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Geography of Electoral Backlash under Temporary Protection in Poland**

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When Refugees Enter Through the Main Door: Geography of Electoral Backlash under Temporary Protection in Poland

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Abstract

We leverage the massive inflow of Ukrainian refugees in Poland after the start of the large-scale Ukrainian-Russian conflict (2022), the largest humanitarian crisis in a European country since World War II, to investigate potential electoral backlashes under the best-case scenario in which the reception system is fully designed for refugee integration (i.e., under EU Temporary Protection). We employ a novel measure of refugees (the share of Ukrainian primary school students in Polish schools), which can better capture refugees' location at a granular (municipal) level compared to alternative measures adopted in previous studies. Our difference-in-differences event-study analysis based on parliamentary and presidential elections held since 2019 shows that Ukrainian refugee inflows increased local support for a mainstream conservative party, while simultaneously reducing the propensity to support more radical alternative-right parties. Drawing on established interdisciplinary scholarship, we discuss potential mechanisms underlying these effects.

Keywords: electoral backlash; refugees; voting; Poland

JEL codes: D72 J15

1 Introduction

Immigration — particularly when it occurs on a large scale — is a phenomenon that can affect the quality of life of native residents and, as a result, may generate social unrest. European experiences of immigration date back several decades and are closely linked to the colonial histories of individual countries. In the past decade, however, immigration to European states has intensified due to armed conflicts and a series of humanitarian, economic, and political crises in various regions of Asia and Africa ([Gehrsitz and Ungerer, 2022](#); [Campo et al., 2024](#); [Caselli, 2024](#)).

As immigration has intensified, concerns have grown that one of its potential consequences may be the strengthening of populist political leaders, and more specifically,

increased voter support for politicians espousing extreme and xenophobic views. Early studies lend plausibility to such a scenario (e.g., [Otto and Steinhardt, 2014](#); [Halla et al., 2017](#); [Vasilakis, 2018](#); [Dustmann et al., 2019](#); [Dinas et al., 2019](#); [Fisunoğlu and Sert, 2019](#); [Bratti et al., 2020](#); [Altındağ and Kaushal, 2021](#); [Campo et al., 2025](#); [Fasani et al., 2025](#)). The findings are, however, mixed, as some studies find support for the “contact theory,” according to which contact and prolonged interactions with refugees improve natives’ attitudes toward them (e.g., [Steinmayr, 2021](#); [Vertier et al., 2023](#); [Achard et al., 2025](#)).¹ This raises further questions: Is the shift toward more radical politicians primarily the result of general social anxieties, or of direct encounters with immigrants? Do different types of immigration (by origin, open vs. selective) provoke distinct electoral behaviors?²

The aim of this study is to examine the relationship between the sudden migration crisis in Poland — caused largely by the influx of refugees from wartorn Ukraine — and changes in the electoral behavior of native residents. More specifically, the analysis focuses on the particular nature of the Ukrainian refugee wave as a catalyst for electoral reactions, and seeks to determine whether these reactions stem from direct interactions with immigrants at the local level, or rather from a deterritorialized response observable.³ Our study covers the period between the 2019 parliamentary elections, held prior to the arrival of Ukrainian refugees, and the most recent presidential elections in 2025.

Compared to the cases examined in the past literature — which has mostly focused on inflows of irregular migrants hosted in reception centers or individuals whose refugee-status recognition is subjected to long and complex procedures — we study Ukrainian refugees displaced by the Russian–Ukrainian war, for whom there was an immediate and broad mobilization across Europe. Ukrainian refugees faced substantially different conditions compared to earlier waves of war refugees (e.g., from Syria) or to economic migrants and asylum seekers without a residence permit hosted in temporary migration centers or group accommodations. In March 2022, the EU invoked the Temporary Protection Directive (2001/55/EC) for the first time in its history in response to Russia’s invasion of Ukraine. This provided immediate protection without the need for individual asylum claims and typically included rights to residence, work, education, social welfare and healthcare across EU member states ([Das, 2025](#)). Ukrainian refugees benefiting from temporary protection were granted immediate access to the labor market and mobility within the EU — something that is not typical of standard asylum seekers under the Dublin system ([Mentzelopoulou and Orav, 2025](#)). Governments rapidly expanded reception capacity for arriving Ukrainians, often bypassing traditional asylum housing systems

¹ See [Cools et al. \(2021\)](#) for a meta-analysis.

² [Mayda et al. \(2022\)](#) find that high-skilled immigrants decrease the share of Republican votes in the U.S., while an inflow of low-skilled immigrants increases it. [Moriconi et al. \(2022\)](#) shows, using data from European elections, that larger inflows of highly-educated immigrants decreased “nationalistic” vote of natives while inflows of low-educated immigrants produced the opposite result.

³ The notion that a given shock becomes more salient to voters the closer it occurs geographically has been widely leveraged in the empirical literature. Several mechanisms may underlie this pattern, including greater exposure through traditional or social media coverage. For example, [Peracchi \(2025\)](#) exploit a quasi-experimental setting in which French authorities introduced militarized controls along the border with Italy, leading to repeated pushbacks of irregular migrants transiting from Italy back into Italian territory. The author shows that local news coverage of migration declined sharply with distance from the border following the introduction of these controls. Social media may also contribute to increasing the salience of immigration among voters. Using Twitter data, [Barilari et al. \(2025\)](#) document that immigration becomes more salient primarily during election periods, even though most arrivals occur months later. This suggests that, in the run-up to elections, informal media may shape — or even distort — public expectations and, consequently, influence voting behavior.

and allowing them to access local housing markets and social services ([Mentzelopoulou and Orav, 2025](#)).

Given that the European response to Ukrainian displacement has been structurally different from previous refugee waves, one may also expect different effects of Ukrainian refugees on natives' voting behavior relative to those documented in earlier studies. Moreover, the inflow of Ukrainian refugees consisted mainly of (often highly educated) women and children ([Deloitte, 2025](#)). From this perspective, we do not expect a major backlash driven by crime-related concerns. However, owing to the extensive financial support Ukrainian refugees received, there may have been concerns among natives about competition for public resources or amenities (cf. [Halla et al., 2017](#)) or the large refugee inflows may have affected natives' preferences towards redistribution ([Alesina et al., 2021](#)). Put differently, our study provides one of the few quasi-experimental assessments of the potential electoral backlash generated by a refugee reception system explicitly designed to foster integration and that does not raise salient concerns related to public safety or crime, arguably a "best-case scenario". This constitutes a relevant contribution to the literature, as existing studies typically document political backlash in contexts where refugees are hosted in large-scale or emergency accommodations (e.g., large reception centres or refugee camps) that provide limited integration services ([Gamalerio et al., 2023](#); [Campo et al., 2024](#); [Rickardsson, 2025](#)). These findings implicitly suggest that deficiencies in reception design and integration policies may be key drivers of political backlash. The EU Temporary Protection regime granted to Ukrainian refugees offers a quasi-natural experiment that allows us to directly test this hypothesis.

Of great importance is the fact that immigration from Ukraine involves a group that is culturally and linguistically close to the Polish population, as well as connected by a shared perception of threat from Russia. In the initial period following the outbreak of war, these conditions prompted massive engagement of Polish citizens in assisting with the absorption and allocation of the refugee wave. Although this situation was not the result of a deliberate migration policy, it may be interpreted as an instance of "selective migration," that is, the admission of migrants with specific characteristics facilitating integration into the host society ([Van de Werfhorst and Heath, 2019](#)). Selective migration is often contrasted in scholarly debate with "open migration," which does not apply admission criteria. The simultaneous occurrence of both types of immigration in Poland allows us to test the hypothesis that selective migration either does not affect electoral behavior, or affects it in a different way compared to open migration.

Poland offers a unique research setting that can make a significant contribution to this field of knowledge. For decades, the country was primarily a source of emigration rather than a destination. A perceptible inflow of foreigners is a relatively new phenomenon, which intensified after 2020 (see [Figure A1](#)). The war in Ukraine played a crucial role, leading within just a few months in 2022 to the settlement of approximately one million refugees in Poland ([Deloitte, 2025](#)). This represents a more intense inflow than in any other European country. Such a rapid process in a country with little prior experience of immigration constitutes a major shock to society and the economy, while also providing an effective opportunity to test hypotheses regarding the impact of migration on changes in native electoral behavior.

The uniqueness of the Polish situation in the period 2020–2025 also lies in the fact that, alongside the influx of refugees from Ukraine, immigration from other countries — particularly from outside Europe — intensified. The causes of this phenomenon are similar to those observed in other European countries that serve as destinations for migrants

from Asia and Africa. Considering only registered immigrants holding residence permits, the number of immigrants from countries other than Ukraine tripled between 2017 and 2025 ([Office for Foreigners, 2025](#)).

This study introduces several important innovations to the existing literature — both that concerning Ukrainian immigration to Poland and that addressing migration crises in other countries. First, it exploits the unique configuration described above, which enables the testing of previously unexamined hypotheses regarding selective migration. Second, it examines immigration and electoral behavior at a highly disaggregated level — namely, at the level of Polish municipalities. To date, all attempts concerning Poland have relied on analyses at the county level (e.g., [Mykhailyshyna and Zuchowski, 2025](#); [Clifton-Sprigg et al., 2025](#)). Existing research from other European countries indicates, however, that an excessive spatial aggregation of data can lead to misleading conclusions regarding the impact of migration on citizens’ voting behavior ([Devillanova, 2021](#)). Third, it employs a distinctive measure of immigrant presence in local communities, based on school data on foreign pupils. These data provide highly reliable information on immigrants’ place of residence and simultaneously relate directly to competition for public services, one of the sources of social anxiety associated with immigration. Indeed, previous studies from other countries demonstrate that a deterioration in access to public services may contribute to a local increase in support for populist political parties ([Fregoni and Leonardi, 2025](#)). Finally, unlike earlier studies of local electoral reactions to immigrant presence in Poland, this research draws on both presidential and parliamentary election results. The combination of data from two types of general elections is crucial, as it allows for the estimation of the long-term effect of migratory inflows (rather than the immediate effect measured shortly after the onset of a crisis), while at the same time enabling the observation of voting behavior at multiple points in time.

Our empirical analysis, based on a municipality-level difference-in-differences (DiD) design, shows that recent inflows of Ukrainian refugees (and migrants) increased support for the main conservative party while reducing support for more extreme alternative-right parties. Interestingly, and against our expectations, our estimated electoral effects of selected refugees under EU Temporary Protection are comparable in magnitude to those reported in the literature for refugees or migrants in general ([Cools et al., 2021](#)). We interpret our findings as reflecting the gradual appropriation of anti-immigration rhetoric by mainstream conservatives. While contextual factors — such as direct contact with migrants and refugees — may have enhanced the political effectiveness of this rhetoric among potential conservative voters, we argue that alternative-right voters are more ideologically driven. As such, their support appears to respond less to local exposure to migrants and more to generalized anxiety about immigration. Ultimately, the mainstream conservative party seems to have absorbed some of the electoral support that might otherwise have accrued to the far-right in municipalities more affected by the refugee shock.

The remainder of the paper is structured as follows: Section 2 describes the institutional background of our research, including the political landscape of Poland, the recent immigration influx, and trends in voting behavior. Section 3 provides information on the data used throughout the analysis. Section 4 illustrates our empirical strategy for assessing the impact of immigration on voting. Section 5 discusses the results and presents some robustness checks. Finally, Section 6 summarizes the main findings.

2 Institutional Background

2.1 The Political Landscape in Poland

Poland is a parliamentary republic in which legislative power is exercised by two chambers of parliament (the Sejm and the Senate), while executive power is vested in the Council of Ministers and the President. Both members of parliament and the President are elected in universal suffrage, with parliamentary terms lasting four years and presidential terms five years. During the period 2019–2025, which is the focus of this study, four nationwide elections were held: parliamentary elections in 2019 and 2023, and presidential elections in 2020 and 2025. The elections of 2019 and 2020 took place before the massive influx of refugees from Ukraine, while the subsequent ones were already held in the context of the migration crisis triggered by Russia’s invasion of Ukraine in 2022.

Poland’s political scene is strongly polarized, with two parties playing the dominant role: the conservative Law and Justice (PiS) and the liberal Civic Platform (PO). These parties, the coalitions they form, or — in the case of presidential elections — the candidates they support, consistently occupy the top two positions in every election. At present (2025), Poland is experiencing cohabitation: the government is formed by a coalition led by PO, while the President is a candidate supported by PiS during the elections.

According to the results of the Chapel Hill Expert Survey (CHES) from 2019, PiS can be characterized as a right-wing party. On the 0–10 left–right scale, it received an average expert score of 7.55. PO, by contrast, appears in this general classification as a centrist party, with an average score of 5.38 (CHES, 2025). However, ideological differences between the main parties are much deeper when, instead of the general left–right division, the so-called GAL–TAN criterion (Green, Alternative and Liberal – Traditional, Authoritarian and Nationalist) is applied. In this spectrum, PiS scored 9.14, while PO scored 4.67. Among the right-wing political parties that fielded candidates in every election between 2019 and 2025, the most radical grouping in Poland is the party called Confederation (“Konfederacja”), which scored 9.53 on the GAL–TAN scale in 2019. Although the score seems close to that attributed to PiS, the issue of ethnic minorities plays a more significant role in Confederation’s program. A CHES sub-survey conducted specifically with regard to minority issues yielded a score of 9.50, compared to the 8.0 obtained by the more mainstream PiS.

2.2 Recent Immigration to Poland

Historically, Poland has been more a source of emigration than a destination for immigration. According to estimates, in 2016 around 2.5 million Polish citizens resided permanently or temporarily abroad, while the resident population of Poland was approximately 38.5 million (Hrynkiewicz et al., 2017). At the same time, since the end of World War II Poland remained a relatively homogeneous country in ethnic terms. According to the 2002 census, only 1.2 percent of Poland’s inhabitants declared a nationality other than Polish (GUS, 2002).

The last decade, however, has brought a significant change in Poland’s role on the global migration map. Between 2017 and 2025, the total number of registered immigrants increased from around 240,000 to 1,950,000 — an eightfold rise. Of this number, as many as 964,000 were war refugees from Ukraine. The number of immigrants from Ukraine without refugee status also rose more than fivefold, to 549,000, which likely reflects a greater tendency to legalize status in the context of ongoing war. At the same

time, the number of immigrants from countries other than Ukraine also tripled, reaching approximately 430,000 in 2025 ([Office for Foreigners, 2025](#)). After two decades of dynamic development within the European Union, Poland has become an attractive destination for economic migration and is increasingly described as an Eastern European “immigration hub” ([Kusek, 2024](#)).

2.3 Is Political Radicalization Evident?

The issue of immigration and the desired policy response has been an important element of recent electoral campaigns in Poland — both in the parliamentary elections of 2023 and the presidential elections of 2025, i.e. contests held after the migration crisis triggered by the war in Ukraine and the intensified inflow of immigrants from other origins. Compared to earlier campaigns, in which migration played a relatively minor role, after 2022 the issue of threats posed by immigrants and the need to protect borders gained priority status in the programs of right-wing parties such as PiS and Konfederacja ([Musiał-Karg and Bértoa, 2025](#)). Moreover, in elections held after 2020, political parties and candidates increasingly used immigration as a tool to attack political opponents rather than to promote social cohesion and intercultural integration ([Wyligala, 2024](#); [Paczeński and Winclawska, 2024](#)).

As for Polish citizens’ attitudes toward Ukrainian war refugees, while in the first months of the war the dominant opinion emphasized unconditional assistance and support, by the second half of 2023 a sense of fatigue with the mass presence of Ukrainians in Poland began to grow, accompanied by attempts to question the support policies directed at them ([Kozyska, 2025](#)). In October 2024, 40 percent of respondents in a nationwide, systematically conducted survey declared opposition to Poland’s acceptance of refugees from Ukraine. By comparison, in March 2022 only 3 percent of respondents gave such an answer to the same question ([CBOS, 2024](#)). In December 2024, 51 percent of respondents in Poland stated that the assistance provided to Ukrainian refugees by the Polish state was excessive ([Centrum-Mieroszewskiego, 2025](#)).

The results of the most recent elections, namely the 2025 presidential elections, appear to reinforce the thesis of advancing nationalist radicalization. Right-wing candidates, who to varying degrees employed nationalist and anti-immigrant themes in their campaigns, obtained a combined 51.5 percent of the vote in the first round (including votes cast for the mainstream right-wing candidate Karol Nawrocki). If attention is limited to non-establishment candidates, thus excluding Nawrocki, the right-wing share amounted to 21.83. A comparable analysis of the 2020 elections yields 50.8 percent for the entire right, but only 7.29 percent when the mainstream candidate is excluded (National Electoral Commission; [PKW, 2020, 2025](#)).

3 Data

To examine the relationship between the inflow of refugees and the radicalization of electoral behavior among the local population, we leverage data from three primary sources: the National Electoral Commission (PKW), which is responsible for conducting all nationwide elections in Poland and announcing their results; the Educational Information System (SIO), which collects data on public schools in Poland, including the attendance of pupils who are citizens of other countries; and the Central Statistical Office (GUS), which

provides data characterizing local communities and economies across all 2,477 municipalities in Poland. This combination of sources makes it possible to conduct estimations at the municipal level, with a higher degree of disaggregation than in previous analyses of electoral geography in the context of immigration in Poland (Mykhailyshyna and Zuchowski, 2025; Clifton-Sprigg et al., 2025). In addition, we made use of unique data on the religious engagement of municipal residents, provided by the Institute of Statistics of the Catholic Church (ISKK).

3.1 Voting

The dominant role of the two main parties, as outlined in Subsection 2.1, positions the remainder of the Polish political scene in relation to these principal actors. In the context of the threat of nationalist radicalization, two types of voter behavior merit particular attention: (i) Support for the strongest conservative-right party (PiS) or its endorsed candidate in presidential elections; (ii) Support for non-mainstream parties or candidates advancing a right-wing program. We will refer to this group as the “alternative-right” (“alt-right,” hereafter).

Consequently, we operationalize the dependent variables as the vote shares obtained by the parties or candidates classified within the categories “main conservative” and “alt-right,” as defined above. In this study, we consider the first rounds of the 2020 and 2025 presidential elections as well as the results of the 2019 and 2023 Sejm elections (the lower chamber of the Polish parliament).⁴ As for the former, we focus on the first round, which is more likely to reflect voters’ true preferences. As for the latter, we rely on the results of the Sejm elections, while disregarding the concurrent Senate (upper chamber) elections. This choice reflects the fact that Sejm elections are conducted under a proportional party-list system, in which a larger number of electoral committees compete, and party affiliation — and thus the ideological orientation of a party — plays a key role in voter choice. By contrast, the single-member district (majoritarian) system used in Senate elections results in weaker party representation and a greater number of independent candidates, who are difficult to classify ideologically. Comparing the support for the main conservative party across elections, in the 2019 and 2020 elections, it was almost identical (43.6 percent and 43.5 percent, respectively). In subsequent elections, PiS (or the presidential candidate supported by the party) achieved lower results — 35.4 percent in 2023 and 29.5 percent in 2025. By contrast, the alternative right gradually gained strength — from 6.8 percent in 2019 to 21.8 percent in the presidential elections of 2025.

The geography of support for the main conservative party remained highly stable throughout the study period (see Figure 1). In each election, the PiS party received the strongest backing in the eastern and south-eastern parts of the country, while support was significantly weaker in the north-western regions and in metropolitan areas. Correlation coefficients between PiS support measured at the municipal level across different years exceed 0.9, and are equally strong both within elections of the same type (e.g., parliamentary) and across different types of contests.

The alt-right category is, for obvious reasons, less stable, as it consists of different groupings and candidates in successive elections. Nevertheless, the territorial distribution of support for them in 2020–2025 (two presidential elections and one parliamentary

⁴ We do not consider the 2015 elections, which were too close to Russia’s annexation of Crimea (2014) and the rise of the conflict. We want to focus on a relatively stable period before the outset of the large-scale Russian invasion.

election) is fairly consistent (see Figure 2). The highest percentage of votes for the alternative right is found in the eastern and south-eastern regions. In the 2019 parliamentary elections, which stand out somewhat in terms of the geographical distribution of results, strong right-wing support also appeared in isolated pockets in the Pomorskie voivodeship (north) and the Lubuskie voivodeship (west).

3.2 Refugees and Immigrants

Most studies examining the impact of immigration on local voting behavior rely on administrative sources — typically population registers classified by nationality or country of birth — to identify the residential distribution of immigrants within a country (Mendez and Cutillas, 2014). In some cases, researchers exploit institutional mechanisms of migrant allocation, such as the spatial distribution of temporary reception centers (Bratti et al., 2020; Campo et al., 2025, 2024), refugee dispersal policies (Dustmann et al., 2019), shift-share instruments based on historical settlements (Barone et al., 2016; Halla et al., 2017; Edo et al., 2019; Mayda et al., 2022), or the availability of building premises suitable for hosting refugees (Steinmayr, 2021; Gamalerio et al., 2023; Fasani et al., 2025).

In the Polish context, analyses of the territorial distribution of Ukrainian refugees have predominantly relied on the PESEL register, in which Ukrainian citizens fleeing the war must enroll to obtain temporary protection, access to the labor market, and social benefits (Mykhailyshyna and Zuchowski, 2025; Clifton-Sprigg et al., 2025). As a measure of immigrant presence, however, PESEL data have important limitations. They include only individuals who crossed the Polish–Ukrainian border after 24 February 2022, excluding approximately 500,000 Ukrainians residing in Poland prior to that date. From the standpoint of everyday social interaction, the legal status of Ukrainian immigrants is indistinguishable to native residents, which makes it analytically necessary to account for the entire immigrant population rather than formally defined refugees alone. Moreover, PESEL data are available only at the county level (380 powiaty), offering limited spatial precision. Finally, administrative registration in a given county does not ensure that a household actually resides, works, or regularly uses public services in that location. Taken together, these features render PESEL an imperfect proxy for capturing sites of routine contact between immigrants and native populations.

To address these limitations, this study employs an alternative indicator of immigrant presence based on educational data. We rely on information on the enrollment of foreign students in primary schools across Poland’s 2,477 municipalities. These data are collected within the SIO, administered by the Ministry of Education. While not previously used to study the electoral consequences of immigration, these data have been employed in analyses of the integration of Ukrainian children into Polish schools (Herbst and Sitek, 2023). Beyond their fine-grained territorial coverage and inclusion of all Ukrainian children regardless of legal status, educational data capture direct and sustained contact between immigrants and natives and can proxy competition for access to local public services. School attendance also represents a particularly sensitive arena of intergroup relations in multicultural settings, which explains the prominence of school-based dynamics in the migration literature (Vega et al., 1995; Walsh et al., 2015; Pratt-Johnson, 2015). In this study, we use only data on the presence of Ukrainian pupils in primary schools, although secondary school data are also available in the SIO. This choice is motivated by the fact that primary schools are present in each of Poland’s 2,477 municipalities, whereas secondary schools are primarily located in urban areas. Consequently, the location of

the secondary school attended does not necessarily correspond to the student’s place of residence. Moreover, at the primary — though not secondary — education level, pupils are formally assigned to a specific local school. While this assignment is not mandatory and parents may choose to enroll their child elsewhere, in practice approximately 95 percent of primary school students in Poland attend a school within their municipality of residence ([GUS, 2023](#)).

The main independent variable in our analysis is defined as the change in the share of Ukrainian students among all primary school students in a municipality between September 2020 and September 2024. The total number of Ukrainian students in Polish primary schools increased from approximately 24 thousand in 2020 to 147 thousand in 2024, rising from 0.8 percent to 4.6 percent of all enrolled students. In 50 municipalities, the share of Ukrainian students reached or exceeded 10 percent. The number of foreign students originating from countries other than Ukraine is clearly lower, but it has also increased sharply — from approximately 9.5 thousand in 2020 to 26 thousand four years later. The spatial distribution of both variables, shown in Figure 3, reveals that the largest increases — both for Ukrainian students and for those originating from other countries — are concentrated in the metropolitan areas of Poland’s largest cities: Warsaw (central Poland), Wrocław (south-west), Poznań (west), and the Tri-City of Gdańsk–Gdynia–Sopot (north). Elsewhere, inflows of immigrant pupils are less systematic, appearing more sporadic and spatially fragmented.

3.3 Other Variables

The results of earlier studies suggest that community characteristics may significantly affect voting, with education, economic status, and local demographic composition playing crucial roles. Evidence from the United States shows that lower-educated and less affluent populations are more likely to support anti-immigration parties in reaction to increased immigration exposure ([Maggio, 2022](#)). Older voters and those living outside urbanized areas are also more likely to radicalize ([Cohen-Marks et al., 2008](#)). Acknowledging evidence from these and many other studies, we seek to determine the extent to which pre-existing municipal characteristics shaped subsequent trajectories of local electoral behavior in Poland. Specifically, we analyze the role of two broad factors: the community’s readiness (capacity) to absorb newcomers, and socio-demographic characteristics of the electorate. For the first factor, the following variables from the Central Statistical Office (GUS) are used: population density and municipal own-source revenue per capita. In terms of the second factor, we considered the share of the population of retirement age (i.e., above 64), the share of the university-educated population and the share of dominicantes, obtained from the GUS. All data comes from 2015, which is well before the immigration crisis, except for two exceptions. Education is based on the 2002 national census. The share of dominicantes is collected by the Institute for Catholic Church Statistics (ISKK): the percentage of parish members attending Sunday Mass in 2019, which, according to the literature, may be related to both bonding and bridging social capital ([Hoyman et al., 2016](#)).

Table A1 reports the estimation sample summary statistics.

4 Empirical Strategy

4.1 Baseline Analysis

Our data span two parliamentary (2019 and 2023) and two presidential elections (2020, 2025). To explore the effect of the local inflow of Ukrainian primary students on voting for the mainstream conservative party or the alt-right parties, we use the following event-study DiD-like model:

$$y_{ict} = \alpha_0 + \sum_{j=2019}^{2025} \alpha_{1,j}(\Delta_{20-24}\text{UKR}_i \times \mathbf{1}(t = j)) + D_i + D_{ct} + \epsilon_{ict} \quad (1)$$

where $j = 2019, 2023, 2025$ and $\mathbf{1}(\cdot)$ is an indicator function; y_{ict} is the share of votes for the mainstream conservative party (or alt-right parties) in the municipality i of county c in the t election; $\Delta_{20-24}\text{UKR}_i$ is the change in the share of Ukrainian primary students between 2020 and 2024 in municipality i , i.e., the size of the migration or refugee *shock* (the intensity of our “treatment” of interest); D_i and D_{ct} are municipality and county-election fixed effects, which capture, *inter-alia* long-term municipality political orientation and election-county specific confounders (e.g., the county unemployment rate or other time-variant county-level characteristics), respectively; ϵ_{ict} is an idiosyncratic error term. Errors are clustered at the municipality level to account for serial correlation in the error term. Our specification is consistent with the idea that large increases in refugee inflows, rather than the long-term stock of migrants and refugees — which was very low and stable in Poland until 2019 (see Figure A1) — are more likely to affect voting behavior, and it aligns with recent findings by Rickardsson (2025).

Our coefficients of interest are the $\alpha_{1,j}$ s, which capture variations in the effect of the Ukrainian refugee inflows over time. In particular, for $t = 2019$, i.e., the “lead” term, the interaction captures the effect of future inflows and is useful to inspect differential pre-trends in voting behavior between municipalities that will host a high number of refugees and those that will not.⁵ Thus, the magnitude and statistical significance of the lead interaction provide useful indications of the potential endogenous location of refugees according to the political attitudes prevailing in the municipalities. The omitted interaction is 2020, which represents the base year, i.e., the year of the last election before the start of the large-scale Russian invasion of Ukraine (2022). Evidence that $\alpha_{1,2019} = 0$ would be consistent with a parallel-trends assumption. By contrast, $\alpha_{1,2023}$ and $\alpha_{1,2025}$ (the “lag” terms) capture dynamic effects of the Ukrainian refugee inflow on the first post-war parliamentary and presidential election, respectively.

Our identification strategy is different from that of recent papers investigating the impact of the refugee crisis in Poland (Mykhailyshyna and Zuchowski, 2025; Clifton-Sprigg et al., 2025), which use instrumental variables (IVs), while it is similar to (Campo et al., 2025, see Figure 2 in their paper), which investigates the effect of a refugee dispersal policy in Italy and Dinas et al. (2019) analyzing votes for the far-right in Greece. An advantage of our DiD specification, compared to IVs, is that it allows for identifying a more general causal parameter, namely the Average Treatment Effect on the Treated

⁵ In what follows, we refer to Ukrainian migrants interchangeably as “migrants” and “refugees.” The two groups cannot be distinguished in our data. Moreover, refugees under EU Temporary Protection enjoyed rights very similar to those of regular migrants and were largely indistinguishable in the eyes of native residents.

(ATT) instead of the Local Average Treatment Effect (LATE), i.e., the effect of refugees on municipalities in which the share of refugees changed because of the instruments.⁶

Although Equation (1) already accounts for some regional confounders via the county-year fixed effects, one possible concern is that the inflow of UKR students might be correlated with some municipality characteristics whose effects on voting are time variant, and that may change because of the same phenomenon generating the refugee inflows. To take an example, while the effect of the demographic structure of the population on voting might have remained almost constant in elections before the Russia-Ukraine war, voting behavior of older (younger) people might have changed after the war towards more far-right voting. In this case, not accounting for such time-variant effects of confounders might generate a spurious correlation with the refugee inflows. It is important to stress that evidence consistent with an unconditional parallel-trends assumption (UPTA) before the war does not discard *per se* the possibility that the relationship between municipal-level characteristics and voting may change after the war. Thus, a safer modelling approach is to control in Equation (1) for interactions between some relevant municipal characteristics measured at baseline (i.e., before the first election in the sample) and year indicators. The new estimated equation is

$$y_{ict} = \alpha_0 + \sum_{j=2019}^{2025} \alpha_{1,j}(\Delta_{20-24}\text{UKR}_i \times \mathbf{1}(t = j)) + \sum_{j=2019}^{2025} \alpha_{2,j}(X_{i0} \times \mathbf{1}(t = j)) + D_i + D_{ct} + \epsilon_{ict} \quad (2)$$

where X_{i0} is a baseline municipality-level characteristic, measured before the 2019 election (the first in the sample), so as it is predetermined. The $X_{i0} \times \mathbf{1}(t = j)$ interactions (or more shortly, $X \times \text{year}$ interactions) control for potential time-varying effects of municipalities' characteristics at baseline. We simultaneously include in the estimated equation interactions between all variables described in Subsection 3.3 and year indicators.

Failing to account for time-variant effects of time-invariant confounders biases the DiD estimates (see, [Zeldow and Hatfield, 2021](#)). To see this formally, consider the two-period DiD setup (pre, $t = 0$, and post, $t = 1$) and let $D_i \in \{0, 1\}$ a dichotomous version of the treatment, e.g., an indicator for the refugee inflows being above (vs. below) the median. Suppose the untreated potential outcome follows

$$Y_{it}(0) = \alpha_i + \lambda_t + \gamma'_t \mathbf{X}_i + u_{it},$$

where $\mathbf{X}_i$ is a vector of time-invariant, pre-treatment municipal characteristics and γ_t is allowed to vary with time. The observed outcome can be written as

$$Y_{it} = Y_{it}(0) + D_i \cdot \tau_t,$$

where τ_t is the average treatment effect on the treated (ATT) at time t (with $\tau_0 = 0$ under no anticipation). Taking first differences yields

$$\Delta Y_i \equiv Y_{i1} - Y_{i0} = (\lambda_1 - \lambda_0) + (\gamma'_1 - \gamma'_0) \mathbf{X}_i + (u_{i1} - u_{i0}) + D_i \cdot \tau_1.$$

⁶ Since the instruments used by researchers are generally related to past migrant or refugee settlements and follow the *enclave* logic, e.g., the distance to historical hotspots of Ukrainian networks ([Mykhailyshyna and Zuchowski, 2025](#)) or shift-share instruments ([Clifton-Sprigg et al., 2025](#)), it is likely that papers using IVs may estimate the effect of refugee inflows in locations characterized by already well-established Ukrainian communities.

The canonical “unconditional” DiD regression omits $\mathbf{X}_i$ and estimates $\hat{\tau}^{\text{DiD}}$ from a regression of ΔY_i on D_i alone. The estimated parameter is in this case

$$\hat{\tau}^{\text{DiD}} = \tau_1 + \underbrace{(E[\mathbf{X}_i \mid D_i = 1] - E[\mathbf{X}_i \mid D_i = 0]) (\gamma'_1 - \gamma'_0)}_{\text{omitted-variable bias}}, \quad (3)$$

where the second term is an omitted-variable bias component. Thus, even if the UPTA holds in the sense that $E[u_{i1} - u_{i0} \mid D_i = 1] = E[u_{i1} - u_{i0} \mid D_i = 0]$, the DiD estimate is biased whenever (i) the distribution of baseline characteristics differs between treated and control municipalities, $E[\mathbf{X}_i \mid D_i = 1] \neq E[\mathbf{X}_i \mid D_i = 0]$, and (ii) those characteristics have time-varying effects on the outcome, i.e., $\gamma_1 \neq \gamma_0$. In our application, the war in Ukraine may change how local socio-economic and demographic features shape voting behavior (i.e., a shift in γ_t), so that, if these factors are not controlled for, the refugee exposure turns out to be mechanically correlated with changes in the influence of municipal characteristics on the vote. Including the $X_i \times 1\{t = 1\}$ interaction (or more generally $X_i \times \text{year}$ interactions in multi-period settings) in Equation (2) directly absorbs $(\gamma'_1 - \gamma'_0)\mathbf{X}_i$, restoring identification of the ATT under the conditional parallel-trends assumption (CPTA). In our setting, changes in γ_t can be interpreted as the “generalized anxiety” effect related to the war or the inflow of refugees in Poland and not to contact with them, which represents the local effect.

In order to compare the effects of Ukrainian refugees with those of other refugees and immigrants — whose numbers also increased rapidly after 2022 — we re-estimated the models in Equations (1) and (2), adding on the right-hand side the 2024-2020 change in the inflow of non-Ukrainian foreign primary students interacted with year indicators.

5 Results

5.1 Main Results

In this section, we discuss the main results of our DiD analysis.

Effect of UKR refugees (and migrants). In what follows, we generally refer to Ukrainian “refugees,” although we do not have direct information on their individual legal status. However, as noted above, the special protection granted to Ukrainian refugees does not materially differentiate them from regular migrants in terms of the rights they enjoy. Table 1 reports the event-study DiD estimates for voting for the main conservative party. Column (1) displays the estimates without $X \times \text{year}$ interactions. The pre-war $\Delta \text{UKR} \times 2019$ interaction is small and close to zero; the coefficient is 0.045, implying an increase of 0.45 pp (slightly less than half a percentage point) in the vote share for the main conservative party associated with a 10-pp increase in the *future* share of Ukrainian refugees at the municipality level. Because this estimate refers to future inflows, it may reflect endogenous location choices by incoming Ukrainian refugees based on past trends in local political preferences — patterns that are not fully captured by municipality fixed effects or by county-by-year interactions included in Equation (1). The p -value for the parallel-trends assumption rejects the null of parallel trends at the 10% significance level. The coefficients on the first and second post-war interaction terms are larger and precisely estimated, at 0.20 and 0.45, respectively. The fact that the effect of refugee inflows on voting behavior strengthens between 2023 and 2025 may indicate that the true impact of

immigration on the political landscape is only now becoming apparent, and that earlier studies, which did not include the 2025 presidential election, were unable to fully capture it.

In column (2), we include the $X \times \text{year}$ interactions. These interactions address two main concerns: first, they may improve the plausibility of the parallel-trends assumption; second, they may capture variation in the impact of municipal characteristics after the war. Two findings stand out from column (2). First, the $\Delta\text{UKR} \times 2019$ coefficient is not statistically significant (the p -value for the parallel-trends test is now 0.614, thus strongly not rejecting the null). Second, some covariates — most notably the share of retired individuals and the percentage of tertiary-educated residents — exhibit coefficients that vary over time. Failing to account for this time variation may confound the estimated effect of refugee inflows with changes in voting behavior driven by observable municipality characteristics, whose effects may themselves be affected by the refugee crisis or the war. Allowing for time-varying effects of municipal characteristics substantially reduces the estimated impact of Ukrainian refugees on voting, to 0.035 and 0.078 in the two post-war elections, of which only the latter is statistically significant at the 10% level.

A potential drawback of the specification in column (2) is that including a large number of interactions between the X s and year dummies reduces degrees of freedom and estimation efficiency. Therefore, in column (3) we replace the $X \times \text{year}$ interactions with interactions between year dummies and the first two principal components (which have Eigenvalues greater than one and explaining about 60% of the total variation) obtained from a Principal Component Analysis (PCA) of the X s.⁷ This specification is more parsimonious and allows us to control for the main characteristics of municipalities, while also potentially capturing non-linear interaction effects among the X s — i.e., cases in which the combination of characteristics, rather than their isolated presence, shapes voting behavior. The PCA analysis and an interpretation of the components are reported in Table 2. Column (3) of Table 1 shows that the $C_k \times \text{year}$ interactions (where C_1 and C_2 denote the first two components) help re-establish parallel trends in the pre-war period compared to the unconditional DiD specification in the first column (the p -value is 0.533). The point estimates of the effect of refugee inflows are more precisely estimated in this specification compared to column (2) and equal to 0.058 and 0.119, statistically significant at the 10% and 5% levels, for the 2023 and 2025 elections, respectively.

Estimates in column (3) imply that a 1-pp increase in the share of Ukrainian refugees increased votes for the main conservative party by 0.12% ($= 0.01 \times 0.058 / 0.5 \times 100$) and 0.24% ($= 0.01 \times 0.119 / 0.5 \times 100$) in the first and second post-war elections, respectively. Comparing instead the difference between the 10th and the 90th percentiles of ΔUKR , rise in voting for the main conservative party is 0.51% in the 2023 election and 1.04% in the 2025 election. While these effects are modest in magnitude, they are in line with those reported in the literature (Cools et al., 2021); they suggest that the local exposure to Ukrainian refugees contributed to increased support for a party that had been losing votes over time. Indeed, average municipal vote shares for the main conservative party declined from 0.54 in the 2019 and 2020 elections to 0.45 and 0.37 in the two subsequent elections. The coefficients on the $C_k \times \text{year}$ interaction terms also reveal that after 2020, support for the conservative party strengthened relatively in urban areas with higher levels of education. This may reflect a tendency toward polarization on the right side of

⁷ Point estimates of the $\Delta\text{UKR} \times \text{year}$ interactions are very similar in magnitude when we include the first three PCA components — which explain 79% of the total variation — interacted with year indicators, and retain the same levels of statistical significance.

the political spectrum in response to new challenges associated with immigrant inflows.

Table 3 reports the results of the regressions with voting for the alt-right. First, comparing columns (1) and (2), we note that controlling for the $X \times \text{year}$ interactions ensures the parallel trends in the pre-war period. As noted in the results for the main conservative party, the covariates exhibit time-varying effects. Also in this case, allowing for time-varying effects of municipality characteristics strongly reduces the coefficients on UKR refugee inflows, which nevertheless remain negative. Interestingly, in column (3), replacing the $X \times \text{year}$ interactions with time-varying coefficients of the first two PCA's components increases precision, and the coefficient on the 2025 election becomes statistically significant at the 5% level, while the p -value for the parallel trends test rejects diverging trends before the war. The estimated effect for the 2025 election entails a 1.06% decrease ($= 0.01 \times (-0.0946/0.0896) \times 100$) for a 1-pp increase in the share of UKR primary students.

Figures A2 and A3 in the Appendix plot the DiD coefficients for voting for the main conservative party and alt-right parties, respectively. In the bottom part of each figure, we report some robustness tests suggested by Rambachan and Roth (2023) to assess whether our estimates are sensitive to deviations from the parallel-trends assumption. Panels (b) and (c), in both figures, report the original estimates with confidence intervals and the confidence intervals when one assumes that violations of the parallel trends in the post-treatment period cannot be larger than 0.5, 1, 1.5 and 2 times that observed in the pre-treatment period. As for conservative voting, both the first and the second election effects are quite robust to violations, remaining significant up to violations that are 1.5 times those observed before the start of the conflict. As for alt-right voting, the second election effect is similarly robust.

Against the backdrop of a general increase in support for alt-right parties in Poland in recent years, the negative relationship between refugee inflows into a municipality and changes in support for the alt-right may appear counterintuitive, although it is consistent with the findings of some earlier studies (Mykhailyshyna and Zuchowski, 2025). At the same time, however, the simultaneous decline in support for far-right formations and the consolidation of the main conservative party in response to the presence of refugees constitutes a novel finding that calls for further interpretation.

One possible explanation for this pattern is the partial appropriation by the mainstream conservative party of a radical discourse previously characteristic of the alt-right. This radicalization of campaign rhetoric was particularly visible during the most recent presidential election (2025), which was not covered in the earlier studies discussed in this article. At present — a few months after the conclusion of the electoral campaign — no scholarly studies have yet addressed the question of radicalization; however, it has been noted by organizations monitoring media content (CEDMO, 2025). To further explore this issue, we employed an experiment using the ChatGPT language model. We deliberately selected this particular tool, as previous studies have demonstrated its strong potential for accurately characterizing political parties (Bol and Bono, 2025). Our approach was inspired by the CHES survey cited earlier in this article, with the difference that expert opinions were replaced by an AI-based analysis of campaign content. We asked the ChatGPT model (version 5.2) to evaluate, on a scale from 0 to 10, the presence of national-populist elements in the campaigns of all candidates in the Polish presidential elections of 2020 and 2025.⁸ For the candidate of the mainstream conservative party (Law

⁸ The precise wording of the prompt (translated from Polish) was: "Evaluate the election campaign of all registered candidates for the President of Poland in 2020 and 2025 in terms of populist-nationalist

and Justice, PiS), the 2020 campaign received a score of 6, while the 2025 campaign was rated at 8, indicating a substantial increase in national-populist content. By contrast, the main challenger from the liberal party (Civic Platform, PO) received a score of 2 in both elections. Notably, the score of 8 assigned to the conservative candidate in 2025 was equal to that of the most radical alt-right candidate in 2020 (Krzysztof Bosak). These results lend support to the claim that, after 2020, the mainstream conservative party significantly radicalized its messaging by incorporating elements of discourse previously associated with the alt-right.

A complementary explanation points, however, to differences in the mechanisms underlying electoral support for conservative parties and the alt-right. Previous research across several countries suggests that far-right voters attach considerable importance to socio-cultural change; however, compared to supporters of mainstream parties, their political behavior is more often driven by fear of change rather than by direct experience of adverse change (McManus, 2021; Im et al., 2022). Therefore, support for the radical right should not be interpreted as a reaction to pauperization or political transformation, but rather as a predominantly ideological choice (Ortiz Barquero et al., 2021). The predominance of ideology-driven voting over experience-based voting may therefore constitute a key criterion distinguishing conservative voters from supporters of anti-system, far-right parties. In the context of our research, the response of the mainstream conservative party to the refugee crisis in Poland has taken the form of a more nationalist electoral message, which was not successful at national scale (as conservatives' share of votes has decreased), but proved to be relatively more effective in areas where the presence of refugees is particularly visible. The possible reason is that conservative voters, to a greater extent than far-right voters, tend to rely on their own experiences rather than on a general ideological line. This mechanism may also help explain why support for the alt-right has increased overall, yet not necessarily in those localities where a heightened inflow of Ukrainian refugees is actually observed. A positive implication of this interpretation is that the perceived presence of refugees does not directly induce radical voting behavior. On the contrary, it may even counteract it. What fosters radicalization is general societal anxiety and susceptibility to populist narratives, which do not necessarily arise from actual exposure to the presumed threat. Importantly in this context, the results of specifications (2) and (3) in Tables 1 and 3 indicate that, following the refugees crisis, urban municipalities with more highly educated populations tended to increase their support for mainstream conservative parties and not for the radical right.

Importantly, the significant positive effect of migration on support for moderately right-wing parties, coupled with the absence of any effect on the electoral performance of more radical groupings, is a finding that has previously been observed in studies of countries other than Poland. Such a pattern was noted, among others, by Dijkstra et al. (2020) in their analysis of voting for pro- and anti-European parties across the European Union as a whole.

Effect of non-UKR migrants. Although our analysis has so far focused on the effect of UKR refugees on voting outcomes, the period after 2020 also saw a sustained increase in immigrants from other origins. Thus, we re-estimated Equation (1) adding also the change in the share of non-UKR immigrants (between 2020 and 2024) with year interactions,

rhetoric. Ratings should be given on a scale from 0 to 10, where 0 indicates a complete absence of populist-nationalist content, and 10 indicates that populist-nationalist content constituted the main element of the candidate's message."

which allows us: (i) to be sure that the UKR refugees in the estimates that we have just commented on were not capturing the general effects of immigration, although we have shown in Section 3 that the two types of migration flows have distinct location patterns; (ii) to contrast the effects of UKR refugees with those of other foreign immigrants.

Table 4 shows the estimates for the effect of UKR refugees on conservative (column (1)) and alt-right (column (2)) voting, obtained from the specification with $C \times \text{year}$ interactions. In both columns, the CPTA holds. Moreover, the $\Delta \text{UKR} \times \text{year}$ point estimates are statistically indistinguishable from those earlier commented on for Tables 1 and 3. Interestingly, the share of non-UKR foreign primary students is never statistically significant at conventional levels.

The absence of a statistically significant effect of the change in the student share coming from foreign countries other than Ukraine on local support for the conservative party and alt-right parties may be attributable to the relatively small scale of these inflows, especially compared to the much larger wave of arrivals from Ukraine. It is important to emphasize that the absence of statistically significant interactions for non-Ukrainian foreign students indicates that their correlation with support for the conservative or alternative-right parties did not change significantly relative to the pre-war elections (2019). One possible explanation is that, although the stock of these students increased substantially in percentage terms, their share of the total student population remained small. Although there is no evidence of a universal minimum percentage “threshold” of immigrant presence that systematically triggers electoral reactions, some empirical analyses lend support to the existence of threshold-type mechanisms. A study from the United States suggests that the threshold share of non-citizen immigrants (measured at the state level) beyond which support for the Republican Party begins to increase is approximately 13.2 percent. Once the share of non-citizen immigrants in a state exceeds this level, the marginal effect of additional non-citizen immigrants on the Republican vote share becomes positive, as immigration emerges as a salient political issue for voters (Mayda et al., 2016). By contrast, evidence from Italy — arguably more relevant to the present context due to geographic proximity and the use of municipal-level data — indicates that such a threshold may be considerably lower, at around 3.3 percent (Bordignon et al., 2019). If the latter estimate is applied mechanically to the situation in Polish municipalities in 2025, the 3.3 percent threshold with respect to Ukrainian pupils in primary schools is exceeded in 20.3 percent of municipalities, whereas the same threshold for immigrant pupils from other countries is reached in only 0.6 percent of municipalities. Thus, despite the growing inflow of immigrants from countries other than Ukraine in recent years, their numbers may still be too small to trigger a meaningful electoral response or to exert a distinct, discernible influence on the configuration of Poland’s political landscape.

Comparison with the past literature. As noted above, our estimated effects are modest in magnitude but broadly consistent with those reported in previous studies, particularly those relying on DiD designs and similarly defined treatment variables (e.g., refugee arrivals per capita or changes in the refugee share of the population). Dinas et al. (2019) report a 0.60 pp increase in the vote share of Golden Dawn in Greek national elections produced by one refugee arrival per capita (equivalent to 0.006 percentage points for a 0.01 increase in per-capita arrivals). Similarly, Edo et al. (2019), using a first-difference specification for French presidential elections, find a 0.4 pp increase in far-right voting following a 1-pp rise in the immigrant share. Our estimated effect on support for the main conservative party falls within this range, while we instead find negative effects on

the vote share of far-right parties. This pattern is noteworthy, as it suggests that the electoral consequences of the arrival of selected refugees granted Temporary Protection in Poland were not fundamentally different in magnitude from those observed in contexts characterized by non-selective refugee inflows or more “open migration” regimes.

Several important differences distinguish our analysis from previous studies on the impact of Ukrainian refugees on voting in Poland — most notably the use of a different identification strategy (DiD instead of IVs), a different measure of local refugee presence, and substantially higher data granularity (municipalities instead of counties). Because of these discrepancies, our results cannot be straightforwardly compared to those in [Mykhailyshyna and Zuchowski \(2025\)](#) or [Clifton-Sprigg et al. \(2025\)](#). Nevertheless, we summarize here the key differences in findings.

[Mykhailyshyna and Zuchowski \(2025\)](#) estimate IVs regressions for each parliamentary election from 2015 to 2023. Focusing on the first post-war parliamentary election (2023), the authors do not find any statistically significant effect of UKR refugees on voting for conservative parties (although the point estimates are negative), but they do find a large negative effect on voting for the far-right when using the Euro 2014 and border instruments. A negative effect of labor immigrants on far-right voting is also found when each of the three instruments is used individually (including the historical instrument based on “Akcja Wisła,” a military operation that led to the resettlement of about 140,000 ethnic Ukrainians in Poland in 1947). Comparability with our results is further complicated by the fact that the set of compliers associated with some of their instruments is hard to characterize. Nonetheless, we simply note here that despite all the above differences, both papers find evidence that Ukrainian migrants led to a reduction in far-right (alt-right in our case) support.

[Clifton-Sprigg et al. \(2025\)](#) focus on the 2019 and 2023 parliamentary elections and use IVs with a shift-share instrument based on past UKR settlement patterns measured in the 2002 Polish Population Census. They find positive effects of UKR refugees on the change of votes for both far-right and right-wing parties. So, the main difference with their results is that we find instead negative effects on voting for the alt-right. In their case, the use of a shift-share instrument makes it easier to infer who the compliers may be — namely, counties where new refugees followed earlier Ukrainian settlement patterns and where Ukrainian communities were already well established. Thus, the differences may stem from the Local Average Treatment Effect (LATE) interpretation of their estimates. In addition, the positive effect they find for the far right may reflect a “saturation” mechanism whereby a backlash emerges once the stock of refugees and migrants reaches a critical threshold.

5.2 Robustness Checks

This section presents robustness checks to assess the sensitivity of our results.

Spatial correlation. Until now, our inference only accounts for serial correlation in the errors by municipality, but not for spatial correlation. For this reason, in [Table A2](#) we computed Conley standard errors ([Conley, 1999](#)) for the estimates in [column \(3\)](#) of [Table 1](#) and [Table 3](#) assuming different distance cutoffs for the spatial correlation (25, 50, 75 and 100 km). Accounting for spatial correlation does not substantially change the significance levels of our estimated effects.

Proxies of Ukrainian refugees and immigrants. Although we have motivated our focus on the share of primary school students in the main analysis, one could argue that the gain in precision from better measuring refugees’ locations comes at a cost. In particular, primary school students: (i) represent only a subset of the Ukrainian migrant/refugee population; and (ii) may have effects on voting that do not generalize to other refugee subpopulations, mainly capturing the “competition for amenities” mechanism.

To address this potential concern, we proceed as follows. For the years around 2019 and 2020, information on the Ukrainian population is available only at the county level and not at the municipal level. We therefore aggregate the data to the county level, weighting municipalities by their 2015 populations, and estimate the following *cross-sectional* regressions:

$$y_{ck} = \alpha_0 + \alpha_1 \text{UKR}_c + \alpha'_2 \mathbf{X}_c + D_r + \epsilon_{ck}, \quad (4)$$

where $k \in \{2019, 2020\}$ denotes the election year; y_{ck} is the vote share of the main conservative party or the alt-right in county c ; D_r are region (Nomenclature of Territorial Units for Statistics Level 2 units, i.e., NUTS-2) fixed effects; and $\mathbf{X}_c$ is a vector of county-level control variables including population size, population density, municipality own-revenue, and the share of the population of retirement age (all measured in 2015 and provided by GUS), as well as the share of university-educated individuals (from the 2002 Population Census) and the share of dominicantes (measured in 2019). The term ϵ_{ck} denotes the error term.

We estimate alternative specifications in which UKR_c is measured either as the share of Ukrainian refugees/migrants in the total population in 2021 (source: Population Census) or as the share of Ukrainian primary school students among all primary school students in 2020 (source: SIO).

The purpose of these regressions is not to estimate causal effects, but rather to compare the magnitude of the conditional correlations between voting outcomes and the two alternative proxies for migrant presence. Since both measures are likely subject to similar sources of endogeneity bias, contrasting the corresponding OLS estimates provides a rough indication of the extent to which relying exclusively on primary school student data may understate the overall association between Ukrainian migration and voting behavior.

The estimates reported in Table A3 reveal several interesting patterns. For conservative voting (columns (1)-(4)), the correlation has the same sign regardless of the proxy used. In both cases, prior to the war, the county-level share of UKR migrants was negatively correlated with support for the main conservative party. The coefficient based on the UKR population is slightly larger and therefore more precisely estimated, although the standard errors are very similar across the two specifications. In terms of magnitude, however, the estimates using the different proxies are very close and statistically indistinguishable.

Columns (5)–(8) reveal the opposite pattern for alt-right voting. Here, before the war, UKR migrants were consistently positively correlated with support for alt-right parties. Notably, the coefficient associated with the share of UKR primary school students is larger and statistically significant at the 5% level.

Thus, from a purely correlational perspective, our proxy for UKR migrant presence in Polish counties based on primary school students closely mirrors the behavior of alternative measures based on the share of UKR immigrants or refugees in the population, similar to those previously used in county-level analyses (Clifton-Sprigg et al., 2025; Mykhailyshyna and Zuchowski, 2025).

Overall, these results indicate stronger support before the war for the alt-right in municipalities with a larger presence of Ukrainian refugees, suggesting that the main conservative parties/conservative candidates may have had incentives to shift the political discourse toward immigration in the following elections in order to attract segments of the anti-immigration electorate.

It is worth emphasizing that the signs of the cross-sectional correlations for the pre-war period reported in Table A3 — positive for the alt-right and negative for the main conservative party — are not inconsistent with the findings from our DiD analysis. Our panel-based DiD estimates, which exploit within-municipality variation across elections, indicate that larger inflows of Ukrainian students at the municipality level progressively shifted electoral support from the alternative right toward the main conservative party.

Sensitivity to excluding major destinations of Ukrainian students. We conduct a sensitivity analysis excluding the top 5%, 10%, and 15% of municipalities ranked by the increase in the share of Ukrainian primary school students (ΔUKR). The results, reported in Table A4, are highly consistent with our baseline estimates. As expected, excluding the most affected municipalities reduces the sample size and leads to some loss of precision, particularly for the estimated effect on conservative voting in the 2023 election. Overall, the findings suggest that our results are not driven by a small number of municipalities that served as primary hubs for Ukrainian migration.

Randomization inference. As noted above, changes in voting behavior at the municipality level may reflect broader country-wide dynamics — such as generalized anxiety toward refugees or shifts in the national political climate — rather than being primarily driven by local exposure to Ukrainian refugees. This raises concerns that our estimated effects may partly capture spurious correlations induced by common trends rather than a causal impact of the local refugee shock.

To address this concern and to assess whether our baseline estimates are distinguishable from what would arise under random assignment, we conduct a randomization inference exercise. Specifically, we randomly reassign the municipality-level change in Ukrainian refugees, ΔUKR , across municipalities and re-estimate the DiD models while controlling for the first two PCA principal components derived from pre-treatment characteristics. Repeating this procedure generates a reference distribution of placebo coefficients that captures the magnitude of effects that could emerge purely by chance. The resulting distributions of coefficients are shown in Figure A4 for conservative voting and Figure A5 for alt-right voting. In both cases, the coefficients obtained in our baseline specifications, indicated by the red vertical line, generally lie far in the tails of the randomization distributions and are substantially larger in absolute value than the vast majority of placebo estimates. The only exception is the interaction term $\Delta\text{UKR} \times 2023$ in the alt-right voting specification, which is also statistically insignificant in the baseline results. Overall, this exercise provides additional support for the interpretation that our main findings are unlikely to be driven by chance assignment or nationwide political shifts alone and instead reflect a meaningful shift in municipalities’ political outcomes induced by the local inflow of Ukrainian refugees.

Recomposition of votes within the right-wing spectrum. Table A5 reports the effects of the inflow of Ukrainian refugees across the full party spectrum: the main conservative party, the alt-right, and a residual category comprising liberal and centrist

parties (e.g., the Civic Coalition and Third Way) as well as left-wing parties (*Lewica*). The results indicate that the electoral gains experienced by conservatives in response to local refugee inflows occurred primarily at the expense of far-right parties. Although the main conservative party also appears to have drawn votes away from centrist/liberal and left-wing parties — whose platforms and campaign rhetoric were largely neutral or supportive of Ukrainian immigration — the corresponding effects are small in magnitude and statistically insignificant.

6 Concluding Remarks

In this study, we analyze the effect of the war-induced inflow of refugees from Ukraine to Poland on the voting behavior of the Polish electorate. We seek to assess whether this sudden immigration shock contributed to the radicalization of the political landscape and to a shift in voter support toward right-wing parties.

To better understand the geography and nature of the potential political backlash, we rely on data at the municipal level, which allow for a spatially disaggregated analysis of both voting outcomes and immigrant presence. As a measure of refugee settlement, we use administrative data on the enrollment of Ukrainian students in Polish public schools. This approach enables us to capture refugees' direct use of public services — and the potential competition for such services that the literature identifies as a key driver of anti-immigration sentiment.

Using a quasi-experimental design, our municipality-level event-study difference-in-differences analysis suggests that the recent inflow of Ukrainian refugees and migrants contributed to a reconfiguration of local electoral support. Rather than fueling support for more radical formations, refugee exposure was associated with greater backing for the mainstream conservative party and a decline in support for alternative-right parties. Taken together, the evidence points to a reallocation of support within the right-wing spectrum, rather than a straightforward process of radicalization. The increase in local support for a conservative party in municipalities that experienced a larger influx of Ukrainian refugees may partly reflect that party's strategic intensification of nationalist rhetoric, which facilitated the direct capture of potential alt-right voters. At the same time, the pattern may be rooted in a more fundamental mechanism identified in the political science literature: compared with supporters of more moderate parties, alt-right voters rely less on personal experience (e.g., contact with migrants) and more on generalized beliefs when making electoral decisions. This interpretation helps explain the Polish case, where support for the radical right — driven by generalized fear — coexists with a negative correlation between such support and the actual local presence of refugees.

The complex relationship between immigration and voting behavior complicates the formulation of unambiguous policy recommendations for political leaders and educators. Consistent with earlier studies and corroborated by our findings, increased support for the radical right appears to be driven primarily by generalized anxiety rather than by direct personal experience with immigrants. In this context, a combination of initiatives aimed at improving general knowledge about other nations, together with programs fostering sustained and meaningful interactions between immigrants and natives, appears warranted — both within school curricula and in adult education. Such an approach may be effective not only in the context of economic immigration from culturally distant regions but also in cases involving refugees from neighboring countries. Our results

indicate that even when refugees are culturally proximate to the native population and their experiences initially elicit genuine compassion and a sincere willingness to provide assistance, their presence may, in the longer term, lead voters to shift toward conservative candidates, though not necessarily toward those espousing the most radical positions.

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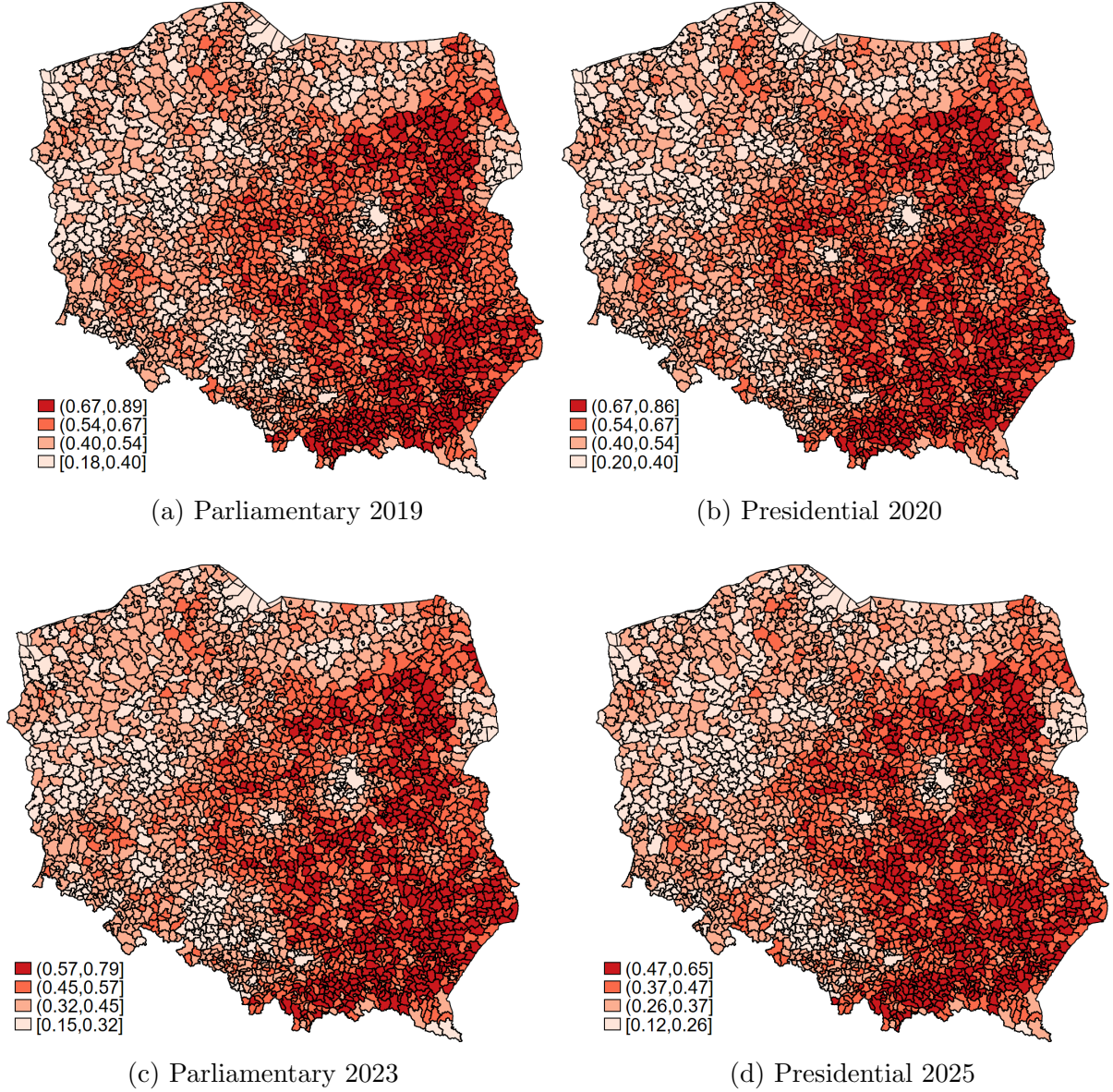
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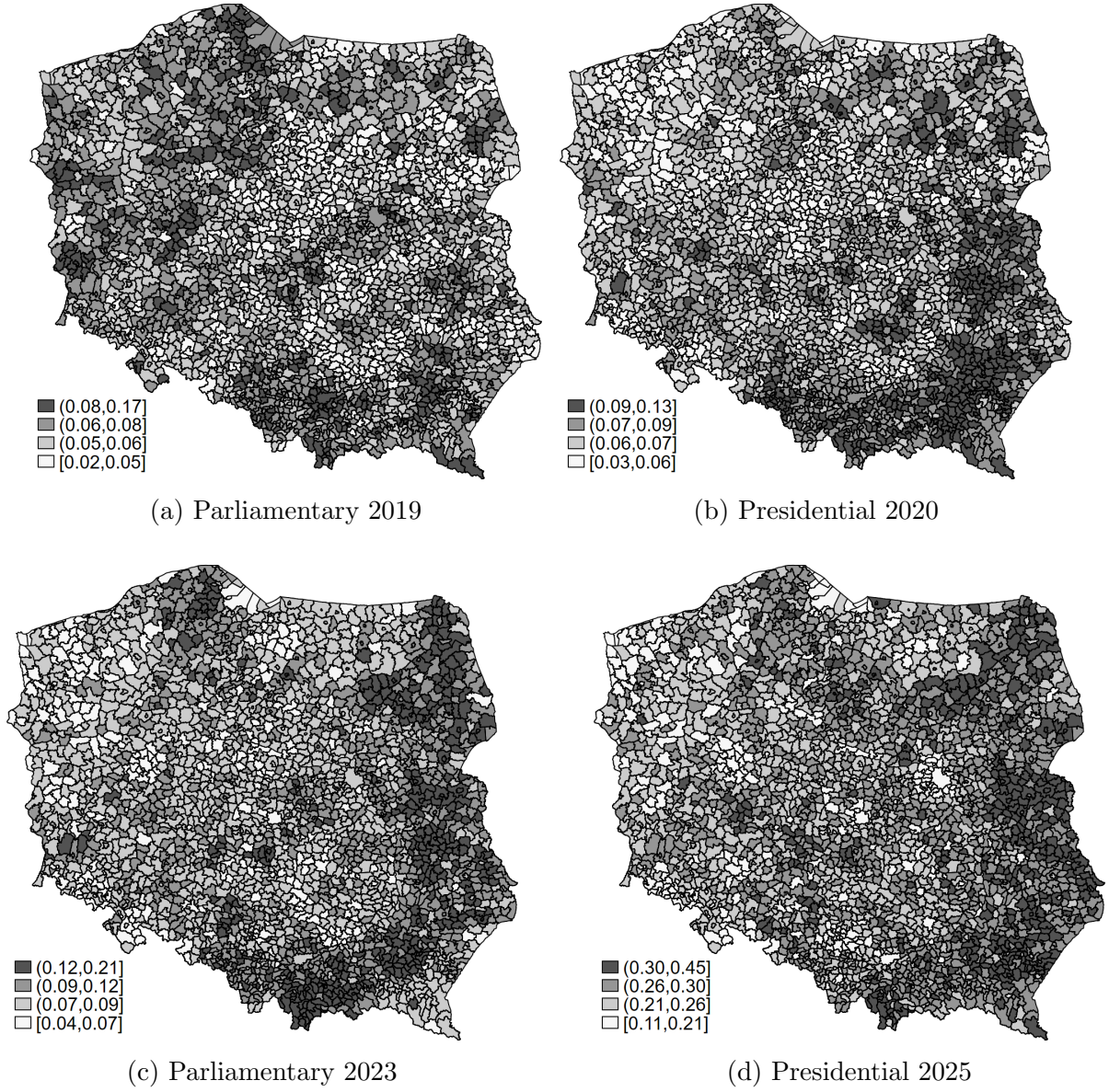
Figures and Tables

Figure 1: Share of votes on conservative party/candidate



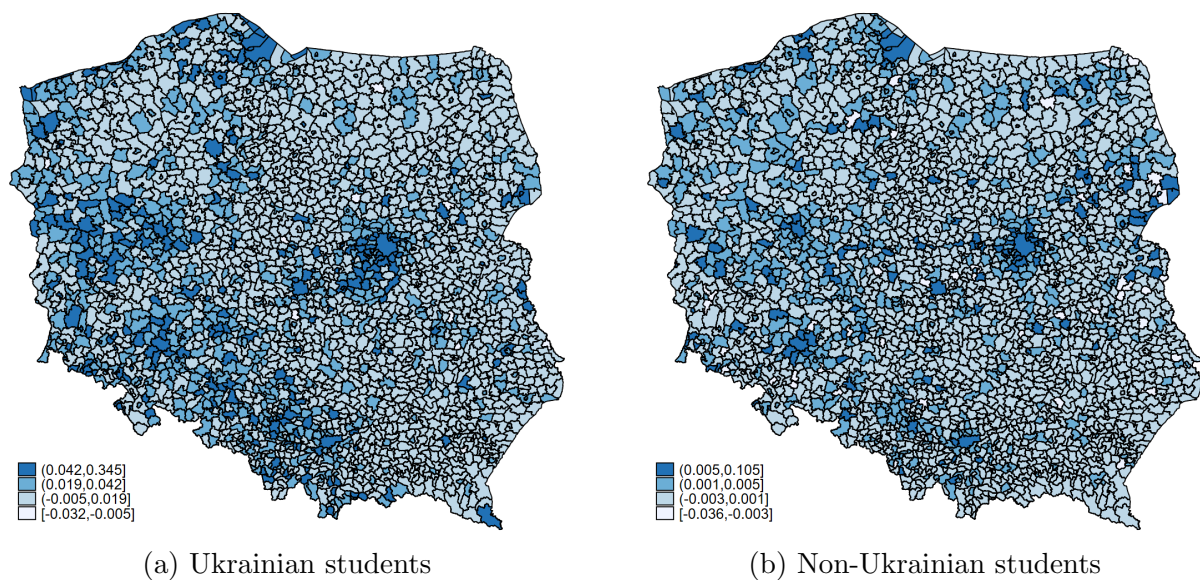
Note. Panels (a) and (c) report the share of votes cast for the Law and Justice (PiS) party in the first round of the parliamentary election (respectively in 2019 and 2023). Panel (b) reports the share of votes cast for Andrzej Duda in the first round of the presidential election in 2020. Panel (d) reports the share of votes cast for Karol Nawrocki in the first round of the presidential election in 2025.

Figure 2: Share of votes on alt-right parties/candidate



Note. Panel (a) reports the joint share of votes cast for “Konfederacja” and “Prawica” in the first round of the parliamentary election in 2019. Panel (b) reports the joint share of votes cast for Krzysztof Bosak, Stanisław Żółtek, Mirosław Piotrowski, and Marek Jakubiak in the first round of the presidential election in 2020. Panel (c) reports the joint share of votes cast for “Konfederacja” and “Polska jest Jedna” in the first round of the parliamentary election in 2023. Panel (d) reports the joint share of votes cast for Artur Bartoszewicz, Grzegorz Braun, Maciej Maciak, and Sławomir Mentzen in the first round of the presidential election in 2025.

Figure 3: Change in the share of foreign students in Poland's primary schools between 2020 and 2024



Note. Panel (a) reports the change (in percentage points) in the share of Ukrainian students in Poland's primary schools between September of 2020 and September of 2024. Panel (b) reports similar change for foreign students of different (non-Ukrainian) nationality.

Table 1: DiD estimates for main conservative voting

VARIABLES	(1)	(2)	(3)
$\Delta \text{UKR} \times 2019$	0.0454* (0.0267)	0.0138 (0.0274)	0.0183 (0.0294)
$\Delta \text{UKR} \times 2023$	0.200*** (0.0392)	0.0351 (0.0288)	0.0582* (0.0304)
$\Delta \text{UKR} \times 2025$	0.453*** (0.0746)	0.0782* (0.0419)	0.119** (0.0487)
Population density $\times 2019$		4.48e-05 (0.000123)	
Population density $\times 2023$		0.000114 (0.000121)	
Population density $\times 2025$		0.000217 (0.000174)	
Municipality's own revenue $\times 2019$		0.00104 (0.000857)	
Municipality's own revenue $\times 2023$		0.000941 (0.000999)	
Municipality's own revenue $\times 2025$		0.00147 (0.00128)	
Share retirement age $\times 2019$		-0.154*** (0.0246)	
Share retirement age $\times 2023$		0.142*** (0.0220)	
Share retirement age $\times 2025$		0.0520* (0.0287)	
Share dominicantes $\times 2019$		0.00881 (0.00588)	
Share dominicantes $\times 2023$		0.00435 (0.00509)	
Share dominicantes $\times 2025$		-0.00732 (0.00545)	
Share university educated $\times 2019$		0.128*** (0.0246)	
Share university educated $\times 2023$		0.238*** (0.0244)	
Share university educated $\times 2025$		0.707*** (0.0410)	
$C_1 \times 2019$			0.00212** (0.000912)
$C_1 \times 2023$			0.00904*** (0.000728)
$C_1 \times 2025$			0.0217*** (0.00170)
$C_2 \times 2019$			-0.00349** (0.00167)
$C_2 \times 2023$			0.00158 (0.00108)
$C_2 \times 2025$			-0.000277 (0.00334)
Observations	9,644	9,644	9,644
R-squared	0.991	0.993	0.993
Municipal FE	Yes	Yes	Yes
County*year interactions	Yes	Yes	Yes
X*year interactions	No	Yes	No
PCA*year interactions	No	No	Yes
Mean vote share before the war	0.540	0.540	0.540
Parallel-trends F -stat	2.894	0.255	0.389
Parallel-trends p -value	0.0890	0.614	0.533

Note. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors (in parentheses) are clustered at the municipality level. Own revenue is in thousands of Polish zloty. The first two PCA components (C_1 and C_2) components are standardized (mean=0, standard deviation=1).

Table 2: Principal Components Analysis (PCA): Eigenvectors

VARIABLES	Comp1	Comp2	Comp3	Comp4	Comp5
Pop. density	0.5838	0.1478	0.2324	-0.3967	0.6527
Municipality's own revenue	0.3548	-0.6653	0.0170	0.6235	0.2061
Share retired pop.	0.2982	0.7179	-0.1863	0.6008	0.0022
Share dominicantes	-0.2564	0.1331	0.9226	0.2543	0.0253
Share university educated	0.6153	-0.0491	0.2445	-0.1683	-0.7286

Note. Components identified by the PCA. Only the first two components are used in the analysis. Component 1 (Comp1) measures “Urban Socio-Economic Development.” Component 1 captures a modern, urban, economically stronger. and more educated environment. Municipalities scoring high on this component tend to be: more densely populated, more educated, fiscally stronger, less religious, slightly older. This is a classic urban–rural socio-economic gradient. High values = “urban, affluent, educated”. Low values = “rural, traditional, lower education”. Component 2 (Comp2) measures “Traditionalism vs. Welfare/Retirement Structure.” Component 2 contrasts: Positive side: municipalities with many retirees, somewhat more traditional social structure, not strongly urban or educated. Negative side: municipalities with higher fiscal revenue, more economically dynamic, fewer retirees. This second component is not about education or urbanization (those belong to Comp1). Instead, it reflects a demographic–fiscal structure: High values = aging, retirement-heavy, fiscally weaker communities. Low values = younger, fiscally stronger municipalities. One can think of it as a “retirement burden vs. fiscal capacity” axis.

Table 3: DiD estimates for alt-right voting

VARIABLES	(1)	(2)	(3)
$\Delta \text{UKR} \times 2019$	0.0373*** (0.0142)	1.18e-05 (0.0137)	0.00241 (0.0137)
$\Delta \text{UKR} \times 2023$	-0.102*** (0.0201)	-0.00985 (0.0138)	-0.0136 (0.0144)
$\Delta \text{UKR} \times 2025$	-0.512*** (0.0788)	-0.0523 (0.0333)	-0.0946** (0.0393)
Population density $\times 2019$		-6.13e-05 (6.26e-05)	
Population density $\times 2023$		-0.000239*** (6.90e-05)	
Population density $\times 2025$		-0.000733*** (0.000194)	
Municipality's own revenue $\times 2019$		-3.73e-06 (0.000184)	
Municipality's own revenue $\times 2023$		-0.000225 (0.000331)	
Municipality's own revenue $\times 2025$		-0.000725 (0.000712)	
Share retirement age $\times 2019$		0.0216** (0.0106)	
Share retirement age $\times 2023$		-0.00788 (0.0164)	
Share retirement age $\times 2025$		-0.139*** (0.0358)	
Share dominicantes $\times 2019$		-0.00951*** (0.00232)	
Share dominicantes $\times 2023$		0.00845*** (0.00296)	
Share dominicantes $\times 2025$		0.00901* (0.00516)	
Share university educated $\times 2019$		0.0705*** (0.0118)	
Share university educated $\times 2023$		-0.155*** (0.0146)	
Share university educated $\times 2025$		-0.795*** (0.0410)	
$C_1 \times 2019$			0.00224*** (0.000351)
$C_1 \times 2023$			-0.00571*** (0.000502)
$C_1 \times 2025$			-0.0269*** (0.00152)
$C_2 \times 2019$			0.000250 (0.000421)
$C_2 \times 2023$			-0.000169 (0.000924)
$C_2 \times 2025$			-0.00257 (0.00294)
Observations	9,644	9,644	9,644
R-squared	0.978	0.988	0.986
Municipal FE	Yes	Yes	Yes
County*year interactions	Yes	Yes	Yes
X*year interactions	No	Yes	No
PCA*year interactions	No	No	Yes
Mean vote share before the war	0.0672	0.0672	0.0672
Parallel-trends F -stat	6.885	7.33e-07	0.0310
Parallel-trends p -value	0.00875	0.999	0.860

Note. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors (in parentheses) are clustered at the municipality level. Own revenue is in thousands of Polish zloty. The first two PCA components (C_1 and C_2) are standardized (mean=0, standard deviation=1).

Table 4: DiD estimates for main conservative and alt-right voting, including non-UKR foreign primary students

VARIABLES	(1) Conservative	(2) Alt-right
$\Delta\text{UKR} \times 2019$	0.0189 (0.0294)	0.00226 (0.0137)
$\Delta\text{UKR} \times 2023$	0.0577* (0.0306)	-0.0152 (0.0144)
$\Delta\text{UKR} \times 2025$	0.122** (0.0489)	-0.0992** (0.0394)
$\Delta\text{non-UKR} \times 2019$	-0.0704 (0.146)	0.0169 (0.0454)
$\Delta\text{non-UKR} \times 2023$	0.0562 (0.142)	0.179 (0.139)
$\Delta\text{non-UKR} \times 2019$	-0.251 (0.256)	0.530 (0.376)
Observations	9,644	9,644
R-squared	0.993	0.986
Municipal FE	Yes	Yes
County*year interactions	Yes	Yes
X*year interactions	No	No
PCA*year interactions	Yes	Yes
Mean vote share before the war	0.540	0.0672
Parallel-trends F -stat	0.414	0.0273
Parallel-trends p -value	0.520	0.869

Note. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors (in parentheses) are clustered at the municipality level.

A Appendix

Table A1: Sample summary statistics

	N. obs.	Mean	Std. Dev.
Share of votes for main conservative party (2019-2025 elections)	9644	0.476	0.144
Share of votes for alt-right parties (2019-2025 elections)	9644	0.122	0.084
Share of votes for left, centrist/liberal and other parties (2019-2025 elections)	9644	0.401	0.136
Change in the share of Ukrainian students in primary schools (2020-2024)	9644	0.018	0.023
Change in the share of non-ukrainian immigrant students in primary schools (2020-2024)	9644	0.001	0.004
Population density per km2 (2015)	9644	1.812	3.877
Municipal own revenues per capita in thousands of Polish zloty (2015)	9644	1.545	1.214
Share of population of post-retirement age (2015)	9644	0.181	0.030
Share of registered catholics attending Sunday mass (2019)	9644	0.412	0.149
Share of population with a higher education degree (2002)	9644	0.050	0.027

Note. Estimation sample summary statistics. See Section 3 for a detailed description of the variables and the data sources.

Table A2: Coefficients and Conley standard errors by distance cutoff (km)

	(1) coef.	(2) s.e. (25-km)	(3) s.e. (50-km)	(4) s.e. (75-km)	(5) s.e. (100-km)
<i>(a) Voting for the main conservative party</i>					
$\Delta\text{UKR} \times 2019$	0.018	0.031	0.034	0.035	0.035
$\Delta\text{UKR} \times 2023$	0.058	0.026	0.028	0.029	0.028
$\Delta\text{UKR} \times 2025$	0.119	0.040	0.044	0.046	0.050
<i>(b) Voting for the alt-right parties</i>					
$\Delta\text{UKR} \times 2019$	0.002	0.026	0.028	0.030	0.031
$\Delta\text{UKR} \times 2023$	-0.014	0.017	0.018	0.017	0.017
$\Delta\text{UKR} \times 2025$	-0.095	0.040	0.044	0.045	0.048

Note. This table reports our main estimates in column (3) of Table 1 and column (3) of Table 3 with standard errors (s.e.) corrected to account for spatial correlation, assuming different distance cutoffs (25-km, 50-km, 75-km and 100-km). Standard errors have been computed using “arbitrary correlation regression” (Stata command `acreg`; [Colella et al., 2023](#)).

Table A3: Correlation between voting and different proxies of UKR immigrants/refugees

VARIABLES	(1) Cons. (2020)	(2) Cons. (2019)	(3) Cons. (2020)	(4) Cons. (2019)	(5) Alt-right (2020)	(6) Alt-right (2019)	(7) Alt-right (2020)	(8) Alt-right (2019)
share of UKR population (2021)	-1.457*** (0.558)	-1.483** (0.590)			0.0636 (0.0872)	0.158 (0.123)		
share of UKR primary school students (2020)			-1.023* (0.555)	-0.947 (0.598)			0.159** (0.0729)	0.272** (0.118)
Observations	380	380	380	380	380	380	380	380
R-squared	0.832	0.824	0.830	0.822	0.656	0.409	0.659	0.416
County FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Control variables	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Mean dep. var.	0.479	0.480	0.479	0.480	0.0719	0.0647	0.0719	0.0647

Note. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors (in parentheses) are robust to heteroskedasticity. This table reports the *cross-sectional* estimates of Equation (4). The column headings report the outcome variable (i.e., voting for a specific party or party group and the year of the election). The analysis is done on county-level data.

Table A4: Sensitivity to excluding major destinations of Ukrainian students

VARIABLES	(1)	(2)	(3)
	Excluding top		
	5%	10%	15%
of UKR student destinations			
<i>(a) Voting for the main conservative party</i>			
$\Delta \text{UKR} \times 2019$	0.0495 (0.0476)	0.0342 (0.0574)	0.0457 (0.0671)
$\Delta \text{UKR} \times 2023$	0.113*** (0.0397)	0.0979** (0.0480)	0.0919 (0.0562)
$\Delta \text{UKR} \times 2025$	0.274*** (0.0570)	0.300*** (0.0667)	0.338*** (0.0776)
Observations	9,256	8,788	8,312
R-squared	0.993	0.992	0.992
Municipal FE	Yes	Yes	Yes
County*year interactions	Yes	Yes	Yes
X*year interactions	No	No	No
PCA*year interactions	Yes	Yes	Yes
Mean vote share	0.546	0.553	0.560
Parallel-trends F -stat	1.082	0.354	0.463
Parallel-trends p -value	0.298	0.552	0.496
<i>(b) Voting for the alt-right parties</i>			
$\Delta \text{UKR} \times 2019$	-0.000955 (0.0197)	-0.00791 (0.0244)	0.00348 (0.0281)
$\Delta \text{UKR} \times 2023$	-0.0651** (0.0271)	-0.0627* (0.0324)	-0.0723** (0.0362)
$\Delta \text{UKR} \times 2025$	-0.210*** (0.0548)	-0.235*** (0.0654)	-0.241*** (0.0755)
Observations	9,256	8,788	8,312
R-squared	0.987	0.987	0.987
Municipal FE	Yes	Yes	Yes
County*year interactions	Yes	Yes	Yes
X*year interactions	No	No	No
PCA*year interactions	Yes	Yes	Yes
Mean vote share	0.0672	0.0672	0.0672
Parallel-trends F -stat	0.00235	0.105	0.0154
Parallel-trends p -value	0.961	0.746	0.901

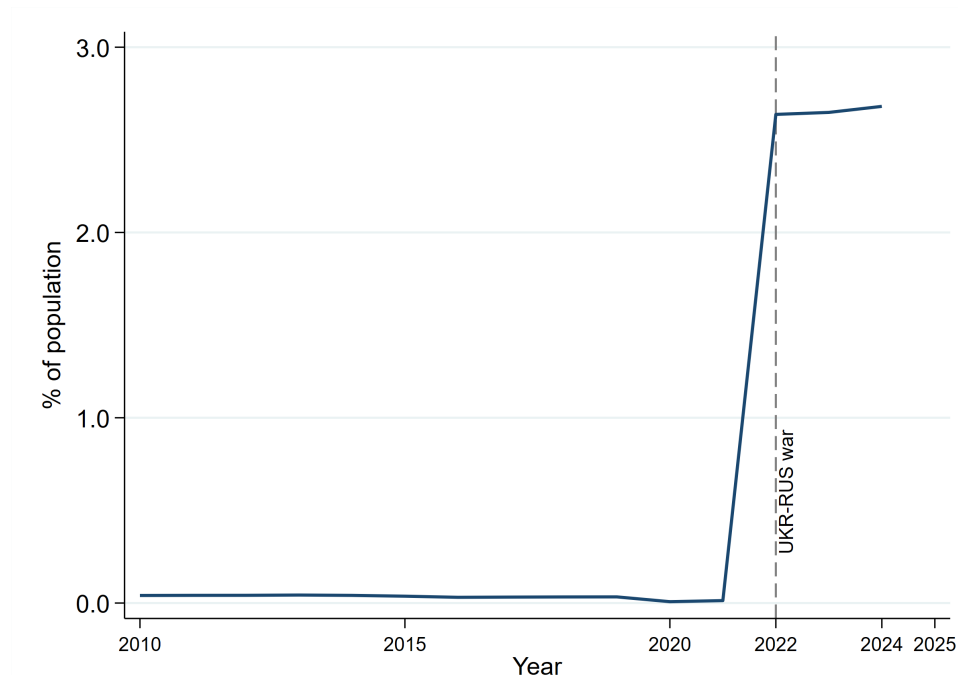
Note. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors (in parentheses) are clustered at the municipality level. This table reports estimates with the same specifications estimated in column (3) of Table 1 and column (3) of Table 3, restricting the sample as indicated in the column headings.

Table A5: Effects on all party groups

VARIABLES	(1) Cons.	(2) Alt-right	(3) Left + Centrist/Liberal + Other
$\Delta\text{UKR} \times 2019$	0.0183 (0.0294)	0.00241 (0.0137)	-0.00765 (0.0366)
$\Delta\text{UKR} \times 2023$	0.0582* (0.0304)	-0.0136 (0.0144)	-0.0446 (0.0280)
$\Delta\text{UKR} \times 2025$	0.119** (0.0487)	-0.0946** (0.0393)	-0.0248 (0.0235)
Observations	9,644	9,644	9,644
R-squared	0.993	0.986	0.991
Municipal FE	Yes	Yes	Yes
County*year interactions	No	No	No
X*year interactions	No	No	No
PCA*year	Yes	Yes	Yes
Mean dep. var. before the war	0.540	0.0672	0.391
Parallel-trend F -stat	0.389	0.0310	0.0436
Parallel-trend p -value	0.533	0.860	0.835

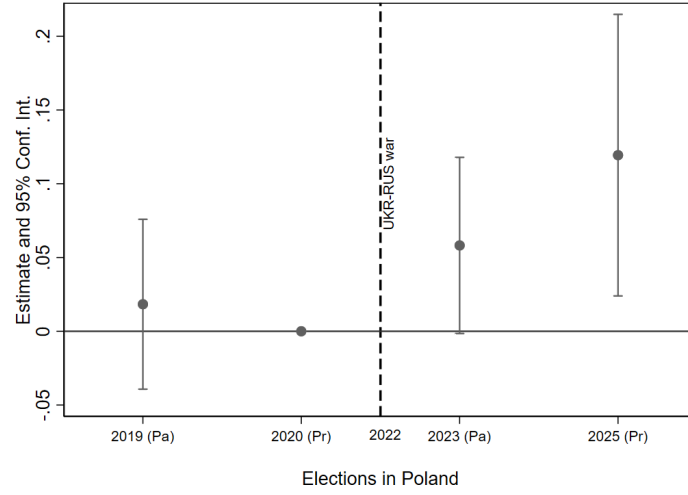
Note. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors (in parentheses) are clustered at the municipality level. This table reports the estimates with the same specifications estimated in column (3) of Table 1 and column (3) of Table 3, in column (1) and (2), respectively, while column (3) reports the effects on voting for parties that had political platforms more neutral or in favor of Ukrainian immigration during the war (i.e., left, centrist/liberal and other parties).

Figure A1: Trend in the percentage of refugees over the Polish population (2010-2024)

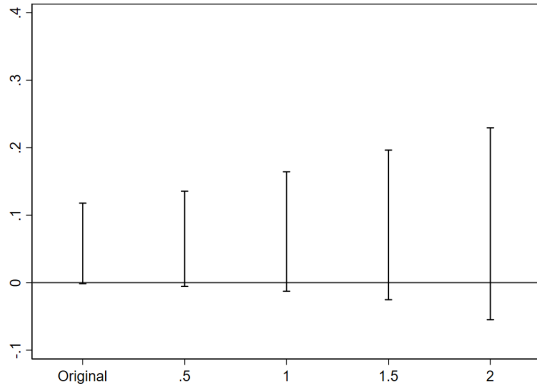


Note. The figure shows the trend in the percentage of refugees over the Polish population (2010-2024). Annual refugee stock data through 2023 are drawn from UNHCR Refugee Population Statistics (country of asylum), accessed via the World Bank World Development Indicators and the FRED mirror (series SMPOPREFGPOL, <https://fred.stlouisfed.org/series/SMPOPREFGPOL>). Figures for 2024 are based on UNHCR operational reporting for Poland, specifically the UNHCR Poland Bi-annual Fact Sheet (September 2024), and are used for descriptive purposes only (<https://www.unhcr.org/europe/sites/europe/files/2024-10/bi-annual-fact-sheet-2024-09-poland.pdf>). Total population figures are taken from the World Bank World Development Indicators (series POPTOTPLA647NWDB, <https://fred.stlouisfed.org/series/POPTOTPLA647NWDB>).

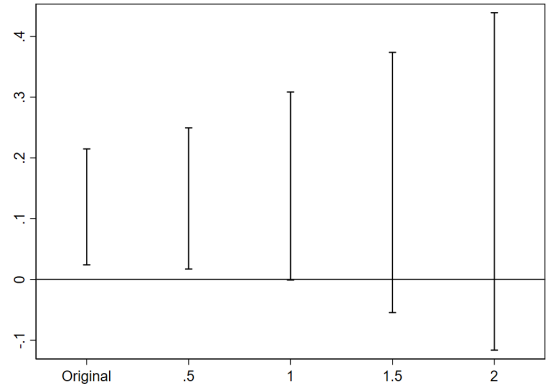
Figure A2: Effect of UKR refugees on conservative voting



(a) Event study DiD



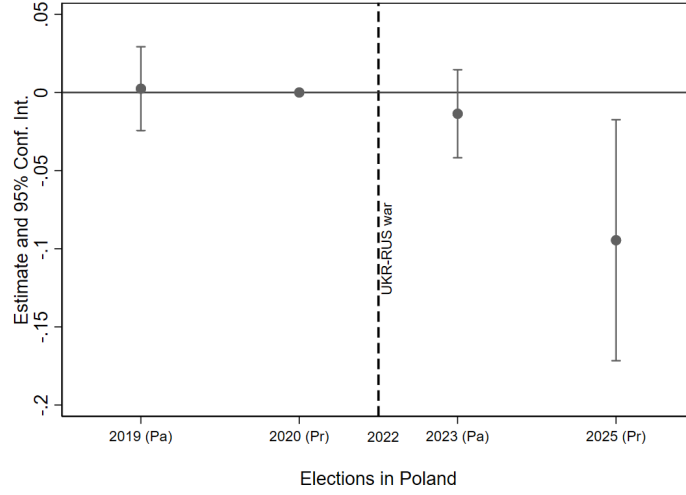
(b) Sensitivity 2023 election interaction



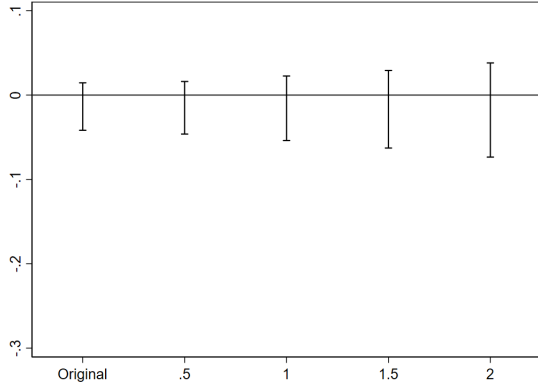
(c) Sensitivity 2025 election interaction

Note. Panel (a) reports the $\Delta\text{UKR} \times \text{year}$ point estimates from Equation (2) replacing the $\mathbf{X} \times \text{year}$ interactions with the first two (PCA) components interacted with the year indicators and 95% confidence intervals. "Pa" and "Pr" stands for Parliamentary and Presidential elections, respectively. Standard errors are clustered at the municipality level. Panel (b) and (c) present robust 95% confidence intervals for violations of parallel trends up to twice as big as the maximum violation in the pre-treatment period, which are obtained following [Rambachan and Roth \(2023\)](#), for the 2023 and 2025 election interactions, respectively.

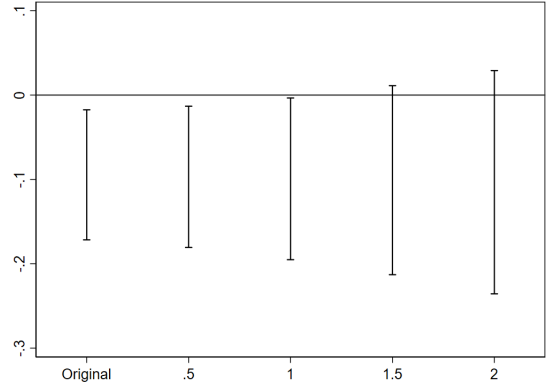
Figure A3: Effect of UKR refugees on alt-right voting



(a) Event study DiD



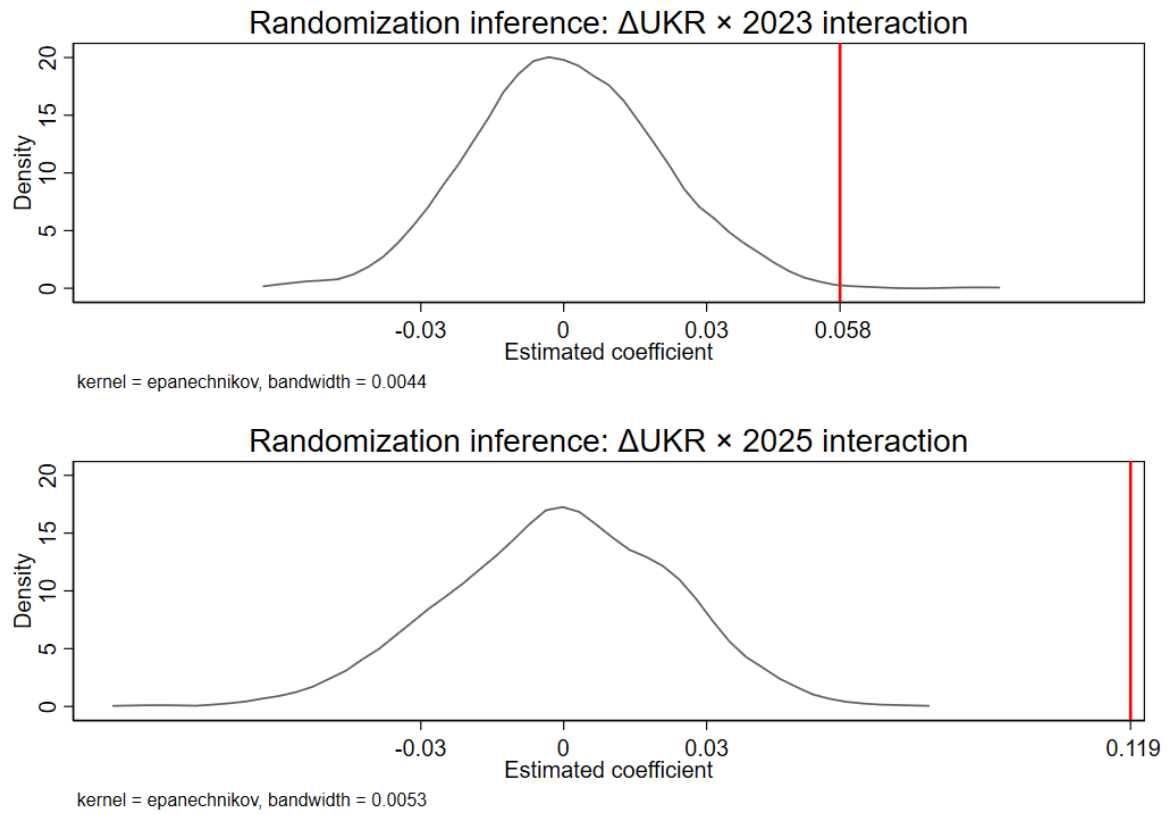
(b) Sensitivity 2023 election interaction



(c) Sensitivity 2025 election interaction

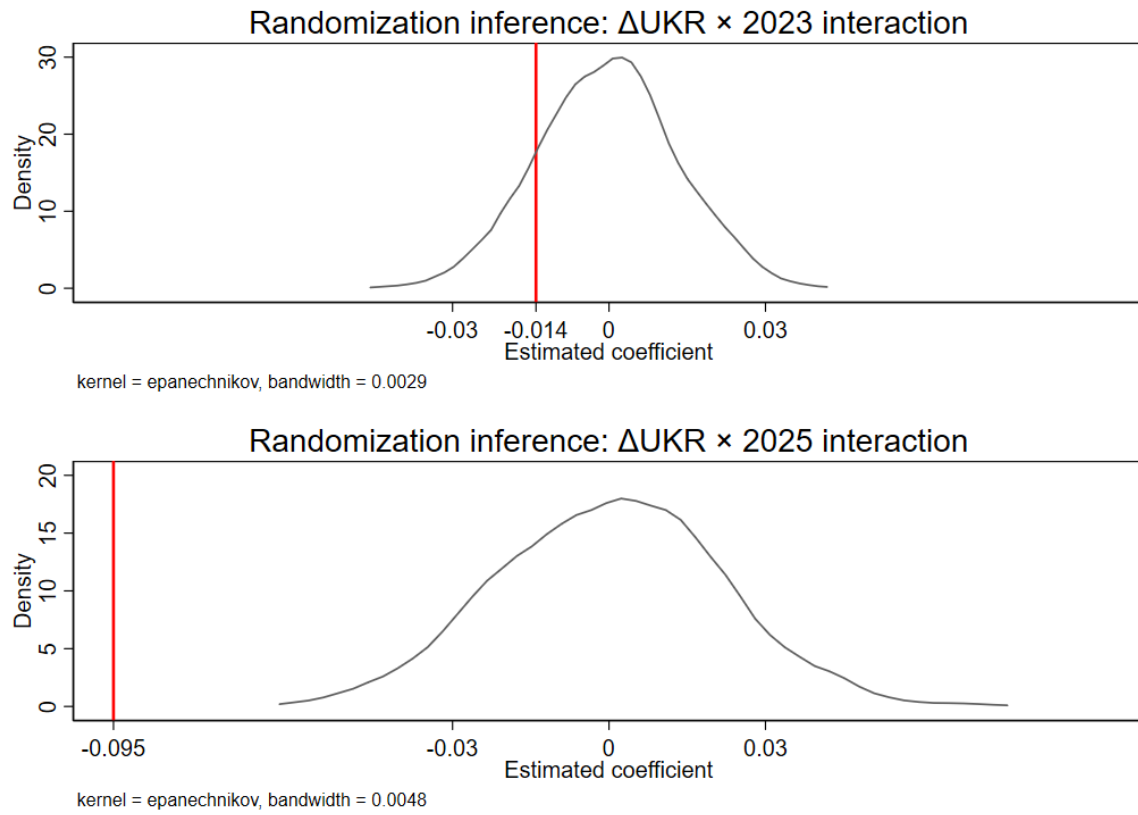
Note. Panel (a) reports the $\Delta\text{UKR} \times \text{year}$ point estimates from Equation (2) replacing the $\mathbf{X} \times \text{year}$ interactions with the first two (PCA) components interacted with the year indicators and 95% confidence intervals. "Pa" and "Pr" stands for Parliamentary and Presidential elections, respectively. Standard errors are clustered at the municipality level. Panel (b) and (c) present robust 95% confidence intervals for violations of parallel trends up to twice as big as the maximum violation in the pre-treatment period, which are obtained following [Rambachan and Roth \(2023\)](#), for the 2023 and 2025 election interactions, respectively.

Figure A4: Randomization inference: Conservative voting



Note. Coefficient distributions obtained with randomization inference (1,000 replications). The red vertical line shows the estimated coefficient in our baseline specification, including the interactions between the first two PCA factors with year indicators, municipality fixed effects, and county-year fixed effects. Inference carried out with the Stata `ritest` command.

Figure A5: Randomization inference: Alt-right voting



Note. Coefficient distributions obtained with randomization inference (1,000 replications). The red vertical line shows the estimated coefficient in our baseline specification, including the interactions between the first two PCA factors with year indicators, municipality fixed effects, and county-year fixed effects. Inference carried out with the Stata `ritest` command.