who were never married (so they cannot be sponsored by their spouse) or married couples who immigrated in the same year (so none of them are likely to be US citizens at the time of immigration).²² Figure A.16 shows that results for this subsample are similar to

²² Information on the year of marriage and on citizenship is available only from 2008 onward. We are thus not able to exploit this additional information for our analysis. In addition, the focus on married couples who immigrated in the same

the main results. The stability of the results suggests that even family migrants without waiting times do not try to synchronize their arrival with labor market conditions in the United States. The large income differences between countries of origin and the United States might explain this behavior.

VII. Conclusion

We analyze how IURs affect the economic integration of immigrants in the United States. Our identification strategy exploits long waiting times for visas that decouple the migration decision of family migrants from economic conditions at the time of arrival.

Economic conditions at arrival generate substantial heterogeneity in the economic integration of family migrants. They also help to explain immigrant-specific phenomena, such as downgrading and reliance on social networks. A 1 percentage point higher IUR has a relatively small and short-lasting effect on employment but decreases annual wage income by about 4% in the short run and 2% in the longer run. These estimates are likely lower-bound estimates of the true effects, as we are not able to precisely identify family migrants in the data. The negative effect on annual wage income is the result of a combination of occupational downgrading, which results in lower hourly wages and a reduction in working hours. Migrant networks mitigate the negative labor market effects of immigrating into a recession and channel migrants into jobs with a higher concentration of fellow countrypeople. Family migrants who arrive at times of high unemployment are also more likely to continue residing with family members. The immobile support received from the family is likely associated with an increase in search frictions, which could explain the persistence of the labor market effects. Our results show that the system of family-sponsored immigration to the United States has important consequences for the economic integration of family migrants.

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year likely also excludes spouses of LPRs (F2A visa holders). Spouses of LPRs (and US citizens) are more likely than other family migrants to adjust their visa status while being in the United States. They might hence have chosen particular economic conditions when they entered the United States on a nonimmigrant visa. In general, adjustment of status is not common among family migrants. For instance, according to the 2015 Yearbook of Immigration Statistics, only 16,783 family migrants did so in 2015. But most of these cases were concentrated among spouses of LPRs and US citizens, whom we exclude from this subsample.

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