# Shifting Paradigms: Exploring Cultural Diversity's Role in Countering Far-Right Support in Chile

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**Abstract**

Does immigration increase far-right support? Migration literature has debated this question for many decades, failing to reach a consensus on whether immigration drives the far-right vote or on the potential channels through which exposure to migrants may fuel it. The recent migration waves from Venezuela and Haiti to Chile (2015-2022) provide an ideal context to address this question, as they allow us to identify the effects of migrants with and without ethnic-cultural similarities with the natives. Using a shift-share IV model, this paper estimates the importance of ethnic-cultural differences as a mechanism by which migration affects the far-right vote. Our findings suggest that engaging with culturally different migrants curbs prejudice and anti-migrant party voting. Specifically, the results show that in the general elections of 2017 and 2021, a 1% increase in the share of culturally different migration yields up to a 5% decrease in far-right voting. Overall, aggregated migration shows no effect on far-right support, suggesting that other factors like crime or economic perceptions drive its current electoral success. However, migration positively affects the vote for more moderate right-wing parties. Migration, particularly from culturally similar populations regarding skill levels, language, and culture, significantly increases support for the traditional right.

## 1 Introduction

In the past few decades, migration has become an important phenomenon of study in political economy and political science due to the significant gain of far-right parties that openly advocate for anti-immigrant public policies in many countries of the developed world (Alesina & Tabellini, 2024; Arzheimer, 2018; Cools, Finseraas, & Rogeberg, 2021; Moriconi, Peri, & Turati, 2022; Otto & Steinhardt, 2014). Despite this growing interest, there is still no academic consensus as to whether immigration drives the far-right vote or the mechanisms by which it does so. While a few argue that contact with the migrant population increases acceptance through shared experiences (Lonsky, 2021; Pagliacci & Bonacini, 2022; Vertier, Viskanic, & Gamalerio, 2023), some find that the anti-immigrant vote is due to fears of labour competition between migrants and natives (Halla, Wagner, &

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Zweimüller, 2017), to competition over resources and social benefits (Barone, D'Ignazio, De Blasio, & Naticchioni, 2016), and to cultural concerns, like migrants changing their neighbourhoods' social, ethnic, or religious composition (Brunner & Kuhn, 2018; Mendez & Cutillas, 2014).

One of the main reasons behind this lack of consensus is that in Western industrialized countries, it is hard to disentangle the economic effects of migration (job competition) from the cultural effects. Most migrants to Europe and the United States come from countries with languages, ethnicities, and cultural traditions different from those of the local population, making it challenging to identify the mechanisms by which migration drives the far-right vote. Moreover, in these countries, migration is a continuum that dates back many decades, combining different types and reasons for migration, which could yield different acceptance levels by the native population. Acceptance of refugees or economic migrants can be different not only because of the empathy of the native population to the origin of the problem (Alrababa'h et al., 2021; Bansak, Hainmueller, & Hangartner, 2016; Newman, Hartman, Lown, & Feldman, 2015), but also in aspects such as the probability of return and rooting in the hosting country (Alrababa'h, Masterson, Casalis, Hangartner, & Weinstein, 2023; Beaman, Onder, & Onder, 2022; Camarena & Hägerdal, 2020; Dustmann & Görlach, 2016). This research uses a novel context to contribute to this research debate by analyzing the recent migration waves happening within Latin America, where, due to exceptional cultural and historical characteristics, it is easier to identify the mechanism behind the anti-immigrant vote.

Since the beginning of the 2010s, two migratory waves from Venezuela and Haiti have shaken a continent that was used to emigrating but not receiving thousands of immigrants fleeing deep economic, social, and political crises. Looking for better life expectations and economic opportunities, until May 2024, 7.7 million Venezuelans (25.5% of its 2015 population[1]) have fled the country, and more than 78% of them have chosen one of their Spanish-speaking neighbours in the region as a destination (R4V, 2024)[2]. Hence, this wave is considered the second-worst migration crisis of this century after the Syrian refugee crisis (United Nations High Commissioner for Refugees, 2022). In the case of Haiti, after the 2010 earthquake, 650 thousands (6.5% of its population) left the country, and 36% of them ended up settling in Chile (UN, 2020)[3]. As shown in Figure 1, these migrations of just a few years, have dramatically changed the demographic composition of many Latin American countries, leaving little room for action to the hosting countries and governments to manage the help and support this population requires (R4V, 2022). As depicted in panel B of Figure 1b, these countries increased their percentage of immigrants up to ten-fold in less than five years, approaching 9% of the population in countries like Chile, where the migrant population was only 1% in 2000.

The Chilean case has the peculiarity that, unlike a large part of the migration phenomena analyzed to date, its migration is rooted in these two distinctive migration waves: i) a Venezuelan migration that shares culture, language, ethnicity, and religion with the

[1]based on WorlBank data estimations of a population of 30.529.716 Venezuelans in 2015 (https://data.worldbank.org/indicator/SP.POP.TOTL?name_desc=true&locations=VE)

[2]R4V is the Inter-Agency Coordination Platform for Refugees and Migrants from Venezuela jointly led by the UN Refugee Agency (UNHCR) and the International Organization for Migration (IOM)

[3]Chile was the leading destination of Haitians in Latin America because, unlike the rest of the countries of Latin America, Chilean immigration legislation allowed until 2018 the legal entry of Haitians for tourism without the requirements of a visa. Once in Chile, they could ask for work permits and residence.

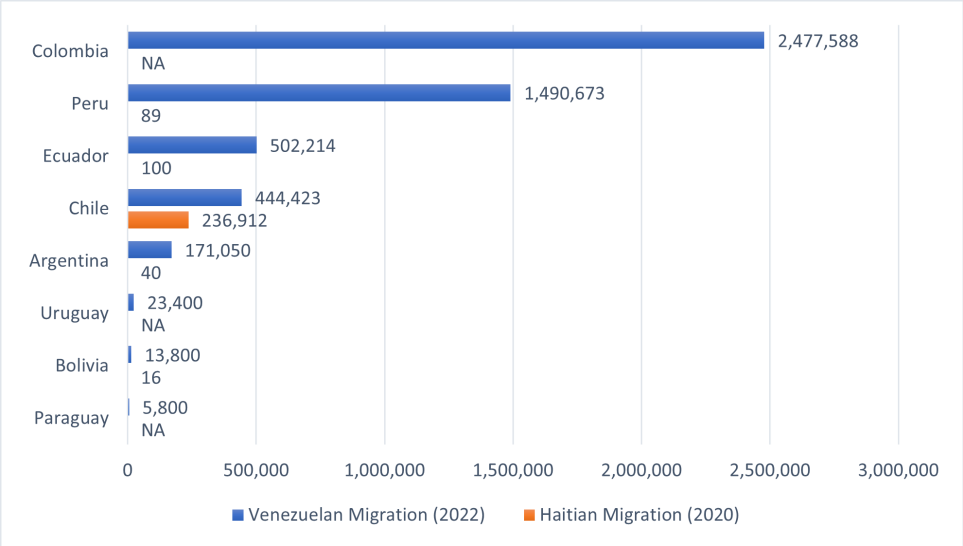
hosting population; and ii) a Haitian migration that does not. Therefore, the clear ethnic-cultural difference between these groups allows us to identify and understand the attitudinal drivers that could be feeding the rise of far-right parties. We rarely observe migrants looking alike to natives and, therefore, with less cultural prejudice than those reviewed by the literature in the USA or Europe. This paper takes advantage of this cultural similarity and separately identifies the cultural mechanisms that literature has not been able to identify. On top of that, Chile allows us to understand this phenomenon in the context of a growing far-right vote that mimics recent examples like the Brexit vote in the UK, the Trump vote in the US, and the Vox surge in Spain. Inexistent before 2015, the far-right vote for the *Partido Republicano de Chile (PR)* has been steadily gaining adepts and increasing voting shares, reaching a solidly 29% of votes in the first round of the presidential election in 2021.

The migration literature has sought to answer how migration fuels anti-immigrant sentiment and far-right vote for decades, particularly in high-migration host democracies, such as European countries and the United States. Four main hypotheses can explain the formation of anti-immigration attitudes in these countries. The first bases anti-migration sentiments and vote on the concerns that natives have of losing jobs to the newcomers, or seeing their salaries reduced due to an expansion in the labour supply (Dustmann & Preston, 2007; Hainmueller & Hiscox, 2010; Halla et al., 2017; Kessler, 2001; Malhotra, Margalit, & Mo, 2013; Mayda, 2006; Scheve & Slaughter, 2001). Many of these research works find that immigrant workers with a particular skill set will fuel rejection among natives with the same skill set. A second hypothesis, of sociotropic nature, theorizes that anti-immigration attitudes are driven by beliefs that migrants will become a burden for governments, forcing them to redirect resources or raise taxes, and affecting natives through the congestion of public goods and services (Barone et al., 2016; Citrin, Green, Muste, & Wong, 1997; Colantone & Stanig, 2018; Facchini & Mayda, 2009; Hainmueller & Hiscox, 2010; Pieroni, Roig, & Salmasi, 2023; Valentino et al., 2019). Others find that nationalism, ethnocentrism, and fears of a local culture reshaping are the main factors behind the rejection of foreigners (Chandler & Tsai, 2001; Dustmann & Preston, 2007; Fetzer, 2000; McLaren, 2003; Mendez & Cutillas, 2014; Sides & Citrin, 2007), particularly if they profess another religion such as Muslim (Adida, Lo, & Platas, 2019; Bansak et al., 2016; Valentino et al., 2019). Finally, direct interactions between natives and foreigners may weaken natives fears of immigration (Allport, 1954), as well as some specific migration causes, like wars or religious persecution, can decrease animosity towards migrants for humanitarian reasons (Alrababa'h et al., 2021; Bansak et al., 2016; Newman et al., 2015).

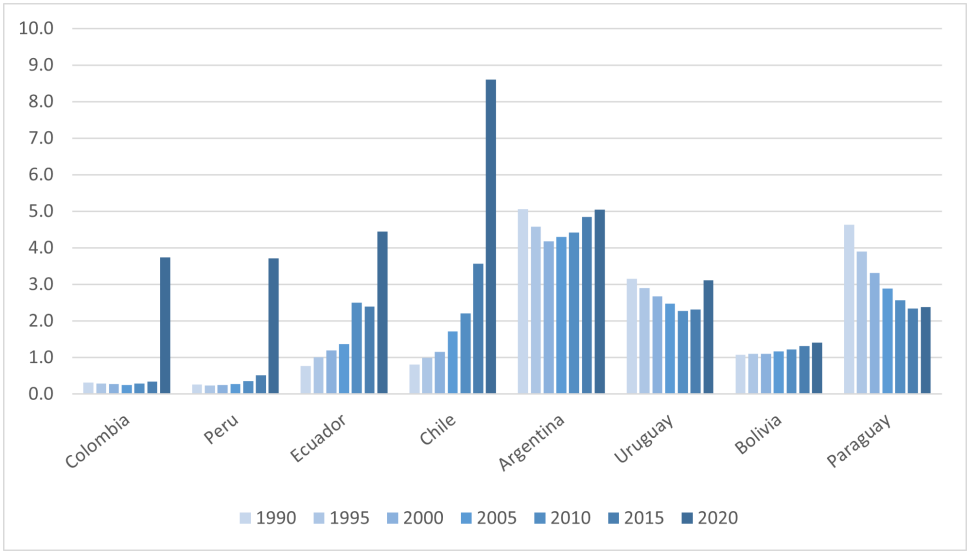
Using administrative data on migratory flows by nationality and electoral results at the municipality level, this research uses a shift-share IV model to estimate whether exposure to migrants at the local level increases electoral behaviour variables such as support for the far-right. Because migrants are not randomly allocated across municipalities, we cannot estimate a causal relationship between migration and voting without an instrumental variables strategy (or a natural experiment setup). Migrants are expected to settle in more welcoming locations, which, in turn, are more likely to vote less for anti-migration parties, possibly skewing the results. To solve this problem, I use an IV strategy devised by Altonji and Card (1991) and Card (2001), where migration is instrumentalized by a combination of past settlement patterns (share) and general changes in migration to Chile during the analysed period (shifts). Specifically, the original migration shares should be

**Figure 1: Distribution of Venezuelan and Haitian migrants by hosting country as of December 2022, and hosting countries migrant percent population from 1995 to 2020**

**(a) Distribution of Venezuelan and Haitian migrants by host country as of December 2022**




**(b) Trend of migrants as percent of population by country, 1990-2020**




*Source*: Information compiled by R4V: Inter-Agency Coordination Platform for Refugees and Migrants from Venezuela jointly led by the UN Refugee Agency (UNHCR) and the International Organization for Migration (IOM) for Panel A. United Nations Department of Economic and Social Affairs, Population Division (2020). International Migrant Stock 2020, for Panel B.

exogenous to the rise of far-right years later, and the shift component depends to the total change in migration during the period, exploiting exogenous events in origin countries that increase the migration outflows and not considering the specific pull factors of a particular municipality.

This study provides evidence supporting recent literature (Lonsky, 2021; Pagliacci & Bonacini, 2022) that challenges the notion of migration driving the rise of the far-right. Instead, our findings reveal that migration diminishes biases within local populations,

leading to decreased support for high anti-migrant rhetoric parties. Our instrumental variable (IV) model, focusing on cultural diversity, demonstrates that greater interaction with culturally different migrants reduces far-right backing by over 5% for every 1% increase in their ratio. At the same time, we found no effect for the culturally similar population. This aligns with Allport's contact theory (Allport, 1954), suggesting that engaging with diverse migrants curbs prejudice and anti-migrant party affiliation. Nevertheless, I am cautious in this assertion since we are not certain about the timing, intensity or quality of this contact, and we cannot ensure Allport's original conditions for "optimal contact" (equal status between groups in contact, common goals and intergroup cooperation, and institutional support). Interestingly, aggregated migration shows no such impact, suggesting other factors like security or economics drive far-right support. However, we do find increasing support for the centre-right coalition, which could be profiting from less harsh stances on migrants while implementing mild anti-migrant policies. Migration, particularly from culturally similar populations regarding skill levels, language, and culture, significantly increases support for the traditional right-wing parties.

These results shed light on the discussion on the mechanisms by which exposure to migration affects support for far-right parties. A better understanding of the mechanisms fueling anti-migrant sentiment is critical for policy implications, especially if governments aim to increase processes of integration and social cohesion, curbing political polarization. Our preliminary results recommend that countries with a high and diverse inflow of migrants, such as Italy, France, or Spain, to prioritize resources into economic factors that may be yielding anti-immigration attitudes and social conflict. Our empirical evidence that contact with migration reduces animosities, and that some other channels might be in play, gives political incentives to design contingency public policies for the arrival of migrants, like boosting public services in critical locations to serve the entire population in optimal conditions. Also, this research contributes to the migration literature by analyzing migration waves between countries from the global south, broadening the literature scope towards cases where migration's socioeconomic and political consequences could be more severe due to a lack of resources or weak institutions. The literature often disregards the fact that 85% of global migration is hosted by developing nations (United Nations High Commissioner for Refugees, 2022), forcing middle-income nations to issue thousands of work permits, offer essential health, education, and housing services, or provide essential arrival support for these migrants and their families, while lacking the sufficient administrative capacity, know-how, and public policy tools to solve the multiple problems accumulating. Documenting the mechanisms tensioning the daily coexistence within these countries is fundamental to ensure better social cohesion and stop the birth of nationalist and/or xenophobic movements that can further destabilize these regions.

# 2 Background

## 2.1 Attitudes towards immigration

The literature has distinguished a set of hypotheses behind anti-immigration attitudes in hosting countries to understand how migration generates attitudinal reactions in the local population and, consequently, in political decisions such as whether to vote or who to vote. According to Hainmueller and Hiscox (2010), we can distinguish between economic, and sociocultural or ethnocentric hypotheses. The first economic hypothesis is the fear

that natives have of losing jobs at the hands of the new migrant population. Based on the Heckscher-Ohlin international trade economic model, several studies analyze how anti-immigration sentiment stems from the fear that skilled and unskilled workers have that migrants will force their real income to drop due to increased labour market competition in the country (Dustmann & Preston, 2007; Hainmueller & Hiscox, 2010; Kessler, 2001; Mayda, 2006; Scheve & Slaughter, 2001). Under this theory, unskilled workers will be less supportive of migration if the migrants are also unskilled and will compete for the same jobs.

Despite how reasonable this proposal seems, empirically, it remains contested. Several works (Hainmueller & Hiscox, 2007, 2010) show evidence that, in general, there is a global preference for skilled workers, regardless of the native's skill set. For Malhotra et al. (2013), the importance is not the skilled-unskilled labour differences but the labour sector we observe in the hosting country. Evaluating the particular case of H1-B visa holders in the US, and being able to control for cultural and social factors of the migrants, they find that "when labor market threat is present, there is a significant association between labor market competition and views on immigration" (Malhotra et al., 2013, p.2). The competition feeling in the labour market does not depend particularly on the natives' skill levels, but on the relative job scarcity in one economic sector vs. the others. Dancygier and Donnelly (2013) analyze how the economic situation of the locals' labour sector shapes the migration attitudes in Europe, finding that natives employed in economic booming sectors (with job expansion) support migration more than individuals employed in depressed sectors, backing the theory that these attitudes are driven by aspects of one's own economic self-interest.

The second economic hypothesis is of a sociotropic nature, where immigration is associated with a greater burden for the State. The literature distinguishes three situations that form pro or anti-immigration sentiments among the native population: i) the tax burden or extra fiscal expenditure that a state must carry out in order to manage and administer the arrival of new migrants (Citrin et al., 1997; Facchini & Mayda, 2009), or concerns about the effect of migration on the general state of the economy (Valentino et al., 2019); ii) the congestion of public goods and services, such as health and education, that citizens experience when they have to share benefits with newcomers (Colantone & Stanig, 2018; Hainmueller & Hiscox, 2010); and iii) the association between crime and migration (based on prejudices), that plays a critical role in forming anti-immigrant attitudes (Ajzenman, Dominguez, & Undurraga, 2023).

The theory behind the first case is that high-SES[4] natives, or those who contribute proportionally more with taxes to the State or welfare system, will be reluctant to accept migration that increases fiscal spending or affects the economy as a whole (Facchini & Mayda, 2009; Hanson, Scheve, & Slaughter, 2007). This increase in spending may be because migrants are likely to increase social welfare costs or reallocate expenses to management services and aid for the newcomers[5]. Others could be motivated by the perception of congestion in public goods and services, especially among the low-SES native population. This could mean that in regions with more significant limitations to public goods provision. Finally, Fasani, Mastrobuoni, Owens, and Pinotti (2019, p.1)

[4]Socioeconomic status

[5]This is an important item in the local budget of some states of the United States, concerning the reception of illegal migrants from Central America (Hanson et al., 2007)

posit that in most countries, "natives are far more concerned that immigrants increase crime, rather than unemployment or taxes", founded on misperceptions about the level of crime (due to an overestimation also of the magnitude of the migrant population). In Europe, this reason is one of the main concerns of natives related to immigration (Bianchi, Buonanno, & Pinotti, 2012) and is one of the main reasons behind the formation of anti-immigration political parties (Dinas & van Spanje, 2011). This should not be an exception in Latin America, since criminal gangs and cartels manage a large part of illegal migration, exacerbating the relationship between migration and crime[6].

The third hypothesis is sociocultural, since it assumes that attitudes towards migrants arise from ethnic and cultural tensions between natives and migrants. In general, the theory behind this hypothesis is that personal values linked to group identities and prejudices "lead citizens to oppose immigration even if it does not represent an economic threat" (Valentino et al., 2019, p.5). An important branch of the literature has studied this phenomenon for several decades, mainly oriented to western industrialized economies (Chandler & Tsai, 2001; Citrin et al., 1997; Dustmann & Preston, 2007; Fetzer, 2000; McLaren, 2003; Sides & Citrin, 2007), and has found an "ample evidence that deeply rooted hostility exists towards immigration groups with largely different cultural and ethnic backgrounds" (Dustmann & Preston, 2007, p.2).

In particular, many of these works have focused on more tangible characteristics of migrants when analyzing anti and pro-immigrant sentiment. While some have focused on the physical characteristics as the main driver (Brader, Valentino, & Suhay, 2008; Lee & Ottati, 2002; Valentino et al., 2019); others focused on language or religion (Chandler & Tsai, 2001; Hainmueller & Hopkins, 2015), mainly on the animosity towards Muslim migration in Christian populated countries (Adida et al., 2019; Bansak et al., 2016; Valentino et al., 2019). Finally, a set of studies focus on the concern that migration could shape a new cultural identity that does not accommodate the native population (Campbell, Wong, & Citrin, 2006; Card, Dustmann, & Preston, 2012; Dustmann & Preston, 2007; Sniderman, Hagendoorn, & Prior, 2004).

Finally, the fourth and last hypothesis suggests a compensating mechanism for the anti-immigrant attitudes activated by the other three reviewed reasons. Based on Allport (1954) contact theory, this literature suggests that direct interactions between natives and foreigners may weaken natives' fears of immigration, resulting in a more respectful and accepting relationship. Also, other humanitarian reasons could enhance pro-immigrant attitudes when migration is due to armed conflict, violence, political persecution, or natural disasters. Natives tolerate these migrants (refugees) more because they are fleeing conflict, and migration is generally forced and involuntary. Consequently, their preference for migration increases (Alrababa'h et al., 2021; Bansak et al., 2016; Newman et al., 2015).

## 2.2 Immigration and far-right vote

A big part of the literature has more directly addressed the question of how migration influences citizens' electoral and political decisions (Barone et al., 2016; Caselli, Fracasso, & Traverso, 2021; Dinas, Matakos, Xefteris, & Hangartner, 2019; Edo, Giesing, Öztunc, & Poutvaara, 2019; Gerdes & Wadensjö, 2008; Mayda, Peri, & Steingress, 2022; Moriconi

[6] "Crisis in Venezuela: How Colombian mafias and armed groups are taking advantage of Venezuelan migrants" https://www.bbc.com/mundo/noticias-america-latina

et al., 2022; Otto & Steinhardt, 2014; Vasilakis, 2018, among others). These are mainly focused on how the increasing inflows of migration since the 1980s have fostered the rise of European far-right movements such as the *National Front* in France or the *Vlaams Belang* in Belgium (Edo et al., 2019), the recent electoral results such as Brexit (Langella & Manning, 2016) or the election of Trump in the United States (Mayda et al., 2022). The main idea is that the local presence of immigrants modifies attitudes towards migrants, and therefore, a part of the hosting population will change their electoral behaviour.

A first early strand in this literature reported how anti-immigrant sentiment is the single most important driver for the far-right vote in single-country and comparative studies (Billiet & De Witte, 1995; Cutts, Ford, & Goodwin, 2011; Kai, 2008; Mayer & Perrineau, 1992; Mughan & Paxton, 2006; Norris, 2005; Van der Brug & Fennema, 2003; Van der Brug, Fennema, & Tillie, 2000). Among these studies, Billiet and De Witte (1995) find that for the general elections in Belgium in 1991, voting for the *Vlaams Blok* was determined by a negative attitude towards immigrants, almost as a single-issue party. Van der Brug et al. (2000) confirm this result in a cross-sectional study of seven European political systems for the 1994 European Parliament elections.

Considering more directly the relationship between voting behaviour and migration, a group of studies analyzes the effect of the presence of migrants in a territory on the electoral behaviour of the native population. While one set of papers finds that the presence of larger migrant communities increases voting for right-wing movements (Golder, 2003; Knigge, 1998; Lubbers, Gijsberts, & Scheepers, 2002; Lubbers & Scheepers, 2002; Swank & Betz, 2003); others find that there is no such relationship (Arzheimer & Carter, 2006; Lucassen & Lubbers, 2012; Norris, 2005; Rydgren, 2008), or even that far-right support grows when minority groups are smaller (Bustikova, 2014). However, many of these papers do not necessarily take into account the sorting problem of migration, and therefore, their ability to identify causality is weak (Cools et al., 2021; Golder, 2016). For example, Golder (2003), in a cross-national study of 19 European countries and 165 national elections, finds that migration levels in countries with stressed labour markets increase the vote share of far-right and populist parties, but it does not consider that migrants are not randomly located in countries or across electoral districts. Instead, it is to be expected that the migrant population will choose cities and neighbourhoods with better living standards, job opportunities, and more acceptable treatment of migrants and, therefore, with different voting behaviours than constituencies most hostile to them.

A new wave of studies corrects this sorting problem using IV techniques proposed by pioneering works such as those by Altonji and Card (1991) and Card (2001), or quasi-experimental and matching strategies allowing causal inference (Dustmann, Vasiljeva, & Piil Damm, 2019; Schaub, Gereke, & Baldassarri, 2021). Within this literature, and for a long time, there has been a broad consensus that the presence of migrants at the local level has a significant and positive effect on the electoral success of far-right parties (Cools et al., 2021), until a recent surge of research started disputing these results (Gessler, Tóth, & Wachs, 2022; Hennig, 2021; Lonsky, 2021; Pagliacci & Bonacini, 2022; Russo, 2021; Schaub et al., 2021; Usta, 2022; Vertier et al., 2023). Notwithstanding, among those who find positive effects of migration on the far-right vote, there is heterogeneity in the effect magnitudes and the driving mechanism. First, one group of studies is not able to identify the mechanism through which migration increases the right-wing vote (Caselli et al., 2021; Dinas et al., 2019; Dustmann et al., 2019; Finseraas & Strøm,

2022; Gálvez-Iniesta & Groizard, 2021; Gerdes & Wadensjö, 2008; Kenny & Miller, 2022; Maggio, 2021; Mehic, 2022; Tomberg, Stegen, & Vance, 2021; Vasilakis, 2018) or suggests that the mechanisms encompass ethnic, cultural, compositional amenities and economic factors (Barone et al., 2016; Edo et al., 2019; Moriconi et al., 2022; Otto & Steinhardt, 2014). Barone et al. (2016) use Italian municipality-level data to causally estimate the effect of a larger immigrant share on center-right votes (*Lega Nord*). They find positive and significant results motivated by cultural diversity and competition in the labour market and for public services. However, they address cultural diversity only based on religious diversity, not considering ethnicity or language. Edo et al. (2019) examine the relationship between migration shares and pro-migrant party support on the French left. The authors find that the arrival of low-skilled migrants drives support for the *National Front* and reduces support on the left, suggesting that low-educated voters are worried about labour market competition. At the same time, the effect is higher considering low-skilled migrants from non-western countries, highlighting the importance of immigrants' cultural backgrounds.

Meanwhile, a large part of these studies points only to economic factors behind the increase in far-right voting, such as competition in labour markets or adverse effects on the welfare state (Bredtmann, 2022; Halla et al., 2017; Mayda et al., 2022; Pieroni et al., 2023; Roupakias & Chletsos, 2020). For the Austrian case, Halla et al. (2017) find that the arrival of migrants in communities explains roughly a tenth of the variation in votes for the *Freedom Party of Austria (FPO)* and suggest that it is due to voters worrying about adverse labour market effects of immigration. Distinguishing itself from the overwhelming evidence for Europe over other countries, Mayda et al. (2022) find for the United States that only low-skilled migration increases the share of the *Republican Party* vote, while high-skilled immigrants decrease the share. This suggests that the main concerns of US voters would be economic factors such as job competition and the financial burden on the state.

A third group of studies suggests that the primary mechanism is ethnic and cultural (Brunner & Kuhn, 2018; Devillanova, 2021; Harmon, 2018; Mendez & Cutillas, 2014; Rozo & Vargas, 2021; Sekeris & Vasilakis, 2016). In order to identify how ethnic factors play a role in driving the far-right vote, Mendez and Cutillas (2014) separately analyze the electoral effects of the Latin American and African migrant population to Spain, noting that migration with similar cultural characteristics increases support for the leftist conglomerate, while the North African to the *Popular Party* (mainstream right party). Similarly, for Switzerland, Brunner and Kuhn (2018) separate the migrants according to an index of ethnic-cultural similarity with the local population, and find it is not so much the overall immigrant share but mainly the presence of immigrants with a different cultural background that affects the voting behaviour of Swiss citizens.

Among the recent studies that find no effect of migration (Gessler et al., 2022; Hennig, 2021; Russo, 2021) or find a negative effect of migration on the far-right vote (Lonsky, 2021; Pagliacci & Bonacini, 2022; Usta, 2022; Vertier et al., 2023), the primary mechanism behind these results seems to be Allport's (Allport, 1954) contact hypothesis. For example, in opposition to Edo et al. (2019), Vertier et al. (2023) find that the migrants' inflow size has a negative effect on the electoral results of the *Front National* for the 2017 parliamentary elections in France. Based on a natural experiment using the dismantling of the Calais "Jungle," an encampment in the North of France, the authors exploit the

fact that between October 2015 and 2016, the government relocated the migrants in other areas of the country in 300 centres. According to the authors, the result differences with Edo et al. (2019) may be because they are considering the effect of relocating small amounts of asylum-seeking population into small communities that had not previously experienced high levels of migration. The analysis shows that municipalities with smaller reception centres drive the main results, which is consistent with the contact hypothesis.

A particular aspect of the studies that support the hypothesis that migration decreases the anti-immigrant vote through the contact mechanism is that, generally, they are based on refugees. As Vertier et al. (2023), Usta (2022) finds a negative effect of Syrian refugee exposure on the vote for the conservative *Justice and Development Party* in Turkey, while Gessler et al. (2022) suggest that there was no significant overall effect of refugee camps on votes of the right-wing as a whole in Hungary, but a redistribution of votes within the right-wing parties. Only two studies found a drop in the far-right vote using general migration and shift-share instrumental variables in Finland and Italy (Lonsky, 2021; Pagliacci & Bonacini, 2022). Pagliacci and Bonacini (2022) analyzes the effect of migrant settlement at the municipal level on the *Lega* vote for the 2019 European elections, finding that *Lega* reduces 0.67% to 0.56% of their vote share per 1% increase in the migrant population. Lonsky (2021) studies the effect of migration on the results of the *Finns Party* on multiple local and general elections, finding that even though migration reduces the average far-right support, the negative effect of immigration is only present in places with large initial exposure to immigrants.

As we can see from all these studies, it is difficult to identify how exposure to migration increases support for far-right parties or decreases the electoral success of pro-immigrant parties. In general, all these studies are focused on industrialized economies such as the United States and Europe, where migration has been a long-standing process and where the cultural and economic effects of pro and anti-immigration attitudes are generally confused. Few exceptions look into the effects in attitudes, turnout, and electoral outcomes of migration waves to less developed economies or with no industrial activity[7] (Adida, 2011; Alrababa'h et al., 2021; Rozo & Vargas, 2021), but no research has investigated far-right support outcomes or mechanisms for this part of the world. The particularity of the recent migratory waves within Latin America, together with the resounding rise of the far-right in just a few years, makes this a propitious scenario to identify ethnic-cultural and economic mechanisms.

## 2.3 Research Questions

Considering the particular case of Chile, the primary goal of this study is to determine if local exposure to migrants is increasing far-right support. While it is logical to believe that the rise of the far-right is related to migration, given that it is a central part of their political discourse, it is not yet clear if exposure to migrants is driving their electoral success. Specifically, we aim to investigate if an increase in the share of migrants in a given territory fuels the anti-immigrant vote. If this is the case, the secondary goal is to understand the role of migrants' ethnicity and culture in this process. We seek to uncover the mechanisms by which exposure to migration influences electoral support for far-right parties in the Chilean context. Is there a cultural (nativism) explanation, or are there other economic or sociotropic reasons, such as fear of job loss? Conversely, we would like

[7] Particularly intensive in low-skilled jobs

to understand the effect of migration on the rest of the political spectrum, particularly on the traditional center-right, which competes electorally with the far-right and therefore adopts some anti-migration stances, and on the center-left and left-wing parties, which mostly advocate a pro-immigration discourse.

Given the significant increase in migration in Chile between 2015 and 2021, and the expectation that migration will mirror the recent rise in electoral success of the far-right in Europe and the US, we will start by testing the following hypothesis using a null significance hypothesis testing framework:

- **Hypothesis 1:** Does increasing the share of migrants in a municipality fuel anti-immigrant vote (particularly far-right vote) in a context different from Europe and the US?

If Hypothesis 1 is supported, and since the chosen context allows for a precise analysis of the attitudinal consequences of culturally similar versus culturally diverse migration, we will also test the following hypotheses:

- **Hypothesis 2: Cultural Differences and Far-Right Voting:** An increase in culturally different immigration (e.g., Haitian) increases support for far-right candidates.
- **Hypothesis 3: Cultural Similarity and Far-Right Voting:** An increase in culturally similar immigration (e.g., Venezuelan) does not increase support for far-right candidates.

# 3 Data and Data description

This empirical research measures the effects of migration at the Chilean municipal level (346 municipalities) on the presidential election results of 2017 and 2021. It relies mainly on three data sources: official immigration data from the Chilean Department of State (National Migration Service), the national 2002 Census from the National Statistics Institute (INE), and the electoral results dataset from the Chilean Electoral Service (SERVEL).

**Migration to Chile** To measure the magnitude of migration at the municipal level, I use individual-level data on 1,826,719[8] visa applications to the Chilean Department of State from 2000 to 2021. This database includes basic official and self-reported demographic statistics such as date of birth, nationality, municipality of residence, gender, education, labor market occupation, and date of application. To account for the initial migration shares in 2000, I use the 2002 Population and Housing Census prepared by INE as a baseline. Conducted in April 2002, the census recorded 187,521 migrants among the 15.1 million Chileans, incorporating variables such as education, age, and year of arrival. Figure 2a shows the stock of migrants for the two main migratory waves of the last 15 years. Following the 2017 election, immigration restrictions were established for the Haitian population, stopping their entry into the country. To compute the shares

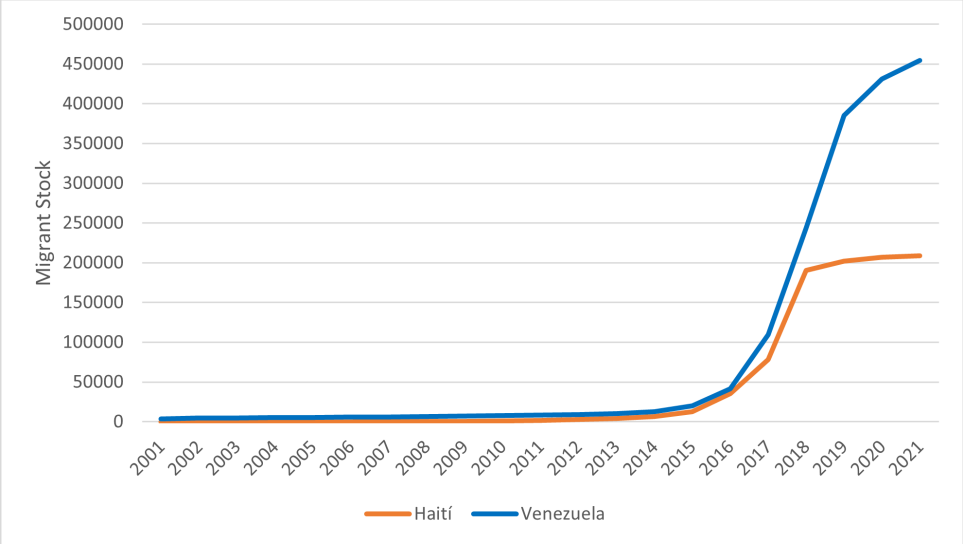
[8]The original dataset contains 2,674,391 visa applications. I removed duplicates based on nationality, date of birth, gender, education, occupation, and professional activity. Also, the dataset contains 749,729 permanent residence permits, granted once a foreigner stays at least 12 months with a temporary visa.

of migrant populations, I use yearly municipal population estimates reported by INE for the years 2000-2035, based on projections from census data.

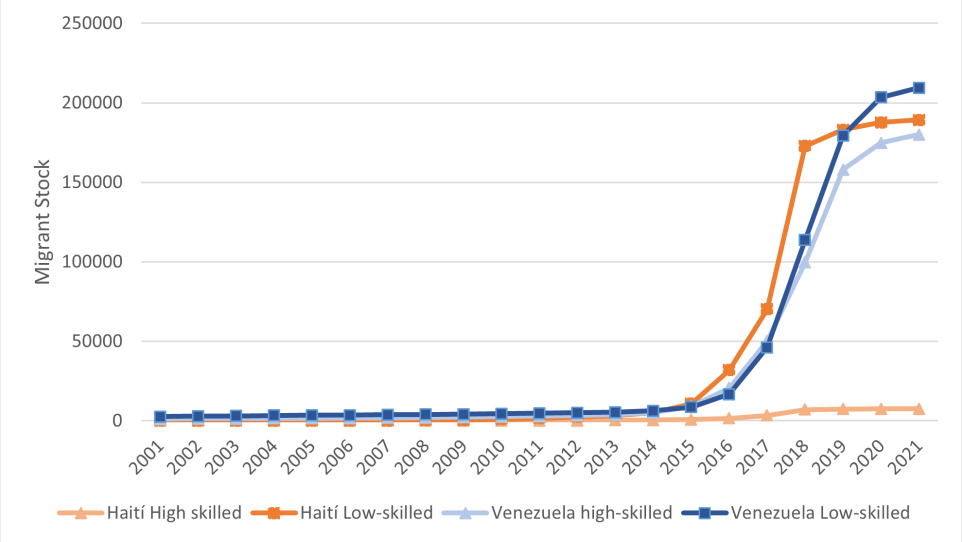
**Skill Level** From the migration database and the 2002 census, I obtained the education levels of each migrant residing in Chile at the time of arrival. Migrants are classified as high-skilled if they have some type of higher education and as low-skilled if they only have primary or secondary education. Given that the migration information collected by the State Department is self-reported, there are 682,067 migrants without educational information. To impute their education, I use the multiple imputation methodology based on Rubin (1987, 2018), where missing data is imputed using other variables such as gender, activity, profession, age, country of origin, year of migration, and place of migration. Figure 2b shows the stock of migrants by skill level and nationality.

**Figure 2: Stock of Migrants by Nationality (2001 - 2021)**

**(a) Venezuelan and Haitian Migrant Stock**




**(b) Migration Stock by skill levels**




*Source*: Author´s calculation using the visa and permanent resident requirement dataset from the Chilean Department of State (National Migration Service) 2000 - 2021.

**Party Ideology** To classify the ideology of Chilean political parties, especially regarding their anti-migrant rhetoric, I use information collected by the Manifesto Project for the 2017 and 2021 Chilean elections. The Manifesto Project analyzes parties' election manifestos, capturing specific policy-related rhetoric and party positions on a left-right index. As shown in Table A.1 in Appendix A, the far-right Republican Party is the only party that had anti-immigration rhetoric in both the 2017 and 2021 elections, with a score of 1.19. In 2021, the center-right also expressed anti-immigration rhetoric with a score of 0.09, while the left-wing parties expressed only positive rhetoric on migrants. For the analysis, we consider the effect of migration on the two right-wing coalitions, as well as on the aggregated right-wing and left-wing parties.

**Election Results** To construct the dependent variables for party voting and turnout, I use the electoral results from the first round of the presidential elections in Chile in 2017 and 2021. These elections marked the first occurrence after the right-wing parties split into a more liberal center-right faction, led in 2017 by presidential candidate Sebastián Piñera, and a far-right coalition, more conservative and closer to the former dictatorial regime of Augusto Pinochet, led by José Antonio Kast[9]. The electoral results for turnout and votes for party coalitions at the municipal level are published by SERVEL for the two aforementioned elections and are presented in Table 1.

**Table 1: Descriptive Statistics by Municipality**

| Variable | 2017 | | | | 2021 | | | |
|---|---|---|---|---|---|---|---|---|
| | Mean | SD | Min | Max | Mean | SD | Min | Max |
| **Outcome** | | | | | | | | |
| Far-right vote (%) | 7.36 | 3.47 | 1.19 | 30.56 | 30.40 | 9.90 | 11.57 | 72.34 |
| Center-right vote (%) | 39.38 | 8.31 | 23.05 | 74.10 | 11.31 | 4.13 | 1.10 | 34.35 |
| Right-wing vote (%) | 46.73 | 8.69 | 26.94 | 81.11 | 41.71 | 11.08 | 15.71 | 85.45 |
| Left-wing vote (%) | 52.30 | 8.68 | 28.70 | 71.97 | 42.23 | 9.32 | 4.42 | 64.04 |
| Turnout (%) | 44.88 | 7.19 | 11.31 | 69.09 | 44.21 | 6.76 | 19.07 | 69.27 |
| **Migration** | | | | | | | | |
| Migration (%) | 3.67 | 6.48 | 0.06 | 61.92 | 5.78 | 8.05 | 0.10 | 62.54 |
| Venezuelan Migration (%) | 0.19 | 0.65 | 0.00 | 8.61 | 0.93 | 1.86 | 0.00 | 20.70 |
| Haitian Migration (%) | 0.24 | 0.48 | 0.00 | 4.41 | 0.77 | 1.00 | 0.00 | 5.95 |
| **Low-skilled Migration** | | | | | | | | |
| Low-skilled Migration (%) | 2.69 | 5.09 | 0.00 | 56.94 | 4.14 | 6.04 | 0.00 | 57.24 |
| Low-skilled Venezuelan Migration (%) | 0.08 | 0.28 | 0.00 | 3.55 | 0.44 | 0.86 | 0.00 | 9.19 |
| Low-skilled Haitian Migration (%) | 0.22 | 0.42 | 0.00 | 3.88 | 0.70 | 0.91 | 0.00 | 5.06 |
| **Controls and Municipal Characteristics** | | | | | | | | |
| Population | 53,235 | 83,186 | 123 | 604,744 | 56,874 | 90,193 | 141 | 655,033 |
| Aging rate | 92.61 | 39.36 | 4.10 | 360.00 | 108.77 | 35.47 | 24.67 | 343.75 |
| Women (%) | 50.07 | 1.77 | 46.36 | 51.89 | 50.07 | 1.81 | 34.75 | 52.23 |

*Source*: Electoral results dataset for the presidential elections of 2017 and 2021 from the Chilean Electoral Service (SERVEL), visa applications dataset from the Chilean Department of State (National Migration Service), Population projections 2000-2035 from the National Statistics Institute (INE). Ageing rate corresponds to the proportion of 60 years old or older, over the young 14 years old or younger in a given municipality. Author's calculation by municipality (346).

[9]The Chilean political system is a presidential system, so the leaders of the coalitions are not necessarily the party presidents, but rather the presidential candidates.

**Municipal Characteristics** To measure heterogeneous effects, perform robustness exercises, and control for relevant variables at the municipal level (with expected effects on electoral outcomes), I consider a set of municipal variables from administrative data sources in Chile. INE provides estimates on demographics at the municipal level, mainly variables related to the age distribution and proportion of women. Among the variables used, I build an aging rate variable, which accounts for the proportion of elderly population in relation to the young population in each municipality. Descriptive statistics of these variables can be seen in Table 1.

# 4 Methods

To understand how exposure to migration affects electoral outcomes, I will implement two identification strategies. First, as a benchmark strategy, I will estimate a linear regression model that allows us to measure the direct effect of migration at the municipal level using time and municipality fixed effects. Second, I will use a Two-Stage Least Square (2SLS) model to ensure a causal relationship using an instrumental variable based on Altonji and Card (1991) and Card (2001).

## 4.1 Benchmark linear model

For the first model, I examine the relationship between direct exposure to migrants and the electoral outcomes of interest at the municipal level for the 2017 and 2021 presidential elections. Specifically, I estimate a linear regression model of the type:

$$y_{mt} = \beta migr_{mt} + \rho_m + \rho_t + \gamma X_{mt} + \epsilon_{mt} \tag{1}$$

Where $y_{mt}$ is the outcome of interest for this research (e.g. far-right vote share) in municipality $m$ for election $t$; $migr_{mt}$ is the stock of migrants over the population in municipality $m$ for election $t$; $\rho_m$ and $\rho_t$ are fixed effects of the municipality and election year; $X_{mt}$ is a set of control variables at the community level, and $\epsilon_{mt}$ is the error term. In this model, our parameter of interest is $\beta$, representing the electoral effect of the percentage increase in migrants in a municipality-specific year.

The set of control variables $X_{mt}$ includes important economic and demographic variables in determining electoral results. These include the age composition of the municipality (ageing ratio), and the share of women.

## 4.2 Causality and IV model

For our benchmark model to measure a causal effect, we should assume no endogeneity in the distribution of migrants across municipalities. Specifically, that there is no correlation between migration and the error term and, therefore, no omitted variables affecting the settlement of migrants in some localities and, at the same time, the electoral results. Immigrants may decide their place of residence based on unobservable characteristics that correlate with voting for parties or coalitions. In particular, they may be attracted to neighbourhoods where migration is welcome, biasing our estimate downward. In turn, pull factors such as a buoyant labour market or absence of unemployment can also be

attractive to migrants, making it difficult to estimate the causal mechanism driving anti-immigrant voting.

To solve this estimation problem, I implement an instrumental variables strategy based on the work of Altonji and Card (1991) and Card (2001). The strategy corresponds to building a shift-share instrument that exploits the supply-push migration component by nationality as a plausibly exogenous variation of the shifts in the immigrant population across municipalities and interact them with the share of immigrants settled in each municipality in the initial period of analysis. The predictive power of the instrument is due to the fact that we can assume that the new migrants of any nationality will settle in the same geographical areas as their predecessors. The shift component exploits the exogenous reasons behind a migratory outflow from the countries of origin, which should be independent of the characteristics across municipalities within the hosting country.

To build this shift-share instrument, we take within-municipality differences of Equation 1 and decompose $\Delta migr_{mt} = migr_{mt} - migr_{mt-1}$ as

$$\Delta migr_{mt} = \sum_{n} \theta^{n}_{m,t-1} \cdot \Delta migr^{n}_{mt} \tag{2}$$

Where $\Delta migr^{n}_{mt}$ is the change of the stock share of immigrants of nationality $n$ in municipality $m$ between the period $t-1$ and $t$, and $\theta^{n}_{m,t-1}$ is the share of immigrant of nationality $n$ in municipality $m$ over total immigrants of nationality $n$ at $t-1$. Specifically:

$$\theta^{n}_{m,t-1} = \frac{MIGR^{n}_{m,t-1}}{\sum_{m} MIGR^{n}_{m,t-1}}$$

Specifically, the Equation 2 decomposition corresponds to the weighted sum of the changes in the share of immigrants of each nationality into destination municipalities $m$. These depend both on supply-push factors in the origin country that affect all municipalities and on demand pull-factors in each municipality that affect all nationalities. To exclude this last effect, the instrument substitutes local migration shift $\Delta migr^{n}_{mt}$ with the change of immigrants of nationality $n$ to the whole country: $\Delta migr^{n}_{t}$ Therefore, the final instrument variable is:

$$\Delta \widehat{migr}_{mt} = \sum_{n} \theta^{n}_{m,t-1} \cdot \Delta migr^{n}_{t} \tag{3}$$

Since we consider Chilean demand pull-factors as a whole, and plausibly exogenous to variation in political outcomes in municipality $m$, the correlation between $\Delta \widehat{migr}_{mt}$ and $\Delta migr_{mt}$ must be solely to supply push-factors in origin countries. This satisfies the exclusion restriction.

For the analysis of electoral outcomes, I use 2013 as the initial period $t-1$ of analysis since it corresponds to an entire electoral cycle prior to the first electoral appearance of a far-right candidate in the 2017 presidential elections, and some years before the

migratory waves from Venezuela and Haiti have started. For the differences in votes and migrant shares, I take the changes between the last two electoral processes of 2021 and 2017. Finally, I consider in the construction of this instrumental variable the 22 Latin American countries and countries with the largest migrant population in Chile in 2021, corresponding to Argentina, Bolivia, Brazil, China, Colombia, Costa Rica, Cuba, Dominican Republic, Ecuador, El Salvador, Guatemala, Honduras, Spain, United States of America, Haiti, Mexico, Nicaragua, Panama, Paraguay, Peru, Uruguay, and Venezuela. Although this is a subgroup of the 164 countries with some share of migrants in Chile, together, they reach 95.30% of the 2021 migrant population in Chile.

Finally, to analyze the effects on voting yielded by the different cultural traits of migration, and similar to what Moriconi et al. (2022) and Mayda et al. (2022), I decompose Equation 3 into two different instrumental variables considering separately the two main migration waves since 2015. To capture the effect produced by a culturally similar migration, I consider only the migration from Venezuela, that is, from a Spanish-speaking population, demographically and ethnically similar to the native population. In the case of culturally different migration, I consider migration from Haiti. In short, I build two instrumental variables of the type:

$$\begin{aligned} \Delta\widehat{migr}^{Ven}_{m(2021-2017)} &= \theta^{Ven}_{m,2013} \cdot \Delta migr^{Ven}_{(2021-2017)} \\ \Delta\widehat{migr}^{Hai}_{m(2021-2017)} &= \theta^{Hai}_{m,2013} \cdot \Delta migr^{Hai}_{(2021-2017)} \end{aligned}$$

As a robustness check, I test other instruments similar to those introduced by Bianchi et al. (2012) and Barone et al. (2016) for Italy. Both instruments yield results consistent with our proposed main instrument. In the case of Bianchi et al. (2012), due to the need for migration information to countries other than Chile (to ensure the exclusion restriction that migration does not depend on Chilean migration pull effects), we are unable to perform the exercise of including only low-skilled migration shares. The explanation of these instruments and their results can be found in Appendix C.

### 4.3 Validity of the IV Shift-Share Model

I intend to analyze the validity of my shift-share strategy based on the recommendations of Goldsmith-Pinkham, Sorkin, and Swift (2020).

## 5 Findings

### 5.1 Benchmark Findings

Before addressing the causal strategy with the instrumental variable, we first analyze the results of our benchmark model of migration and electoral effects of Equation 1. Table 2 shows the OLS regression estimates for our two-way fixed effects model for the 346 municipalities in Chile and our electoral outcomes of interest.

Table 2 presents the initial model without controls (Model 1 for every party family), revealing that a higher percentage of migrants at the municipal level significantly decreases electoral support for the far-right. Specifically, a 1% increase in the migrant ratio leads to

**Table 2: Two-way fixed effects model: Vote share & General Migration**

| Dependent Variable % Vote | Far-right | | Center-right | | Right-wing parties | | Left-wing parties | |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Immigration ratio | -1.021*** | -0.754*** | 0.549* | 0.278 | -0.472*** | -0.477*** | -0.401** | -0.409** |
| | (0.254) | (0.274) | (0.302) | (0.321) | (0.170) | (0.171) | (0.190) | (0.187) |
| Demographic Controls | | Yes | | Yes | | Yes | | Yes |
| Municipal FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Year FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Mean vote | 18.8 | | 25.34 | | 44.22 | | 47.27 | |
| | (13.71) | | (15.5) | | (10.26) | | (10.31) | |
| Mean immigration ratio | 0.047 | | 0.047 | | 0.047 | | 0.047 | |
| | (0.0738) | | (0.0738) | | (0.0738) | | (0.0738) | |
| Observations | 692 | 692 | 692 | 692 | 692 | 692 | 692 | 692 |
| R-squared | 0.888 | 0.894 | 0.934 | 0.938 | 0.416 | 0.463 | 0.747 | 0.766 |
| Number of id | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 |

*Notes*: Results of a 2WFE model regression at the respondent for 346 municipalities (Equation 1). The definition of variables are in the data and methodological sections. The dependent variable is the average vote in a given municipality in 2017 and 2021. Regressions include demographic controls such as the ageing rate, and the ratio of women in each municipality; and year and municipality fixed effects. Robust standard errors clustered at the municipality level in parentheses.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

a notable reduction in far-right votes. However, this effect diminishes when other demographic controls, such as the percentage of women and the ageing ratio in the municipality, are included in the model. Conversely, the impact of migration on the centre-right coalition, while in the opposite direction, is not statistically significant. Overall, a 1% increase in migration negatively affects the vote share of right-wing parties, primarily driven by a decline in far-right support. Similarly, among left-wing parties, migration reduces electoral outcomes, suggesting that neighbourhoods receiving more migrants might retaliate against mainstream parties by supporting independent or minor party candidates.

To understand how the heterogeneous characteristics of migration affect electoral outcomes in Chile, I modify Equation 1 to capture differences between culturally similar and dissimilar migrations. Specifically, I differentiate between changes in the stock of ethnically diverse migrants (Haitian migrants) and those ethnically and culturally similar to the native population (Venezuelan migrants).

Table 3 displays the results considering these heterogeneous effects and two types of migratory classifications. Models (1, 3, 5, 7) include only different migration shares as regressors with municipality and year fixed effects, while Models (2, 4, 6, 8) also incorporate control variables. Except for the far-right, an increase in culturally similar migration boosts electoral support for both left-wing and right-wing parties by approximately the same proportion. Among right-wing parties, this electoral boost is primarily received by the centre-right, with its vote share increasing by more than 1% for each percentage increase in migration. In contrast, the presence of culturally different migration produces more dissimilar results. The far-right's vote share decreases by over 2% with a 1 percentage point increase in Haitian migration, a vote likely absorbed by the more traditional centre-right parties.

The negative effect of migration on the votes for the most vociferous anti-migration party, particularly when the migration is culturally different from the natives, supports Allport's contact theory (Allport, 1954). As natives share neighbourhoods and life experiences with

**Table 3: Two-way fixed effects model: Vote share & Heterogeneous Migration**

| Dependent Variable % Vote | Far-right | | Center-right | | Right-wing parties | | Left-wing parties | |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Venezuelan Immigration ratio | -1.053*** | -0.847** | 1.499*** | 1.199*** | 0.446*** | 0.352** | 0.508*** | 0.490*** |
| | (0.297) | (0.328) | (0.298) | (0.324) | (0.124) | (0.175) | (0.113) | (0.186) |
| Haitian Immigration ratio | -2.103*** | -2.089*** | 2.338*** | 2.112*** | 0.235 | 0.0232 | 1.082*** | 0.950** |
| | (0.718) | (0.694) | (0.599) | (0.597) | (0.386) | (0.382) | (0.411) | (0.417) |
| Demographic Controls | | Yes | | Yes | | Yes | | Yes |
| Municipal FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Year FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Mean vote | 18.8 | | 25.34 | | 44.22 | | 47.27 | |
| | (13.71) | | (15.5) | | (10.26) | | (10.31) | |
| Mean immigration ratio | 0.047 | | 0.047 | | 0.047 | | 0.047 | |
| | (0.0738) | | (0.0738) | | (0.0738) | | (0.0738) | |
| Observations | 692 | 692 | 692 | 692 | 692 | 692 | 692 | 692 |
| R-squared | 0.886 | 0.895 | 0.941 | 0.942 | 0.403 | 0.449 | 0.749 | 0.766 |
| Number of id | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 |

*Notes*: Results of a 2WFE model regression at the respondent for 346 municipalities (Equation 1). The definitions of variables are in the data and methodological sections. The dependent variable is the average vote in a given municipality in 2017 and 2021. Regressions include demographic controls such as the ageing rate, and the ratio of women in each municipality; and year and municipality fixed effects. Robust standard errors clustered at the municipality level in parentheses.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

a larger immigrant population, their prejudices diminish, leading to reduced support for politicians with anti-immigrant rhetoric. When the migration is even more noticeable (culturally different), the benefits of integration increase. However, this benchmarking exercise does not account for potential migration sorting issues that could influence these results. New migrant populations might be settling in communities more welcoming to migration, thus less likely to vote for far-right parties or more inclined to support left-wing parties with pro-migration stances. These factors could bias our results, so I derive an instrumental variables model to address identification and causality problems.

## 5.2 IV Estimation Findings

Table 4 presents the results of our instrumental variables (IV) model using the shift-share method, which addresses the migration sorting issue inherent in the two-way fixed effects model from the previous section. When considering the overall effect of migration on electoral outcomes, the IV model does not replicate the simple OLS model's results, particularly for right-wing and left-wing party aggregations. This discrepancy suggests that migrant sorting may be influencing the observed negative impact of immigration on voting patterns: migrants tend to settle in more welcoming areas.

Being the effects of migration on far-right the literature (and ours) main outcome of interest, the IV model's findings indicate that increased migration at the municipal level does not generally explain the rise of the far-right but rather supports the center-right[10]. A 1% increase in migration correlates with a 0.9% increase in the centre-right vote, which is substantial given that this coalition's average national vote share was 11.3%. Additionally, migration also boosts the vote share for left-wing parties by nearly 0.5%.

[10]I found similar results when considering other model specifications and instruments, particularly when using as a base year for the instrument the year 2009. Please review these results in Appendix C and D.

**Table 4: 2SLS Model 2017-2021: Vote share & General Migration changes**

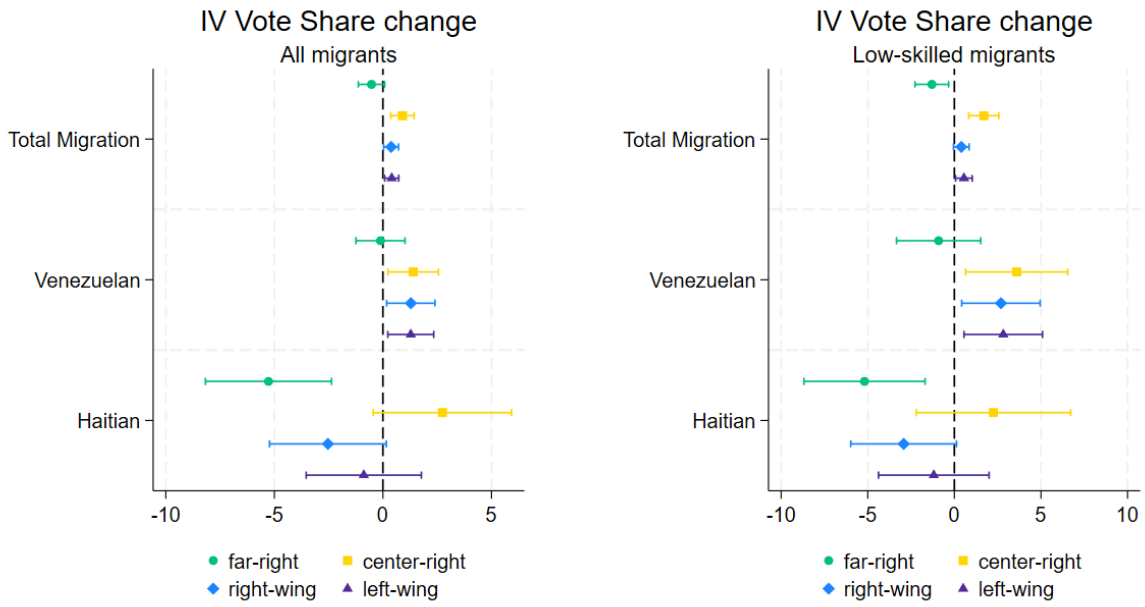
| Dependent Variable % vote | Far-right | | | Center-right | | | Right-wing parties | | | Left-wing parties | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | **OLS** | **IV** | **IV - control** | **OLS** | **IV** | **IV - control** | **OLS** | **IV** | **IV - control** | **OLS** | **IV** | **IV - control** |
| Immigration ratio change | -0.611*** | -0.525 | -0.410 | 0.348** | 0.901*** | 0.813*** | -0.264** | 0.376* | 0.403** | -0.220 | 0.406** | 0.420*** |
| | (0.149) | (0.368) | (0.305) | (0.156) | (0.326) | (0.287) | (0.124) | (0.207) | (0.168) | (0.138) | (0.197) | (0.155) |
| Controls | | | Yes | | | Yes | | | Yes | | | Yes |
| Observations | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 |
| Mean vote 2021 | 30.4 | 30.4 | 30.4 | 11.31 | 11.31 | 11.31 | 41.71 | 41.71 | 41.71 | 42.23 | 42.23 | 42.23 |
| | (9.90) | (9.90) | (9.90) | (4.13) | (4.13) | (4.13) | (11.08) | (11.08) | (11.08) | (9.32) | (9.32) | (9.32) |
| **First Stage Regressions** | | | | | | | | | | | | |
| Immigration ratio change (inst) | | 1.83e-06*** | 1.85e-06*** | | 1.83e-06*** | 1.85e-06*** | | 1.83e-06*** | 1.85e-06*** | | 1.83e-06*** | 1.85e-06*** |
| | | (5.15e-07) | (4.95e-07) | | (5.15e-07) | (4.95e-07) | | (5.15e-07) | (4.95e-07) | | (5.15e-07) | (4.95e-07) |
| F-stat | | 12.64 | 14.04 | | 12.64 | 14.04 | | 12.64 | 14.04 | | 12.64 | 14.04 |
| Part. R2 | | 0.177 | 0.190 | | 0.177 | 0.190 | | 0.177 | 0.190 | | 0.177 | 0.190 |

*Notes*: Results of OLS and IV estimates on the differences between 2017 and 2021 across 346 municipalities. The definition of variables are in the data and methodological sections. The dependent variable is the difference of the average political party vote in a given municipality between 2017 and 2021. The variables Immigration ratios are the change of immigrants divided by the municipality 2021 population; the instruments are built according to Equation 3 in the methods section. For the models with controls, I include the ageing rate and the ratio of women in each municipality during 2021. 2SLS coefficients are reported in the top panel under the heading IV. The bottom panel reports first-stage regressions with the Kleibergen-Paap Wald F statistic. Robust standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

A direct effect on the far-right vote emerges when considering migration types based on cultural traits. Figure 3a decomposes the instrumental variable to assess the electoral impact of Haitian and Venezuelan migration separately[11]. These results confirm those from our benchmark model. Increased contact with culturally different migrants (Haitians) significantly reduces far-right support. Specifically, a 1% rise in the Haitian migrant ratio decreases far-right support by more than 5%[12], with no significant impact on other parties' vote shares. Conversely, Venezuelan migration (culturally similar) positively affects all political coalitions except the far-right. For centre-right parties, a 1% increase in the Venezuelan population correlates with a more than 1.3% rise in votes. This suggests that moderate anti-immigration rhetoric adopted by the centre-right successfully attracts voters at the far-right's expense[13]. Pro-immigration left and centre-left parties also see electoral gains, albeit smaller than those of the centre-right.

**Figure 3: 2SLS Model 2017-2021: Vote share & Heterogeneous Migration changes**

**(a) All types of migrants** **(b) Low-skilled migrants**




*Notes*: Results of IV estimates on the differences between 2017 and 2021 across 346 municipalities. The definition of variables is in the data and methodological sections. The dependent variable is the difference of the average party vote in a given municipality between 2017 and 2021. The instruments are built according to Equation 3 in the methods section. The regression table is on Appendix B.

## 5.3 Low-skilled migration

Figure 2 clearly distinguishes the different skill levels of Haitian and Venezuelan migrants. While Venezuelan migrants in Chile have comparable high-skilled and low-skilled groups, the majority of Haitian migrants are low-skilled. Thus, this section examines how low-skilled migration specifically affects native voting decisions. Figure 3b highlights the

[11]Regression table is on Appendix B.

[12]It is important to remember (review Table 1) that, on average, this population is no more than 0.81% at the municipal level, so it is difficult for the ratio of this population to increase by that magnitude in the cycle of four years between elections.

[13]See Appendix A for party Manifests related to migratory topics.

electoral impact of exclusively low-skilled migration on the 2021 election results[14]. Focusing on low-skilled migrants emphasizes mechanisms beyond cultural differences. Studies by Halla et al. (2017), Moriconi et al. (2022), and Mayda et al. (2022) suggest that natives may feel more threatened by low-skilled migrants due to labour market competition and competition for public goods and services.

The electoral effects of low-skilled migration are particularly pronounced with Venezuelan migrants. A 1% increase in the low-skilled Venezuelan population raises right-wing party votes by 2.5% and centre-right votes by 3.5%. Although we cannot explain with certainty the mechanism behind this steep increase, specifically due to the growth of the Venezuelan population, we can hypothesize that factors such as better job placement facilitated by shared language might heighten competition fears among natives. For the far-right, the electoral impact of exclusively low-skilled migration is similar to general migration, with their vote share decreasing only with increased Haitian migration.

### 5.4 Turnout

Migration may also affect voter turnout in two distinct ways. On the one hand, the inaction of historically dominant parties (left-wing and centre-right parties in Chile) in addressing migratory waves can generate voter disaffection, potentially decreasing turnout and favouring protest parties or independent candidates through this effect (Barone et al., 2016). On the other hand, migration can stimulate turnout, either due to voter discontent or by activating a larger pro-immigration base (Lonsky, 2021).

Table 5 examines migration's impact on turnout in Chile. Our IV model shows that a 1% increase in migrant share leads to only a 0.3% rise in turnout at the municipal level, with no significant effects when distinguishing between culturally similar and different migration. This weak turnout effect suggests that changes in centre-right and left-wing votes and the decline in far-right votes are not mainly driven by changes in turnout.

Overall, these findings elucidate the mechanisms by which migration influences municipal voting patterns. Unlike many studies that find a clear effect of migration on support for anti-migrant parties, our results show that overall migration does not directly impact the Chilean far-right. Instead, an increase in ethnically and culturally different migrants significantly reduces far-right support, aligning with Allport's contact theory (Allport, 1954), which posits that contact with culturally different migrants reduces prejudice and support for anti-immigrant parties. However, migration does affect centre-right parties, which, although not primarily focused on migration issues, borrow some anti-migration discourses from the far-right[15]. Specifically, the labour competition posed by low-skilled migrants, who can easily adapt to the local reality due to shared language and culture, can drive electoral interest in moderate parties with anti-migrant stances.

## 6 Discussion and Conclusion

For years, migration literature has examined the rising electoral success of anti-immigrant parties and far-right movements in Europe and the US, often attributing it to large inflows of migrants and refugees. The driving mechanisms behind this trend are thought to

[14]For the detailed effect of general low-skilled migration on voting outcomes, see Appendix E.
[15]See Appendix A for party Manifests related to migratory topics

**Table 5: 2SLS Model 2017-2021: Turnout & Migration changes**

| Dependent Variable | Turnout change 2017 - 2021 | | | Turnout change 2017 - 2021 | | |
|---|---|---|---|---|---|---|
| | OLS | IV | IV - control | OLS | IV | IV - control |
| Immigration ratio change | 0.257*** | 0.327* | 0.349* | | | |
| | (0.0723) | (0.176) | (0.197) | | | |
| Imm ratio change Venezuela | | | | 0.294** | 0.600 | 0.783 |
| | | | | (0.142) | (0.395) | (0.516) |
| Imm ratio change Haiti | | | | -0.185 | -0.559 | -1.341 |
| | | | | (0.276) | (1.115) | (1.336) |
| Controls | | | Yes | | | Yes |
| Observations | 346 | 346 | 346 | 346 | 346 | 346 |
| Mean Turnout 2021 | 44.21 | 44.21 | 44.21 | 44.21 | 44.21 | 44.21 |
| | (6.76) | (6.76) | (6.76) | (6.76) | (6.76) | (6.76) |
| **First Stage Regressions** | | | | | | |
| Immigration ratio change (inst) | | 1.83e-06*** | 1.85e-06*** | | | |
| | | (5.15e-07) | (4.95e-07) | | | |
| Imm ratio change Venezuela (inst) | | | | | 1.33e-06* | 1.36e-06* |
| | | | | | (7.86e-07) | (7.84e-07) |
| Imm ratio change Haiti (inst) | | | | | 7.50e-07*** | 7.32e-07*** |
| | | | | | (2.34e-07) | (2.37e-07) |
| F-stat | | 12.64 | 14.04 | | 19.35 | 16.29 |
| Part R2 | | 0.177 | 0.190 | | | |
| Part R2 Instrument 1 | | | | | 0.193 | 0.194 |
| Part R2 Instrument 2 | | | | | 0.0825 | 0.0787 |

*Notes*: Results of OLS and IV estimates on the differences between 2017 and 2021 across 346 municipalities. The dependent variable is the difference of the average turnout in a given municipality between 2017 and 2021. The variables Imm ratios are the change of immigrants divided by the municipality 2021 population; the instruments are built according to Equation 3 in the methods section. For the models with controls, I include the ageing rate and the ratio of women in each municipality during 2021. 2SLS coefficients are reported in the top panel under the heading IV. The bottom panel reports first-stage regressions with the Kleibergen-Paap Wald F statistic. Robust standard errors are in parentheses.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

include fears among locals that immigrants will: (i) harm their prospects in the labour market through job competition or downward wage pressure, (ii) reshape the social fabric and alter customs and traditions, (iii) increase crime rates, or (iv) become a burden on public services due to their vulnerability. These hypotheses have been empirically studied, primarily in developed countries, considering demographic factors, migrant characteristics, and the humanitarian reasons for migration, but scant research has been done elsewhere, despite 85% of the displaced population being hosted in developing countries (United Nations High Commissioner for Refugees, 2022). For developing countries, with lesser resources and weaker political institutions, any migration phenomenon is another source for social and political conflicts. In this paper we provide empirical evidence of the political consequences in a developing country context, especially in a case where migrants and natives share language, culture, history, and religion.

The recent emergence of anti-migrant vote and sentiment in Chile, seemingly due to the Venezuelan and Haitian diasporas that have displaced around 7.5 million migrants in the continent, allows us to contribute to the migrant and far-right literature from an ideal context: recent and significant migration waves, where a large share of the migrants and locals share common demographic characteristics, and in a democratic country with stable institutions and data reliability. This paper also contributes at the local level by understanding the mechanisms and reasons behind the formation of these negative feelings

towards foreigners so that policymakers can design public policies focused on providing support and tools to new migrants, without neglecting the distrust of the locals and the social cohesion that we need to strengthen democracy.

In this paper, I show evidence that supports recent literature (Lonsky, 2021; Pagliacci & Bonacini, 2022) indicating that migration does not drive the rise of the far-right but rather reduces prejudices in the local population, decreasing support for far-right parties. Using a shift-share instrumental variables model, we find that an increase in the migrant population does not affect the vote for the Chilean far-right party (*Republican Party of Chile*). Our identification strategy, which separates the exposure effects of different types of migration, shows that contact with culturally different migrants yields a more benevolent electoral attitude towards migration. Specifically, in our IV model, a 1% increase in culturally different migrants decreases support for the far-right by over 5%. We found no effect for culturally similar migrants, supporting Allport's contact theory (Allport, 1954) that contact with culturally different migrants reduces prejudice and support for anti-migrant parties.

However, migration does positively affect the vote for more moderate right-wing parties, which, although not primarily focused on migration, borrow some anti-migration rhetoric from the far-right. Migration, particularly from culturally similar populations regarding skill levels, language, and culture, significantly increases support for the traditional right. Left-wing parties also benefit from migration, likely by activating a larger pro-immigration base. Since we found no aggregated migration effects on the far-right vote, other factors such as security issues or economic concerns must explain its support.

To fully understand the influence of migration on the far-right's electoral success, future research should address several questions. First, it is crucial to assess other factors contributing to native fears of migration, particularly job competition, especially in economically depressed regions or areas with high job competition. Second, further research should explore how proximity to migration crises affects local attitudes. For instance, countries closer to migration sources, such as those near the Colombian-Venezuelan border, may have different responses compared to those farther away, like Argentina or Uruguay. This proximity effect is also relevant in Europe, where countries like Greece and Italy experience daily illegal sea crossings and may be more inclined to accept higher migration rates than their northern peers. The South American context, where language and religion are less relevant, provides an excellent setting to study this interaction.

Finally, the role of the media in shaping anti-immigrant attitudes and voting deserves investigation. Media coverage can create false perceptions about the number of migrants and their impact on issues like unemployment and crime. Employing strategies similar to those of Alesina, Miano, and Stantcheva (2023) in the Latin American context could help identify the media's influence in driving anti-immigrant sentiment and voting behaviour.

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# Appendix

## A Chilean Coalitions and Immigrant Rhetoric

**Table A.1: Manifesto Project Party Characteristics**

| Coalition | Right - left Index | Immigration Negative Score | | | | Immigration Positive Score | | | |
|---|---|---|---|---|---|---|---|---|---|
| | Average | 2009 | 2013 | 2017 | 2021 | 2009 | 2013 | 2017 | 2021 |
| Left | -35.37 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.03 | 0.27 | 0.07 |
| Center-Left | -21.04 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.47 | 0.00 | 0.15 |
| Center-Right | -11.26 | 0.00 | 0.00 | 0.00 | 0.09 | 0.00 | 0.00 | 0.00 | 0.32 |
| Far-right | 0.40 | | | 0.88 | 1.19 | | | 0.00 | 0.00 |

*Source*: Manifesto Project Version 2024a. Parties by coalition are: (a) Frente Amplio, Apruebo Dignidad and Partido Progresista for the Left Coalition; (b) Concertación de Partidos por la Democracia y Nueva Mayoria (Socialist Party and Christian Democrat Party) for the center-left Coalition; (c) Chile Vamos and Chile Podemos más (Renovación Nacional y Unión Demócrata Independiente) for the center-right Coalition; and (d) Partido Social Cristiano y Partido Republicano for the far-Right Coalition.

## B Main models regression tables

**Table B.2: 2SLS Model 2017-2021: Vote share & Heterogeneous Migration changes**

| Dependent Variable % vote | Far-right | | | Center-right | | | Right-wing parties | | | Left-wing parties | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control |
| Imm ratio change Venezuela | -0.599*** | -0.111 | -0.155 | 0.909*** | 1.398** | 1.329** | 0.310*** | 1.286* | 1.174** | 0.327*** | 1.283** | 1.131** |
| | (0.215) | (0.688) | (0.621) | (0.233) | (0.709) | (0.662) | (0.0863) | (0.677) | (0.579) | (0.0872) | (0.643) | (0.509) |
| Imm ratio change Haiti | -1.985*** | -5.271*** | -4.113** | 2.190*** | 2.739 | 2.464 | 0.205 | -2.532 | -1.648 | 1.044*** | -0.881 | 0.182 |
| | (0.637) | (1.766) | (1.663) | (0.534) | (1.939) | (1.815) | (0.352) | (1.635) | (1.572) | (0.376) | (1.613) | (1.506) |
| Controls | | | Yes | | | Yes | | | Yes | | | Yes |
| Observations | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 |
| Mean vote 2021 | 30.4 | 30.4 | 30.4 | 11.31 | 11.31 | 11.31 | 41.71 | 41.71 | 41.71 | 42.23 | 42.23 | 42.23 |
| | (9.90) | (9.90) | (9.90) | (4.13) | (4.13) | (4.13) | (11.08) | (11.08) | (11.08) | (9.32) | (9.32) | (9.32) |
| **First Stage Regressions** | | | | | | | | | | | | |
| Imm ratio change Venezuela (inst) | | 1.33e-06* | 1.36e-06* | | 1.33e-06* | 1.36e-06* | | 1.33e-06* | 1.36e-06* | | 1.33e-06* | 1.36e-06* |
| | | (7.86e-07) | (7.84e-07) | | (7.86e-07) | (7.84e-07) | | (7.86e-07) | (7.84e-07) | | (7.86e-07) | (7.84e-07) |
| Imm ratio change Haiti (inst) | | 7.50e-07*** | 7.32e-07*** | | 7.50e-07*** | 7.32e-07*** | | 7.50e-07*** | 7.32e-07*** | | 7.50e-07*** | 7.32e-07*** |
| | | (2.34e-07) | (2.37e-07) | | (2.34e-07) | (2.37e-07) | | (2.34e-07) | (2.37e-07) | | (2.34e-07) | (2.37e-07) |
| F-stat | | 19.35 | 16.29 | | 19.35 | 16.29 | | 19.35 | 16.29 | | 19.35 | 16.29 |
| Part R2 Instrument 1 | | 0.193 | 0.194 | | 0.193 | 0.194 | | 0.193 | 0.194 | | 0.193 | 0.194 |
| Part R2 Instrument 2 | | 0.0825 | 0.0787 | | 0.0825 | 0.0787 | | 0.0825 | 0.0787 | | 0.0825 | 0.0787 |

*Notes*: Results of OLS and IV estimates on the differences between 2017 and 2021 across 346 municipalities. The definition of variables are in the data and methodological sections. The dependent variable is the difference of the average party vote in a given municipality between 2017 and 2021. The variables Imm ratios are the change of immigrants divided by the municipality 2021 population; the instruments are built according to Equation 3 in the methods section. For the models with controls, I include the ageing rate and the ratio of women in each municipality during 2021. 2SLS coefficients are reported in the top panel under the heading IV. The bottom panel reports first-stage regressions with the Kleibergen-Paap Wald F statistic. Robust standard errors in parentheses.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

# C IV Results with alternative instruments

As a robustness check, in this section, I evaluate the main results using two alternative instruments to the one used in this article and in mainstream migration literature, such asMendez and Cutillas (2014), and Mayda et al. (2022).

## C1 Excluding Destiny Municipality

Unlike the instruments based on Altonji and Card (1991) and Card (2001), this instrument corrects for the possibility that migration inflows by nationality may be correlated with local pull-factors. Therefore, instead of using migratory inflows to Chile as shifts across municipalities, we took migration net of municipality $m$'s contribution to the total. In this way, we ensure that migratory events are not due to specific conditions in Chilean localities but to conditions in other regions of Chile.

To build this shift-share instrument, we take within-municipality differences of Equation 1 and decompose $\Delta \text{migr}_{mt} = \text{migr}_{mt} - \text{migr}_{mt-1}$ as:

$$\Delta \text{migr}_{mt} = \sum_{n} \theta^{n}_{m,t-1} \cdot \Delta \text{migr}^{n}_{mt} \tag{4}$$

Where $\Delta \text{migr}^{n}_{mt}$ is the change of the stock share of immigrants of nationality $n$ in municipality $m$ between the period $t-1$ and $t$, and $\theta^{n}_{m,t-1}$ is the share of immigrants of nationality $n$ in municipality $m$ over total immigrants of nationality $n$ at $t-1$. Specifically:

$$\theta^{n}_{m,t-1} = \frac{\text{MIGR}^{n}_{mt-1}}{\sum_{m} \text{MIGR}^{n}_{mt-1}}$$

The decomposition in Equation 4 corresponds to the weighted sum of the changes in the share of immigrants of each nationality into destination municipalities $m$. These depend both on supply-push factors in the origin country that affect all municipalities and on demand pull-factors in each municipality that affect all nationalities. To exclude the latter effect, the instrument substitutes local migration shift $\Delta \text{migr}^{n}_{mt}$ with the change of immigrants of nationality $n$ in destination municipalities other than $m$: $\Delta \text{migr}^{n}_{m't}$. Therefore, the final instrument variable is:

$$\Delta \widehat{\text{migr}}_{mt} = \sum_{n} \theta^{n}_{m,t-1} \cdot \Delta \text{migr}^{n}_{m't} \tag{5}$$

Since demand pull-factors in other municipalities are plausibly exogenous to variation in political outcomes in municipality $m$, the correlation between $\Delta \widehat{\text{migr}}_{mt}$ and $\Delta \text{migr}_{mt}$ must be solely due to supply-push factors in origin countries. This satisfies the exclusion restriction.

By using this instrument, these are my main results:

**Table C.3: 2SLS Model 2017-2021: Vote share & General Migration changes**

| Dependent Variable % vote | Far-right | | | Center-right | | | Right-wing parties | | | Left-wing parties | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control |
| Immigration ratio change | -0.611*** | -0.541 | -0.422 | 0.348** | 0.945*** | 0.850*** | -0.264** | 0.404* | 0.428** | -0.220 | 0.436** | 0.445*** |
| | (0.149) | (0.389) | (0.324) | (0.156) | (0.339) | (0.300) | (0.124) | (0.222) | (0.178) | (0.138) | (0.210) | (0.164) |
| Controls | | | Yes | | | Yes | | | Yes | | | Yes |
| Observations | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 |
| Mean vote 2021 | 30.4 | 30.4 | 30.4 | 11.31 | 11.31 | 11.31 | 41.71 | 41.71 | 41.71 | 42.23 | 42.23 | 42.23 |
| | (9.90) | (9.90) | (9.90) | (4.13) | (4.13) | (4.13) | (11.08) | (11.08) | (11.08) | (9.32) | (9.32) | (9.32) |
| **First Stage Regressions** | | | | | | | | | | | | |
| Immigration ratio change (inst) | | 1.93e-06*** | 1.96e-06*** | | 1.93e-06*** | 1.96e-06*** | | 1.93e-06*** | 1.96e-06*** | | 1.93e-06*** | 1.96e-06*** |
| | | (6.13e-07) | (5.91e-07) | | (6.13e-07) | (5.91e-07) | | (6.13e-07) | (5.91e-07) | | (6.13e-07) | (5.91e-07) |
| F-stat | | 9.868 | 10.96 | | 9.868 | 10.96 | | 9.868 | 10.96 | | 9.868 | 10.96 |
| Part. R2 | | 0.169 | 0.183 | | 0.169 | 0.183 | | 0.169 | 0.183 | | 0.169 | 0.183 |

*Notes*: Results of OLS and IV estimates on the differences between 2017 and 2021 across 346 municipalities. The dependent variable is the difference of the average political party vote in a given municipality between 2017 and 2021. The variables Imm ratios are the change of immigrants divided by the municipality 2021 population; the instruments are built according to Equation 4 in the methods section. For the models with controls, I include the ageing rate and the ratio of women in each municipality during 2021. 2SLS coefficients are reported in the top panel under the heading IV. The bottom panel reports first-stage regressions with the Kleibergen-Paap Wald F statistic. Robust standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

**Table C.4: 2SLS Model 2017-2021: Vote share & Heterogeneous Migration changes**

| Dependent Variable % vote | Far-right | | | Center-right | | | Right-wing parties | | | Left-wing parties | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control |
| Imm ratio change Venezuela | -0.599*** | -0.107 | -0.156 | 0.909*** | 1.520** | 1.439** | 0.310*** | 1.413** | 1.283** | 0.327*** | 1.404** | 1.228** |
| | (0.215) | (0.774) | (0.698) | (0.233) | (0.753) | (0.703) | (0.0863) | (0.714) | (0.610) | (0.0872) | (0.675) | (0.533) |
| Imm ratio change Haiti | -1.985*** | -5.315*** | -4.147** | 2.190*** | 2.465 | 2.220 | 0.205 | -2.850 | -1.928 | 1.044*** | -1.172 | -0.0564 |
| | (0.637) | (1.957) | (1.831) | (0.534) | (2.051) | (1.918) | (0.352) | (1.763) | (1.680) | (0.376) | (1.719) | (1.581) |
| Controls | | | Yes | | | Yes | | | Yes | | | Yes |
| Observations | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 |
| Mean vote 2021 | 30.4 | 30.4 | 30.4 | 11.31 | 11.31 | 11.31 | 41.71 | 41.71 | 41.71 | 42.23 | 42.23 | 42.23 |
| | (9.90) | (9.90) | (9.90) | (4.13) | (4.13) | (4.13) | (11.08) | (11.08) | (11.08) | (9.32) | (9.32) | (9.32) |
| **First Stage Regressions** | | | | | | | | | | | | |
| Imm ratio change Venezuela (inst) | | 1.27e-06 | 1.31e-06 | | 1.27e-06 | 1.31e-06 | | 1.27e-06 | 1.31e-06 | | 1.27e-06 | 1.31e-06 |
| | | (8.24e-07) | (8.26e-07) | | (8.24e-07) | (8.26e-07) | | (8.24e-07) | (8.26e-07) | | (8.24e-07) | (8.26e-07) |
| Imm ratio change Haiti (inst) | | 7.90e-07*** | 7.71e-07*** | | 7.90e-07*** | 7.71e-07*** | | 7.90e-07*** | 7.71e-07*** | | 7.90e-07*** | 7.71e-07*** |
| | | (2.45e-07) | (2.48e-07) | | (2.45e-07) | (2.48e-07) | | (2.45e-07) | (2.48e-07) | | (2.45e-07) | (2.48e-07) |
| F-stat | | 13.96 | 15.99 | | 13.96 | 15.99 | | 13.96 | 15.99 | | 13.96 | 15.99 |
| Part R2 Instrument 1 | | 0.175 | 0.176 | | 0.175 | 0.176 | | 0.175 | 0.176 | | 0.175 | 0.176 |
| Part R2 Instrument 2 | | 0.0838 | 0.0801 | | 0.0838 | 0.0801 | | 0.0838 | 0.0801 | | 0.0838 | 0.0801 |

*Notes*: Results of OLS and IV estimates on the differences between 2017 and 2021 across 346 municipalities. The dependent variable is the difference of the average political party vote in a given municipality between 2017 and 2021. The variables Imm ratios are the change of immigrants divided by the municipality 2021 population; the instruments are built according to Equation 4 in the methods section. For the models with controls, I include the ageing rate and the ratio of women in each municipality during 2021. 2SLS coefficients are reported in the top panel under the heading IV. The bottom panel reports first-stage regressions with the Kleibergen-Paap Wald F statistic. Robust standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

**Table C.5: 2SLS Model 2017-2021: Vote share & Low-Skilled Migration Changes**

| Dependent Variable % vote | Far-right | | | Center-right | | | Right-wing parties | | | Left-wing parties | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control |
| Immigration ratio change (low-skilled) | -0.966*** | -1.340** | -1.075** | 0.436 | 1.773*** | 1.627*** | -0.530*** | 0.433 | 0.552** | -0.436** | 0.595* | 0.702** |
| | (0.236) | (0.605) | (0.503) | (0.265) | (0.544) | (0.494) | (0.178) | (0.287) | (0.242) | (0.203) | (0.305) | (0.274) |
| Controls | | | Yes | | | Yes | | | Yes | | | Yes |
| Observations | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 |
| Mean vote 2021 | 30.4 | 30.4 | 30.4 | 11.31 | 11.31 | 11.31 | 41.71 | 41.71 | 41.71 | 42.23 | 42.23 | 42.23 |
| | (9.90) | (9.90) | (9.90) | (4.13) | (4.13) | (4.13) | (11.08) | (11.08) | (11.08) | (9.32) | (9.32) | (9.32) |
| **First Stage Regressions** | | | | | | | | | | | | |
| Immigration ratio change (inst) | | 1.93e-06*** | 1.92e-06*** | | 1.93e-06*** | 1.92e-06*** | | 1.93e-06*** | 1.92e-06*** | | 1.93e-06*** | 1.92e-06*** |
| | | (5.47e-07) | (5.54e-07) | | (5.47e-07) | (5.54e-07) | | (5.47e-07) | (5.54e-07) | | (5.47e-07) | (5.54e-07) |
| F-stat | | 12.49 | 11.96 | | 12.49 | 11.96 | | 12.49 | 11.96 | | 12.49 | 11.96 |
| Part. R2 | | 0.146 | 0.151 | | 0.146 | 0.151 | | 0.146 | 0.151 | | 0.146 | 0.151 |

*Notes*: Results of OLS and IV estimates on the differences between 2017 and 2021 across 346 municipalities. The dependent variable is the difference of the average political party vote in a given municipality between 2017 and 2021. The variables Imm ratios are the change of immigrants divided by the municipality 2021 population; the instruments are built according to Equation 4 in the methods section. For the models with controls, I include the ageing rate and the ratio of women in each municipality during 2021. 2SLS coefficients are reported in the top panel under the heading IV. The bottom panel reports first-stage regressions with the Kleibergen-Paap Wald F statistic. Robust standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

**Table C.6: 2SLS Model 2017-2021: Vote share & Low-Skilled Heterogeneous Migration changes**

| Dependent Variable % vote | Far-right | | | Center-right | | | Right-wing parties | | | Left-wing parties | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control |
| Immigration ratio change (Venezuela, low-skilled) | -1.401*** | -0.977 | -1.041 | 1.911*** | 3.923** | 3.732** | 0.510*** | 2.946** | 2.691** | 0.685*** | 3.084** | 2.744** |
| | (0.492) | (1.652) | (1.489) | (0.491) | (1.884) | (1.755) | (0.178) | (1.452) | (1.241) | (0.185) | (1.443) | (1.167) |
| Immigration ratio change (Haiti, low-skilled) | -2.170*** | -5.128** | -3.836* | 2.378*** | 1.852 | 1.608 | 0.208 | -3.276 | -2.228 | 1.086*** | -1.531 | -0.273 |
| | (0.667) | (2.330) | (2.099) | (0.562) | (2.859) | (2.636) | (0.380) | (2.000) | (1.902) | (0.411) | (2.065) | (1.923) |
| Controls | | | Yes | | | Yes | | | Yes | | | Yes |
| Observations | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 |
| Mean vote 2021 | 30.4 | 30.4 | 30.4 | 11.31 | 11.31 | 11.31 | 41.71 | 41.71 | 41.71 | 42.23 | 42.23 | 42.23 |
| | (9.90) | (9.90) | (9.90) | (4.13) | (4.13) | (4.13) | (11.08) | (11.08) | (11.08) | (9.32) | (9.32) | (9.32) |
| **First Stage Regressions** | | | | | | | | | | | | |
| Immigration ratio change (Venezuela, inst) | | 1.48e-06* | 1.51e-06* | | 1.48e-06* | 1.51e-06* | | 1.48e-06* | 1.51e-06* | | 1.48e-06* | 1.51e-06* |
| | | (8.76e-07) | (8.77e-07) | | (8.76e-07) | (8.77e-07) | | (8.76e-07) | (8.77e-07) | | (8.76e-07) | (8.77e-07) |
| Immigration ratio change (Haiti, inst) | | 7.46e-07*** | 7.30e-07*** | | 7.46e-07*** | 7.30e-07*** | | 7.46e-07*** | 7.30e-07*** | | 7.46e-07*** | 7.30e-07*** |
| | | (2.33e-07) | (2.36e-07) | | (2.33e-07) | (2.36e-07) | | (2.33e-07) | (2.36e-07) | | (2.33e-07) | (2.36e-07) |
| F-stat | | 19.19 | 18.53 | | 19.19 | 18.53 | | 19.19 | 18.53 | | 19.19 | 18.53 |
| Part. R2 (Instrument 1) | | 0.181 | 0.180 | | 0.181 | 0.180 | | 0.181 | 0.180 | | 0.181 | 0.180 |
| Part. R2 (Instrument 2) | | 0.0745 | 0.0720 | | 0.0745 | 0.0720 | | 0.0745 | 0.0720 | | 0.0745 | 0.0720 |

*Notes*: Results of OLS and IV estimates on the differences between 2017 and 2021 across 346 municipalities. The dependent variable is the difference of the average political party vote in a given municipality between 2017 and 2021. The variables Imm ratios are the change of immigrants divided by the municipality 2021 population; the instruments are built according to Equation 4 in the methods section. For the models with controls, I include the ageing rate and the ratio of women in each municipality during 2021. 2SLS coefficients are reported in the top panel under the heading IV. The bottom panel reports first-stage regressions with the Kleibergen-Paap Wald F statistic. Robust standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

### C2 Migrant Shocks to Other Countries

Unlike the previous instruments, this instrument corrects for the possibility that migration inflows by nationality may be correlated with local pull-factors. Therefore, instead of using migratory inflows to Chile as shifts across municipalities, the measure of exogenous supply-push factors is based on bilateral migration flows from the country of origin to destination countries other than Chile. This ensures that migratory events are not due to specific conditions in Chilean localities but to conditions in other countries.

To build this shift-share instrument, introduced by Bianchi et al. (2012), we take within-municipality differences of Equation 1 and decompose $\Delta \text{migr}_{mt} = \text{migr}_{mt} - \text{migr}_{mt-1}$ as:

$$\Delta \text{migr}_{mt} \approx \sum_n \theta^n_{m,t-1} \cdot \Delta \ln(\text{MIGR}^n_{mt}) - \Delta \text{pop}_{mt} \tag{6}$$

Where $\Delta \ln(\text{MIGR}^n_{mt})$ is the log change of the stock of immigrants of nationality $n$ in municipality $m$ between the period $t-1$ and $t$. $\Delta \text{pop}_{mt}$ is the change in the municipality population, and $\theta^n_{m,t-1}$ is the share of immigrants of nationality $n$ over total immigrants residing in $m$ at $t-1$. Specifically:

$$\theta^n_{m,t-1} = \frac{\text{MIGR}^n_{mt-1}}{\sum_n \text{MIGR}^n_{mt-1}}$$

The decomposition in Equation 6 corresponds to the weighted sum of the log changes in immigrants of each nationality into destination municipalities $m$. These depend both on supply-push factors in the origin country that affect all municipalities and on demand pull-factors in each municipality that affect all nationalities. To exclude local effects, the instrument substitutes $\Delta \ln(\text{MIGR}^n_{mt})$ with the log change of immigrants of nationality $n$ in destination countries other than Chile, $\Delta \ln(\text{MIGR}^n_t)$. Therefore, the final instrument variable is:

$$\widehat{\Delta \text{migr}}_{mt} = \sum_n \theta^n_{m,t-1} \cdot \Delta \ln(\text{MIGR}^n_t) \tag{7}$$

Since demand pull-factors in destination countries other than Chile are plausibly exogenous to variation in political outcomes across Chilean municipalities, the correlation between $\Delta \hat{\text{migr}}_{mt}$ and $\Delta \text{migr}_{mt}$ must be solely due to supply-push factors in origin countries. This satisfies the exclusion restriction.

By using this instrument, and since we do not have migration to other countries by skill levels, we can only provide the effect on general migration. Here are the results:

**Table C.7: 2SLS Model 2017-2021: Vote Share & General Migration Changes (based on Bianchi et al. 2012 instrument)**

| Dependent Variable % vote | Far-right | | | Center-right | | | Right-wing parties | | | Left-wing parties | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control |
| Immigration ratio change | -0.611*** | -0.0956 | -0.114 | 0.348** | 0.919** | 0.813** | -0.264** | 0.824* | 0.699** | -0.220 | 0.775* | 0.621*** |
| | (0.149) | (0.534) | (0.432) | (0.156) | (0.440) | (0.370) | (0.124) | (0.459) | (0.318) | (0.138) | (0.396) | (0.233) |
| Controls | | | Yes | | | Yes | | | Yes | | | Yes |
| Observations | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 |
| Mean vote 2021 | 30.4 | 30.4 | 30.4 | 11.31 | 11.31 | 11.31 | 41.71 | 41.71 | 41.71 | 42.23 | 42.23 | 42.23 |
| | (9.90) | (9.90) | (9.90) | (4.13) | (4.13) | (4.13) | (11.08) | (11.08) | (11.08) | (9.32) | (9.32) | (9.32) |
| **First Stage Regressions** | | | | | | | | | | | | |
| Immigration ratio change (inst) | | 0.506** | 0.549** | | 0.506** | 0.549** | | 0.506** | 0.549** | | 0.506** | 0.549** |
| | | (0.248) | (0.235) | | (0.248) | (0.235) | | (0.248) | (0.235) | | (0.248) | (0.235) |
| F-stat | | 4.155 | 5.436 | | 4.155 | 5.436 | | 4.155 | 5.436 | | 4.155 | 5.436 |
| Part. R2 | | 0.0866 | 0.107 | | 0.0866 | 0.107 | | 0.0866 | 0.107 | | 0.0866 | 0.107 |

*Notes*: Results of OLS and IV estimates on the differences between 2017 and 2021 across 346 municipalities. The dependent variable is the difference of the average political party vote in a given municipality between 2017 and 2021. The variables Imm ratios are the change of immigrants divided by the municipality 2021 population; the instruments are built according to Equation 6 in the methods section. For the models with controls, I include the ageing rate and the ratio of women in each municipality during 2021. 2SLS coefficients are reported in the top panel under the heading IV. The bottom panel reports first-stage regressions with the Kleibergen-Paap Wald F statistic. Robust standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

**Table C.8: 2SLS Model 2017-2021: Vote Share & Heterogeneous Migration Changes (based on Bianchi et al. 2012 instrument)**

| Dependent Variable % vote | Far-right | | | Center-right | | | Right-wing parties | | | Left-wing parties | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control |
| Imm ratio change (Venezuela) | -0.599*** | -0.111 | -0.155 | 0.909*** | 1.398** | 1.329** | 0.310*** | 1.286* | 1.174** | 0.327*** | 1.283** | 1.131** |
| | (0.215) | (0.688) | (0.621) | (0.233) | (0.709) | (0.662) | (0.0863) | (0.677) | (0.579) | (0.0872) | (0.643) | (0.509) |
| Imm ratio change (Haiti) | -1.985*** | -5.271*** | -4.113** | 2.190*** | 2.739 | 2.464 | 0.205 | -2.532 | -1.648 | 1.044*** | -0.881 | 0.182 |
| | (0.637) | (1.766) | (1.663) | (0.534) | (1.939) | (1.815) | (0.352) | (1.635) | (1.572) | (0.376) | (1.613) | (1.506) |
| Controls | | | Yes | | | Yes | | | Yes | | | Yes |
| Observations | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 |
| Mean vote 2021 | 30.4 | 30.4 | 30.4 | 11.31 | 11.31 | 11.31 | 41.71 | 41.71 | 41.71 | 42.23 | 42.23 | 42.23 |
| | (9.90) | (9.90) | (9.90) | (4.13) | (4.13) | (4.13) | (11.08) | (11.08) | (11.08) | (9.32) | (9.32) | (9.32) |
| **First Stage Regressions** | | | | | | | | | | | | |
| Imm ratio change (Venezuela, inst) | | 0.624* | 0.639* | | 0.624* | 0.639* | | 0.624* | 0.639* | | 0.624* | 0.639* |
| | | (0.370) | (0.369) | | (0.370) | (0.369) | | (0.370) | (0.369) | | (0.370) | (0.369) |
| Imm ratio change (Haiti, inst) | | 2.317*** | 2.260*** | | 2.317*** | 2.260*** | | 2.317*** | 2.260*** | | 2.317*** | 2.260*** |
| | | (0.724) | (0.731) | | (0.724) | (0.731) | | (0.724) | (0.731) | | (0.724) | (0.731) |
| F-stat | | 19.35 | 16.29 | | 19.35 | 16.29 | | 19.35 | 16.29 | | 19.35 | 16.29 |
| Part R2 (Instrument 1) | | 0.193 | 0.194 | | 0.193 | 0.194 | | 0.193 | 0.194 | | 0.193 | 0.194 |
| Part R2 (Instrument 2) | | 0.0825 | 0.0787 | | 0.0825 | 0.0787 | | 0.0825 | 0.0787 | | 0.0825 | 0.0787 |

*Notes*: Results of OLS and IV estimates on the differences between 2017 and 2021 across 346 municipalities. The dependent variable is the difference of the average political party vote in a given municipality between 2017 and 2021. The variables Imm ratios are the change of immigrants divided by the municipality 2021 population; the instruments are built according to Equation 6 in the methods section. For the models with controls, I include the ageing rate and the ratio of women in each municipality during 2021. 2SLS coefficients are reported in the top panel under the heading IV. The bottom panel reports first-stage regressions with the Kleibergen-Paap Wald F statistic. Robust standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

## D IV Results with an alternative base year (2009)

**Table D.9: 2SLS Model 2017-2021: Vote share & General Migration changes**

| Dependent Variable % vote | Far-right | | | Center-right | | | Right-wing parties | | | Left-wing parties | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control |
| Immigration ratio change | -0.611*** | -0.484 | -0.364 | 0.348** | 1.007** | 0.895** | -0.264** | 0.522* | 0.531** | -0.220 | 0.503** | 0.492** |
| | (0.149) | (0.448) | (0.370) | (0.156) | (0.409) | (0.356) | (0.124) | (0.275) | (0.221) | (0.138) | (0.252) | (0.193) |
| Controls | | | Yes | | | Yes | | | Yes | | | Yes |
| Observations | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 |
| Mean vote 2021 | 30.4 | 30.4 | 30.4 | 11.31 | 11.31 | 11.31 | 41.71 | 41.71 | 41.71 | 42.23 | 42.23 | 42.23 |
| | (9.90) | (9.90) | (9.90) | (4.13) | (4.13) | (4.13) | (11.08) | (11.08) | (11.08) | (9.32) | (9.32) | (9.32) |
| **First Stage Regressions** | | | | | | | | | | | | |
| Immigration ratio change (inst) | | 1.59e-06*** | 1.63e-06*** | | 1.59e-06*** | 1.63e-06*** | | 1.59e-06*** | 1.63e-06*** | | 1.59e-06*** | 1.63e-06*** |
| | | (5.64e-07) | (5.40e-07) | | (5.64e-07) | (5.40e-07) | | (5.64e-07) | (5.40e-07) | | (5.64e-07) | (5.40e-07) |
| F-stat | | 7.953 | 9.134 | | 7.953 | 9.134 | | 7.953 | 9.134 | | 7.953 | 9.134 |
| Part. R2 | | 0.131 | 0.144 | | 0.131 | 0.144 | | 0.131 | 0.144 | | 0.131 | 0.144 |

*Notes*: Results of OLS and IV estimates on the differences between 2017 and 2021 across 346 municipalities. The dependent variable is the difference of the average political party vote in a given municipality between 2017 and 2021. The variables Imm ratios are the change of immigrants divided by the municipality 2021 population; the instruments are built according to Equation 3 in the methods section considering 2009 as the base year. For the models with controls, I include the ageing rate and the ratio of women in each municipality during 2021. 2SLS coefficients are reported in the top panel under the heading IV. The bottom panel reports first-stage regressions with the Kleibergen-Paap Wald F statistic. Robust standard errors are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

**Table D.10: 2SLS Model 2017-2021: Vote share & Heterogeneous Migration changes**

| Dependent Variable % vote | Far-right | | | Center-right | | | Right-wing parties | | | Left-wing parties | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control |
| Imm ratio change (Venezuela) | -0.599*** | -0.165 | -0.203 | 0.909*** | 1.648** | 1.554** | 0.310*** | 1.483** | 1.351** | 0.327*** | 1.504** | 1.329** |
| | (0.215) | (0.800) | (0.732) | (0.233) | (0.828) | (0.772) | (0.0863) | (0.745) | (0.650) | (0.0872) | (0.732) | (0.600) |
| Imm ratio change (Haiti) | -1.985*** | -6.559*** | -4.919** | 2.190*** | 3.259** | 2.880** | 0.205 | -3.300 | -2.038 | 1.044*** | -0.986 | 0.561 |
| | (0.637) | (2.280) | (2.203) | (0.534) | (1.384) | (1.370) | (0.352) | (2.116) | (1.967) | (0.376) | (1.537) | (1.277) |
| Controls | | | Yes | | | Yes | | | Yes | | | Yes |
| Observations | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 |
| Mean vote 2021 | 30.4 | 30.4 | 30.4 | 11.31 | 11.31 | 11.31 | 41.71 | 41.71 | 41.71 | 42.23 | 42.23 | 42.23 |
| | (9.90) | (9.90) | (9.90) | (4.13) | (4.13) | (4.13) | (11.08) | (11.08) | (11.08) | (9.32) | (9.32) | (9.32) |
| **First Stage Regressions** | | | | | | | | | | | | |
| Imm ratio change (Venezuela, inst) | | 1.22e-06 | 1.26e-06 | | 1.22e-06 | 1.26e-06 | | 1.22e-06 | 1.26e-06 | | 1.22e-06 | 1.26e-06 |
| | | (8.14e-07) | (8.19e-07) | | (8.14e-07) | (8.19e-07) | | (8.14e-07) | (8.19e-07) | | (8.14e-07) | (8.19e-07) |
| Imm ratio change (Haiti, inst) | | 5.76e-07*** | 5.53e-07*** | | 5.76e-07*** | 5.53e-07*** | | 5.76e-07*** | 5.53e-07*** | | 5.76e-07*** | 5.53e-07*** |
| | | (8.68e-08) | (8.54e-08) | | (8.68e-08) | (8.54e-08) | | (8.68e-08) | (8.54e-08) | | (8.68e-08) | (8.54e-08) |
| F-stat | | 50.74 | 33.51 | | 50.74 | 33.51 | | 50.74 | 33.51 | | 50.74 | 33.51 |
| Part R2 (Instrument 1) | | 0.135 | 0.137 | | 0.135 | 0.137 | | 0.135 | 0.137 | | 0.135 | 0.137 |
| Part R2 (Instrument 2) | | 0.0458 | 0.0425 | | 0.0458 | 0.0425 | | 0.0458 | 0.0425 | | 0.0458 | 0.0425 |

*Notes*: Results of OLS and IV estimates on the differences between 2017 and 2021 across 346 municipalities. The dependent variable is the difference of the average political party vote in a given municipality between 2017 and 2021. The variables Imm ratios are the change of immigrants divided by the municipality 2021 population; the instruments are built according to Equation 3 in the methods section considering 2009 as the base year. For the models with controls, I include the ageing rate and the ratio of women in each municipality during 2021. 2SLS coefficients are reported in the top panel under the heading IV. The bottom panel reports first-stage regressions with the Kleibergen-Paap Wald F statistic. Robust standard errors are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

**Table D.11: 2SLS Model 2017-2021: Vote share & Low-Skilled Migration Changes**

| Dependent Variable % vote | Far-right | | | Center-right | | | Right-wing parties | | | Left-wing parties | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control |
| Immigration ratio change (low-skilled) | -0.966*** | -1.356* | -1.042* | 0.436 | 1.981*** | 1.797*** | -0.530*** | 0.625* | 0.754** | -0.436** | 0.671* | 0.779** |
| | (0.236) | (0.744) | (0.609) | (0.265) | (0.710) | (0.640) | (0.178) | (0.356) | (0.305) | (0.203) | (0.380) | (0.344) |
| Controls | | | Yes | | | Yes | | | Yes | | | Yes |
| Observations | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 |
| Mean vote 2021 | 30.4 | 30.4 | 30.4 | 11.31 | 11.31 | 11.31 | 41.71 | 41.71 | 41.71 | 42.23 | 42.23 | 42.23 |
| | (9.90) | (9.90) | (9.90) | (4.13) | (4.13) | (4.13) | (11.08) | (11.08) | (11.08) | (9.32) | (9.32) | (9.32) |
| **First Stage Regressions** | | | | | | | | | | | | |
| Immigration ratio change (inst) | | 1.50e-06*** | 1.49e-06*** | | 1.50e-06*** | 1.49e-06*** | | 1.50e-06*** | 1.49e-06*** | | 1.50e-06*** | 1.49e-06*** |
| | | (4.72e-07) | (4.77e-07) | | (4.72e-07) | (4.77e-07) | | (4.72e-07) | (4.77e-07) | | (4.72e-07) | (4.77e-07) |
| F-stat | | 10.05 | 9.694 | | 10.05 | 9.694 | | 10.05 | 9.694 | | 10.05 | 9.694 |
| Part. R2 | | 0.102 | 0.106 | | 0.102 | 0.106 | | 0.102 | 0.106 | | 0.102 | 0.106 |

*Notes*: Results of OLS and IV estimates on the differences between 2017 and 2021 across 346 municipalities. The dependent variable is the difference of the average political party vote in a given municipality between 2017 and 2021. The variables Imm ratios are the change of immigrants divided by the municipality 2021 population; the instruments are built according to Equation 3 in the methods section considering 2009 as the base year. For the models with controls, I include the ageing rate and the ratio of women in each municipality during 2021. 2SLS coefficients are reported in the top panel under the heading IV. The bottom panel reports first-stage regressions with the Kleibergen-Paap Wald F statistic. Robust standard errors are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

**Table D.12: 2SLS Model 2017-2021: Vote share & Low-Skilled Heterogeneous Migration changes**

| Dependent Variable % vote | Far-right | | | Center-right | | | Right-wing parties | | | Left-wing parties | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control |
| Imm ratio change (Venezuela, low-skilled) | -1.401*** | -1.215 | -1.251 | 1.911*** | 4.409** | 4.186** | 0.510*** | 3.194** | 2.935** | 0.685*** | 3.341** | 3.006** |
| | (0.492) | (1.807) | (1.647) | (0.491) | (2.111) | (1.983) | (0.178) | (1.537) | (1.359) | (0.185) | (1.554) | (1.317) |
| Imm ratio change (Haiti, low-skilled) | -2.170*** | -6.378*** | -4.466** | 2.378*** | 2.663 | 2.239 | 0.208 | -3.715 | -2.227 | 1.086*** | -1.315 | 0.477 |
| | (0.667) | (2.231) | (2.105) | (0.562) | (2.047) | (1.971) | (0.380) | (2.394) | (2.195) | (0.411) | (1.858) | (1.560) |
| Controls | | | Yes | | | Yes | | | Yes | | | Yes |
| Observations | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 |
| Mean vote 2021 | 30.4 | 30.4 | 30.4 | 11.31 | 11.31 | 11.31 | 41.71 | 41.71 | 41.71 | 42.23 | 42.23 | 42.23 |
| | (9.90) | (9.90) | (9.90) | (4.13) | (4.13) | (4.13) | (11.08) | (11.08) | (11.08) | (9.32) | (9.32) | (9.32) |
| **First Stage Regressions** | | | | | | | | | | | | |
| Imm ratio change (Venezuela, inst) | | 1.41e-06* | 1.45e-06* | | 1.41e-06* | 1.45e-06* | | 1.41e-06* | 1.45e-06* | | 1.41e-06* | 1.45e-06* |
| | | (8.53e-07) | (8.58e-07) | | (8.53e-07) | (8.58e-07) | | (8.53e-07) | (8.58e-07) | | (8.53e-07) | (8.58e-07) |
| Imm ratio change (Haiti, inst) | | 5.00e-07*** | 4.81e-07*** | | 5.00e-07*** | 4.81e-07*** | | 5.00e-07*** | 4.81e-07*** | | 5.00e-07*** | 4.81e-07*** |
| | | (7.27e-08) | (7.21e-08) | | (7.27e-08) | (7.21e-08) | | (7.27e-08) | (7.21e-08) | | (7.27e-08) | (7.21e-08) |
| F-stat | | 55.35 | 30.68 | | 55.35 | 30.68 | | 55.35 | 30.68 | | 55.35 | 30.68 |
| Part R2 (Instrument 1) | | 0.131 | 0.131 | | 0.131 | 0.131 | | 0.131 | 0.131 | | 0.131 | 0.131 |
| Part R2 (Instrument 2) | | 0.0386 | 0.0363 | | 0.0386 | 0.0363 | | 0.0386 | 0.0363 | | 0.0386 | 0.0363 |

*Notes*: Results of OLS and IV estimates on the differences between 2017 and 2021 across 346 municipalities. The dependent variable is the difference of the average political party vote in a given municipality between 2017 and 2021. The variables Imm ratios are the change of immigrants divided by the municipality 2021 population; the instruments are built according to Equation 3 in the methods section considering 2009 as the base year. For the models with controls, I include the ageing rate and the ratio of women in each municipality during 2021. 2SLS coefficients are reported in the top panel under the heading IV. The bottom panel reports first-stage regressions with the Kleibergen-Paap Wald F statistic. Robust standard errors are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

## E IV Results with low-skilled migration

**Table E.13: 2SLS Model 2017-2021: Vote share & Low-skilled Migration changes**

| Dependent Variable % vote | Far-right | | | Center-right | | | Right-wing parties | | | Left-wing parties | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control |
| Immigration ratio change (low-skilled) | -0.966*** | -1.298** | -1.040** | 0.436 | 1.701*** | 1.563*** | -0.530*** | 0.403 | 0.523** | -0.436** | 0.553* | 0.664** |
| | (0.236) | (0.590) | (0.488) | (0.265) | (0.527) | (0.477) | (0.178) | (0.274) | (0.231) | (0.203) | (0.293) | (0.265) |
| Controls | | | Yes | | | Yes | | | Yes | | | Yes |
| Observations | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 |
| Mean vote 2021 | 30.4 | 30.4 | 30.4 | 11.31 | 11.31 | 11.31 | 41.71 | 41.71 | 41.71 | 42.23 | 42.23 | 42.23 |
| | (9.90) | (9.90) | (9.90) | (4.13) | (4.13) | (4.13) | (11.08) | (11.08) | (11.08) | (9.32) | (9.32) | (9.32) |
| **First Stage Regressions** | | | | | | | | | | | | |
| Immigration ratio change (inst) | | 1.82e-06** | 1.80e-06*** | | 1.82e-06** | 1.80e-06*** | | 1.82e-06** | 1.80e-06*** | | 1.82e-06** | 1.80e-06*** |
| | | (4.73e-07) | (4.80e-07) | | (4.73e-07) | (4.80e-07) | | (4.73e-07) | (4.80e-07) | | (4.73e-07) | (4.80e-07) |
| F-stat | | 14.80 | 14.05 | | 14.80 | 14.05 | | 14.80 | 14.05 | | 14.80 | 14.05 |
| Part. R2 | | 0.150 | 0.154 | | 0.150 | 0.154 | | 0.150 | 0.154 | | 0.150 | 0.154 |

*Notes*: Results of OLS and IV estimates on the differences between 2017 and 2021 across 346 municipalities. The dependent variable is the difference of the average far-right vote in a given municipality between 2017 and 2021. The variables Imm ratios are the change of immigrants divided by the municipality 2021 population; the instruments are built according to Equation 3 in the methods section. For the models with controls, I include the ageing rate and the ratio of women in each municipality during 2021. 2SLS coefficients are reported in the top panel under the heading IV. The bottom panel reports first-stage regressions with the Kleibergen-Paap Wald F statistic. Robust standard errors are in parentheses.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

**Table E.14: 2SLS Model 2017-2021: Vote share & Low-skilled Migration changes from Haiti and Venezuela**

| Dependent Variable % vote | Far-right | | | Center-right | | | Right-wing parties | | | Left-wing parties | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control | OLS | IV | IV - control |
| Imm ratio change Venezuela (low-skilled) | -1.401*** | -0.908 | -0.966 | 1.911*** | 3.599** | 3.437** | 0.510*** | 2.690* | 2.471** | 0.685*** | 2.824** | 2.529** |
| | (0.492) | (1.478) | (1.340) | (0.491) | (1.794) | (1.675) | (0.178) | (1.377) | (1.181) | (0.185) | (1.378) | (1.120) |
| Imm ratio change Haiti (low-skilled) | -2.170*** | -5.188** | -3.902** | 2.378*** | 2.263 | 1.980 | 0.208 | -2.925 | -1.922 | 1.086*** | -1.187 | 0.0133 |
| | (0.667) | (2.126) | (1.922) | (0.562) | (2.710) | (2.506) | (0.380) | (1.856) | (1.786) | (0.411) | (1.941) | (1.835) |
| Controls | | | Yes | | | Yes | | | Yes | | | Yes |
| Observations | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 | 346 |
| Mean vote 2021 | 30.4 | 30.4 | 30.4 | 11.31 | 11.31 | 11.31 | 41.71 | 41.71 | 41.71 | 42.23 | 42.23 | 42.23 |
| | (9.90) | (9.90) | (9.90) | (4.13) | (4.13) | (4.13) | (11.08) | (11.08) | (11.08) | (9.32) | (9.32) | (9.32) |
| **First Stage Regressions** | | | | | | | | | | | | |
| Imm ratio change Venezuela (inst) | | 1.53e-06* | 1.55e-06* | | 1.53e-06* | 1.55e-06* | | 1.53e-06* | 1.55e-06* | | 1.53e-06* | 1.55e-06* |
| | | (8.22e-07) | (8.20e-07) | | (8.22e-07) | (8.20e-07) | | (8.22e-07) | (8.20e-07) | | (8.22e-07) | (8.20e-07) |
| Imm ratio change Haiti (inst) | | 7.10e-07*** | 6.94e-07*** | | 7.10e-07*** | 6.94e-07*** | | 7.10e-07*** | 6.94e-07*** | | 7.10e-07*** | 6.94e-07*** |
| | | (2.23e-07) | (2.26e-07) | | (2.23e-07) | (2.26e-07) | | (2.23e-07) | (2.26e-07) | | (2.23e-07) | (2.26e-07) |
| F-stat | | 20.90 | 15.73 | | 20.90 | 15.73 | | 20.90 | 15.73 | | 20.90 | 15.73 |
| Part R2 Instrument 1 | | 0.199 | 0.198 | | 0.199 | 0.198 | | 0.199 | 0.198 | | 0.199 | 0.198 |
| Part R2 Instrument 2 | | 0.0735 | 0.0708 | | 0.0735 | 0.0708 | | 0.0735 | 0.0708 | | 0.0735 | 0.0708 |

*Notes*: Results of OLS and IV estimates on the differences between 2017 and 2021 across 346 municipalities. The dependent variable is the difference of the average party vote in a given municipality between 2017 and 2021. The variables Imm ratios are the change of immigrants divided by the municipality 2021 population; the instruments are built according to Equation 3 in the methods section. For the models with controls, I include the ageing rate and the ratio of women in each municipality during 2021. 2SLS coefficients are reported in the top panel under the heading IV. The bottom panel reports first-stage regressions with the Kleibergen-Paap Wald F statistic. Robust standard errors are in parentheses.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$