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Political backlash to refugee settlement: Cultural and economic drivers☆

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ABSTRACT

The 2015 refugee crisis in Europe has fueled anti-immigrant sentiment in host regions, with potential unintended consequences for refugee integration. We examine the heterogeneity of political backlash across Italian municipalities post-crisis and assess the concomitant role of economic *vs* socio-cultural factors in "welcoming" refugees (i.e., the supply side of integration). By leveraging the quasi-random dispersal policy and using causal forests, we find that refugee exposure has a significantly higher impact on anti-immigration backlash in more affluent areas and those with more bonding social capital. Conversely, areas with more bridging social capital, as measured by meaningful intergroup contact with former immigrants (e.g., mixed marriages), show less political backlash. We exploit this pattern of heterogeneity to evaluate counterfactual resettlement policies that minimize backlash. Results show that economic factors alone are insufficient to stem local discontent, while the socio-cultural dimension of host communities is crucial for the design of effective refugee resettlement programs.

1. Introduction

The rising inflows of immigrants and refugees, particularly from the Global South, into advanced countries have revealed new social and political concerns, such as populist anti-immigration sentiments traced back to both economic and cultural threats (Hainmueller and Hopkins, 2014; Halla et al., 2017; Guriev and Papaioannou, 2022). The "European refugee crisis", with its unexpected inflows of more than 1.5 million refugees in 2015 alone, has fueled public hostility and the electoral success of far-right parties advocating stricter immigration policies (e.g., Hangartner et al., 2019; Dustmann et al., 2017; Dinas et al., 2019; Campo et al., 2024).[1] Yet, while refugee migration appears to trigger backlash on average, growing evidence suggests that this effect hides a high degree of heterogeneity across receiving local communities (Damm and Rosholm, 2010; Dustmann et al., 2019; Steinmayr, 2021). For instance, political reactions to refugee exposure appear to be significantly harsher in rural than in urban areas. This spatial variation, though, is not sufficient to grasp the underlying mechanisms through which refugee hosting mitigates or exacerbates natives' concerns.

Understanding why some locations are more hostile to refugees than others has important policy implications. The integration pro-

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[1] Between 2014 and 2017, a record 3.5 million refugees applied for asylum in the EU-28 countries (Eurostat, 2020), most fleeing war and terror in Syria and social unrest in North Africa and the Near East (Afghanistan, Iraq, and Yemen). The arrival of asylum seekers in Europe in 2015 marked the largest annual flow of asylum seekers since 1985 (Pew Research Center). This crisis has put some EU Member States under severe pressure regarding their national capacities to host and manage asylum seekers in a fully-fledged reception system (UNHCR, 2016). Integrating refugees is currently a critical political goal in many European countries.

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cess involves acceptance, interactions, and exchanges with the majority group in the receiving areas where refugees settle upon arrival (Ager and Strang, 2008; Danzer and Yaman, 2013). Hence, the host community significantly affects the supply side of integration by providing or denying opportunities (Bisin and Tura, 2019; Fouka, 2022; Davis et al., 2024). Political backlash to refugee settlement can harden hostility and xenophobic sentiments against ethnic minorities. Most importantly, this backlash may translate into actual policies restricting opportunities for refugee integration and undermining any efforts and demand for integration. Therefore, matching refugees to host communities is crucial for designing effective integration policies.

In this paper, we assess how local conditions at the time of refugees' settlement influence anti-immigration preferences in receiving areas. We identify the causal effects of a wide range of socio-economic factors at the local level, isolating the in-group and out-group dimensions. We provide evidence of contrasting political effects driven by economic *vs* socio-cultural channels, including novel measures of social capital and intergroup interactions. We build on this pattern of heterogeneity and evaluate counterfactual resettlement schemes based on a matching framework that assigns refugees to locations to minimize anti-immigrant backlash. Our counterfactual exercises show that relative to the initial dispersal policy, optimal resettlement policies induce, on average, a significant reassignment of refugees from wealthy and socially connected locations to less affluent but more culturally integrated local areas. Our results contribute to the policy debate on managing and integrating refugees in host communities by examining local responses to ethnic diversity. While understudied in natural settings, this supply-side perspective offers novel and important policy implications for refugee reception programs.

We focus on Italy, a country highly exposed to the European refugee crisis, where asylum seekers have been distributed quasi-randomly across municipalities under the dispersal policy (Campo et al., 2024; Chamber of Deputies, 2017).[2] This provides us with a unique research design to study the causal effects of local pre-determined economic and socio-cultural conditions on the outcomes of refugee reception, which is otherwise impossible due to the sorting of newcomers across locations. Moreover, Italy is an interesting context to identify various socio-cultural mechanisms since it displays extensive granular-level variation in economic and non-economic factors across nearly 8000 municipalities while presenting homogeneous policies and institutions throughout the country.

We combine unique administrative data on refugee centers opened between 2014 and 2018 with electoral data and a rich set of pre-crisis municipality-level characteristics, including novel measures of social capital and intergroup interactions compiled using different administrative sources (see below). Our identification strategy relies on a difference-in-differences framework, comparing changes in the share of anti-immigration votes between the pre- and post-crisis period (2013–2018) across municipalities with different shares of hosted refugees. We provide evidence that, due to the design of the dispersal policy, the assignment of refugees to municipalities during the crisis is orthogonal to voters' political preferences and unrelated to contemporaneous or historical socio-economic and political trends.

We estimate local political backlash to refugee exposure (treatment effect) by comparing the electoral support for anti-immigration parties, i.e., the *League* and *Brothers of Italy (BoI)*, in the 2013 and 2018 national elections. Although these right-wing parties frequently raised the issue of immigration and minority groups, the refugee crisis has made immigration and ethnic diversity even more salient issues in their political agenda. Our political outcome is also a good proxy for negative attitudes toward immigration. As a validation exercise, we show that voters supporting the *League* and *BoI* display stronger negative attitudes and more stringent behavior toward ethnic diversity compared to voters of other parties along the political spectrum. As a matter of fact, this agenda translated into restrictive refugee integration policies enforced by the *League*-led government that took office after the 2018 national election, as we document in Section 2.2.

We estimate heterogeneous effects using both linear interaction models and the causal forest algorithm (Athey and Imbens, 2016; Athey et al., 2019), which allows us to capture the high-dimensional combination of local predictors (Conditional Average Treatment Effect or CATE). Political backlash is significantly heterogeneous across receiving municipalities, ranging from 0.02 percentage points (p.p.) in the 5th percentile of the distribution to more than 0.13 p.p. in the 95th percentile. We characterize the heterogeneity of the treatment effects across economic and non-economic local mechanisms and find three core sets of results. First, backlash is higher in better-off areas, as measured by income, activity, and employment rate. This result aligns with a 'welfare dependency' argument, such that (richer) natives may be reluctant to support refugees through the general welfare state (Facchini and Mayda, 2009; Dustmann et al., 2019).

Second, drawing from the seminal work by Putnam (1993), a substantial body of literature has pointed out that social capital, conceived in terms of civic engagement and pro-social behavior, fosters mutual support and cooperation within a community (Portes, 2000; Guiso et al., 2008). However, communities with dense social ties may not necessarily be better equipped to manage the challenges posed by immigration and increasing diversity. The refugee crisis, with its unexpected inflows of culturally diverse asylum seekers, increases the salience of cultural-ethnic boundaries. By using different proxies for social capital, i.e., referenda turnout rate, blood donation, and volunteering rates, we find that all these measures exacerbate voters' backlash. This finding aligns with the *bonding* notion of social capital, identified as the set of exclusive connections formed within a homogeneous group or a community (Coleman, 1990; Woolcock, 1998; Putnam et al., 2000; Portes, 2000).[3,4]

We also focus on a complementary dimension of social interactions, i.e., those across groups. More specifically, we identify the role of *meaningful* social interactions across native and former immigrants, defined as *bridging* social capital, in our natural setting. We collect various measures for the frequency of positive intergroup contact and integration at the municipality level, i.e., intermarriage rate, naturalization rate, residential integration, and foreign-born elections to the local office. According to the "contact hypothesis", *meaningful* interactions across members of different groups are expected to reduce prejudice and hostility (Allport,

[2] Between 2014 and 2017, an average of 150,000 people reached Italian shores each year via Mediterranean routes (UNHCR, 2018), accounting for about 18 percent of all first-time applicants in the EU-28 (Eurostat, 2020). The refugee crisis was a sudden shock both in terms of scale and ethnic composition of the refugee population. Indeed, the new influx of refugees strikingly exceeded prior values with an average of 25,000 annual applications; see Figure A1.

[3] Bonding social capital is, by choice or necessity, inward-looking and tends to reinforce exclusive identities and homogeneous groups (such as clubs) and to create strong in-group loyalty, but possibly also strong out-group antagonism (Onyx and Bullen, 2000). Bridging networks, instead, are outward-looking and encompass people across diverse social cleavages, such as civil rights movements, ecumenical religious organizations, and youth service groups. As argued in Putnam et al. (2000), bonding social capital is good for "getting by", but bridging social capital is crucial for "getting ahead". In the context of the refugee crisis, *bonding* social capital reinforces exclusive identities and possibly strengthens out-group antagonism (Satyanath et al., 2017).

[4] Several contributions have pointed out that social capital contributes to economic development and good institutions (Guiso et al., 2008, 2011; Algan and Cahuc, 2013). Yet, evidence is scarce and mixed about the relationship between social capital, political ideology, and voting. On the one hand, some contributions point out a negative association between voting for populist parties and the strength of civil society both in Europe and the USA (Boeri et al., 2021; Giuliano and Wacziarg, 2020). On the other hand, in the historical context of the Nazi party's rise to power, Satyanath et al. (2017) show that social capital, measured by the density of associations in German towns, stimulated Nazi Party membership and electoral success.

1954). Consistently, our results suggest that backlash is significantly lower in communities with higher *bridging* social capital at baseline. These findings, overall, point to the potential for *bridging* social capital to stem discontent and anti-immigrant attitudes, suggesting that sustained experience of *meaningful* cross-group interactions mitigates public discontent. Interestingly, though, we find higher discontent in areas with a higher share of former immigrants. This evidence upholds that pure out–group exposure in the past, without positive intergroup contact with natives, may trigger backlash.

Our findings raise concerns about the unintended consequences of refugee dispersal policies in many Western countries. The mismatch between refugees and local communities might hamper the long-term integration of minorities. Hence, we exploit heterogeneous estimates of local responses from CATE and evaluate alternative resettlement policies that aim to assign refugees to locations while minimizing anti-immigrant backlash. We show that *ceteris paribus*, these optimal resettlement schemes ensure a sizable reduction in backlash compared to a purely random dispersal policy—ranging from 34 to 120 percent under different capacity constraints. More precisely, the predicted reduction in anti-immigrant backlash grows along with refugee concentration across locations. The reduction in backlash is driven by both the reallocation of treated municipalities and the different shares of refugees assigned per municipality (both extensive and intensive margins). We quantify and describe the mismatch in refugee assignment and show that optimal policies lead to, on average, a significant reassignment of refugees from rich and socially bonded areas to less affluent but more culturally integrated municipalities. Finally, we quantify the pivotal role of socio-cultural characteristics, beyond economic factors, in shaping more inclusive resettlement programs. Our analysis indicates that by overlooking the socio-cultural structure, assignment policies are less effective in mitigating anti-immigrant backlash. This underscores the need for a comprehensive approach that takes into account not only economic factors but also the socio-cultural fabric of the communities involved.

This paper adds to the flourishing literature on the political effects of immigration, showing that exposure to ethnic minorities triggers the electoral success of far-right and anti-immigrant parties (e.g. Halla et al., 2017; Hangartner et al., 2019; Edo et al., 2019; Alesina and Tabellini, 2023) and lowers support for redistributive policies (Dahlberg et al., 2012; Alesina et al., 2021, 2023). In a related paper, Campo et al. (2024) show that the increase in support for right-wing anti-immigration parties in Italy, which runs in parallel to a decline in support for center-left parties, is not linked to any material economic effect (e.g., native income or local public spending effects) of refugee exposure.[5] Yet, more nuanced evidence is emerging comparing rural and urban areas or locations with different educational compositions of residents (e.g., Dustmann et al., 2019; Mayda et al., 2022). For instance, to explain the urban-rural divide, Dustmann et al. (2019) use ESS survey data to show that support for anti-immigrant parties is associated with natives' views and attitudes toward refugees and contact with immigrants (proxied by the likelihood of having immigrant friends or colleagues). Moreover, while transient refugee exposure increases far-right voting, persistent or meaningful contact appears to reduce anti-immigration sentiment (e.g. Steinmayr, 2021; Dinas et al., 2019; Achard et al., 2022; Bursztyn et al., 2021; Asimovic et al., 2022).[6]

Our contribution to the existing literature on refugee immigration is twofold. First, we complement the analysis of local heterogeneity in anti-immigration responses along three novel perspectives. We examine the role of both economic and socio-cultural conditions, introducing local in-group and out-group dimensions. Hence, we use fine-grained panel data from a nationwide natural setting, which allows for simultaneous comparison of these underlying mechanisms within a systematic framework during a salient moment of crisis. Eventually, we leverage recent advancements in causal forest algorithms to effectively untangle the contribution of each local mechanism and capture potential interactions among predictors. This analysis is crucial to discriminate between bonding and bridging social capital measures, which, while not mutually exclusive, may operate in opposite directions (Schuller, 2007).

A second contribution of our analysis is that we leverage our heterogeneous estimates to evaluate counterfactual policies, matching refugees to locations. Building on simulations, we discuss potential avenues for designing inclusive resettlement policies. Government policies dealing with the management and allocation of refugees should consider both economic and non-economic local characteristics to mitigate hostility and conflict and to promote integration.

Finally yet importantly, a large literature examines how local characteristics impact refugees' economic integration, with variations in labor market performance linked to different local conditions such as ethnic networks and population density (Edin et al., 2003; Damm, 2009; Martén et al., 2019; Eckert et al., 2019; Battisti et al., 2022; Fasani et al., 2022), economic and educational status (Damm and Rosholm, 2010; Godøy, 2017; Ahrens et al., 2023), hostility toward out-group members (Jaschke et al., 2022). Most studies focus on policies affecting immigrants' labor market convergence to native outcomes, i.e., the *demand* side of integration, questioning the effectiveness of random policies by simulating novel allocations to improve refugees' labor market opportunities (Godøy, 2017; Bansak et al., 2018; Andersson and Ehlers, 2020). Our contribution centers on the *supply* side of integration, as public hostility toward minorities in receiving areas may contribute to suboptimal integration outcomes (Damm and Rosholm, 2010; Arendt et al., 2022; Fasani et al., 2022). By interpreting integration as an equilibrium outcome (Bisin and Tura, 2019; Fouka, 2022), our counterfactual evaluation studies, for the first time, the implications of refugee dispersal policies on natives' welcoming of ethnic minorities. Indeed, the supply-side analysis provides novel insights into the design of refugee reception programs and represents a necessary step toward a comprehensive general equilibrium model that includes both demand and supply components.

2. Background and data

2.1. The refugee-reception system and the dispersal policy

The number of asylum applications in Italy remained limited and stable until 2013, with an average of 25,000 applications per year. Starting in 2014, the uprisings in the Middle East and the escalating conflict in Syria led to an increase in the influx of asylum seekers to Europe, with an average influx of 150,000 arrivals per year and the number of asylum applications exceeding 130,000 in 2017, which represents 18 percent of all first-time applicants in the EU-28 (Eurostat, 2020).[7] In early 2018, refugee flows began to decline following the Italy-Libya Memorandum of Understanding, a financial agreement to crack down on asylum seekers in exchange for foreign aid funding. Figure A1 illustrates this trend. Compared to pre-crisis levels, this surge in asylum seekers represents a large and unexpected demand shock that overwhelmed the existing refugee reception system (SPRAR scheme).[8] To deal with this lack of capacity, in 2014, the government set up a complementary reception

[5] Campo et al. (2024) also explore the role of political propaganda – proxied by right-wing rallies – as a post-treatment booster of anti-immigration voting. These rallies took place after refugee allocation and just before elections, and they represent a form of political supply rather than a local contextual factor.

[6] Research in social psychology and political science has examined if intergroup contact can reduce prejudice and negative attitudes toward out-groups (Paluck et al., 2019; Tropp, 2012; Mousa, 2020). Yet, many of these studies take place in a laboratory (or lab in the field) or rely on surveys.

[7] During the crisis, refugees reached the Italian borders mainly through the Central Mediterranean Route, from Sub-Saharan African countries (mainly Nigeria, Gambia, and Senegal) and from the Middle East and East Asia (mainly Pakistan and Bangladesh), see Table A2 for more details.

[8] Before the crisis, Italy's refugee-reception scheme was entirely carried out by the System for the Protection of Asylum Seekers and Refugees (SPRAR). After

system following a spatial dispersal policy (see below). These Temporary Reception Centers (*Centri di Accoglienza Straordinaria* – CAS) due to the lack of capacity of the SPRAR scheme, quickly replaced the previous reception scheme as the go-to system hosting on average 75–80 percent of asylum seekers in Italy. CAS centers only provide basic reception services (food and accommodation) and limited resources prevent investment in integration programs (Campo et al., 2024).

Since data on refugees' redistribution in CAS released by the Home Office do not cover the entire period between 2014 and 2018, we create a unique and harmonized dataset tracking the list of CAS opened at the municipality level, their capacities, timelines, and the number of hosted refugees every year from 2014 to 2019; see Campo et al. (2024) for further details about the data collection process. The final sample counts 92 out of 106 Prefectures (provinces), reporting refugee data for 6965 out of 7918 Italian municipalities. We provide evidence of the absence of selective attrition in our sample by running balance tests on pre-treatment characteristics.[9]

The refugee resettlement across CAS reception centers followed a dispersal policy plan (*Piano Nazionale di Riparto*) to reduce their concentration in a few locations and share the burden of reception and hospitality across the national territory.[10] The resettlement scheme is conducted in two steps: first across provinces and second across municipalities within provinces.[11] First, the Home Office centrally redistributes refugees to each province according to the resident population, with an allotment plan of 2.5 refugees per 1000 inhabitants.[12]

In the second step, the allocation of CAS centers within the province is coordinated by Prefectures. Provincial government offices open public bids that assign the management of reception centers to cooperatives, NGOs, or private operators based on tender cost schemes. Crucial to our research design, economic operators propose and decide the allocation of CAS centers without consultation with local municipalities. Hence, municipal authorities did not influence the redistribution process. They had no control over the number and characteristics of assigned refugees, nor the timing of allocation. In Section 3, we provide robust evidence showing that the reallocation scheme implied by the dispersal policy is indeed orthogonal to a broad set of baseline local features. An extensive description of these procedures is provided in Campo et al. (2024).

In our sample, CAS centers hosted 37,000 refugees in 2014 and up to 144,000 in 2017. In parallel, the number of municipalities hosting a CAS tripled over the same period, with the maximum number of CAS centers observed at the end of 2017. Figure A3 plots the distribution of the share of refugees across municipalities in these years. Some CAS centers are housed in former group accommodation buildings, but around 85 percent are divided across networks of private apartments (Chamber of Deputies, 2017). Panel A of Table 1 reports summary statistics. On average, CAS centers host about 20 refugees, with a decreasing capacity as the dispersal policy was implemented over time.[13]

2.2. Political background

We focus on two consecutive parliamentary elections in 2013 and 2018, which, importantly for our design, took place before and after the refugee crisis.[14] Thus, our outcome of interest measures the change in political preferences for anti-immigration parties throughout the crisis. We gather electoral data from the Italian Home Office reporting the number of votes obtained by each political party and the number of voters per municipality in each national election.

The two main anti-immigration parties are the *League* and the *Brothers of Italy* (*BoI*).[15] Table 1 shows the electoral outcomes for the anti-immigration front in 2013 and 2018. In 2013, immediately before the refugee crisis, the *League* and *BoI* jointly accounted for about 8 percent of votes. The results of the 2018 election marked a moment of stark discontinuity with the previous political arena. The anti-immigration front gained substantial support, reaching 25.5 percent of votes, and took the lead within the center–right area across the country; see Figure A4.[16]

We identify anti-immigration parties by focusing on the salience of the anti-immigration arguments in political agendas following the Manifesto Project (Volkens et al., 2020). This data repository extrapolates election-specific information about parties' positions on various issues, including immigration and multiculturalism (see Table A4). The political agenda of both the *League* and *BoI* includes negative references to diversity, aversion to multiculturalism, and support for restrictive immigration policies. These arguments evolved and became more pronounced with the escalation of the refugee crisis, and in 2018, the appeal for cultural homogeneity became a pillar of the *League*'s program. Their propaganda during the 2018 election strongly emphasized the risk of a demographic and cultural change due to immigration (referring to ethnic substitution), fed a climate of xenophobia and social hostility, and triggered anti-European and anti-globalization sentiments.[17]

In line with these programmatic items, both parties strongly opposed the reception of immigrants crossing the Mediterranean Sea, voted against the reform of the Dublin system in the European Parliament, and opposed any attempt to reform the current laws limiting immigration quotas to Italy. In particular, a few months after taking office as Minister of the Interior in 2018, the leader of the *League* Matteo Salvini implemented a restrictive reform of the refugee reception scheme. The reform downsized the financial and administrative resources for reception by cutting the budget for CAS management by more than 20 percent, abolishing humanitarian protection shrinking the pool of people eligible for reception, and limiting access to integration services (labor market active policies, language training and psychological support)

a preliminary phase of identification and assistance conducted at main disembarkation sites (hotspots), asylum seekers and refugees are hosted in SPRAR centers. Funded by the national government, SPRAR reception centers are set up voluntarily by municipality administrators and managed by local authorities on a nonprofit basis. Consequently, political orientation of each municipality and its administrative capacity affect the limited and uneven distribution of SPRAR centers across the country. In December 2014, only 5 percent of municipalities were hosting a SPRAR center. For a full description of SPRAR reception scheme see Campo et al. (2024).

[9] Table A3 confirms that out-of-sample municipalities in non-responding provinces are not systematically different from in-sample municipalities along economic, demographic, political, or institutional characteristics.

[10] Similar dispersal policies have also been enforced in other European countries (e.g., Sweden, Denmark, Switzerland, and Germany).

[11] In Italy, provinces correspond to NUTS-3 level administrative units, while municipalities correspond to NUTS-5 administrative units.

[12] Figure A2 plots the number of assigned refugees per province over the pre-policy province population in 2013, uncovering a robust positive correlation. The regression slope is equal to 2.3 (s.e. 0.0001), in line with the allotment plan, with an R-squared equal to 0.85.

[13] Overall, the share of CAS centers hosting more than 100 refugees is 3 percent, on average (and never exceeded 3.5 percent), reflecting the granular dispersion of refugees across municipalities due to the policy.

[14] Every five years, Italian voters elect the members of the two chambers of the national parliament, i.e., the *Chamber of Deputies* and the *Senate*. All adult Italian citizens over 18 years old are entitled to vote for the election of the members of the Chamber of Deputies, while only those over 25 are eligible to vote for the Senate.

[15] The anti-immigration front also includes several extreme-right and neo-fascist movements, such as *Casa Pound*, despite their limited electoral impact (around 1 percent of votes overall).

[16] Concurrently, the *5 Stars Movement* (5SM) populist party continued to grow, becoming the most-voted party in both chambers of the national parliament, while the center-left coalition experienced a considerable loss of support (see Campo et al. (2024) for more details).

[17] See the Hate Barometer collected by Amnesty International Italia during the early 2018 electoral campaign (https://www.amnesty.it/barometro-odio/).

Table 1
Summary statistics.

	Count	Mean	Sd	Min	Max
Panel A. Refugees					
Share of refugees in 2017	6891	0.40	1.60	0	61.31
Number of refugees in 2017	6891	20.81	96.75	0	4000
Avg number of refugees 2014–2017	6891	13.63	73.26	0	4000
Municipality with CAS 2014–2017	6891	0.43	0.50	0	1
Avg number refugees per CAS	2562	23.14	84.78	0.400	4000
Municipality with CAS, more 1 year	6891	0.31	0.46	0	1
Municipality with CAS, more 100 refugees	6891	0.03	0.17	0	1
Municipality with SPRAR	6891	0.10	0.30	0	1
Share of refugees in SPRAR 2017	6891	0.07	0.54	0	17.49
Avg share refugees in SPRAR 2014–2017	6891	0.05	0.39	0	12.92
Panel B. Electoral outcomes					
Vote share for anti-immigration parties in 2013	6891	8.44	7.36	0	56.52
Vote share for *League* in 2013	6891	5.98	7.22	0	47.83
Vote share for *BoI* in 2013	6891	2.08	2.58	0	42.28
Change in vote share for anti-immigration parties	6891	17.55	7.88	−25.00	53.67
Change in vote share for *League*	6891	14.67	7.07	−8.155	45.53
Change in vote share for *BoI*	6891	1.92	2.88	−30.08	40.57
Panel C. Municipality characteristics					
Economic drivers					
Income per capita (log)	6891	9.33	0.26	8.034	10.26
Activity rate	6891	49.86	6.31	19.33	77.11
Employment rate	6891	44.91	7.61	18	74.02
Tertiary education rate	6891	7.44	2.78	0	29.06
Population over 65 (%)	6891	24.51	5.38	7.425	61.68
Social capital					
Electoral participation referenda	6891	48.98	6.30	21.33	72.49
Association density (%)	6885	9.21	8.22	0	95.59
AVIS branch in 2010	6891	0.38	0.48	0	1
Intergroup contact					
Share of immigrants	6891	524.38	3205.85	0	194991
Residential segregation index	6891	21.82	10.66	0	98.25
Naturalization rate	6891	13.71	10.37	0	100
Intermarriage rate	6869	10.97	8.11	0	76.92
Elected Foreign-born admin.	6891	0.35	0.48	0	1
Elected Non-EU15 born admin.	6891	0.14	0.34	0	1

Notes: This table shows summary statistics for electoral outcomes in Panel A, for refugee assignment and dispersal policy in Panel B, and municipality local characteristics in Panel C. Table A1 reports the definition of all variables of interest and data sources.

excluding asylum seekers whose request for protection was pending approval (Openpolis and ActionAid, 2019).[18]

We validate our electoral outcome showing that voters supporting anti-immigration parties display stronger negative attitudes and more stringent behavior toward ethnic diversity. We exploit individual-level data from the European Social Survey (ESS) collected in Italy starting from the second half of 2018, right after the national election, providing self-reported voting and political preferences as well as attitudes and behavior against immigration.[19] Descriptively, 80 percent of respondents who voted for anti-immigration parties viewed immigration as bad for the country's economy, and 82 percent believed the presence of immigrants undermines the country's cultural life. Fig. 1 shows that individuals who declare to vote for (or be close to) the *League* and *BoI* are more likely to (i) support restrictive immigration policies from different backgrounds and (ii) perceive immigration as bad for the economy and socio-cultural environment. On the contrary, political support for Forza Italia and 5SM, potentially expressing welfare and employment considerations, does not correlate with attitudes and behavior against immigration.[20]

2.3. Local heterogeneity: economic and socio-cultural dimensions

To explore the role of local contextual factors, we complement our data with a rich set of municipal characteristics observed before the refugee policy launch. We consider three different dimensions of the local context. First, we examine standard economic prosperity and human capital measures at the municipality level. Then, we focus on socio-cultural factors, including fine-grained social capital variables. We distinguish between the bonding and the bridging dimensions of social capital by considering, for the latter, variables that proxy for the actual integration of former immigrants and intergroup contact at the municipality level. For each of these dimensions, we consider a broad set of pre-crisis indicators from multiple data sources, as described below.

Panel C of Table 1 summarizes the variables we include in our empirical analysis. On the economic dimension, we resort to aggregate data from the Ministry of Finance on taxable gross income earned by residents to compute municipality per-capita income. We also consider several labor market indicators, i.e., employment and activity rates, retrieved from the latest available Census before the refugee crisis (2011). Our sample municipalities have a mean (median) per capita income

[18] The Law n.132/2018 known as "Salvini decree" was finally approved by the Italian Parliament in November 2018. In the subsequent months, the Italian government took further restrictive anti-immigration initiatives like the refugee boat policy, barring NGO rescue boats from docking.

[19] ESS data are representative of voting preferences, with an average vote share for anti-immigration parties of 24 percent compared to 25.5 percent at the national election. Summary statistics are reported in Table A5. Unfortunately, ESS data do not report individuals' municipality of residence, preventing us from exploiting these measures as additional outcomes in our analysis.

[20] The estimates partial out survey year and region fixed effects, and individual characteristics.

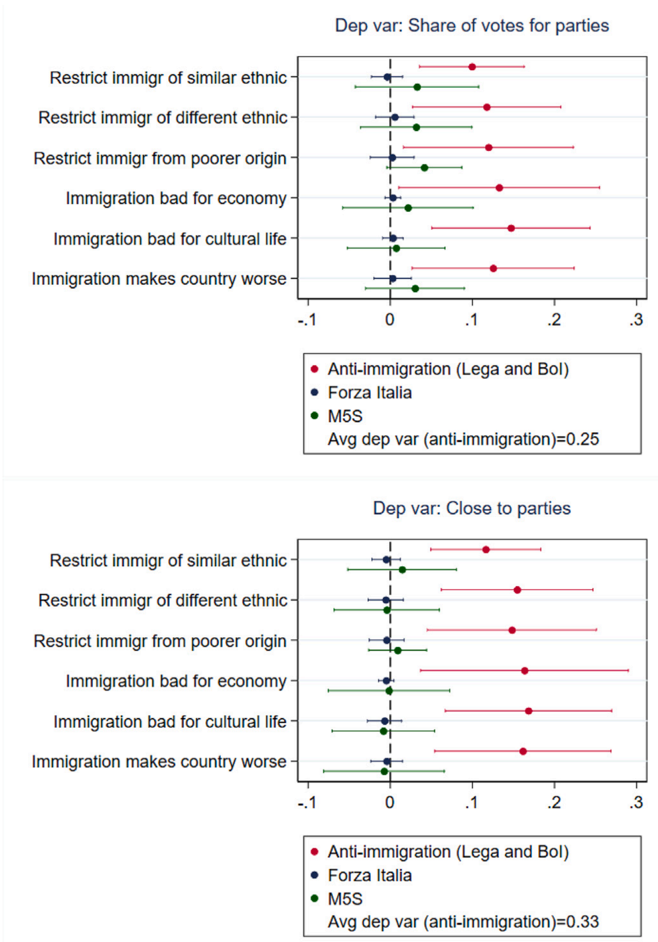

Fig. 1. Validation ESS-political preferences and attitudes against immigration. *Notes:* This Figure shows OLS estimates of the correlation between self-reported political preferences and attitudes and behaviour against immigration. The dependent variables in panel A include an indicator for whether the individual declare to vote for *League* and *BoI*, or alternatively Forza Italia and M5S in the last election in 2018, and in panel B an indicator for whether the individual declare to be close to *League* and *BoI*, or alternatively Forza Italia and M5S. The main explanatory variables are indicators equal to one if the individual declares to (i) support restrictive immigration policies of similar ethnic groups, (ii) of immigrants from different backgrounds, (iii) of immigrants from poorer origin, and (iv) perceive immigration as bad for the economy and v) socio-cultural environment, and (vi) believe immigration makes the country worse. Estimates partials out survey year and region fixed effects, and individual characteristics including gender, age, age square, education, marital status, and having a child. Summary statistics for the main variables are reported in Table A5. *Source*: ESS data, Italy (2018).

of 11,652 (12,123) and an average employment rate of 45 percent. We measure human capital from 2011 Census data; the share of the population with tertiary-education is 7.44 percent.

On bonding social capital indicators, we refer to the measures commonly adopted in the literature, i.e., electoral participation in referenda, blood donation, and association density of non-profit organizations (Putnam et al., 2000; Guiso et al., 2004; Cartocci, 2007; Durante et al., 2025).[21] We conceive social capital as the broad set of values and connections that foster cooperation and mutual support within a group. Indeed, frequent interactions among individuals in a group tend to produce a norm of generalized trust and reciprocity (Coleman, 1988, 1990; Putnam et al., 2000; Guiso et al., 2004).

We retrieved data at the municipality level on electoral participation in referenda from the Historical Archive of the Ministry of Interior.[22] We observe an average turnout of 49 percent, with considerable heterogeneity not only across provinces but also across municipalities within the same province.[23] In addition, we use registry data from AVIS, the main blood volunteers' association in Italy, to construct an indicator for the presence of an AVIS branch in the municipality; 38 percent of municipalities in our sample have an active AVIS branch at baseline.[24] We also include the percentage of residents volunteering in non-profit institutions (mean 9.2 percent) as a further measure of social capital, exploiting data from the 2011 Census of non-profit institutions.

As for measures of bridging social capital, we construct a set of indicators of intergroup contact by using several administrative micro-level data sources from the Italian Statistical Institute (ISTAT, ADELE Laboratory). We aim to measure the importance and strength of connections between natives and former immigrants, capturing different degrees of interethnic contact as well as the extent to which immigrants are socially integrated into local communities. We exploit registry data on the universe of marriages formed in Italy from 1998 to 2012 to measure the intermarriage rate, i.e., the number of marriages between immigrants and natives over total marriages celebrated from 1998 to 2012 (ISTAT, ADELE Laboratory). We further use Census division level data from the 2011 census to construct a residential segregation index (Duncan index of immigrants' residential segregation). Using the same source of data we define a naturalization index that measures the share of naturalized immigrants over the total number of immigrants living in Italy for at least 10 years, i.e., those potentially eligible to apply for citizenship.[25] In addition, we include former immigration waves (i.e., the share of first and second-generation regular immigrants) to describe municipality-level ethnic networks. It is worth noting that, while intermarriage is a well-documented proxy for intergroup relations and acceptance of other groups, former shares of immigrants may also capture 'meaningless' exposure without contact.

Finally, we take advantage of the Local Administrators Registry 2007–2013 to identify those municipalities where at least one foreign-born administrator was elected to the municipality board; 36 percent of our sample municipalities elected at least one foreign-born administrator over the period considered, although only 14 percent elected someone born in non-EU15 countries.

3. Estimation strategy and identification

3.1. Estimation

We estimate how local characteristics affect the political response to refugees' exposure with a two-period difference-in-differences (DiD) specification that compares changes in anti-immigration vote shares across municipalities with different levels of exposure to refugee settlement. We examine how this double-difference between 2013 and

[21] Additional survey measures from WVS or ESS are available at the individual level. However, the geographical level of granularity is at the province and not the municipality level, corresponding to the unit of observation in our identification strategy.

[22] We compute the average turnout considering those referenda for which data are available at the municipality level, i.e., 1974, 2009, 2011, 2016, and 2020.

[23] For electoral participation in referenda, the standard deviation between provinces is 5.15, while the within-province standard deviation reaches 4.16.

[24] AVIS administration provides us with the list of active branches from 2010 to 2015 (see https://www.avis.it/it). This dummy variable is quite 'broad' as an indicator and it may capture many different local features including municipality size or remoteness. This is to say this variable may have significant measurement error. Yet, a continuous variable for blood donation at the municipality level does not exist or is not available.

[25] We exclude from this computation foreign-born residents who obtained Italian citizenship through the faster procedure that is accorded to those married to Italian citizens.

2018 varies systematically with local pre-determined characteristics by estimating the following fixed effects model:

$$Y_{jt} = \alpha + \beta Refugee\ Share_{jt} + \gamma Refugee\ Share_{jt} \times Z_{j0} + \mu_j + \delta_t + \varepsilon_{jt}, \tag{1}$$

where the outcome variable Y_{jt} represents our backlash measure, i.e., the vote share (over the total number of voters) for anti-immigration parties in municipality j at time t. We observe vote shares at the national elections in 2013 (before refugee-crisis) and 2018 (after refugee-crisis). Our measure of refugee exposure, $Share\ of\ Refugees_{jt}$, is defined as the share of refugees assigned to municipality j at time t over the resident population in 2013.[26] Specifically, we consider *only* refugees hosted by the CAS reception system, which was rapidly created in 2014 as described in Section 2. Therefore our explanatory variable, $Share\ of\ Refugees_{jt}$, takes value zero in the 2013 pre-crisis year in all municipalities j, while it is equal to the sum of the capacity of all CAS centers in a municipality at the end of 2017 (two months before the elections) in the post-crisis year, representing the change in refugee exposure over time.[27] By considering the assigned rather than the actual number of refugees living in municipality j, our estimated treatment effects represent intention-to-treat effects. In principle, refugees have the option to relocate from their assigned location. However, legal provisions limiting access to reception facilities, legal advice, and economic support exclusively within the center of the initial assignment strongly discourage such relocations. Furthermore, compliance with the assigned location is bolstered by the protracted Italian asylum application process, lasting from 18 to 24 months (Chamber of Deputies, 2017). In a descriptive analysis, we exploit additional data on refugee location in 51 out of 92 provinces, uncovering a strong positive correlation of 0.96 between the assigned and actual refugee numbers.

To investigate the contribution of municipal characteristics, we interact our measure of refugee exposure, $Share\ of\ Refugees_{jt}$, with the vector Z_{j0} of pre-determined characteristics at the local level j at baseline time 0, i.e., before the crisis. To ease interpretation, the Z_{j0} variables are standardized with mean of zero and standard deviation of one.

Estimation (1) includes both municipality and time fixed effects. Municipality fixed effects, μ_j, capture time-invariant local observables and unobservables. Importantly, they absorb any static determinant of voting behavior, including the local historical presence of anti-immigration or extreme right parties, cross-sectional variation in the duration of refugee reception, the geographical municipality area, local infrastructure, as well as cultural and social norms. Time fixed effects, δ_t, instead, account for common shocks in a given year. ϵ_{jt} is an idiosyncratic error component. Standard errors are clustered at the municipality level.[28] Thus, our identification strategy exploits the within-municipality variation between the national elections in 2013 and 2018, before and after the refugee crisis respectively.

The main effect β identifies the impact of refugee exposure on the vote share for anti-immigration parties when $Z_{j0} = 0$. The parameter of interest γ, identifies the differential backlash effect based on local characteristics at baseline, Z_{j0}. Hence, the anti-immigration political response differs for various values of Z_{j0}, which we measure along both economic and socio-cultural municipal dimensions.

3.2. Identification strategy

Our identification prerequisite is that, conditional on time-invariant municipality effects μ_j, refugee assignment to CAS centers is not determined by past local political preferences nor by municipality-level shocks that simultaneously affect refugee allocation and voting behavior. In what follows, we provide evidence in support of our identification.

Overall, we document the lack of systematic correlation between refugee exposure and political outcomes at the local level before the refugee crisis. As a first piece of evidence, Table 2 reports correlations from the univariate cross-sectional regression of the share of votes for parties along the political spectrum and voters' participation on the share of refugees. Panel A of Table 2 shows that electoral outcomes in national elections in 2013, both for the Chamber of Deputies and the Senate, are not predictive of refugee assignment. We adjust p-values for multiple hypothesis testing to control for Family-wise Error Rate (FWER) by group of variables (column 4). We further restrict these tests by considering only within-province variation in local political outcomes, and results are robust to including province fixed effects in our specification (column 5).

As a second piece of evidence, we show that the allocation of refugees is orthogonal to long-term trajectories in local political preferences for anti-immigration parties. Specifically, we estimate the political response to refugee exposure by exploiting longitudinal variation in political election outcomes from 2001, 2006, 2008, and 2013 (before the crisis), and 2018 (after the crisis) via a fully flexible DiD model. We present the coefficients of the interactions between the time-invariant measure of the share of refugees assigned to municipality j over the resident population and the full set of year fixed effects β_t in Fig. 2, along with their corresponding confidence intervals. Notably, the figure shows no evidence of long-term differential trends in political preferences for the anti-immigration front prior to the refugee crisis across municipalities with varying levels of refugee exposure. At the same time, we document a positive and statistically significant effect of refugee exposure in 2018, following the refugee settlement. Specifically, a 1 percentage point increase in the refugee share leads to about 0.168 percentage point increase in anti-immigration votes. It is worth noting that the anti-immigration front evolved remarkably over this period. For example, *BoI*, which did not exist until 2012, emerged and gained consensus quickly, while others disappeared from the political arena. Fig. 2 considers the vote shares obtained by those parties that presented anti-immigration stances in each election. As a further check, we also focus on the *League*, the only party existing throughout the period. Appendix Figure A5 confirms the absence of political pre-trends for this specific party.

Identification of γ in Eq. (1) also requires that the assignment of refugees is not systematically correlated with local municipal characteristics in the pre-treatment period. Similarly to our previous exercise, we show that refugee exposure is balanced with respect to local institutional outcomes. Panel B of Table 2 shows that refugee exposure is, first of all, unrelated to the presence of a SPRAR reception center in the municipality, nor to its capacity, confirming that the two systems are managed by different and independent authorities and respond to different incentive schemes. We also collect data from municipal elections from 2008 and 2012 and show that municipalities that elected a mayor belonging to anti-immigration parties are not systematically different in terms of refugee assignment. Moreover, refugee assignment is orthogonal to the municipality's welfare generosity (overall and toward immigrants specifically) and the quality of local institutions, proxied by an indicator for municipality administrations being under receivership in 2007–13. Eventually, we show that refugee assignment is independent

[26] We have no precise data on refugee administrative registrations. While some municipalities registered CAS refugees among the resident population (generally the rule since it guarantees access to essential health and social services), others did not. Depending on the number of assigned refugees, this registration issue may generate inconsistencies in the population size across municipalities. For this reason, we compute the share of refugees over the total municipal population in 2013, just before the introduction of the CAS system.

[27] As robustness check, we also use the maximum share of asylum seekers hosted at any point in time between 2014 and 2017 as our end-line observation. Results are qualitatively the same (available upon request).

[28] Results are robust when clustering standard errors at the province level to account for the fact that the party may run the same candidate in various municipalities within the province (available upon request).

Table 2
Balancing tests.

Baseline variables:	(1) Share of refugees in 2017	(2) Std. err.	(3) *p*-value	(4) *p*-value FWER	(5) *p*-value FWER w/ prov FE
A. Political outcomes					
Chamber of deputies:					
Anti-immigration (%)	−0.005	0.058	0.935	1.000	0.544
League (%)	−0.012	0.057	0.833	0.998	0.600
BoI (%)	0.005	0.015	0.749	0.997	0.997
PDL (%)	−0.002	0.051	0.968	1.000	0.997
M5S (%)	−0.104	0.055	0.061	0.372	0.094
Center-left (%)	0.110	0.071	0.125	0.528	0.125
Election turnout (%)	−0.079	0.066	0.237	0.733	0.145
Senate:					
Anti-immigration (%)	0.011	0.067	0.869	0.998	0.467
League (Nord) (%)	0.005	0.062	0.937	1.000	0.560
BoI (%)	0.004	0.018	0.830	0.998	0.989
PDL (%)	−0.014	0.050	0.775	0.997	0.996
M5S (%)	−0.089	0.053	0.098	0.492	0.112
Center-left (%)	0.147	0.085	0.087	0.472	0.145
Election turnout (%)	−0.089	0.067	0.187	0.640	0.115
B. Institutional context					
Municipality hosted a SPRAR	−0.001	0.001	0.288	0.870	0.983
Share of refugees in SPRAR (%)	0.000	0.002	0.907	0.968	0.983
Municipality under receivership 2007–2013	−0.002	0.002	0.250	0.870	0.787
Municipality expenditure (log)	−0.017	0.010	0.084	0.694	0.783
Municipality expenditure for immigration services (log)	−0.008	0.015	0.573	0.968	0.983
Votes for Lega candidate (%) in latest municipality elections	−0.006	0.004	0.115	0.738	0.787
Lega mayor in charge	−0.013	0.008	0.114	0.738	0.787
Mafia presence 1982–2013	0.002	0.004	0.686	0.968	0.787
Mafia crime rate 2004–2013	0.000	0.000	0.618	0.968	0.787
Crime rate 2004–2013	0.002	0.002	0.239	0.870	0.849
C. Economic and demographic characteristics					
Income per capita (log)	−0.0001	0.002	0.718	0.838	0.283
Activity rate	−0.191	0.062	0.003	0.043	0.007
Employment rate	−0.160	0.071	0.027	0.163	0.007
Rent prices sqm. (log)	−0.023	0.016	0.168	0.521	0.155
Tertiary education rate	−0.024	0.021	0.263	0.607	0.660
Population over 65 (%)	0.201	0.063	0.002	0.043	0.011
Natives net migration flows (%)	−0.008	0.015	0.603	0.838	0.660
D. Social capital					
Referenda turnout	0.020	0.058	0.733	0.739	0.429
Volunteers (% pop.)	0.164	0.092	0.078	0.203	0.147
AVIS branch	−0.014	0.003	0.000	0.003	0.004
E. Intergroup contact					
Share of immigrants (% pop.)	−0.037	0.044	0.401	0.880	0.591
Residential segregation index	0.082	0.089	0.356	0.880	0.914
Naturalization rate	0.063	0.106	0.554	0.896	0.914
Intermarriage rate	0.000	0.001	0.783	0.945	0.914
Foreign-born administrators	−0.005	0.003	0.128	0.555	0.717
Non-EU15 born administrators	0.000	0.003	0.903	0.945	0.943

Note: This table shows balance tests of pre-treatment local municipality-level characteristics on refugee exposure. Each row reports OLS estimates and standard errors from the univariate cross-sectional regression of the share of refugees on local pre-treatment variable in column 1 and 2. Column 3 reports the *p*-value of these regressions. Column 4 reports *p*-values adjusted for multiple hypothesis testing by group of outcomes using the free step-down resampling method with 10,000 bootstrap repetitions (Westfall and Young 1993) to control the family-wise error rate (FWER). Column 5 reports *p*-values adjusted for multiple hypothesis testing (as before) from the univariate cross-sectional regressions including province fixed effects.

of criminal activity at the local level, and more specifically from Mafia infiltration.[29]

Finally, in panels C to E of Table 2, we report balance tests for the vector of our predetermined local economic and socio-cultural drivers Z_{j0}. Refugee exposure is not significantly correlated with our economic and human capital indicators. The only exceptions are represented by the activity rate and the share of the population over 65 years old. These two measures are strongly correlated (with a correlation of about 60 percent) since they both reflect the age structure of the resident population.[30] Statistical imbalance is due to the fact that the Italian population is unevenly distributed across locations by age, i.e., younger residents tend to concentrate in larger urban locations, whereas elderly people are

[29] We identify Mafia presence following Dipoppa (2021). The indicator combines three sources of data: the list of goods, properties, and firms seized from Mafias from 1982 to 2013; an indicator for city councils dissolved due to Mafia infiltration from 1991 to 2013; and Mafia-related victims (from VittimeMafia.it). In addition, we measure the intensity of Mafia infiltration as the number of Mafia related criminal episodes occurring from 2004 to 2013 at the local level, following Pinotti (2015) and Alesina et al. (2018).

[30] The activity rate is computed as the share of individuals over 15 years old who are either working or actively looking for a job. Thus, municipalities with a higher fraction of the elderly population report lower activity rates.

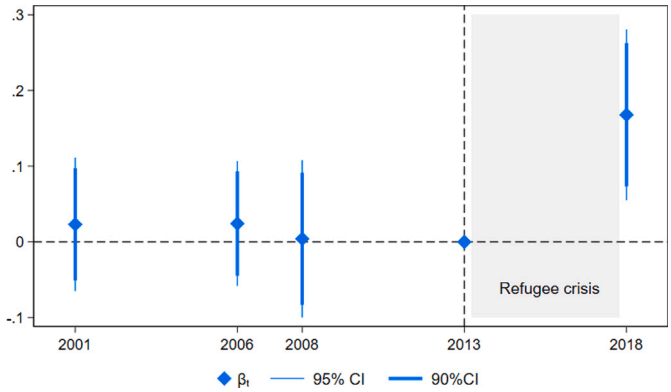

Fig. 2. Trends in election results for anti-immigration parties (2001–2018).*Notes:* This Figure shows the effects of refugee exposure on vote shares for anti-immigration parties, computed as in equation (1). The outcome variable is the share of votes for anti-immigration parties in the political elections in 2001, 2006, 2009, 2013 (before the crisis) and 2018 (after the crisis). Anti-immigration parties include: Northen League, Alleanza Nazionale, Forza Nuova, Fiamma Tricolare in 2001; Northen League, Alleanza Nazionale, Alleanza Sociale Mussolini, Fiamma Tricolare in 2006; Northen League, Forza Nuova, Alleanza Sociale Mussolini, Fiamma Tricolare–Destra Sociale in 2008; Northen League, Brothers of Italy, Forza Nuova, Casa Pound in 2013. The main explanatory variables are interactions between the time-invariant treatment variable $RefugeeShare_j$ (i.e., the share of refugees assigned to municipality j) and a full set of year fixed effects. We take year 2013 as reference. Estimation includes municipality and year fixed effects. The Figure plots the estimated coefficients and associated confidence intervals, based on standard errors clustered at the municipality level.

randomly dispersed across all areas. Hence, refugees' assignment induced by the dispersal policy mechanically mirrors the residential dispersion of elderly people. Importantly, for our context of analysis, even if the opportunity cost of hosting refugees might differ across locations, we show that refugee assignment is orthogonal to rent prices (per sqm) and natives' migration flows in panel C of Table 2. We also find no correlation with either social capital or intergroup contact indicators, with the only exception of the presence of an AVIS branch, hindered by measurement error issues in panels D and E of Table 2.[31] Similarly, as before, column (5) of Table 2 shows that the coefficients of the balance tests are robust after introducing province fixed effects. The only difference concerns employment rates showing a significant negative correlation with the share of assigned refugees. However, this element is in line with the demographic trend discussed above.

4. Results

We report estimates of Eq. (1) in Fig. 3 and in Tables A6–A9 in Appendix, where column (1) reports the impact of refugee assignment on vote share for anti-immigration parties, while columns (2) and (3) show results for the *League* and *BoI* in turn. Treatment effects are expressed in terms of relative differences from the baseline outcome in the pre-crisis period.[32]

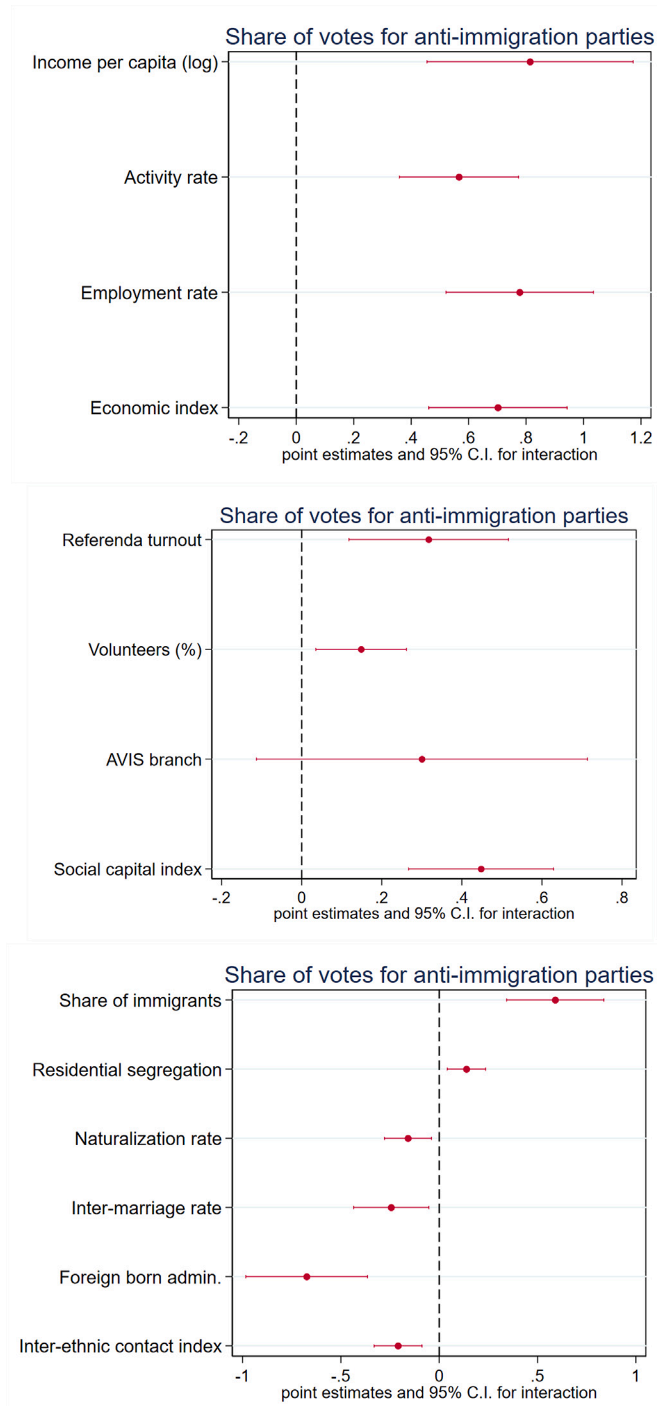

Fig. 3. Political Backlash by Local Characteristics.*Notes:* This figure shows the estimated effect on vote share for anti-immigration parties of refugee exposure interacted with local economic characteristics, social capital and intergroup contact measures at baseline. Effects are estimated using the fixed effects model in (1). Regressions include municipality and time fixed effects. The graph plot the estimated coefficients and associated confidence intervals, based on standard errors clustered at the province level. Point estimates are reported in Tables A6, A8, and A9.

In the first panel in Fig. 3, we report interaction effects of our measure of refugee exposure with a set of economic baseline characteristics at the municipality level, Z_{j0}, i.e., income per capita, activity rate, employment rate and an index combining these measures. We find that

[31] As mentioned in Section 2.3, we measure municipal blood donation with an indicator for local AVIS branch supply. Therefore, this dummy variable is quite 'broad' as an indicator, and it may capture many different local features, including municipality size or remoteness. For this reason, related to measurement error, we may observe some imbalance in the share of refugees across AVIS-flagged municipalities. A continuous variable for blood donation at the municipality level is not available.

[32] Figure A6 shows estimates for national elections for the Senate, reporting similar results despite involving a selected and slightly older electoral pool (adults above 25 years old).

while an increase 9 in the share of refugees increases vote shares for anti-immigration parties, this effect is significantly larger in economically better-off municipalities. Specifically, when the (log) income per capita increases by one standard deviation at the mean, the effect of one p.p. increase in the share of assigned refugees shifts from 0.16 to 0.97 p.p. (see Table A6). This effect is statistically significant and non-negligible in magnitude as it approximately amounts to 10 percent of the average vote share for anti-immigration parties in the 2013 pre-crisis election. We document a similar pattern while testing the heterogeneity of the treatment effect across other local economic factors and the principal component index combining all these dimensions together. This supports the welfare dependency argument, suggesting that wealthier natives may be particularly averse to supporting refugees through the general welfare state (Facchini and Mayda, 2009; Dustmann et al., 2019) and exhibit weaker preferences for redistribution (Dahlberg et al., 2012; Alesina et al., 2023). *A priori*, the outcome of this heterogeneity analysis was ambiguous. Indeed, if low-skilled natives perceive refugee immigration as increasing competition in the labor market, we would expect an opposite result, i.e., higher backlash in lower-income areas. Our findings thus suggest that the welfare dependency argument prevails over labor market concerns, which is in line with expectations given the legal and factual constraints that limit refugee participation in the labor market.

Overall, variation in political responses to refugee exposure is driven by a significant change in the share of votes for *League*, while the change in vote share for *BoI* is orthogonal to local economic characteristics.[33] We investigate additional socio-demographic determinants regarding population size, age structure, and local educational standing in Table A7. Results show that a higher population size as well as a higher share of tertiary-educated municipality-residents significantly lower the impact of refugee exposure on vote share for anti-immigration parties.

Next, we explore the heterogeneity in anti-immigration political responses to social capital (second panel in Fig. 3). We find that social capital significantly exacerbates political backlash to refugee migration; see Table A8. One standard deviation shift at the mean of the turnout rate at referenda increases the effect of one p.p. increase in the share of refugees from 0.06 to 0.38 p.p., about 4 percent of the average vote share for anti-immigration parties at baseline. Results are consistent when considering additional measures of social capital commonly adopted in the literature, such as the number of volunteers belonging to non-profit organizations and the presence of a blood donation center in the municipality (AVIS branch). In this latter case, political responses are similar in magnitude but noisier due to the measurement issues we discussed above. Finally, we combine these different measures of social capital into a principal component index and results using this synthetic indicator are in line with our main conclusion. The evidence suggests that standard measures of social capital proxy for within-group connections, which may be inward-looking and create strong in-group loyalty (Putnam et al., 2000; Portes, 2000). Thus, in contexts characterized by marked cultural-ethnic boundaries, *bonding social capital* reinforces exclusive identities and possibly out-group antagonism (see also Satyanath et al., 2017). The unexpected surge in refugee inflows during the crisis, coupled with administrative hardships and fierce propaganda against immigration, makes cultural-ethnic boundaries more salient and exacerbates the in-group/out-group distinction.[34]

Finally, we explore the heterogeneity in anti-immigration political responses to interethnic connections, in line with the 'contact hypothesis', which posits that social interaction between different groups can be pivotal in reducing intergroup bias (Allport, 1954; Mousa, 2020). We exploit various proxies to capture the frequency and strength of interactions between natives and former immigrants, i.e., rates of intermarriage, naturalization, residential integration, and foreign-born candidates elected to local office; see the third panel in Fig. 3.

First of all, the backlash effect of refugee migration appears to be significantly higher in municipalities with higher share of former immigrants, representing a long-term exposure to cultural diversity; see also panel B in Table A9.[35] By contrast, the backlash is significantly lower in municipalities with higher evidence of integration of former immigrants. For instance, if the intermarriage rate increases by one standard deviation at the mean, then the effect of a one p.p. increase in the share of refugees reduces the vote share for anti-immigration parties by 0.24 p.p. A similar pattern is observed in municipalities with foreign-born elected administrators. Results align with our conclusions when we combine the above measures into a principal component index. Our findings suggest that *bridging social capital* and meaningful intergroup contact between natives and immigrants can mitigate anti-immigration backlash. On the other hand, ethnic exposure *per se*, as measured by the share of immigrants at the local level, does not necessarily entail meaningful cooperation. In line with the contact hypothesis, intergroup interactions under particular conditions of actual integration (e.g., similar status, common goals, and support from social and institutional authorities) can reduce prejudice and negative attitudes toward immigrant minorities.

4.1. Robustness

We run a set of additional estimations to assess the robustness of our findings. We assess that our results are not driven by selective attrition in our sample; see Table A3 and the discussion in Section 2.1. Moreover, the results are not driven by selection into voting participation since refugee exposure does not predict the willingness to vote (voter turnout) for either the Chamber of Deputies or the Senate; see Table A10.

We prove our results to be robust to alternative specifications. Figure A7 shows that results remain unchanged when including a time-varying control for the share of population over 65, the only local factor that displays statistical imbalance at baseline.[36] Importantly for our design, findings are robust to excluding municipalities with SPRAR centers; see Figure A8. In addition, we trim our outcome variable to account for the fact that the change in vote share for anti-immigration parties is by definition bounded, and may therefore have a different effect on municipalities with strong preferences for anti-immigration parties at baseline, i.e., in the top percentiles of the anti-immigration distribution. Figure A9 shows that results are robust to excluding the sample above the 90th percentile of the outcome distribution at baseline.

To quantify our estimates, we run placebo regressions. Specifically, we construct a counterfactual scenario by randomly reassigning the share of refugees across municipalities within the same province. We replicate this counterfactual random assignment exercise for thousand times and, with each replication, we estimate the backlash effect of

[33] The *League* and *BoI* have similar programs and target the same electorate. Yet, in 2018, *BoI* got only 4 percent of the votes and ran as a junior far-right partner in a conservative coalition with Matteo Salvini's *League*. Hence, the bulk of backlash heterogeneity comes from the *League*, the major right-wing party at the time. Notably, the situation flipped in 2022 when, during new elections, *BoI* drew most of its newly found support from *League* voters and barged into power in coalition with a more depowered *League* party. This suggests that heterogeneity is likely to come from shifting votes rather than hard-core identity votes.

[34] Other works have explored the "dark side" of social capital. Satyanath et al. (2017) study the downfall of democracy in interwar Germany and show that areas with denser networks saw a more rapid rise of the Nazi Party. They show that the result that social capital undermined Germany's first democracy and aided the rise of the Nazi movement holds in areas with unstable governments and weak political contexts; see also Acemoglu et al. (2014).

[35] One standard deviation increase at the mean of the share of immigrants increases the effect of one p.p. increase in the share of refugees from 0.42 to 1 p.p., about 7 percent of the average vote share for anti-immigration parties at baseline.

[36] In addition, our results are robust to the introduction of other time-varying controls, i.e., the (log) income per capita, the (log) municipality per-user expenditure for public services, and the total resident population. Results are available upon request.

refugee exposure with Eq. (1). Panel A in Figure A10 reports the distribution of placebo effects of refugee exposure; the average exposure effect is precisely estimated around zero, with mean of 0.01 and a rejection rate of 12 percent. Panel B proposes a similar counterfactual exercise where we randomly redistribute the absolute number of refugees and calculate the exposure based on the local population. In this case, again, the average exposure effect is centered around zero with mean of 0.02.

Finally, we investigate the presence of non-linear effects on the vote share of anti-immigration parties with refugee exposure. Figure A11 presents non-parametric Nadaraya-Watson estimates of a first-difference model, with confidence intervals. The estimated effect is flat throughout the refugee share distribution, especially at the bottom of the distribution where the density is the highest. Overall, there is no evidence to support a significant non-linear backlash effect. Additional analyses also indicate that our effects are not driven by larger municipalities (see Figure A12) or municipalities hosting large reception centers or major governmental hotspots.[37]

4.2. Conditional average treatment effects (CATE)

To conduct consistent heterogeneity analysis for the political response to refugee exposure along all local dimensions, i.e., economic and socio-cultural, we estimate conditional average treatment effects (CATE) across municipalities using causal forest estimators (Athey and Imbens, 2016; Athey and Wager, 2019; Athey et al., 2019). Causal forest is a supervised machine learning technique based on several iterations across randomly-selected subsets of observations and covariates, which are used to construct causal trees. The algorithm builds each tree by recursively creating partitions of the data. The splitting point at each node is the value of one of the covariates that maximizes treatment effect heterogeneity. Essentially, the algorithm is trained to look for areas in the covariate space where the effect differs the most. This process is iterated until the observations are grouped into "leaves" with similar treatment effects. The causal forest eventually aggregates across the ensemble of estimated trees and provides estimates of treatment effects that are robust to subsample selection (Athey et al., 2023). Unlike the specification in Eq. (1), we estimate a first-differences model to partial out municipality fixed effects, hence computationally speeding up the procedure.[38]

Let β_j denote the estimates of the conditional average treatment effects on the change in vote shares for anti-immigration parties for all municipalities j in our sample. Figure A10 reports the distribution of β_j. Refugee exposure increases anti-immigration backlash in almost all municipalities (98 percent). The response varies considerably, ranging from a decrease of 0.07 to an increase of 0.16 p.p., twice the sample mean of CATE at 0.08 p.p.

To investigate treatment heterogeneity across baseline characteristics, we separately compare the average value of each characteristic for municipalities below and above the median distribution of predicted treatment effects. Results are reported in Table 3, together with standardized differences in the mean and p-values adjusted for multiple hypothesis testing (List et al., 2019). Overall, the standardized differences are significantly different from zero for all of the baseline characteristics, except for tertiary education rate, suggesting that observables explain variation in political responses. Consistent with our previous conclusions, municipalities with an above-median predicted backlash also have higher mean (log) per capita income, employment, and activity rate, as well as a higher share of the population above 65 years old. Moreover, larger municipalities exhibit lower responses, in line with the evidence on rural-urban gap in attitudes toward immigrants (Dustmann et al., 2019). We plot the average predicted effects across deciles of the baseline observables in Figure A14.

The results for social capital provide mixed evidence. On one hand, above-median predicted CATE is associated with higher social capital, i.e., higher referenda turnout rate and higher probability of hosting a blood donor center. On the other hand, NGO association density is higher among municipalities with below-median predicted backlash. Interestingly, from Figure A14, the predicted CATE follows a hump-shaped pattern for NGO association density, with the peak at the fourth decile of the distribution. Since association density refers to volunteers from any type of organization, this measure might include an out-group element deriving from previous positive interactions with immigrant minorities. Therefore, while referenda turnout and the presence of blood donor centers strictly capture the within-group or *bonding* component of social capital, association density may also encompass a *bridging* component that entails intergroup synergies and may plausibly attenuate backlash.

Consistent with linear estimates, Table 3 shows that the share of immigrants at the municipality level exacerbates political backlash to refugee exposure. Conversely, municipalities with below-median CATE display higher residential segregation (23.6 vs. 20), suggesting that mere intergroup spatial contact does not mitigate the anti-immigration backlash. The results for the remaining intergroup contact indicators – interethnic marriages, naturalization rate, and the presence of foreign-born local administrators – indicate that the baseline level of migrant integration attenuates negative political responses.

Finally, Fig. 4 shows that the anti-immigration backlash increases significantly with social capital conditional on any income level (panel A). Instead, the backlash varies non-linearly with different measures of socio-cultural integration. In particular, we estimate higher backlash at the middle of the income distribution in places with poor cultural integration, i.e., with significant exposure to immigration (panel B) or with a limited intermarriage rate (panel C). This evidence suggests that resettlement policies based on economic factors only – as is the case in most real-world scenarios – may not take advantage of the interactions with other local characteristics beneficial for refugee inclusion. We delve into the consequences of refugees' mismatch in Section 5 via counterfactual simulations.

5. Resettlement schemes and policy implications

The heterogeneous results discussed so far question random dispersal policies that may hamper minorities' integration due to a potential mismatch between refugees and receiving areas. In this section, we study alternative resettlement schemes to assign refugees to locations, accounting for local differences in backlash. We exploit CATE estimates of local responses to evaluate resettlement policies that aim to minimize anti-immigration backlash. Finally, we present a counterfactual policy evaluation.

In this matching exercise, we focus on the supply side of refugee integration, i.e., the acceptance (or lack thereof) of refugee minorities by natives in their societies. This supply component is fundamental for mitigating discrimination in the labor market and fostering social inclusion. This dimension complements the demand side of integration, i.e., the refugee's undertaking to integrate into the hosting society. The results of this analysis add to the existing literature by providing novel insights into the design of refugee resettlement policies. This partial equilibrium analysis is a complementary addition and a necessary step toward a more

[37] Results available upon request.

[38] To avoid over-fitting, we adopt the "honest" approach which entails dividing each training sample into two parts: half of the observations are used to grow the tree, i.e., performing the sample splits, while the other half is used to estimate the treatment effects. We train the causal forest algorithm to build 100,000 trees and we set to 20 the minimum number of observations in a leaf. The `causal_forest` function of the `R` package `grf` (generalized random forest) has a default value for the minimum leaf size equal to 5 observations. However, because our treatment variable is continuous, we decide to use a substantially higher value of this parameter to improve the precision of our estimates.

Table 3
Predicted CATE of refugee exposure.

Baseline characteristics	(1)	(2)	(3)	(4)
	Predicted treatment effects		Std. diff.	MHT p-value
	Below median	Above median	(1)–(2)	(1)–(2)
Income (log)	9.302	9.361	−0.227	0.001
Employment rate	44.389	45.434	−0.138	0.001
Activity rate	49.574	50.147	−0.091	0.001
Population	7867.521	6571.607	0.048	0.114
Share over 65	22.637	23.482	−0.153	0.001
Tertiary education rate dummy	0.492	0.507	−0.030	0.249
Referenda turnout	46.977	50.979	−0.670	0.001
NGO associations density	10.806	7.612	0.396	0.001
Blood donor centre	0.353	0.397	−0.092	0.001
Share of immigrants	5.543	6.882	−0.311	0.001
Residential segregation index	23.554	20.094	0.329	0.001
Naturalization rate	14.163	13.262	0.087	0.001
Intermarriage rate	0.116	0.104	0.147	0.001
Foreign-born local administrators	0.379	0.338	0.087	0.001

Note: This table shows average baseline local characteristics for municipalities with, respectively, below and above median Conditional Average Treatment Effect (CATE) of refugee exposure, in columns (1) and (2). For each municipality, CATE is estimated via a causal forest algorithm. Column (3) report the standardized difference, while column (4) reports p-values testing for differences across groups, while accounting for multiples hypothesis testing, as in List et al. (2019).

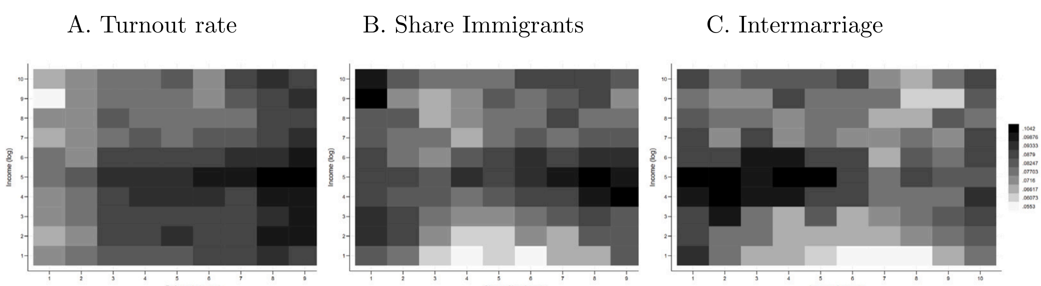

Fig. 4. CATE of Refugee Exposure, by Pairs of Characteristics. *Notes:* This Figure shows the average predicted Conditional Average Treatment Effects (CATE) over income and turnout rate (panel A), income and share of former immigrants (panel B), and income and intermarriage (panel C). For each municipality, CATE is estimated via a causal forest algorithm.

comprehensive general equilibrium model encompassing both demand and supply components.

5.1. The assignment problem

Setup. We propose a matching exercise to assign $i \in I$ refugees to $j \in J$ locations (municipalities), within each province $p \in P$. Refugees are all equal, whereas locations are heterogeneous in their response to refugee exposure β_j, as a result of pre-determined local differences in economic prosperity, as well as social and cultural structure. We estimate β_j in Section 4.2, as the change in the share of votes for anti-immigration parties due to a one percent increase in the share of assigned refugees within each municipality. Supported by validation results on ESS data in Section 2, we interpret this change in voters' preferences for anti-immigration parties as an expression of anti-immigration backlash, that is, a hardening of negative attitudes and behavior toward ethnic minorities induced by refugee exposure.

It should be noted that the local estimates, β_j, are obtained under an allocation scheme that is held invariant in the data. Consequently, the validity of the conceptual exercise that we pursue in this section, where we design various allocation schemes given the estimated β_j's, depends on whether these schemes do not induce voting reactions to the allocation policy change per se. For example, in alternative settings where policies allocate refugees to only a few provinces or target only a few municipalities within the national territory, this may be perceived as particularly unfair and thus trigger a backlash, not only because of the reception of refugees but also because of the perceived unfairness of the policy itself. While in principle our exercise rests on this restrictive assumption, our simulations address such potential limitations as we consider fairly neutral allocation schemes that lead to realistic refugee dispersion. Indeed, we maintain the aggregate number of refugees assigned to each province constant (as defined by the initial allotment plan) and we optimally reassign refugees across municipalities within the province.[39] As a result, the assignment schemes we consider do not manifest any uneven distribution of refugees, such as along the North-South or rural-urban dimension. Moreover, we propose alternative resettlement policies subject to various capacity constraints. In this respect, we trust that the results emerging from some specific counterfactual scenarios will guide our understanding of the design of inclusive reforms.

Matching problem. Given I refugees and J locations, $I \times J$ matches are potentially observed. A *matching* defines who (i) is matched to which location (j). Specifically, a matching is a measure μ_{ij} over the $I \times J$ space, such that $\mu_{ij} = 1$ if refugee i is assigned to location j, and zero otherwise. By considering a many-to-one matching framework, each refugee is assigned to only one location, but many refugees might be assigned to the same location. In this particular context, we assume all refugees entering the host country are assigned to a given location, hence nobody remains unmatched. The matching measure thus satisfies the following

[39] In this respect it is important that the β_j estimates refer to national elections where local – municipality level – considerations are relatively subdued.

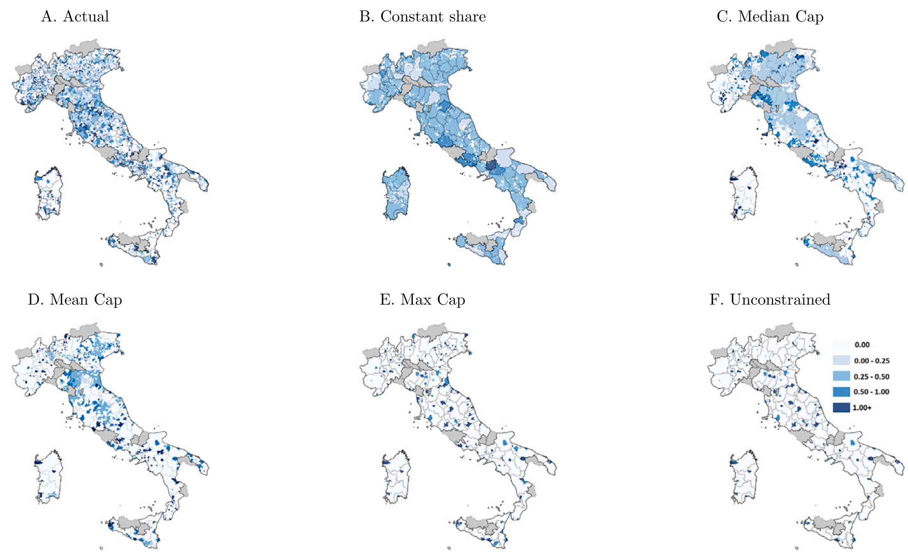

Fig. 5. Actual vs Counterfactual Simulated Distributions of Refugees. *Notes:* This Figure shows the distribution of the share of refugees assigned to Italian municipalities by the actual dispersal policy (panel A), and the optimal distribution of the share of refugees implied by the matching problem in (3), subject to diverse capacity constraints (panel B-F).

feasibility constraints:

$$\sum_{J} \mu_{ij} = 1 \quad \forall i \in I, \qquad \sum_{I} \mu_{ij} \leq \bar{M}_j \quad \forall j \in J, \tag{2}$$

where $\bar{M}_j$ (potentially) describes a refugee quota per location j, with $\bar{M}_j \leq I$.[40]

We take a social planner perspective, with the objective to minimize the probability of failure of resettlement policies, i.e., the *total* anti-immigration backlash for each province. Thus, the objective is to minimize the sum of the products between the marginal anti-immigration effect of one p.p. increase in the share of assigned refugees, β_j, and the share of refugees assigned to a given location j over the resident population, $\sum_I \mu_{ij}/Pop_j$. We consider the anti-immigration backlash as the best proxy available for opposition to the resettlement program and, ultimately, to refugees' integration. It is worth mentioning that the objective function of our analysis does not alter political preferences *per se*, but seeks to minimize the effect on preferences for anti-immigration parties *only due to* refugee assignment.

The *optimal* matching μ_{ij} turns out to be the solution to the following (social) minimization problem over all potential matches:

$$\min_{(\mu_{ij} \geq 0)} \quad \sum_{j \in J} \sum_{i \in I} \frac{\mu_{ij}}{Pop_j} \beta_j \tag{3}$$

subject to feasibility constraints in (2), and possibly different capacity constraints. We consider progressively less stringent (exogenous) capacity constraints, up to the extreme (unrealistic) assignments imposing full refugee concentration. Specifically, we impose for each municipality a limit to the number of assigned refugees equal to (i) the observed average share at province level (constant share); (ii) the median capacity observed within province; (iii) the mean capacity observed within province; (iv) the max capacity observed within province; (v) no capacity constraints.

From the problem in (3), some considerations emerge. First, the optimal matching from (3) trade-offs the contribution of two components: (i) the local characteristics affecting anti-immigration responses and (ii) the size of the resident population in j. For instance, if the anti-immigration responses were homogeneous across municipalities, $\beta_j = \bar{\beta} \quad \forall j$, the model would simply assign refugees to the most populated municipalities within each province. Vice versa, if the anti-immigration responses were independent of the local population, the model would assign refugees to the municipalities with lower marginal anti-immigration return β_j. Second, the response to the marginal refugee exposure is constant within municipality and equal to β_j/Pop_j. Hence, in absence of capacity constraints, the optimal matching simply assigns all refugees to the municipality with the lowest marginal effect of β_j/Pop_j, within each province of reference.

5.2. Counterfactual resettlement policies

In this section, we evaluate the random dispersal policy implemented in Italy by comparing it with different counterfactual assignments. In the comparison, we keep the aggregate number of refugees assigned to each province constant, as defined by the policy allotment plan of 2.5 refugees per 1000 inhabitants. Hence, we evaluate counterfactual assignments of refugees across municipalities within province.

Based on β_j estimates from causal forest analysis, we solve for the optimal matching μ_{ij} in (3) across municipalities in all provinces p, subject to different capacity constraints. Results in Fig. 5 report the share of refugees assigned to each location (municipality) under the observed random dispersal policy (panel A) and the counterfactual assignments (panels B to F). As a result of different capacity constraints, the optimal refugee assignment moves from a scattered and widespread refugee exposure across municipalities to a narrow and targeted allocation, reaching a single municipality assignment per province under the unconstrained assignment. Table 4 reports summary statistics for optimal assignment policies. In order, the assignment under the constant share rule leads to a refugee dispersion of 80 percent, i.e., 80 percent of municipalities receive at least one refugee. The refugee dispersion lowers, on average, to 37 percent under the median capacity policy (in line with the dispersion rate observed under the random dispersal policy of 38 percent) and shrinks to 1 percent under the unconstrained policy.

[40] The matching design in Eq. (2) considers the municipality as the unit of analysis for counterfactual relocation. As such, its policy implications cannot be extended to the relocation of refugees across smaller geographical sub-units within each municipality.

Table 4
Summary statistics-counterfactual assignments.

	Count	Mean	Sd	Min	Max
Panel A. Actual refugee distribution					
Share of refugees	6891	0.40	1.60	0.00	61.31
Share of refugees no zero	2624	1.05	2.46	0.01	61.31
Number of refugees no zero	2624	54.70	150.79	1.00	4000
Municipality with CAS	6891	0.38	0.49	0.00	1.00
Panel B. Simulated refugee distribution-Constant share					
Share of simulated refugees	6891	0.27	0.19	0.00	1.14
Share of simulated refugees no zero	5512	0.34	0.15	0.02	1.14
Number of simulated refugees no zero	5512	26.04	69.54	1.00	2151
Municipality with simulated CAS	6891	0.80	0.40	0.00	1.00
Mismatch rate	6891	0.48	0.50	0.00	1.00
Treated (at baseline) to control	6891	0.03	0.17	0.00	1.00
Control (at baseline) to treated	6891	0.45	0.50	0.00	1.00
Panel C. Simulated refugee distribution-Median capacity					
Share of simulated refugees	6891	0.18	0.51	0.00	22.06
Share of simulated refugees no zero	2540	0.49	0.75	0.03	22.06
Number of simulated refugees no zero	2540	56.51	149.28	1.00	3573
Municipality with simulated CAS	6891	0.37	0.48	0.00	1.00
Mismatch rate	6891	0.31	0.46	0.00	1.00
Treated (at baseline) to control	6891	0.16	0.37	0.00	1.00
Control (at baseline) to treated	6891	0.15	0.35	0.00	1.00
Panel D. Simulated refugee distribution-Mean capacity					
Share of simulated refugees	6891	0.09	0.52	0.00	22.06
Share of simulated refugees no zero	796	0.78	1.33	0.02	22.06
Number of simulated refugees no zero	796	180.33	378.53	5.00	4923
Municipality with simulated CAS	6891	0.12	0.32	0.00	1.00
Mismatch rate	6891	0.32	0.47	0.00	1.00
Treated (at baseline) to control	6891	0.29	0.46	0.00	1.00
Control (at baseline) to treated	6891	0.03	0.16	0.00	1.00
Panel E. Simulated refugee distribution-Max capacity					
Share of simulated refugees	6891	0.14	1.59	0.00	61.35
Share of simulated refugees no zero	168	5.59	8.55	0.06	61.35
Number of simulated refugees no zero	168	854.43	859.10	11.00	4969
Municipality with simulated CAS	6891	0.02	0.15	0.00	1.00
Mismatch rate	6891	0.37	0.48	0.00	1.00
Treated (at baseline) to control	6891	0.36	0.48	0.00	1.00
Control (at baseline) to treated	6891	0.01	0.08	0.00	1.00
Panel F. Simulated refugee distribution-Unconstrained					
Share of simulated refugees	6891	0.38	8.23	0.00	456.85
Share of simulated refugees no zero	91	28.44	66.15	0.23	456.85
Number of simulated refugees no zero	91	1577.42	983.27	385.00	5240
Municipality with simulated CAS	6891	0.01	0.11	0.00	1.00
Mismatch rate	6891	0.37	0.48	0.00	1.00
Treated (at baseline) to control	6891	0.37	0.48	0.00	1.00
Control (at baseline) to treated	6891	0.00	0.05	0.00	1.00

Note: This table shows summary statistics for refugee assignment across Italian municipalities implied by the actual dispersal policy (panel A), and the optimal distribution of the share of refugees implied by the matching problem in (3), subject to diverse capacity constraints (panel B-F).

We report the variation in anti-immigration backlash implied by the optimal assignment policies in Fig. 6. Three results emerge. First, regardless of the constraints, all refugee assignments guarantee a sizable reduction in anti-immigration backlash compared to the random dispersal policy, ranging from a backlash drop of 34 percent under the constant share policy (corresponding to a reduction from 2.4 to 1.60 in average backlash) to a drop of more than 90 percent under the mean capacity reassignment (from 2.4 to 0.26 in backlash). Second, the anti-immigration backlash even reverses when imposing none or minimal capacity constraints, resulting from a change in pro-immigration attitudes for a non-zero proportion of municipalities throughout Italy. Finally, we provide evidence of a policy trade-off between a reduction in anti-immigration backlash and refugee geographical dispersion, highlighted in the second vertical axis of Fig. 6. Indeed, the lower the refugee dispersion, the more effective the policy is in stemming anti-immigration backlash.

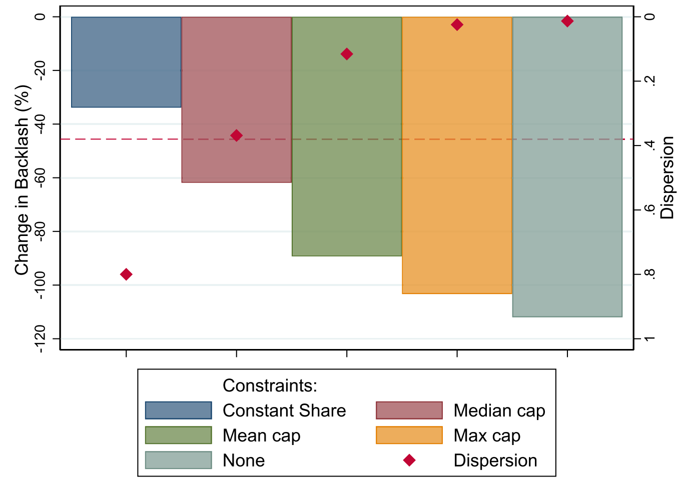

Fig. 6. Change in Backlash and Refugee Dispersion under Optimal Assignments. *Notes:* This Figure shows the change in anti-immigration backlash and the change in refugee dispersion across Italian municipalities computed under the optimal assignment policies (for diverse capacity constraints) with respect to the actual dispersal policy in place. The red dashed line reports the average dispersion rate under the random dispersal policy. (For interpretation of the references to colour in this figure legend, the reader is referred to the web version of this article.)

The reduction in anti-immigration backlash generated under different resettlement schemes results from a change in both the intensive and extensive margins of refugee assignment. For instance, we observe a significant difference in the intensity of refugee exposure for municipalities treated at baseline and, at the same time, a change in refugee assignment across municipalities. Precisely, we quantify the mismatch rate in the assignment as the share of municipalities treated under the random policy that would not have been treated under the optimal allocation and, vice versa, the share of municipalities not treated under the random policy that would have been treated under the optimal allocation. Table 4 reports the results. While under some policy constraints, the degree of refugee dispersion mechanically translates into a high mismatch rate; there are two non-trivial cases worth discussing. The refugee allocation under the median capacity policy, while keeping constant the dispersion rate compared to baseline, implies a mismatch rate of 31 percent, with 15 percent of municipalities becoming treated under the optimal assignment; see panel C of Table 4. Similarly, the allocation with mean capacity constraints implies a mismatch rate of 32 percent, with 29 percent of municipalities that would not have been treated under the optimal allocation.

We investigate how the change in refugee assignment is correlated with local economic and socio-cultural characteristics. Table 5 reports the results of balance tests for the vector of our predetermined local economic and socio-cultural drivers Z_{j0} on the change in refugee exposure under the optimal constrained policies, conditional on positive treatment. We focus on two intermediate assignments subject to median and mean capacity constraints. Optimal policies induce, on average, a significant re-assignment of refugees from rich and socially connected areas to less affluent but more culturally integrated municipalities.

One might argue that our assumption of a linear backlash effect β_j is restrictive. However, we tested for non-linear responses and provide evidence supporting our model assumption in Section 4.1. The variation in refugee share we can observe in the data might provide limited support for non-linear effects when the assignment implies a high refugee concentration in a few locations. While non-linearity might bias our backlash estimates upward in the extreme case of highly concentrated policies (e.g., in panel F of Fig. 5), we show a significant reduction in backlash even in the opposite case of highly dispersed policies (e.g., in panels B and C of Fig. 5) when non-linearity issues become negligible.

Table 5
Correlation between change in refugee share and baseline characteristics.

	(1)	(2)	(3)	(4)
Exp. variable:	Δ Share of refugees			
Capacity constraint:	Median cap		Mean cap	
Economic index PCA	−1.36**	0.028	−1.39**	0.023
	(0.61)		(0.60)	
Income per capita (log)	−0.05**	0.028	−0.06**	0.014
	(0.02)		(0.02)	
Activity rate	−0.95**	0.048	−0.90**	0.036
	(0.47)		(0.42)	
Employment rate	−1.27**	0.018	−1.37**	0.017
	(0.53)		(0.56)	
Bonding social capital index PCA1	−0.79*	0.078	−1.16**	0.041
	(0.44)		(0.56)	
Average turnout	−1.17*	0.055	−1.30**	0.042
	(0.60)		(0.63)	
Association density (%)	0.20	0.649	−0.40*	0.099
	(0.43)		(0.24)	
AVIS branch in 2010	−0.08**	0.046	−0.06***	0.006
	(0.04)		(0.02)	
Bridging social capital index PCA	0.29	0.271	0.28*	0.094
	(0.26)		(0.17)	
Intermarriage rate	1.14**	0.042	1.05**	0.030
	(0.55)		(0.48)	
Naturalization rate	0.23	0.488	0.20	0.502
	(0.33)		(0.29)	
Share of foreign born	−0.65*	0.077	−0.64**	0.033
	(0.36)		(0.29)	
Observations	2003		678	

Note: This table shows OLS estimates of the effect of the change in refugee share on local economic and socio-cultural characteristics. Columns (1)–(2) and (3)–(4) present the results considering the optimal refugee assignment subject to median and mean capacity constraints, respectively. Columns (1) and (3) report coefficient estimates and standard errors, while columns (2) and (4) report p-values. Standard errors clustered at the province level are reported in parentheses. Significance levels: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

5.3. The contribution of culture

How important are socio-cultural *vs* economic characteristics for refugee assignment? We provide an answer to this question by investigating an alternative resettlement scheme that assigns refugees to locations *only* according to economic and population heterogeneity across municipalities. By neglecting the importance of the socio-cultural structure, we show that the policy is less effective in mitigating anti-immigration backlash.

In many Western countries, refugees have until recently been relocated according to the resident population or, eventually, based on economic characteristics (Hatton, 2013, 2016). The leading example is the Dublin regulation's proposed reform, setting the criteria for refugee assignment across European countries during the refugee crisis based on population, GDP per capita, and unemployment. Similar policies have been enacted to assign refugees within countries based on a quota system according to the resident population, as in Germany, the Netherlands, the UK, Norway, Denmark, and Sweden (Andersson et al., 2018; Dumont et al., 2016). The principle of solidarity and fair sharing of responsibility guides these assignment schemes. However, there is no evidence of their longer-term impact on immigration acceptance.

We assess the effect of these assignment schemes, simulating the backlash when refugees are assigned based on resident population and economic factors only. Specifically, we estimate conditional average treatment effects across municipalities, as in Section 4.2, accounting for heterogeneity in observed economic drivers (income per capita, employment, and activity rate) and the resident population at baseline. Thus, we obtain a new β_j^{ec} distribution. Figure A15 reports the correlation between our main β_j estimates computed conditioning on economic, social capital, and intergroup contact drivers and β_j^{ec} estimates limiting heterogeneity to economic drivers and population only. The correlation is about 0.19; the high volatility between the two estimates potentially leads to significant differences in refugee assignment.

Based on β_j^{ec}, we solve for the matching μ_{ij}^{ec} in (3) for all provinces p, subject to the same capacity constraints as above. We compare the change in anti-immigration backlash implied by initial assignment policies and simulated ones in Fig. 7 in turn. We keep as a reference the observed dispersal policy. On the one hand, assignment policies under the β_j^{ec} heterogeneity reduce anti-immigration backlash with respect to the dispersal policy in place (right panel). On the other hand, however, neglecting the role of the socio-cultural structure, this alternative scheme is less effective in minimizing anti-immigration preferences. For instance, the differential change in anti-immigration backlash amounts to 13 and 25 percent, under the median and mean capacity constraints. Interestingly, the differential change between the two matching schemes is even amplified under minimal capacity constraints. This result suggests that the mismatch in refugee assignment due to the difference between β_j and β_j^{ec} is more prominent under policy rules, leading to a narrow dispersion rate. Thus, while more concentrated policies are more efficient in minimizing anti-immigration backlash in cases of complete information on the true distribution of treatment effects, they are also riskier in cases of incomplete information about the relevant dimensions that matter for heterogeneity.

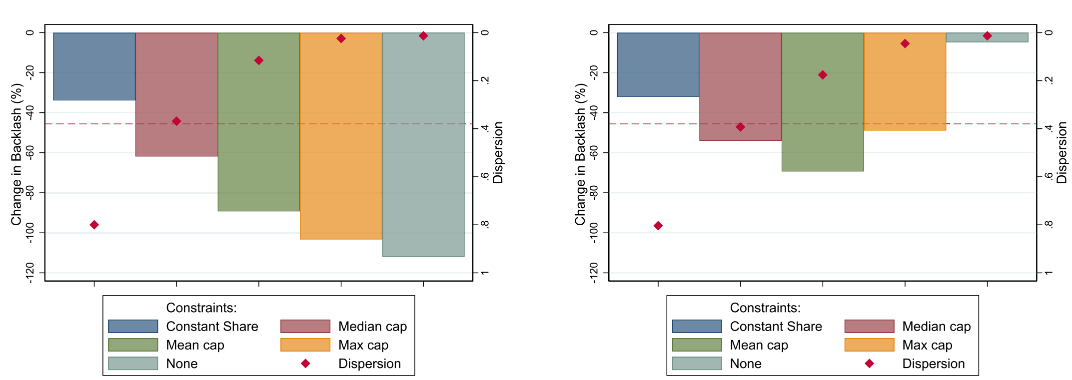

Fig. 7. Counterfactual Assignment–Comparison Between Different Assignment Criteria. *Notes:* This Figure shows the change in anti-immigration backlash and the change in refugee dispersion computed under the optimal assignment policies (for diverse capacity constraints) considering β_j and β_j^{ec} estimates accounting on economic drivers and population only. The red dashed line reports the average dispersion rate under the actual dispersal policy. Summary statistics are reported in Table A11. (For interpretation of the references to colour in this figure legend, the reader is referred to the web version of this article.)

6. Conclusions

Since the peak of the 2015 record of migration inflows to the EU, challenges to reception systems and efficient rules for refugee distribution have surged. Clearly, refugee exposure may engender adverse native responses, but whether these adverse effects can be prevented by restructuring the receiving conditions is an open question. Indeed, local characteristics of receiving areas may significantly impact refugee integration through the supply or denial of opportunities.

This paper investigates the heterogeneity in political response to refugee exposure across Italian municipalities and evaluates counterfactual assignment policies that minimize local backlash. We focus on the recent refugee crisis in Italy (2014–2017), during which unexpected inflows of about 150,000 asylum seekers per year from Northern Africa and the Near East fueled native hostility and increased the salience of ethnic boundaries. Within this setting, we harmonize a wide range of administrative data at a granular level and estimate the role of local economic and cultural mechanisms, distinguishing between economic prosperity, social capital, and intergroup interactions. By leveraging the quasi-random assignment of refugees across municipalities and controlling for potential unobserved local heterogeneity, we find that the impact of refugee exposure on anti-immigration backlash is significantly higher in more affluent areas and contexts with more bonding social capital. On the contrary, the anti-immigration political response is mitigated in areas marked by meaningful intergroup contact with former immigrants. Sizable heterogeneity in political responses across municipalities is also estimated via causal forest algorithm, putting into question the effectiveness of random allocation policies in limiting the political costs of refugee reception.

We exploit this pattern of heterogeneity by baseline local-level characteristics to evaluate novel resettlement schemes. We propose a matching framework to assign refugees to locations, accounting for local differences in backlash. We show that targeted resettlement schemes significantly reduce anti-immigration backlash compared to the dispersal policy in place, subject to different capacity constraints. Simulation results suggest that government policies dealing with the management and allocation of refugees should consider both economic and non-economic local characteristics to foster integration and stem hostility and backlash. Thus, socio-cultural factors may complement traditional dispersal policies based on population size, economic conditions, or social housing availability.

This is crucial for policy evaluation for two reasons. First, it seems feasible and desirable to leverage the extant stock of positive cross-cutting contact and bridging social capital to stem backlash and spur the supply of integration opportunities by natives toward refugees. Second, accurately identifying locations where mere refugee exposure triggers backlash can guide targeting programs promoting contact and meaningful interactions (Enos, 2017). Indeed, since these programs require grassroots initiatives that are generally costly and difficult to scale up in large natural contexts, improving the targeting strategy is of fundamental importance (Mousa, 2020).

Declaration of competing interest

We declare that there are no relevant or material financial interests related to the research described in the paper "Political Backlash to Refugee Settlement: Cultural and Economic Drivers" coauthored by Francesco Campo, Sara Giunti, Mariapia Mendola and Giulia Tura.

Appendix A. Supplementary data

Supplementary data to this article can be found online at doi:10.1016/j.jpubeco.2025.105467.

Data availability

The authors do not have permission to share data.

References

Acemoglu, D., Reed, T., and Robinson, J.A., 2014. Chiefs: economic development and elite control of civil society in Sierra Leone. J. Polit. Econ. 122 (2), 319–368.

Achard, P., Albrecht, S., Cettolin, E., Ghidoni, R., and Suetens, S., 2022. The effect of exposure to ethnic minorities on ethnic preferences.

Ager, A. and Strang, A., 2008. Understanding integration: a conceptual framework. J. Refug. Stud. 21 (2), 166–191.

Ahrens, A., Beerli, A., Hangartner, D., Kurer, S., and Siegenthaler, M., 2023. The Labor Market Effects of Restricting Refugees' Employment Opportunities.

Alesina, A., Miano, A., and Stantcheva, S., 2023. Immigration and redistribution. Rev. Econ. Stud. 90 (1), 1–39.

Alesina, A., Murard, E., and Rapoport, H., 2021. Immigration and preferences for redistribution in Europe. J. Econ. Geogr. 21, 925–954.

Alesina, A., Piccolo, S., and Pinotti, P., Jul 2018. Organized crime, violence, and politics. Rev. Econ. Stud. 86 (2), 457–499.

Alesina, A. and Tabellini, M., 2023. The political effects of immigration: culture or economics? J. Econ. Lit.

Algan, Y. and Cahuc, P., 2013. Trust, institutions and Economic development. Handb. Econ. Growth 1.

Allport, G.W., 1954. The nature of prejudice.

Andersson, T. and Ehlers, L., 2020. Assigning refugees to landlords in Sweden: efficient, stable, and maximum matchings. Scand. J. Econ. 122 (3), 937–965.

Andersson, T., Ehlers, L., and Martinello, A., 2018. Dynamic refugee matching.

Arendt, J.N., Dustmann, C., and Ku, H., 2022. Refugee migration and the labor market: lessons from 40 years of post-arrival policies in Denmark. Oxf. Rev. Econ. Policy 38 (3), 531–556.

Asimovic, N., Ditlmann, R., and Samii, C., 2022. Estimating the effect of intergroup contact over years: evidence from a youth program in Israel.

Athey, S. and Imbens, G., 2016. Recursive partitioning for heterogeneous causal effects. Proc. Natl. Acad. Sci. 113 (27), 7353–7360.

Athey, S., Simon, L.K., Skans, O.N., Vikstrom, J., and Yakymovych, Y., 2023. The heterogeneous earnings impact of job loss across workers, establishments, and markets. arXiv preprint arXiv:2307.06684.

Athey, S., Tibshirani, J., and Wager, S., 2019. Generalized random forests. Ann. Stat. 47 (2), 1148–1178.

Athey, S. and Wager, S., 2019. Estimating treatment effects with causal forests: an application. Observational Studies 5 (2), 37–51.

Bansak, K., Ferwerda, J., Hainmueller, J., Dillon, A., Hangartner, D., Lawrence, D., and Weinstein, J., 2018. Improving refugee integration through data-driven algorithmic assignment. Science 359 (6373), 325–329.

Battisti, M., Romiti, A., and Peri, G., 2022. Dynamic effects of Co-ethnic networks on immigrants' economic success. Econ. J. 132 (641), 58–88.

Bisin, A. and Tura, G., 2019. Marriage, fertility, and cultural integration in Italy. Working Paper 26303, National Bureau of Economic Research.

Boeri, T., Mishra, P., Papageorgiou, C., and Spilimbergo, A., 2021. Populism and civil society. Economica 88 (352), 863–895.

Bursztyn, L., Chaney, T., Hassan, T.A., and Rao, A., 2021. The immigrant next door: exposure, prejudice, and altruism. Natl. Bur. Econ. Res.

Campo, F., Giunti, S., and Mendola, M., 2024. Refugee crisis and right-wing populism: evidence from the Italian dispersal policy. European Economic Review, 168, 104826.

Cartocci, R., 2007. Mappe del Tesoro: Atlante del Capitale Sociale in Italia, vol. 168. Il Mulino.

Chamber of Deputies, 2017. Relazione sul sistema di protezione e di accoglienza dei richiedenti asilo. Commissione parlamentare di inchiesta. https://inchieste.camera.it/migranti/documenti.html?leg=17&leglabel=xvii%20legislatura

Coleman, J.S., 1988. Social capital in the creation of human capital. Am. J. Sociol. 94, S95–S120.

Coleman, J.S., 1990. Foundations of Social Theory. Harvard University Press.

Dahlberg, M., Edmark, K., and Lundqvist, H., 2012. Ethnic diversity and preferences for redistribution. J. Polit. Econ. 120 (1), 41–76.

Damm, A.P., 2009. Ethnic enclaves and immigrant labor market outcomes: quasi-experimental evidence. J. Labor Econ. 27 (2), 281–314.

Damm, A.P. and Rosholm, M., 2010. Employment effects of spatial dispersal of refugees. Rev. Econ. Househ. 8 (1), 105–146.

Danzer, A.M. and Yaman, F., 2013. Do ethnic enclaves impede immigrants' integration? Evidence from a quasi-experimental Social-interaction approach. Rev. Int. Econ. 21 (2), 311–325.

Davis, C.A., López-Peña, P., Mobarak, A.M., and Wen, J.Y., 2024. Refugees are hosted in highly vulnerable communities. in: AEA Papers and Proceedings, vol. 114. pp. 75–79.

Dinas, E., Matakos, K., Xefteris, D., and Hangartner, D., 2019. Waking up the Golden Dawn: does exposure to the refugee crisis increase support for extreme-right parties? Polit. Anal. 27 (2), 244–254.

Dipoppa, G., 2021. How criminal organizations expand to strong states: migrant exploitation and political brokerage in Northern Italy.

Dumont, J.-C., Liebig, T., Peschner, J., Tanay, F., and Xenogiani, T., 2016. How are refugees faring on the labour market in Europe? A first evaluation based on the 2014 EU labour force survey ad hoc module. European Commission DG Employment Working Paper (1).

Durante, R., Mastrorocco, N., Minale, L., and Snyder, J.M., 2025. Unpacking Social capital. Econ. J. 135 (667), 773–807.

Dustmann, C., Fasani, F., Frattini, T., Minale, L., and Schönberg, U., 2017. On the economics and politics of refugee migration. Econ. Policy 32 (91), 497–550.

Dustmann, C., Vasiljeva, K., and Piil Damm, A., 2019. Refugee migration and electoral outcomes. Rev. Econ. Stud. 86 (5), 2035–2091.

Eckert, F., Walsh, C., Hejlesen, M., et al, 2019. The return to big city experience: evidence from Danish refugees. Oppor. Incl. Growth Inst. Work Pap. Fed. Reserv. Bank Minneap. 24.
Edin, P.-A., Fredriksson, P., and Aslund, O., 2003. Ethnic enclaves and the Economic success of immigrants. Evidence from a natural experiment. Q. J. Econ. 118 (1), 329–357.
Edo, A., Giesing, Y., Öztunc, J., and Poutvaara, P., 2019. Immigration and electoral support for the far-left and the far-right. Eur. Econ. Rev. 115, 99–143.
Enos, R.D., 2017. the Space Between Us: Social Geography and Politics. Cambridge University Press.
Eurostat, 2020. Asylum and managed migration database, EU 27, 2008-2019.
Facchini, G. and Mayda, A.M., 2009. Does the welfare state affect individual attitudes toward immigrants? Evidence across countries. Rev. Econ. Stat. 91 (2), 295–314.
Fasani, F., Frattini, T., and Minale, L., 2022. (The struggle for) refugee integration into the labour market: evidence from Europe. J. Econ. Geogr. 22 (2), 351–393.
Fouka, V., 2022. Assimilation. Oxford Handbook of Historical Political Economy.
Giuliano, P. and Wacziarg, R., 2020. Who voted for Trump? Populism and Social capital. Technical report, National Bureau of Economic Research.
Godøy, A., 2017. Local labor markets and earnings of refugee immigrants. Empir. Econ. 52 (1), 31–58.
Guiso, L., Sapienza, P., and Zingales, L., 2004. The role of social capital in financial development. Am. Econ. Rev. 94 (3), 526–556.
Guiso, L., Sapienza, P., and Zingales, L., 2008. Social capital as good culture. J. Eur. Econ. Assoc. 6 (2–3), 295–320.
Guiso, L., Sapienza, P., and Zingales, L., 2011. Civic Capital as the Missing Link Handbook of Social Economics: North-Holland, vol. 1 pp. 417–480.
Guriev, S. and Papaioannou, E., 2022. The political economy of populism. J. Econ. Lit. 60 (3), 753–832.
Hainmueller, J. and Hopkins, D.J., 2014. Public attitudes toward immigration. Annu. Rev. Polit. Sci. 17, 225–249.
Halla, M., Wagner, A.F., and Zweimüller, J., 2017. Immigration and voting for the far right. J. Eur. Econ. Assoc. 15 (6), 1341–1385.
Hangartner, D., Dinas, E., Marbach, M., Matakos, K., and Xefteris, D., 2019. Does exposure to the refugee crisis make natives more hostile? Am. Polit. Sci. Rev. 113 (2), 442–455.
Hatton, T.J., 2013. Refugee and asylum migration. Int. Handb. Econ. Migr. 453–467.
Hatton, T.J., 2016. Refugees, asylum seekers, and policy in OECD countries. Am. Econ. Rev. 106 (5).
Jaschke, P., Sardoschau, S., and Tabellini, M., 2022. Scared straight? Threat and assimilation of refugees in Germany. Technical report, National Bureau of Economic Research.
List, J.A., Shaikh, A.M., and Xu, Y., 2019. Multiple hypothesis testing in experimental economics. Exp. Econ. 22 (4), 773–793.
Martén, L., Hainmueller, J., and Hangartner, D., 2019. Ethnic networks can foster the Economic integration of refugees. Proc. Natl. Acad. Sci. 116 (33), 16280–16285.
Mayda, A.M., Peri, G., and Steingress, W., 2022. The political impact of immigration: evidence from the United State. Am. Econ. J. Appl. Econ. 14 (1), 358–389.
Mousa, S., 2020. Building Social cohesion between christians and muslims through soccer in post-ISIS Iraq. Science 369 (6505), 866–870.
Onyx, J. and Bullen, P., 2000. Measuring Social capital in five communities. J. Appl. Behav. Sci. 36 (1), 23–42.
Paluck, E.L., Green, S.A., and Green, D.P., 2019. The contact hypothesis Re-evaluated. Behav. Public Policy 3 (2), 129–158.
Pinotti, P., 2015. The Economic costs of organised crime: evidence from southern Italy. Econ. J. 125 (586), F203–F232.
Portes, A., 2000. The two meanings of Social capital. in: Sociological forum, vol. 15. Springer, pp. 1–12.
Putnam, R.D., 1993. The prosperous community. Am. Prospect 4 (13), 35–42.
Putnam, R.D., et al, 2000. Bowling Alone: The Collapse and Revival of American Community. Simon and Schuster.
Satyanath, S., Voigtländer, N., and Voth, H.-J., 2017. Bowling for fascism: social capital and the rise of the nazi party. J. Polit. Econ. 125 (2), 478–526.
Schuller, T., 2007. Reflections on the use of social capital. Rev. Soc. Econ. 65 (1), 11–28.
Steinmayr, A., 2021. Contact versus exposure: refugee presence and voting for the far right. Rev. Econ. Stat. 103 (2), 310–327.
Tropp, L., 2012. the Oxford Handbook of Intergroup Conflict. Oxford University Press.
UNHCR, 2018. Mediterranean situation–UNHCR data portal.
Volkens, A., Burst, T., Krause, W., Lehmann, P., Matthieß, T., Merz, N., Regel, S., Weßels, B., Zehnter, L., 2020. The manifesto data collection. Manifesto project (MRG/CMP/MARPOR). Version 2020b.
Woolcock, M., 1998. Social capital and Economic development: toward a theoretical synthesis and policy framework. Theory And Society 27 (2), 151–208.