The role of short-term and longer-term immigration trends on voting for populist radical right parties in Europe

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Abstract

The success of populist radical right parties (PRRPs) in Europe has, in part, been attributed to growing immigration, but previous findings have found an inconsistent relationship between immigration and voting for PRRPs. We address previous inconsistencies by suggesting a time-focused perspective on intergroup relations. We disentangle short-term from longer term immigration trends and argue that a recent increase in immigration should predict PRRP support. With time, however, citizens will adapt to these demographic changes and voting for PRRPs could decline. We drew on official immigration records and representative data from the European Social Survey, capturing the voting behavior of 75,874 individuals from 15 European countries between 2002 and 2014. We found that a recent increase in immigration predicted more PRRP voting, and this relationship was strengthened under conditions of higher economic strain and inequality. In contrast, sustained immigration in the longer term was not related with PRRP votes.

Keywords

immigration, voting behavior, populist radical right parties, social inequalities, country wealth

During the first decades of the 21st century, immigration increased in all European countries (Eurostat, 2019) and populist radical right parties (PRRPs), who highlight the topic of immigration (Norris, 2005), increased their influence in national parliaments (European Parliament, 2017). Scholarly work has shown, however, that the association between immigration and PRRP endorsement is not consistent, with some studies finding positive (e.g., Arora, 2019), but others finding negative (e.g., Biggs & Knauss, 2011) correlations. In the present study, we aim to reconcile conflicting results by proposing a time-focused perspective. We argue that the effects of immigration can be better understood by conceptualizing intergroup relations as dynamic. That is, while initial immigration can create a novel situation demanding reactions that are often exploited by PRRPs, with time individuals adjust to these changes and the appeal of PRRPs' anti-immigration narratives diminishes. We tested this idea with a multicountry analysis of 12 years of nationally representative data in Europe, including multiple robustness checks considering the role of key moderators such as adverse economic conditions and social inequalities.

Immigration and Votes for PRRPs

PRRPs endorse beliefs in popular sovereignty, anti-elitism, and anti-pluralism. Populist actors claim to give a voice to "the people" (Mudde, 2010) and characterize the so-called elite or establishment as evil and corrupt (Jagers & Walgrave, 2007). Within this broad agenda, PRRPs granted native citizens privileges and framed nonnatives as a threat (Arzheimer, 2015), responsible for increased criminality and the loss of national identity (Rydgren, 2008).

A Lack of Consensus About the Association Between Immigration and Support for PRRPs

Research on this topic has been multidisciplinary, with most work emerging from social psychology and political science. Previous studies found that the presence of ethnic minorities in municipalities or districts is associated with anti-immigrant attitudes (e.g., Green et al., 2010; Wagner et al., 2006). This is consistent with other research showing that voting for PRRPs across Europe is predicted by an increase in the size of the immigrant population (e.g., Arora, 2019; Coffé et al.,

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2007; Halla et al., 2017; Patana, 2018). Although these findings have been supported mostly by electoral and municipal data, scholars have also found a positive association between immigration and PRRP support in multicountry comparative research on voting behavior (Arzheimer, 2009; Arzheimer & Carter, 2006; Kessler & Freeman, 2005; Lubbers et al., 2002). Despite the significance of this body of work, conflicting results have been found. For example, parallel research has shown that the percentage of ethnic minorities in Germany is associated with less prejudice (Wagner et al., 2006). It has been similarly documented that the presence of more immigrants in a community or country predicts lower support for PRRPs (e.g., Biggs & Knauss, 2011; Charitopoulou & García-Manglano, 2018; Della Posta, 2013; van der Waal et al., 2013). Further, Stockemer and Amengay (2015) failed to find an association between immigration and voting for the Front National (the PRRP party in France). Similar studies found no significant relationship between immigration and voting for parties endorsing anti-immigration beliefs (e.g., Arzheimer & Carter, 2006; Lubbers & Scheepers, 2010). In fact, a recent meta-analysis by Amengay and Stockemer (2019) revealed that the significance and direction of the relationship between immigration and PRRP voting is inconsistent. This has also been found in Pottie-Sherman and Wilkes's (2017) meta-analysis focusing on immigration and anti-immigrant prejudice. This lack of consensus suggests that current models are not adequate to explain PRRP voting (for a similar argument, see Arzheimer, 2009).

A Time-Focused Perspective

To address the discrepancies above, we sought to improve understanding of the outcomes of intergroup relations by considering the dynamic relationships between groups over time (MacInnis & Page-Gould, 2015; Ramos et al., 2019). Immigration is typically approached as a variable having a linear association with outcomes, but reactions to immigration might change as native citizens adapt to these changes. For example, there is the perspective that individuals prefer to be among those who are similar to them (i.e., homophily; McPherson et al., 2001) and have a predisposition to endorse homogeneity and stability (Caporael, 1997). Research has also shown that people may approach out-groups with caution (Stephan & Stephan, 2000) and, at least initially, intergroup interactions are characterized by increased stress, anxiety, and outgroup avoidance (for a meta-analysis, see Toosi et al., 2012). These perspectives suggest an *ingroup orientation*, which indicates that individuals may tend to react negatively to newcomers, along the lines defined by studies showing that immigration is associated with PRRPs and anti-immigration support.

A strong ingroup orientation is, however, incompatible with the demographic changes occurring in modern societies and it requires adaptation from individuals (Crisp & Meleady, 2012). And indeed, despite the ingroup orientation highlighted above, individuals may be motivated to engage in contact with out-groups (i.e., xenophilia; Stürmer et al., 2013). In fact, when this outgroup orientation results in increased intergroup contact, it allows access to more diverse resources and knowledge (Bar-Yosef, 2002), while improving intergroup relations (Allport, 1954; Christ et al., 2014). Thus, increasing immigration should provide increased opportunities for natives to have contact with other groups, which is expected to lead to more trust and social cohesion (Schmid et al., 2014).

A few recent studies have now recognized the significance of correlates of ingroup and outgroup orientations in explaining the association between immigration and PRRP support (e.g., Green et al., 2016; Savelkoul et al., 2017). These studies, however, do not consider the evolving nature of intergroup relations in this context. Here, we argue that an ingroup orientation may be triggered immediately, while an outgroup orientation will take longer, requiring a certain degree of adaptation from majority groups. This is because some negative mechanisms stemming from an ingroup orientation (e.g., perceived threat) are likely to be triggered only by recent increases in immigration rather than stable levels of immigration. Moreover, the beneficial effects of an outgroup orientation should only be apparent in the long run, after short-term challenges have been negotiated. This reasoning is consistent with work by Ramos and colleagues (2019) showing that, in Europe, short-term increases in religious diversity were associated with a dip in well-being, but 6–8 years afterward these initial negative effects were mitigated. Immigration creates new demands and natives require time to adapt.

The Present Study

We propose that after a recent increase in immigration (i.e., a short-term immigration trend), members of majority groups are more vulnerable to the PRRPs' political agenda and will show greater endorsement of these parties. This will be in line with studies showing a positive association between immigration and support for PRRPs (e.g., Arora, 2019; Patana, 2018). However, the PRRPs' focus on immigration should become less appealing, with time, once majority groups have adapted to these demographic changes and then PRRPs will lose support. This prediction gleans support from studies showing either no association or a negative association between immigration and PRRP support (e.g., Della Posta, 2013).

Immigration has been found to be the most significant predictor of PRRP voting, but other factors such as unemployment, country wealth, and social inequalities have also been highlighted in the literature (e.g., Amengay & Stockemer, 2019; Arzheimer & Carter, 2006; Mols & Jetten, 2016; Stockemer & Amengay, 2015). Some authors have considered the interaction of these factors with immigration in the pursuit of more predictive power in their models. For example, Stockemer and Amengay (2015) predicted that immigration would be more strongly associated with PRRP voting during economic recession. Golder (2003) found an interaction between immigration and unemployment, revealing that PRRP voting increases the most when there are high levels of both unemployment and immigration (see also, Arzheimer, 2009). Social inequalities, which are associated with greater social distance and higher status anxiety

(Sánchez-Rodríguez et al., 2018), and reluctance to distribute welfare to immigrants (van Der Waal et al., 2013), might also interact with immigration to impact on PRRP voting. By testing the interaction of immigration with unemployment, country wealth, and social inequalities, we offer a more thorough test of our argument that short-term effects of immigration should differ from its longer term effects. This is because short-term increases in immigration should be associated with PRRP voting in contexts of both greater structural strain (i.e., lower country wealth and higher social inequalities) and unemployment. Given that, in the longer-term, we predict either a negative or no association between immigration and PRRP support, the moderation effects above should not be significant for longer term trends in immigration.

To test these predictions, we conducted secondary analysis of extensive cross-national European survey data. We advance the literature that has relied largely on cross-sectional attitudinal data by performing a multiwave analysis of voting behavior with seven waves of data, covering a 12-year period. With this approach, we are able to test the effects of short-term increases in immigration and compare them with longer trends comprising more than a decade of data.

Method

We relied on data of national representative samples that were collected for the European Social Survey (ESS, 2019). The ESS follows a cross-sectional multiwave design and we drew on seven waves of ESS data (2002–2014). Given that we were interested in examining change over time, we excluded from our analysis countries without multiple waves of data. Moreover, scholars have advised that including data from countries with extremely low support for PRRPs will bias results (Arzheimer & Carter, 2006; Lubbers et al., 2002). Including data from countries where PRRPs have struggled to gain traction will provide no variability and is likely to introduce floor effects. To address this issue, we followed a common procedure in the voting literature and included data only from countries in which at least one PRRP had received more than 0.5% of votes in each wave (see Werts et al., 2013). Applying these criteria, our analyses were based on data from $N = 75{,}874$ residents in 15 countries, suggesting our analysis were sufficiently powered. Sample characteristics for each country are provided in Table 1.

For our analysis, we only drew on data from individuals who had voted in the past elections and, therefore, individuals who were 18 years old or older. Our models focused on the perspectives of majority group members and not the views of immigrants. Immigrants were not sufficiently numerous for a subsample analysis (5% of the total sample).

Measures

Voting for PRRPs. Respondents specified which *party they had voted for* in their country's previous election. Votes for PRRPs were coded as "1" and votes for any other party as "0." All "don't know," "refuse to answer," and "no responses" were recorded as missing values (the proportion of these responses was below 1% and the total percentage of missing values was 5%).

PRRPs were identified following Werts and colleagues' (2013) research. Since Werts et al.'s work, some more recent PRRPs were established and were included in our analyses. These included the "Alternative for Germany" in Germany, the "Congress of the New Right" and "Law and Justice" in Poland, the "True Finns" in Finland, and "Fidez" in Hungary. In total, we examined votes for 28 PRRPs across 15 countries (see Online Supplementary Material, Table S1).

Immigration. Information about the *number of immigrants* was combined with *population size* in each country to calculate the proportion of immigrants in the 15 countries and seven waves of ESS data. The number of immigrants were retrieved from Eurostat data (Eurostat, 2019), reporting yearly numbers of incoming migrants from European and non-European countries. Population size for each country and wave was taken from the World Bank Repository (2019). These data are summarized in Online Supplementary Material, Table S2. The distinction between short-term and longer term immigration was achieved with our modeling strategy described below.

Country-level controls and moderators. To specify the countries' level of *economic strain* at each measuring point, we relied on data indicating the gross domestic product (GDP, per capita in current US$) as well as the unemployment rate. This information was extracted from the World Bank Data Repository (2019; Online Supplementary Material, Tables S3 and S4). We operationalized the *level of social inequality* in each country using ESS data about respondents' education distributions. To correct for sampling bias within the ESS and guarantee that the distributions of educational attainment were nationally representative, we relied on the statistical weights provided by the survey. Social inequality was then computed using a dissimilarity index (Massey & Denton, 1988), in which "0" indicates maximum equality and "1" indicates maximum inequality (Online Supplementary Material, Tables S5). We computed our own measure, successfully used previously (e.g., Ramos et al., 2019), which preserves the full sample of countries rather than other measures of social inequality (e.g., the GINI index) which are not available in several countries and waves in our data set.

Individual-level controls. We controlled for the following individual-level demographic characteristics: gender (1 = *female*, 0 = *male*), age, employment status (1 = *self-employed*, 2 = *other*, with the reference category "employee"), number of years of education, whether respondents lived in an urban or rural environment (1 = *a big city* to 5 = *farm or home in the countryside*), and belonging to a religious faith (1 = *yes*, 0 = *no*; see Arzheimer & Carter, 2006). In addition, we took into account two factors that have been previously associated with support of PRRPs: individuals' perceived deprivation (Mols & Jetten, 2016; Pettigrew, 2017), operationalized as respondents' perceptions of their current income (1 = *living comfortably* to 4 = *finding it very difficult*),

Table 1. Sample Characteristics Per Country Across Waves Based on the Maximum Available Sample According to Our Inclusion Criteria.

	Countries														
Variables	Austria	Belgium	Bulgaria	Denmark	Finland	France	Germany	Greece	Hungary	The Netherlands	Norway	Poland	Slovakia	Slovenia	Switzerland
N	6,456	7,140	3,660	7,843	10,253	9,672	11,656	7,193	6,637	9,260	7,907	8,604	5,391	5,420	6,621
Sex (%)															
Male	46.7	48.6	41.1	50.5	48.8	46.2	50.9	44.2	44.4	45.8	52.2	47.3	41.6	45.9	47.6
Female	53.3	51.4	58.9	49.5	51.2	53.8	49.1	55.8	55.6	54.2	47.8	52.7	58.4	54.0	52.3
Age (years)															
M	45.4	46.8	52.5	48.8	49.27	49.6	48.3	47.7	48.3	50.1	46.1	45.3	48.1	47.1	48.5
SD	18.1	18.9	17.4	18.4	19.0	12.5	18.2	18.4	18.6	17.9	18.3	18.9	17.8	18.8	18.5
Education (years)															
M	12.3	12.8	11.3	13.0	13.0	12.5	13.4	10.9	12.3	13.4	13.5	12.1	12.7	11.7	12.0
SD	3.1	3.8	3.5	5.1	4.3	4.1	3.4	4.4	3.7	4.2	3.9	3.5	3.1	3.7	3.7
Generalized trust (0–10)															
M	5.1	5.04	3.3	6.9	6.6	4.5	4.8	4.0	4.5	5.9	6.7	4.2	4.1	4.2	5.7
SD	2.4	2.2	2.6	2.0	1.9	2.2	2.3	2.4	2.4	2.0	1.8	2.4	2.5	2.5	2.2
Belonging to a religion (%)															
Yes	72.2	41.4	77.8	58.9	56.3	48.8	55.8	91.0	56.2	39.4	53.9	90.8	77.8	53.8	67.5
No	27.8	58.6	22.2	41.1	43.7	51.2	44.2	9.0	43.8	60.6	46.1	9.2	22.2	46.2	32.1
Employment status (%)															
Employee	87.2	86.0	92.3	90.1	85.5	89.5	87.8	66.1	92.1	86.8	89.8	82.6	91.1	89.2	85.4
Self-employed	10.9	12.3	5.9	9.2	11.8	9.5	10.9	31.2	6.2	11.5	8.9	16.3	8.4	7.4	12.5
Working family business	1.9	1.8	1.7	0.7	2.7	1.0	1.3	2.7	1.7	1.7	1.3	1.1	0.5	3.3	2.1
Interest in politics (1–4)															
M	3.7	2.3	2.4	2.9	2.5	2.5	2.7	2.1	2.1	2.7	2.5	2.3	2.3	2.3	2.6
SD	0.9	0.9	0.9	0.8	0.8	0.9	0.9	0.9	0.9	0.8	0.8	0.8	0.8	0.9	0.9
Household income (1–4)															
M	3.2	3.1	1.8	3.6	3.1	3.1	3.1	2.3	2.4	3.31	3.5	2.8	2.6	3.2	3.4
SD	0.7	0.9	0.8	0.6	0.7	0.7	0.8	0.9	0.8	0.8	0.7	0.7	0.8	0.8	0.8
Size of area (1–5)															
M	3.1	2.8	3.6	3.2	2.9	3.1	3.1	3.8	3.2	3.1	2.9	3.2	3.0	2.7	2.7
SD	1.2	1.2	1.3	1.2	1.4	1.2	1.1	1.3	1.2	1.2	1.3	1.2	1.1	1.1	1.0

and levels of generalized trust (Berning & Ziller, 2017; 0 = *you can't be too careful about others* to 10 = *most people can be trusted*). We also included a measure of political interest (1 = *very interested* to 4 = *not at all*). Answers to perceived economic status, interest in politics, and living environment were all reverse coded, so that higher scores reflected higher values of these measures.

Results

Analytical Approach

As in ESS each wave includes a new set of randomly drawn respondents, we performed a repeated cross-sectional procedure and accounted for the different waves in which countries were surveyed by fitting a three-level multilevel linear regression model (see Equation a). Respondents (i) were nested within waves (t), which, in turn, were nested within countries (j). The number of immigrants, x, was considered as a characteristic of specific country waves indexed as tj.

$$y_{itj} = \beta_0 + \beta_1 x_{itj} + \beta_2 x_{tjM} + \beta_3 \bar{x}_j + \mu_j + \mu_{tj} + e_{itj}. \quad \text{(a)}$$

Our predictions were assessed at the country level and we used the structural equation modeling framework to create a latent dependent variable at the country level based on individual-level responses. This dependent variable is a latent continuous variable and, as such, results at the country level are reported using regression coefficients. At the individual level, the dependent variable is dichotomous and odds ratios are reported. The independent variable was decomposed into two parts. First, we calculated the mean score of proportion of immigrants, x_j, across all ESS waves (from 2002 to 2014) for each country. This procedure provides a between-country coefficient that is time-invariant and taps into the role of long-term immigration trends that differ between countries (x_j). Second, we subtracted countries' number of immigrants in a specific wave (x_{tj}), from the mean number of immigrants in the respective country across the study period, which yields a longitudinal component $x_{tjM.}$ This within-country coefficient captures time-variant dynamics, that is, recent changes in immigration that indicate how much each country has changed in each wave from its overall levels (i.e., the means across all waves) of immigration. The between-country coefficient represents the average proportion of immigrants across the 12 years of data and it is used in our analysis as a time-invariant component, representing sustained levels of immigration over this longer period of time (for an identical methodological approach, see Easterlin et al., 2010). In Equation a, x_{itj} represents individual-level control variables, μ_j and μ_{tj} denote country and country-wave specific heterogeneities, and e_{itj} reflects an error term.

Analyses

We performed our analyses with Mplus 8.0 (Muthén & Muthén, 1998–2017) and used full information maximum-likelihood estimates with robust standard errors to better deal with missing data and deviations from a normal distribution. We estimated random intercepts and all country-level variables were group-mean centered. Effect sizes are reported as the proportion of unexplained variance.

We tested our main prediction that increases in immigration are associated with PRRP voting in the short term, but not in the longer term, using two models. In our first model (Model 1), we allowed PRRP votes to be predicted by both within-country changes in immigration (time-variant component) and between-country immigration (time-invariant component), with within-country changes indicating short-term changes and between-country immigration indicating sustained levels of immigration over the 12-year period of the analyzed data. This model was tested again with a quadratic term of between-country immigration predicting an inverted U-shape (Model 1a). This tested whether an increase in immigrants is associated with more PRRP voting, but, after a certain point, receiving more immigrants will be associated with fewer PRRP voting. Evidence of an inverted U-shaped curve would provide support for the prediction that countries having sustained high levels of immigration for a longer period would show a negative association between immigration and PRRP support. In Model 2, we examined whether high levels of immigration in the past attenuate the positive relationship between short-term increases in immigration and PRRP votes. We tested this model by allowing the between-country and within-country coefficients to interact. Finding that high sustained levels of immigration attenuate the association between short-term immigration and PRRP would provide support for our argument. It would suggest that in societies that already contain a relatively larger proportion of immigrants, native citizens may have already adapted to these changes and are not sensitive to the PRRP political agenda. Finally, we tested the interactions of short-term changes in immigration with GDP (Model 3), unemployment (Model 4), and social inequalities (Model 5). The code to reproduce our analyses is available online (https://osf.io/jfbpz/?view_only=cdf1d0d9b1d548d7988bd1afde79b27f). The contextual-level data are available in the online appendix (Tables S1–S5) and the individual-level data is available from the ESS website upon registration. Table S6 in the Online Appendix provides the zero-order correlations between our key variables.

Main analysis. In our first model, we introduced the between-country (time-invariant, longer term trends) and within-country (time-variant, short-term changes) coefficients of immigration as predictors of voting for PRRPs (Model 1, Table 2). A short-term increase in immigration within countries was associated with more PRRP voting, $\beta = 0.16$, $SE = .065$, $p = .016$, 95% CI [.05, .26] (Figure 1, Panel A). Longer term trends of immigration were not statistically related with votes for PRRPs, $\beta = 0.14$, $SE = .188$, $p = .465$, 95% CI [−.17, .45]. This pattern of results was robust considering individual-level covariates as well as average country wealth, level of unemployment, and social inequality. To provide further support for our results, we conducted a series of robustness checks, controlling for country-level variables that could

Table 2. The Relation Between Recent Changes and Long-Term Trends in Immigration and Populist Radical Right Party Votes.

Coefficients	Variables	No Controlsβ (SE), 95% CI	Model 1Odds Ratio (OR)/β (SE), 95% CI	Model 1aOR/β (SE), 95% CI
Individual-level coefficients	Sex	—	0.663 [0.59, 0.75]	
	Age	—	0.991 [0.99, .1.00]	
	Education	—	0.930 [0.92, 0.95]	
	Size of town	—	0.925 [0.89, 0.96]	
	Generalized trust	—	0.891 [0.86, 0.93]	
	Subjective income	—	0.888 [0.81, 0.98]	
	Political interest	—	0.862 [0.81, 0.92]	
	Religion belonging	—	0.872 [0.71, 1.07]	
	Employment (ref: employee)		—	
	Self-employed	—	1.063 [0.89, 1.27]	
	Other	—	1.008 [0.90, 1.13]	
Country-level coefficients	Immigration (between-country)	.272 (.115)*, [.08, .46]	0.137 (.188) [−0.17, 0.45]	0.119 (.192) [−0.20, 0.43]
	Immigration (within-country)	.172 (.071)*, [.06, .29]	0.156 (.065)* [0.05, 0.26]	0.391 (.118)** [0.20, 0.59]
	Gross domestic product (GDP; between-country)	—	0.298 (.193) [−0.02, 0.61]	0.311 (.183) [0.01, 0.61]
	GDP (within-country)	—	−0.020 (.053) [−0.11, 0.07]	−0.006 (.056) [−0.10, 0.09]
	Inequality (between-country)	—	0.076 (.131) [−0.14, 0.29]	0.076 (.128) [−0.13, 0.29]
	Inequality (within-country)	—	−0.022 (.041) [−0.09, 0.05]	0.001 (.046) [−0.08, 0.08]
	Unemployment (between-country)	—	−0.006 (.241) [−0.40, 0.39]	0.007 (.242) [−0.39, 0.41]
	Unemployment (within-country)	—	−0.203 (.199) [−0.53, 0.13]	−0.206 (.196) [−0.53, 0.12]
	Immigration (between-country)^2	—	—	−0.247 (.123)* [−0.45, −0.05]
Unexplained variance		.896 (.062)	0.787 (.117)	0.778 (.111)
Fit indicators	Loglikelihood	2.1904	5.6401	5.3182
	Akaike information criterion	42,199	40,034	40,035
	Bayesian information criterion	42,236	40,219	40,229
Sample size	Countries; country waves respondents	15; 6777,011	15; 6775,874	15; 6775,874

Note. Model 1a extends Model 1 by including the quadratic term of long-term immigration trends. *OR*s are reported at the individual-level, together with confidence intervals. At the country level, standardized regression coefficients are reported, together with standard errors and confidence intervals.
*p < .050. **p < .010. ***p < .001.

potentially interfere with our findings. These include government stability, inflation, and migrant integration policies. In addition, to address the potential issue of having a limited number of countries in our multilevel model, we fitted a multilevel model with individuals nested in country waves, controlling for country with dummy variables. In all these analyses, our results were maintained (for more details, see Online Supplementary Material, Table S7).

Next, we introduced the squared product of the between-country coefficient of immigration as a predictor variable (Model 1a, Table 2). Results highlighted a significant quadratic regression term, $\beta = -0.25$, $SE = .123$, $p = .045$, 95% CI [−0.45, −0.05]. That is, higher levels of immigration over the study period were initially positively associated with voting for PRRPs, but after a certain level of immigration, the relationship was negative (Figure 1, Panel B) such that higher long-term levels of immigration predicted fewer PRRP votes.

Additional moderation analyses. In Model 2 (Table 3), we included the interaction term between long-term (between-country coefficient) and short-term (within-country coefficient) immigration trends as a predictor, which was significant, $\beta = -0.25$, $SE = .123$, $p = .055$, 95% CI [−0.45, −0.05]. Multilevel simple slope analysis (Bauer & Curran, 2005) indicated that for countries with high average levels of past immigration (i.e., 1 standard deviation [*SD*] above the mean), a recent increase in immigration was not significantly associated with more PRRP votes, $b = .23$, $p = .814$. Conversely, for countries with low average levels of past immigration (i.e., 1 *SD* below the mean), a larger recent increase in the immigrant population predicted more voting for PRRPs, $b = .46$, $p < .001$ (Figure 2, Panel A). We then tested our models with interactions with country wealth (GDP), unemployment, and social inequalities (Models 3–5, Table 3). These additional models should provide a robustness check of our main prediction by considering the role of significant country characteristics. Model 3 demonstrated that the interaction term between changes in the size of the immigrant population and average country GDP was significant, $\beta = -0.29$, $SE = .146$, $p = .050$, 95% CI [−0.53, −0.05]; and the direction of the results (Figure 2, Panel B) was in line with our predictions. That is, when the average GDP was high (i.e., 1 *SD* above the mean), there was no association between a recent increase in immigration and PRRP voting, $b = -.16$, $p = .640$. In contrast, when the GDP was low (i.e., 1 *SD* below the mean), there was a positive correlation between

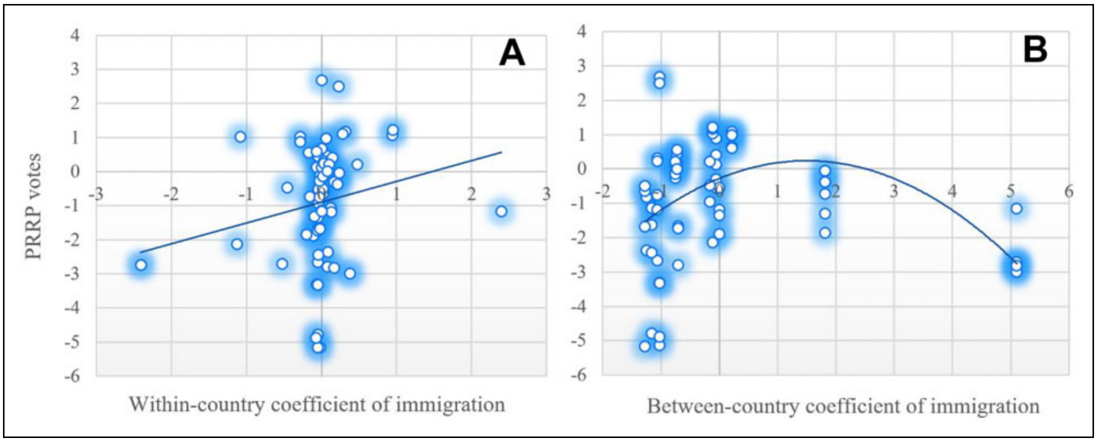

Figure 1. Linear (Panel A) and quadratic (Panel B) relationships between changes in immigration (within-country), as well as average levels of immigration (between-country) and populist radical right party votes.

short-term increases in immigration and support for PRRPs, $b = .97$, $p = .026$. In Model 4, a significant interaction between changes in immigration and average unemployment in a country, $\beta = 0.26$, $SE = .116$, $p = .024$, 95% CI [0.07, 0.45], indicated that in countries where unemployment was high (i.e., 1 *SD* above the mean), a recent rise in immigration predicted stronger support for PRRPs, $b = .71$, $p < .001$. Where unemployment was low (i.e., 1 *SD* below the mean), the relationship between short-term immigration and voting for PRRPs did not differ significantly from zero, $b = .16$, $p = .566$ (Figure 2, Panel C).

Finally, a significant interaction term between changes in immigration and the level of social inequality in a country (Model 5), $\beta = 1.73$, $SE = .805$, $p = .032$, 95% CI [0.40, 3.05], showed that in countries where levels of social inequality were high (i.e., 1 *SD* above the mean), a recent increase in immigration was associated with more PRRP votes, $b = .42$, $p < .001$. Where the level of social inequality was low (i.e., one 1 *SD* below the mean), there was no significant relationship between short-term immigration and voting for PRRPs, $b = .32$, $p = .080$ (Figure 2, Panel D). In Models 2–5, the between-country coefficient did not predict PRRP voting and did not interact with GDP, unemployment rate, and social inequalities (*p*s > .121).

Discussion

With an analysis of the voting behavior of citizens from 15 European countries, over a 12-year period, we found support for our dynamic perspective. While initial immigration creates a novel situation evoking reactions that are manifested in support for PRRPs, with time, the political agendas of these parties lose traction. Evidence supporting these findings was four-fold. The first analysis showed that PRRP voting is associated with short-term increases in immigration, but not with longer term sustained levels of immigration. This was further supported, in a subsequent analysis, by a significant quadratic coefficient revealing that support for PRRPs is associated with immigration until a certain tipping point after which countries with higher sustained levels of immigration start showing less support for these parties. A second model provided further evidence by indicating that short-term increases in immigration are associated with PRRP support in countries with lower average immigration, but not in countries having higher sustained levels of immigration. Finally, the third analysis showed that the effect of short-term increases in immigration on support for PRRPs is stronger under conditions of economic hardship and social inequalities. Longer term levels of immigration did not predict PRRP voting.

With these findings, we reconcile conflicting results found in a recent meta-analysis (e.g., Amengay & Stotckemer, 2019). We do so by showing that different effects of immigration on PRRP support depend on the specific snapshot of time being captured in a study. That is, if researchers capture a recent increase in immigration, then support for PRRPs may be observed, but a scenario of a sustained trend in immigration may be reflected in a lack of support for PRRPs. Indeed, tendencies for homophily (McPherson et al., 2001) and ingroup protection (Bowles, 2009) may initially lead to negative reactions toward newcomers, especially immigrants from unknown groups. This is in line with the work of Major et al. (2018) demonstrating that when White Americans were informed that non-White groups would become the majority in the future, participants reported stronger support for Donald Trump and anti-immigrant policies. As immigration increases, perceptions of realistic and symbolic threat may be enhanced (Outten et al., 2012) and PRRPs appear to offer effective solutions to deal with these concerns (Green et al., 2016). These findings are consistent with research showing that immigration is associated with PRRP voting (e.g., Lubbers et al., 2002). However, once there are possibilities for newcomers to participate in the host society, opportunities for a more positive outgroup orientation arise. The intergroup contact opportunities emerging in these contexts are expected to improve trust and reduce perceptions of threat (Schmid et al., 2014). This will not, however, be immediate. It will typically take some time before these contact opportunities

Table 3. Contextual Moderators of the Relation Between Recent Changes in Immigration and Populist Radical Right Party Votes.

Coefficients	Variables	Interaction With Long-Term Immigration Model 2 β (*SE*), 95% CI	Interaction With Gross Domestic Product (GDP) Model 3 β (*SE*), 95% CI	Interaction With Unemployment Model 4 β (*SE*), 95% CI	Interaction With Social Inequality Model 5 β (*SE*), 95% CI
Country-level coefficients	Immigration (between-country)	.12 (.192) [−.20, .43]	.120 (.189) [−.19, .43]	.117 (.191) [−.20, .43]	0.120 (.193) [−.20, .44]
	Immigration (within-country)	.391 (.118)** [.20, .59]	.424 (.165)* [.15, .70]	−.095 (.139) [−.32, .13]	−1.560 (.774)* [−2.83, −.29]
	GDP (between-country)	.311 (.183) [.01, .61]	.295 (.190) [−.02, .61]	.300 (.189) [−.01, .61]	0.323 (.185) [.02, .63]
	GDP (within-country)	−.006 (.056) [−.10, .09]	.005 (.056) [−.09, .10]	.001 (.057) [−.09, .10]	−0.030 (.058) [−.13, .07]
	Inequality (between-country)	.076 (.128) [−.13, .29]	.080 (.130) [−.14, .29]	.078 (.129) [−.14, .29]	0.082 (.130) [−.13, .30]
	Inequality (within-country)	.001 (.046) [−.08, .08]	−.007 (.043) [−.08, .06]	−.009 (.044) [−.08, .06]	−0.016 (.050) [−.10, .07]
	Unemployment (between-country)	.007 (.242) [−.39, .41]	−.006 (.240) [−.40, .39]	−.008 (.239) [−.40, .39]	0.006 (.243) [−.40, .41]
	Unemployment (within-country)	−.206 (.196) [−.53, .12]	−.196 (.199) [−.53, .13]	−.200 (.198) [.07, .45]	−0.216 (.193) [−.54, .10]
	Immigration (Within) × Immigration (Between)	.247 (.123)* [−.45, −.05]	—	—	—
	Immigration (Within) × Inequality (Between)	—	—	—	1.726 (.805)* [.40, 3.05]
	Immigration (Within) × GDP (Between)	—	−.286 (.146)* [−.53, −.05]	–	—
	Immigration (Within) × Unemployment (Between)	—	—	.263 (.116)* [.07, .45]	—
Unexplained variance (*SE*)		.778 (.111)	.777 (.111)	.782 (.114)	0.772 (.116)
Fit indicators	Loglikelihood	5.3182	5.3209	5.3216	5.3303
	Akaike information criterion	40,035	40,035	40,035	40,034
	Bayesian information criterion	40,229	40,229	40,229	40,228
Sample size	Countries; country waves respondents	15; 6775,874	15; 6775,874	15; 6775,874	15; 6775,874

Note. Individual-level coefficients are omitted from this table as they were the same as in the main model (Table 1).
$^{*}p < .050$. $^{**}p < .010$. $^{***}p < .001$.

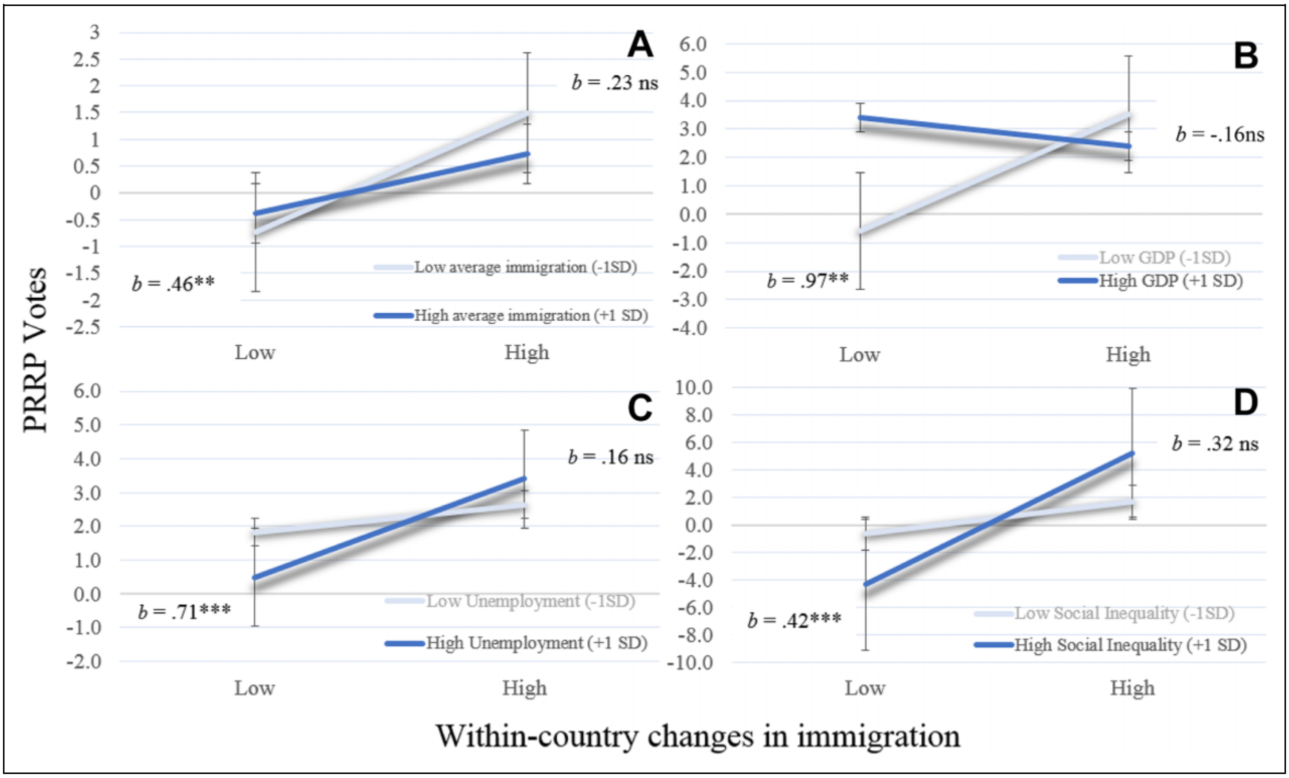

Figure 2. Interactions of changes in immigration with average immigration (Panel A), country wealth (Panel B), unemployment (Panel C), and social inequality (Panel D). *Note*. The plotted "low" and "high" values of changes in immigration flow indicate minimum and maximum observed scores of this variable. ns = nonsignificant slope. $^{**}p < .010$. $^{***}p < .001$.

are established and able to provide the basis for meaningful exchanges between groups in society (MacInnis & Page-Gould, 2015). Likewise, Ramos et al. (2019) argued that, in the context of religious diversity, the benefits of intergroup contact are only observed after some initial challenges have been overcome. Apart from intergroup contact, there are cognitive adjustments that are essential and play a role in the longer term. Evidence for this notion comes from a longitudinal study in the United States showing that undergraduate students attending ethnically diverse colleges changed their stereotypes about ethnic out-groups, reducing the perceived difference between groups during their degrees (Bai et al., 2020). These changes in perceptions should promote an environment in which more positive intergroup contacts should emerge (Allport, 1954). Taken together, these findings are in line with work showing no association between immigration and PRRP voting (e.g., Stockemer & Amengay, 2015).

Although the present study emphasizes that PRRP voting is reduced over time, the impact of a temporary increase in the support and popularity of PRRPs should not be neglected. On the one hand, increasing immigration may create the conditions for perceived threat and elicit support for PRRPs. On the other hand, politicians often persuade voters to be fearful of certain societal events and the theme of immigration can be used to gather support for PRRPs (Rydgren, 2007). Moreover, PRRPs promote anti-pluralist and nativist narratives that can incite intergroup conflict as well as inequalities—thus creating an environment fueling further threat and fear of immigrants. Such environments may dictate the relationships between groups and create tensions that are often not easily resolved, promoting deleterious conditions for immigrant groups (e.g., segregation and discrimination). This could be extremely damaging not only for newcomers but also for societies that may be entrenching long-term problematic intergroup relations.

Notwithstanding the robust findings of our study, a limitation of our data is its repeated cross-sectional data structure, which precludes causal inferences. The potential for reverse causality—whereby countries in which PRRPs receive more votes rapidly attract a larger number of immigrants—seems, however, unlikely and lacks supportive evidence. There is, nevertheless, the possibility that countries that have more PRRP votes have characteristics that might attract immigrants (e.g., higher GDP), but we controlled for these variables (see also Online Supplementary Material, Table S7). Moreover, our study examines the specific association between immigration and PRRP voting. Although previous scholarship has suggested that immigration is the most significant predictor of PRRP voting, to more broadly understand and predict PRRP voting, research should focus on a wider range of variables and perspectives (e.g., analysis capturing the interplay between the reasons why people become attracted to PRRPs and the actions of PRRPs to increase support for their political aims; see Mols & Jetten, 2020).

Conclusion

The relationship between immigration and voting for PRRPs changes over time. While initial changes can be associated with support for these parties, with time this support declines and the appeal of PRRPs' strong anti-immigration narratives loses strength. This finding emphasizes the vulnerability of majority groups—contexts of increasing immigration become relevant for majority groups and PRRPs capitalize on this opportunity. With time, however, these parties lose support, but one should never underestimate the societal effects of narratives inciting hatred between groups.

Authors' Note

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Supplemental Material

The supplemental material is available in the online version of the article.

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