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"The mother of all political problems"? On asylum seekers and elections in Germany

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“The mother of all political problems”? On asylum seekers and elections in Germany

Lukas Tomberg*, Karen Smith Stegen[†], Colin Vance[‡]

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Abstract

Drawing on panel data from six elections between 1998 and 2017 in Germany, we estimate the causal effect of immigration – described by Germany’s interior minister as the “mother of all political problems” – on electoral support for the far right and the far left. Our identification strategy is underpinned by focusing on a particular category of immigrants, asylum seekers, who are administratively allocated across Germany according to pre-defined quotas. We find that the presence of asylum seekers has a polarizing effect, increasing vote shares for both the far right and far left. For the right, the magnitude of this effect is found to be independent of the unemployment level. For the left, the positive effect of asylum seekers tapers off with increases in unemployment, eventually becoming negative. The results suggest that the confluence of high unemployment and high immigration would tilt the electoral landscape in Germany to the right.

JEL codes: D72, J15, K37, P16.

Keywords: Asylum seekers, foreigners, voting outcomes, fractional response, instrumental variables

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1 Introduction

“We will manage.”¹ So spoke Chancellor Angela Merkel in August of 2015 in response to growing public anxiety following her decision to open Germany’s borders to over a million refugees. Two years later, her right-of-center party, the Christian Democrats (CDU), recorded their worst postwar result in national elections, with heavy losses to both the far left and the far right. On the left, the Greens and The Left performed strongly, collectively garnering 18% of the vote, a slight increase over their share in the previous national election held in 2013. On the right, the anti-immigrant party Alternative for Germany (AfD) won 12.6% of the vote, catapulting it into the Bundestag as the country’s third largest party. The reverberations of this upset have extended well beyond the election. It took nearly six months as well as multiple concessions to political rivals to form a government that still faces periodic threats of collapse owing to internal tension over the immigration issue.

Germany is not alone. In several countries of the European Union (EU), parties on far ends of the political spectrum, particularly the right, have seemingly leveraged the refugee issue to their advantage in recent elections. Austria, France, the Netherlands and Hungary have all seen vote shares for the far right exceed 10% in the most recent parliamentary elections. Given the prospect of continued refugee inflows from ongoing strife in the Middle East and Africa, along with the political turmoil visited upon the EU as it struggled to accommodate the last wave, it is natural to probe the implications for election outcomes. Drawing on county-level data from Germany, the present paper takes up this question, applying a fractional response model for panel data (Papke and Wooldridge 2008) to explore the relationship between asylum seekers and vote shares for the far left and the far right over six elections between 1998 and 2017. To test the hypothesis that the influence of asylum seekers is conditioned on economic conditions, our specification additionally includes an interaction term to allow for differential effects according to the unemployment level.

While several studies have examined how the presence of foreigners affects elections, analyses that specifically address the causal role of asylum seekers are scant. We are aware of only two studies that tackle this issue (Steinmayr 2016, Dustmann et al. 2019). In line with the contact hypothesis, Steinmayr (2016) finds that hosting refugees decreases support for the far right among a sample of Austrian communities from 2015. Conversely, drawing on Danish election data over the 1986-1998 interval, Dustmann et al. (2019) identify a positive effect of refugees on vote shares for the far right in the majority of municipalities, notwithstanding a small but negative effect seen in the most urbanized ones. The former result is in line with other studies examining the influence of foreigners on support for right-leaning parties, which generally find a positive association (Barone et al. 2016, Becker et al. 2016, Brunner and Kuhn 2018, Halla et al. 2017, Otto and Steinhardt 2014, Sekeris and Vasilakis 2016). Relatively fewer studies

¹Translated from the German phrase “Wir schaffen das,” as reported in the newspaper Zeit Online (Hildebrandt and Ulrich 2015).

have examined the effect of foreigners on support for the left, and the existing evidence is mixed. While Otto and Steinhardt (2014) identify a negative relationship between foreigners and votes for Germany's Green party, Gerdes and Wadensjö (2008) find that a pro-immigrant party on the Danish left benefits from their presence. Dustmann et al. (2019) conversely find evidence for a negative effect of refugees on center-left parties, but one that is again conditional on location; in highly urbanized areas they find the effect to be positive.

The question of causality is an issue that looms large in identifying these influences, particularly as regards the likely endogeneity of location choice: To the extent that immigrants choose where to live based on the political leanings and socioeconomic circumstances of the host region, the estimated effect of immigration on voting outcomes is subject to bias. We address this issue by distinguishing two types of immigrants, those who have been granted permission to reside in Germany, referred to here as foreigners, and those who have an application for refugee status pending, referred to as asylum seekers. Because foreigners enjoy freedom of movement and thus may self-select into particular regions, we employ an instrumental variable approach to control for their influence.

Asylum seekers, by contrast, are subject to strict rules that govern where they are settled, thereby circumventing biases arising from endogenous location choice. Upon arrival in Germany, asylum seekers are allocated by the authorities first across states and then within states across regions according to pre-defined quotas. As suggested by Glitz (2012) in his analysis of the impact of immigrants on the German labor market, this exogenously given dispersal policy can be regarded as a quasi-experiment, one that obviates the threat to identification that is otherwise posed by self-selection into particular locations. More recently, Piopiunik and Ruhose (2017) and Dehos (2017) employ the same logic to estimate the effect of immigration on crime in Germany. As in these studies, and similar to the identification strategy applied in Dustmann et al.'s analysis of Danish electoral outcomes, we leverage the exogenously given location restrictions of the German allocation system to identify the causal impact of asylum-seekers on elections.

Two variants of the model are specified to further tighten the estimated magnitude of the effects. The first includes county-level fixed effects to control for the influence of time-invariant unobservables, while the second omits fixed effects and instead includes the lagged dependent variable (LDV) to directly control for past election results. As demonstrated by Guryan (2001; 2004) and elaborated by Angrist and Pischke (2008), the virtue of jointly referencing these two models is that they serve to bracket the causal effect of interest, with the fixed effect model yielding the upper bound and the LDV model the lower bound given positive selection bias.

Among our main insights, we find that for both the far right and the far left, increases in asylum seekers have a statistically significant effect in increasing vote shares, with a counterfactual exercise revealing a particularly strong effect in the 2017 election, when the refugee issue took on greater salience

in public discourse. We also find a positive association between the unemployment rate and vote shares. On the right, this association is direct, with the evidence indicating that the unemployment rate has no mediating impact on asylum seekers. On the left, we also find a direct positive association of asylum seekers with the vote share, but one that is predicated on a low unemployment rate; at a high level of unemployment, the association becomes negative. We conclude that high unemployment coupled with high immigration would increase vote shares for the right and decrease them for the left.

2 Asylum seekers and the mediating role of unemployment

Beyond estimating the direct impact of asylum seekers on voting outcomes, our aim is to additionally account for the role of local socio-economic conditions in mediating voter sentiments about their presence. The unemployment level, in particular, has been identified as an important factor that potentially drives voters to the margins of the political spectrum, but the exact nature of the relationship between unemployment and asylum seekers remains murky. Much of the work in this area focuses on support for the far right. While Knigge (1998) and Lubbers et al. (2002) find no direct association between the unemployment rate and voting for radical-right parties, many other studies find a positive effect (Golder 2003, Arzheimer 2009, Halla et al. 2017). This is consistent with the idea – encapsulated in relative deprivation theory (Davis 1959, Crosby 1976) – that deteriorating socioeconomic conditions engender a sense of grievance that attracts people to the far right under the perception that they are disadvantaged relative to others (De Witte and Klandermans 2000, Koopmans 2005, Rydgren 2007). As these ‘others’ often include minorities, one extension of this reasoning is that the presence of minority groups, such as asylum seekers, strengthens the positive impact of the unemployment rate on support for the far right (Golder 2003).

Theorizing from political psychology offers an alternative – though not mutually exclusive – prediction rooted in the notion of *relative gratification* (Grofman and Muller 1973), whereby conditions of economic prosperity foment support for the far right out of anxiety that gains to one’s status or privilege are under threat of being stripped away. Support for this prediction is not only found in observational studies of voting behavior, which identify a negative correlation between the unemployment level and support for the far right (Knigge 1998, Arzheimer and Carter 2006), but also from experimental studies (Mols and Jetten 2016, Guimond and Dambrun 2002). As in the case of relative deprivation, the psychology underpinning relative gratification may exacerbate a negative reaction to asylum seekers to the extent they are seen as a source of competition. Mols and Jetten (2016), for example, report evidence from Australia that respondents’ endorsement of an anti-immigration message was stronger when they were confronted with the prospect of economic prosperity rather than economic downturn. Similarly, Guimond and Dambrun (2002) provide evidence from an experiment with university students in France

showing that respondents exposed to relative gratification treatments exhibited increased levels of generalized prejudice and expressed a greater willingness to support restrictive immigration policies.

Thus, while both relative deprivation and relative gratification lead to the expectation of a positive effect of asylum seekers on support for the far right, the associated expectations with respect to the effects of unemployment as well as the interaction of unemployment with asylum seekers are divergent. If voters are motivated by relative deprivation, higher unemployment is not only expected to directly increase support for the far right but also magnify the positive effect of asylum seekers on such support. Conversely, if voters are motivated by relative gratification, higher unemployment is expected to weaken support for the far right and additionally moderate the positive effect of asylum seekers.

Of course, it is also possible that voters' motivations are compartmentalized, such that their reactions to asylum seekers are independent of economic conditions. Social identity theory posits that support for the far right may instead reflect voter concerns over cultural integration and an effort to preserve group identity. Sniderman et al. (2004), for example, conclude from experimental evidence from a survey in The Netherlands that economic considerations are of only secondary importance in evoking exclusionary reactions to migrants, with the perceived threat to national identity playing a stronger role. This sentiment is reflected in the word *Überfremdung* (over alienation), a term commonly used by German far right movements to describe threats to national identity due to immigration.

While there has been less theoretical work on electoral support for the far left, there is broad agreement that such support falls into two categories: traditional working class voters who seek redistribution and 'new left' voters who prioritize values (Alonso and Fonseca 2012). Redistribution voters are generally supportive of immigrants, but as the level of immigration rises, they may become concerned about the effect on the social safety net and on wages (Keith and McGowan 2014). As Freeman (2006) and Lidén and Nyhlén (2014) argue, parties on the left are consequently torn between appealing to voters for whom economic threats may make them more critical of immigration, and voters whose wealth buffers them from such concerns and leads to a more steady support of immigration.

With regard to such value-oriented voters, Inglehart's (1977) thesis of the 'silent revolution' of value change explains how conditions of affluence and physical security, which have widely prevailed in industrialized countries since the end of World War II, reorient people's priorities from satisfying basic material needs to an emphasis on quality of life and higher order values, such as multiculturalism. However, even among well-off voters, this post-materialist reorientation might be subject to temporal lapses. In situations of extreme hardship and crises, some of these voters might re-emphasize economic concerns (Inglehart and Welzel 2005). Thus, it follows that voter support for far left parties sympathetic to immigrants would be conditional on favorable economic conditions. This tension is seen in the results of a conjoint experiment conducted across 15 European countries by Bansak et al. (2016). They find that while humanitarian reasons play a more important role in increasing the probability of

accepting asylum seekers among respondents on the left than on the right, those on the left respond more negatively to asylum seekers when economic opportunities are cited as the reason for migrating, thereby providing some support for the hypothesis that sympathy is conditional.

The present paper contributes to the above literature by drawing on quasi-experimental evidence to assess the causal impact of immigrants on voting patterns. Unlike many of the studies reviewed above, which use individual-level data, we explore here the relationship between variables aggregated at the county-level. Although the use of such data prevents us from drawing inferences on individual behavior, its main virtue – contrary to the hypothetical nature of experiments or public opinion surveys – is to reveal evidence on observed election outcomes. Linking this evidence to particular theoretical predictions is complicated by the fact that the theories themselves are not mutually exclusive, so that the aggregated results may reflect multiple overlapping impact channels that cannot be precisely disentangled. By isolating differential effects of asylum seekers according to the unemployment level, however, our analysis affords insight into which channels dominate at the aggregate level, thereby complementing empirical work based on individual-level experimental evidence.

3 Data and descriptive statistics

The data for this paper was assembled from publically available sources that are accessible online. Variables are measured at the county level using boundary definitions from the year 2010.² Temporal coverage begins in 1998 and includes each year in which a national election was held – 1998, 2002, 2005, 2013 and 2017. With 405 counties recorded over six years, the data forms a balanced panel comprising 2,430 observations.

Data on election results was obtained from the Federal Returning Officer (2018). The data contains county-level information on the number of eligible voters, the number who voted, and the number of votes going to each party for each election year. This information is recorded for two types of ballots, the first vote and the second vote (*Erststimme* and *Zweitstimme*). The first vote allows voters to directly elect a local representative, either party-affiliated or independent, from each of the voting districts to the parliament. This vote is intended to ensure that all regions have representation. The second vote, which is the focus of the present analysis, is cast for a party, whose representatives are sent to parliament. This vote is more consequential, as it determines the final distribution of seats in parliament by party. In turn, the party with the most seats traditionally selects the chancellor, whose final approval is subject to an up or down vote by the parliament.³

²To account for territorial reforms in the state of Mecklenburg-Vorpommern as of 2011, we used population-based weights provided by Germany's Federal Office for Building and Regional Planning (BBSR 2018), which serve to convert values to the county boundaries of 2010.

³There are occasionally attempts to buck tradition. In the 2005 election, Gerhard Schroeder tried unsuccessfully to be designated chancellor even though his party, the Social Democrats, won a smaller share of seats in the Bundestag than Angela Merkel's party, the Christian Democrats.

To calculate the dependent variables, we aggregate votes for each party in two categories, far left and far right, and then calculate the respective shares by county. The designation of parties along the left-right spectrum is complicated by the fact that no universally accepted criteria or methods exist to inform the process. In the case of Germany, the difficulty is particularly pronounced in designating the Greens, a party that drew its initial support in the early 1980s from activists on the left championing environmental politics, civil rights, and women’s participation (Müller-Rommel 1985), but one that has since moved to the center on issues such as technological modernization and the bolstering of police forces. Thus, Falck et al.’s (2014) analysis of voting behavior excludes the Greens from the measure of the vote share of left-fringe parties. Conversely, Otto and Steinhardt (2014) focus exclusively on the Greens as a counterpoint to their analysis of far-right voting in Hamburg, arguing that the Greens is the only party in Germany to have consistently promoted liberal immigration and integration policies unconditionally. For our main estimates, we include the Greens in the calculation of the far left share, grouping them with other parties characterized by internationalism, anti-capitalist rhetoric, socialism, and protectionism, including The Left and the German Communist Party. To gauge robustness, we also estimate a model that defines the dependent variable as the vote share exclusively for the Greens.

The designation of far right parties is less ambiguous since, as Arzheimer and Carter (2006) note, a general consensus has emerged that such parties share a number of ideological features comprising some combination of racism, xenophobia, nationalism, and a desire for a strong state. The National Democratic Party of Germany and the Republicans are two such parties that we include in our measure of the far right vote share. We also include the Alternative for Germany (AfD) in the measure, but only for the 2017 election. For the 2013 election, the year the AfD was formed, we exclude the party from the calculation of the far right share. This exclusion owes to the content of the party’s original platform, which targeted the dissolution of the European currency union (Arzheimer 2015).⁴ The 2013 platform is devoid of anti-immigrant sentiment and indeed briefly references the desirability of bringing in qualified labor from abroad (AfD 2013). It was only after just missing the 5% of the vote needed for representation in the Bundestag in 2013 that the AfD pivoted to an overtly anti-immigrant tone, which coincided with the refugee crisis of 2015. This transformation of the AfD’s message from technocratic to nativist resulted in several of its founding members disassociating themselves, even as the party’s fortunes improved markedly with the 2017 election.

Data on asylum seekers was taken from the Federal Statistical Office (Destatis 2018b), which posts a yearly online spreadsheet that separately tabulates the number of all categories of immigrants – including asylum seekers – by the state and county in Germany in which they reside. Upon arrival, asylum seekers are allocated across states according to annually adjusted quotas based on the state’s tax revenue and population. The criteria for allocating asylum seekers across counties within states varies, but

⁴Excluding the AfD from the far right also follows the party’s positioning in the RILE index (Budge et al. 2001, Klingemann et al. 2006), a widely used source for situating parties along the political spectrum, which positioned the AfD in the center in 2013.

most states employ either fixed allocation keys or flexible quotas based on the population size and, in some cases, the land area. To construct our main explanatory variable, designated as *asylum seekers*, we extracted entries having the status of an asylum seeker from the spreadsheet and divided this by the county's population in the corresponding year. To allow for the possibility that responses to asylum seekers may be conditioned on the number of foreigners already living in a county, we created a second explanatory, designated as *foreigners*, by summing all remaining immigrants in a county who do not have the status of an asylum seeker and dividing by the population.

Figure 1: Spatial distribution of foreigners and asylum seekers

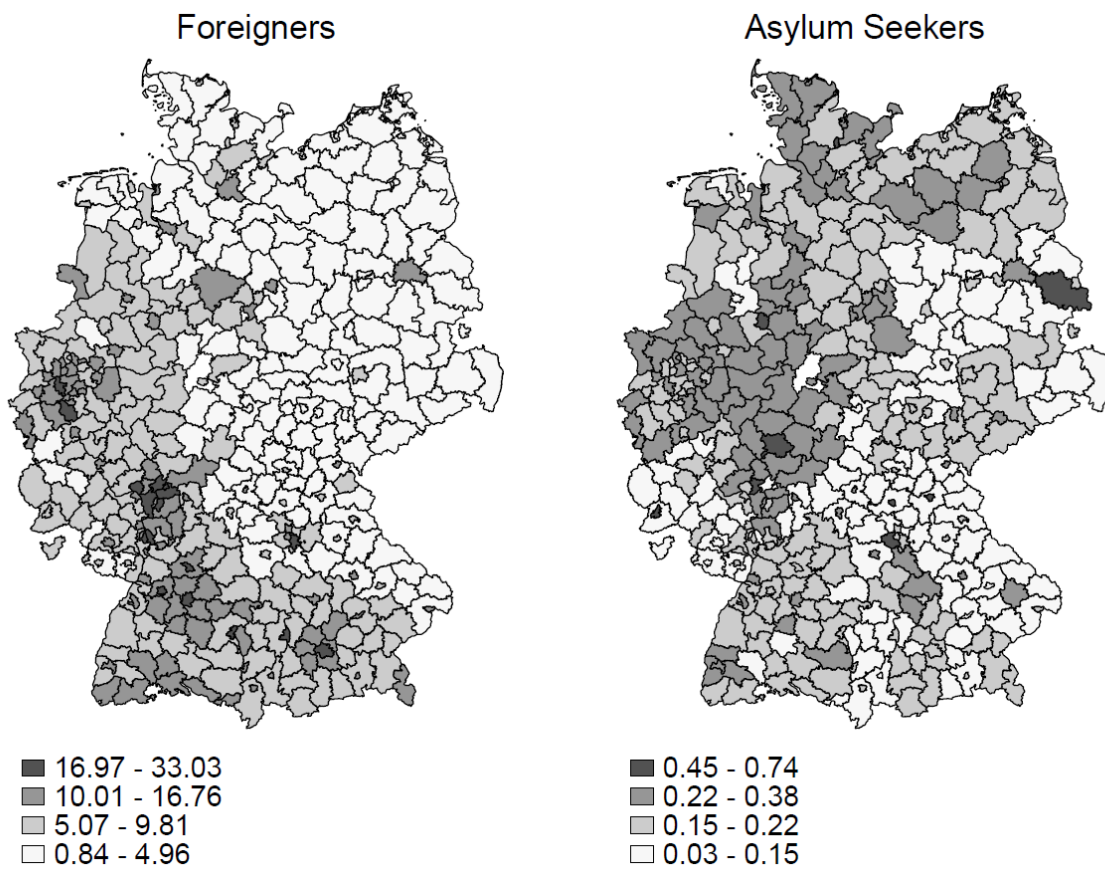
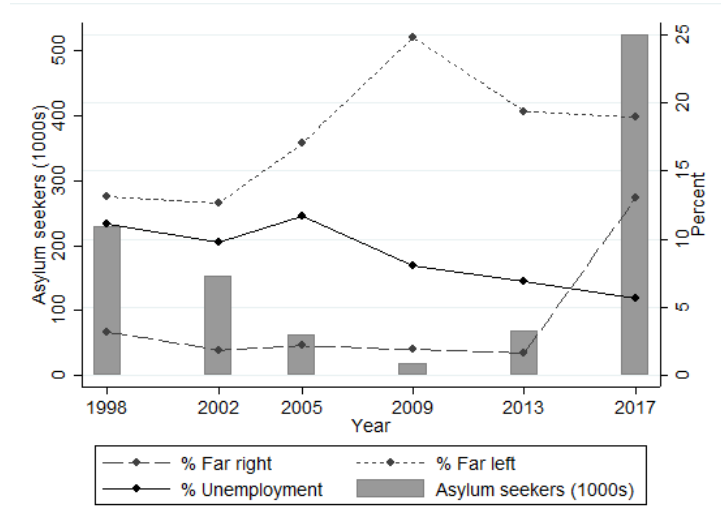


Figure 1 presents time averages of the per capita spatial distribution of foreigners and asylum seekers across German counties. For foreigners, a spatial sorting process is evident, with most concentrated in the western part of the country and in the capital Berlin. The overall share of foreigners in the West is 8.6%, over threefold the share of 2.3% in the East. Asylum seekers, by contrast, are seen to be more evenly dispersed throughout the country. In the West, their overall share is 0.21%, somewhat higher than the share of 0.18% in the East, which likely reflects the higher tax revenue collected by western states. There is, however, also evidence of individual counties where asylum seekers have a particularly high concentration. While self-selection cannot account for this pattern given that asylum seekers must follow administrative directives on where they settle, the pattern does force us to consider the

Figure 2: Asylum-seekers, unemployment and vote shares in each election year



possibility that the directives themselves are subject to political manipulation, an issue we address below.

Figure 2 presents the temporal evolution of the key variables under consideration in the analysis: the vote shares for the far right and the far left, the total number of asylum-seekers, and the unemployment rate, the latter of which was obtained from Germany's Federal Employment Agency (2018).⁵ The far left is seen to have had a consistently stronger following in Germany than the right, even in the most recent election. Nevertheless, a pronounced ascendancy of the far right since 2013 is evident. This ascendancy coincides with the sharp increase in asylum seekers, but also with a fall in the unemployment rate, which dropped by over a percentage point between 2013 and 2017. These aggregate trends are thus consistent with the theory of relative gratification, which would interpret the upsurge in support for the far right in 2017 as reflecting an attempt by voters to secure gains in economic prosperity threatened by the influx of asylum seekers.

The descriptive statistics on the full set of dependent and independent variables that enter the model are presented in Table 1. In addition to our main explanatory variables, we include controls for the share of foreigners, population density (persons per square kilometer), education, and age structure, all of which are derived from time-varying, county-level data available from the Federal Statistical Office (Destatis 2018a;b;c). Similar to Barone et al. (2016), education is measured as the share of high school students who graduate with a university-qualifying degree (*Abitur*), while the age structure is captured by the share of residents over 65 years old.

⁵Although the plots indicate contemporaneous measurement of these three variables, the figures for unemployment and asylum applications and are recorded in December of the preceding year.

Table 1: Summary statistics

| | | Mean | Std. Dev. |
|----------------|--|-------|-----------|
| Far right | Vote share for the far right | 0.043 | 0.051 |
| Far left | Vote share for the far left | 0.168 | 0.083 |
| Asylum seekers | Asylum seekers per resident | 0.002 | 0.002 |
| Foreigners | Foreigners per resident | 0.071 | 0.051 |
| Unemployment | Unemployment rate | 0.093 | 0.049 |
| Pop density | Population density | 0.518 | 0.674 |
| Share Abitur | Share of students graduating with Abitur | 0.259 | 0.105 |
| Share old | Share of resident older than 65 | 0.194 | 0.03 |
| No. obs. | | 2430 | |

4 Methodology

Virtually all studies that investigate election outcomes using regional data on vote shares employ a linear estimator.⁶ One drawback of this approach is that it neglects the bounded nature of the dependent variable, potentially resulting in the same problems that afflict the linear probability model (Papke and Wooldridge 2008). These include non-normality of the errors, incorrect statistical inferences, and non-sensible predictions. More disconcerting, Horrace and Oaxaca (2006) demonstrate that the unbiasedness and consistency of the linear estimator in the case of a bounded dependent variable demands a strong condition that is unlikely to hold, namely that the predictions given by the population parameters and the data matrix fall strictly in the zero to one range.

To deal with these issues, we estimate a fractional probit model (Papke and Wooldridge 1996; 2008) using two specifications. The first of these controls for unit-level fixed effects, with the model expressed as:

$$E(\text{vote}_{it} | \mathbf{X}_{it}, \bar{\mathbf{X}}_i) = \Phi(\boldsymbol{\beta}^T \mathbf{X}_{it} + \boldsymbol{\zeta}^T \bar{\mathbf{X}}_i + \sum_{j=1} \sum_{n=1} \gamma_{jn} s_j y_n), \quad (1)$$

where the dependent variable, *vote*, measures the vote share for either the far right or far left in county *i* and year *t*, and Φ represents the cumulative density function of the standard normal distribution. The vector \mathbf{X}_{it} contains explanatory variables, the key ones of interest being *asylum seekers*, *unemployment*, and their interaction. To avoid the incidental parameter problem that often plagues non-linear panel models, fixed effects are controlled for using the Chamberlain-Mundlak device, by which unit-level time-averages of the explanatory variables, $\bar{\mathbf{X}}_i$, are included (Papke and Wooldridge 2008). The specification is completed with controls for time-varying effects across each of Germany's sixteen states via the interaction of state and year dummies, $\sum_{j=1} \sum_{n=1} \gamma_{jn} s_j y_n$. $\boldsymbol{\beta}$, $\boldsymbol{\zeta}$, and γ_{jn} are a set of parameters and parameter vectors to be estimated. Model 1 can be estimated by fitting the data to a Bernoulli distribution and using the Maximum-Likelihood estimator, which is analogous to estimating a Probit

⁶An exception is Otto and Steinhardt (2014), who report having found negligible difference between the estimates of a linear model and a non-linear fractional response model.

model (Papke and Wooldridge 2008).

Although the specification in Model 1 conditions on state-level time-varying effects, county-level fixed effects, and a suite of time-varying county-level variables, the existence of unobserved dynamic processes at the county level that affect the outcome variable and are correlated with the distribution of asylum-seekers cannot be completely ruled out. As asylum-seekers have no say in where they are settled, the corresponding threat to identification is not from sorting, but rather from possible influences on the administrative process that determines settlement. For example, political parties may establish regional strongholds based on past electoral performance, and this power may be used to influence how asylum seekers are allocated, even in the face of official quotas that steer the process. One solution to this issue, discussed at length by Keele and Kelly (2006), is to use a specification that controls for past election outcomes by including a lagged dependent variable (LDV):

$$E(\text{vote}_{it} | \mathbf{X}_{it}, \text{vote}_{i(t-1)}) = \Phi(\boldsymbol{\beta}^T \mathbf{X}_{it} + \zeta \text{vote}_{i(t-1)} + \sum_{j=1} \sum_{n=1} \gamma_{jn} s_j y_n), \quad (2)$$

where $y_{i(t-1)}$ measures the vote share from the preceding election. Distinguished from Model 1, Model 2 directly captures dynamic effects in election outcomes. However, it does not condition on county fixed effects, as so doing would yield estimates that are not consistent owing to correlation between the LDV and the error term (Angrist and Pischke 2008). Nevertheless, Models 1 and 2 have a useful bracketing property. Given a negative correlation between asylum seekers and the error term, as would be expected if political forces in regions of strong far right support conspire against accepting refugees, the fixed effect model of the far right share gives an upper bound estimate while the LDV model gives a lower bound (see the discussion of bounding in Guryan (2001)). Conversely, given a positive correlation between asylum seekers and the error, which could be expected in regions where pro-immigrant parties exercise political clout, the effect from the LDV model of the far left share will exceed that of the fixed effects model. Consequently, as Angrist and Pischke (2008) suggest, referencing the two models jointly serves to bracket the causal effect of interest.

A final empirical issue concerns the possible endogeneity of one of the control variables, *foreigners*, who, contrary to asylum seekers, face no restrictions on where they may settle. Assuming that the prevailing level of xenophobic hostility is plausibly one factor that deters foreigners from locating in a particular county, it follows that the estimated effect of *foreigners* on the vote share could be biased. Failure to address this endogeneity may moreover bias the remaining estimates. Among studies that use regional election data, a common solution to this difficulty is to draw on an instrumental variable (IV), one typically based on historical settlement patterns. We adapt this course by using a so-called shift share instrument that was originally employed by Altonji and Card (1991) and Card (2001). As elaborated in the Appendix, the instrument is based on the settlement pattern of foreigners from a

specific home-country in a baseline year t_0 , which is 1999 in our application.

Following Papke and Wooldridge (2008), we implement the control function approach to incorporate IVs into the fractional panel data model. This entails first estimating a linear first stage model of the endogenous variable, $foreigners_{it}$, on the shift share instrument, the other exogenous variables, X_{it} and the corresponding unit-level averages, \bar{X}_i . The residuals \hat{v}_{it2} obtained from this model are then included as an additional regressor to the second stage models in Equations 1 and 2.⁷ In the LDV case, the time averages of the regressors are replaced with the second stage lagged dependent variable. A test for the endogeneity of $foreigners_{it}$ is then easily obtained as an asymptotic t-statistic on \hat{v}_{it2} (Papke and Wooldridge 2008). For the remaining explanatory variables, the two-step nature of the control function approach requires that the standard errors in the second stage be adjusted for the first stage estimation, which is accomplished in the present application by bootstrapping.

5 Results

5.1 Determinants of voting outcomes: Baseline estimates

This section presents the estimates from the fixed effect (FE)- and lagged dependent variable (LDV) models of the share of votes for the far right and far left. The estimation sample covers the years beginning with 2002 to accommodate the loss of the year 1998 in the LDV model.⁸ All models employ the control function approach to allow the effect of *foreigners* to be endogenous. The estimates from the first stage models are presented in the Appendix. We note here that these estimates support the strength of the shift share variable as an instrument for the foreigner share. The associated F-statistics are in all models higher than 100, thereby far exceeding Staiger and Stock's (1997) well-known rule of thumb according to which the F-statistic should exceed 10 to reject weak identification.

Tables 2 and 3 present the coefficient estimates from the second stage models. As the magnitude of the coefficients are not immediately interpretable, the discussion below refers to the average partial effects (APEs), also presented either in the table or graphically for the models with the interactions in the final two columns. The APEs are calculated by differentiating Models 1 and 2 with respect to the variable *asylum seekers*, which yields an estimate of the marginal effect for each observation in the data (see the Appendix for details). The columns labeled APE in the table present the mean of these observation-specific marginal effects.

Turning first to the results for the far right in Table 2, all four models indicate a positive and statistically significant association between increases in the density of asylum seekers and the vote share for

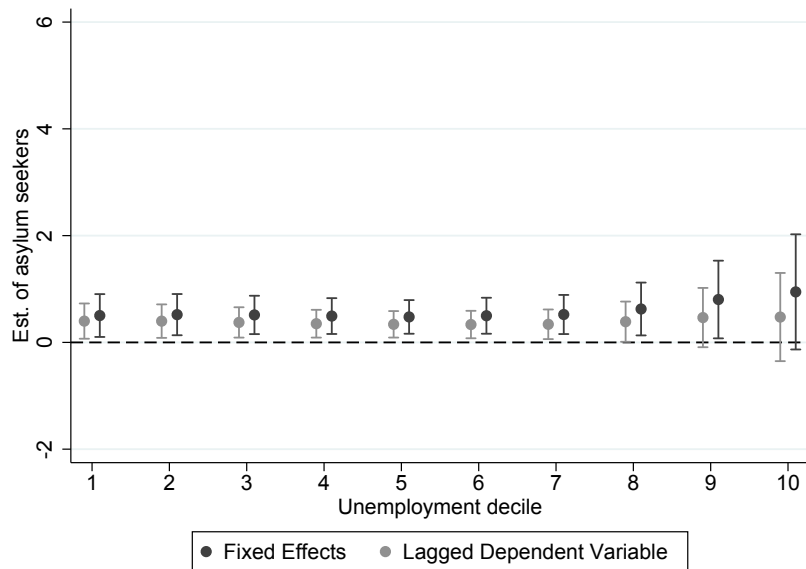
⁷The inclusion of the time average of the endogenous variable follows from (Bluhm et al. 2018). This approach facilitates the differentiation between endogeneity due to unobserved heterogeneity in contrast to endogeneity because of time-varying omitted variables (Lin and Wooldridge 2017).

⁸Including the year 1998 in the FE model does not have a substantial bearing on the results.

Table 2: Fractional probit coefficients for the far right

| | FE | | LDV | | FE | |
|--|---------------------|---------------------|---------------------|---------------------|---------------------|---------------------|
| | Coef. | APE | Coef. | APE | Coef. | LDV Coef. |
| Asylum seekers per cap | 7.005** (2.336) | 0.576** (0.192) | 4.899** (1.817) | 0.402** (0.148) | 6.059 (3.597) | 5.444 (3.078) |
| Foreigners per cap | 2.039** (0.615) | 0.168** (0.050) | 0.054 (0.129) | 0.004 (0.011) | 2.073** (0.635) | 0.056 (0.128) |
| Unemployment rate | 0.546* (0.272) | 0.045* (0.023) | 0.779** (0.146) | 0.064** (0.012) | 0.539* (0.272) | 0.799** (0.155) |
| Population density | -0.443** (0.114) | -0.036** (0.009) | -0.006 (0.009) | -0.000 (0.001) | -0.443** (0.114) | -0.006 (0.009) |
| Share Abitur | -0.153 (0.095) | -0.013 (0.008) | -0.290** (0.047) | -0.024** (0.004) | -0.156 (0.096) | -0.290** (0.047) |
| Share old | 0.594 (0.480) | 0.049 (0.039) | 0.390** (0.139) | 0.032** (0.012) | 0.621 (0.491) | 0.390** (0.139) |
| Lagged dependent variable | — | — | 11.376** (0.497) | 0.932** (0.040) | — | 11.380** (0.498) |
| Unemployment \times asylum seekers | — | — | — | — | 13.671 (44.864) | -8.822 (37.580) |
| Constant | -0.853** (0.074) | — | -1.174** (0.040) | — | -0.861** (0.083) | -1.175** (0.040) |
| Coef. of $\hat{\nu}_2$ (2nd stage S.E.) | -0.955* (0.462) | — | 0.166 (0.165) | — | -0.985* (0.486) | 0.161 (0.167) |
| F-Stat of IV in 1st stage | 118.20 | — | 154.77 | — | 118.20 | 154.77 |
| County fixed effects | ✓ | — | — | — | ✓ | — |
| State \times year dummies | ✓ | — | ✓ | — | ✓ | ✓ |
| No. obs. | 2025 | | | | | |

Note: Clustered standard errors are reported in parentheses. ** and * indicate statistical significance at the 1% and 5% level, respectively.

Figure 3: Average partial effects for the far right

the far right. With reference to the FE results in columns 1 & 2, the associated APE suggests that a one percent point increase in *asylum seekers* is associated with a 0.58 percent point increase in the vote share. The estimated APE from the LDV model is, as expected, somewhat smaller at 0.40, which is consistent with the possibility that settlement across counties is subject to a negative selection process owing to local political resistance. Nevertheless, even this more conservative estimate is considerably higher in

magnitude than estimates from models of the far right presented elsewhere in the literature. In their study of Hamburg, Otto and Steinhardt (2014, p. 71) report that a “a one percentage point increase in the foreigner share increases right-wing parties’ vote share by 0.23 percentage points,” while Halla et al. (2017) report a corresponding estimate for Austria of 0.16 percentage points. Both studies, however, use a more expansive definition that includes all immigrants, with the upshot being that with the more narrow definition of asylum seekers used here, a percentage point increase is a larger quantity relative to the mean.

In fact, the estimate of *foreigners* from the FE model presented in Table 2 is in line with these earlier studies. The estimated APE from column 2 is 0.17, falling squarely within the range of estimates of Otto and Steinhardt (2014) and Halla et al. (2017). The fully robust t-statistic on $\hat{\nu}_{it2}$ in this model is 2.22, providing some evidence against the null hypothesis that the foreigner share is conditionally strictly exogenous. By contrast, in the LDV model in column 2, we fail to reject the hypothesis of exogeneity, and also fail to reject the hypothesis that the effect of foreigner share is zero.

As with asylum seekers, the estimated effect of the unemployment rate is found to be positive and statistically significant in all four models. The APE corresponding to the FE model in column 2 suggests that a percentage point increase in the unemployment rate is associated with a 0.05 percent point increase in the vote for the far right. The estimated APE from the LDV model in column 4 is somewhat higher, at 0.06. Thus, notwithstanding the coupled trends of lower unemployment and increased support for the far right documented at the national level in Figure 2, the estimates from the econometric model suggest an opposite association at the county level, consistent with the notion that relative deprivation attracts voters to the far right in the face of high unemployment.

To address the question of whether unemployment and the presence of asylum seekers mediate each other’s effect on the vote share for the far right, the models in columns 5 and 6 include the interaction of the two variables, which, for both the FE and LDV model, is found to be statistically insignificant. Further insight is shown by a plot of the APEs of *asylum seekers* and the associated 95% confidence intervals over different levels of unemployment, defined by dividing the counties into ten evenly large groups based on the time-averaged unemployment rates and then calculating the APEs separately for each of these ten groups. This division avoids creating artificial combinations of covariate values that fall outside the sample, and has the added benefit of accounting for the skewness of the unemployment distribution. The plot in Figure 3 confirms the results from Models 1 and 2: Across all deciles, the estimated APE of *asylum seekers* varies moderately – between roughly 0.41 and 0.76 – for the FE model and 0.40 and 0.48 for the LDV model. This corroborates the finding of Golder (2003), who, based on a linear FE model spanning 19 countries, concludes that higher levels of immigration increase the electoral support of right-leaning populist parties irrespective of the unemployment rate.

Table 3 presents estimates from the model of the vote share for the far left. We again see positive

Table 3: Fractional probit coefficients for the far left

| | FE | | LDV | | FE | LDV |
|--|---------------------|---------------------|---------------------|--------------------|------------------------|-----------------------|
| | Coef. | APE | Coef. | APE | Coef. | Coef. |
| Asylum seekers per cap | 2.629** (0.897) | 0.653** (0.224) | 3.406** (1.089) | 0.844** (0.271) | 9.671** (1.603) | 7.921** (1.612) |
| Foreigners per cap | 0.663 (0.486) | 0.165 (0.132) | 0.343** (0.077) | 0.085** (0.019) | 0.832 (0.525) | 0.356** (0.078) |
| Unemployment rate | -0.157 (0.155) | -0.039 (0.039) | 0.116 (0.083) | 0.029 (0.021) | -0.055 (0.154) | 0.223* (0.091) |
| Population density | 0.051 (0.065) | 0.013 (0.017) | 0.003 (0.004) | 0.001 (0.001) | 0.054 (0.065) | 0.003 (0.004) |
| Share Abitur | 0.012 (0.042) | 0.003 (0.011) | 0.130** (0.025) | 0.032** (0.006) | 0.012 (0.040) | 0.128** (0.025) |
| Share old | -1.232** (0.304) | -0.306** (0.074) | 0.033 (0.103) | 0.008 (0.025) | -1.300** (0.310) | 0.029 (0.103) |
| Lagged dependent variable | — | — | 3.441** (0.092) | 0.853** (0.021) | — | 3.441** (0.092) |
| Unemployment \times asylum seekers | — | — | — | — | -113.418** (21.726) | -67.022** (21.867) |
| Constant | -0.949** (0.088) | — | -1.942** (0.042) | — | -0.958** (0.097) | -1.947** (0.042) |
| Coef. of $\hat{\nu}_2$ (2nd stage S.E.) | -0.301 (0.431) | — | -0.153 (0.124) | — | -0.343 (0.465) | -0.182 (0.125) |
| F-Stat of IV in 1st stage | 118.20 | | 153.99 | | 118.20 | 153.99 |
| County fixed effects | ✓ | | | | ✓ | |
| State \times year dummies | ✓ | | ✓ | | ✓ | ✓ |
| No. obs. | | | | 2025 | | |

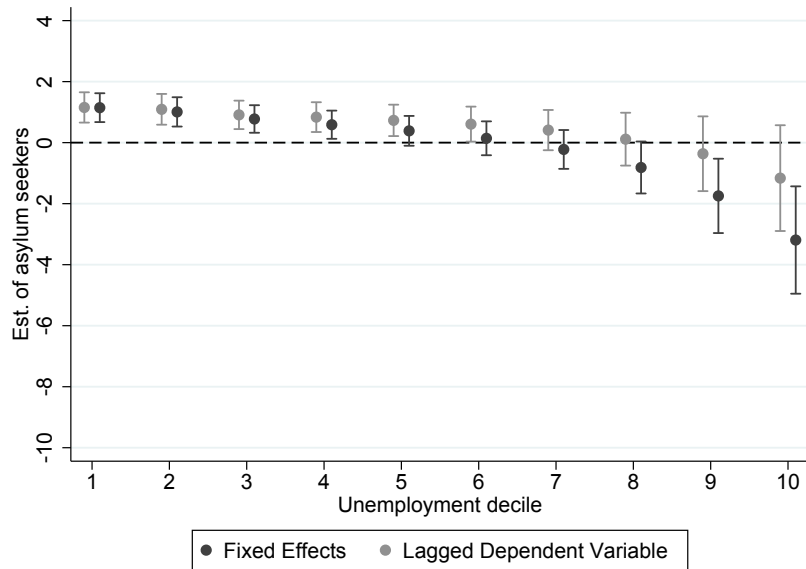
Note: Clustered standard errors are reported in parentheses. ** and * indicate statistical significance at the 1% and 5% level, respectively.

coefficient estimates on asylum seekers across all four models, with the APEs being of somewhat higher magnitude than the corresponding estimates from the models of the far right share. The estimated APE corresponding to the coefficient in the FE model in column 2 suggests that a one percent point increase in *asylum seekers* is associated with a 0.65 percent point increase in the vote share for the far left. As would be expected given positive selection of settlement, this estimated effect is now bounded from above by the LDV estimate of 0.84.

Contrasting with the negative association found by Otto and Steinhardt (2014), we also find a positive coefficient on *foreigners*, but only in the LDV model, which indicates that a one percent point increase in foreigners is associated with an 0.085 percent point increase in electoral support for the far left. Expressed as an elasticity, the estimate is equal to 0.037, which is of similar magnitude to the estimate of 0.044 obtained by Gerdes and Wadensjö (2008) in their model of the pro-immigrant Socialist People's Party in Denmark. In none of the models of Table 3 do we reject the null hypothesis that *foreigners* is conditionally strictly exogenous.

Neither the FE- nor the LDV model in columns 1 through 4 indicate a statistically significant direct effect of the unemployment rate on the vote share for the far left. However, the models including the interaction terms in columns 5 and 6 provide evidence for an indirect effect of the unemployment rate that mediates the positive effect of the share of asylum seekers. Referring to the plot of APEs in Figure 4, a pattern is seen wherein the effect of asylum seekers is positive and statistically significant

Figure 4: Average partial effects for the far left



at low levels of unemployment, below the 50th decile. As unemployment increases, this positive effect weakens, eventually becoming negative in the FE model and not statistically different from zero in the LDV model. Consistent with Inglehart’s (1977) notion of post-materialism, this pattern suggests that concerns about asylum seekers and unemployment are linked on the left, with support for pro-immigrant parties predicated on a low to moderate level of unemployment.

5.2 Robustness checks

Halla et al. (2017) point to the possibility that the effect of immigrants on election outcomes could be subject to non-linearities. For example, it is conceivable that at low densities, the presence of immigrants has no bearing on voting, with an effect only detected once the density passes a certain threshold. It is also plausible that the effect increases but at a decreasing rate, eventually reaching a tipping point at which it levels off and potentially turns negative. Given our separate focus on asylum seekers and foreigners, yet a third possibility is that the two variables mediate each other’s effects: Regions with a high density of foreigners may have a different response to an influx of asylum seekers than regions with a low density of foreigners.

To explore these possibilities, we estimate additional models that include a quadratic specification of the variable *asylum seekers* as well as models that include the interaction of *asylum seekers* with *foreigners*. Table 4 presents the FE and LDV variants of the model using the quadratic specifications. In the models of the far right, the coefficient estimates on *asylum seekers* and its square are statistically significant, indicating a positive effect that increases at a decreasing rate. Evaluated at the mean level of unemployment, the average partial effect reaches a tipping point of zero when the density of asy-

Table 4: Fractional probit coefficients with quadratic specification

| | Far right | | Far left | |
|--|-------------------------|-------------------------|------------------------|------------------------|
| | FE | LDV | FE | LDV |
| Asylum seekers per cap | 17.575** (5.084) | 15.501** (4.383) | 10.840** (2.389) | 11.678** (2.312) |
| Foreigners per cap | 2.087** (0.633) | 0.029 (0.129) | 0.803 (0.526) | 0.356** (0.077) |
| Unemployment rate | 0.440 (0.273) | 0.793** (0.156) | -0.059 (0.154) | 0.227* (0.091) |
| Population density | -0.451** (0.116) | -0.004 (0.010) | 0.051 (0.065) | 0.003 (0.004) |
| Share Abitur | -0.140 (0.095) | -0.284** (0.048) | 0.013 (0.041) | 0.129** (0.025) |
| Share old | 0.586 (0.494) | 0.371*** (0.140) | -1.321** (0.309) | -0.026 (0.103) |
| Lagged dependent variable | — (0.092) | 11.488** (0.507) | — (0.092) | 3.439** (0.092) |
| Unemployment \times asylum seekers | 10.241 (43.263) | -6.103 (38.705) | -112.776** (21.717) | -69.712** (21.310) |
| Asylum seekers per cap ² | -776.919** (246.426) | -652.417** (227.469) | -63.246 (114.164) | -239.038* (116.912) |
| Constant | -2.129** (0.085) | -2.376** (0.045) | -1.483** (0.092) | -1.657** (0.028) |
| Coef. of $\hat{\nu}_2$ (2nd stage S.E.) | -1.013* (0.489) | 0.152 (0.167) | -0.326 (0.460) | -0.180 (0.124) |
| F-Stat of IV in 1st stage | 118.20 | 154.77 | 118.20 | 153.99 |
| County fixed effects | ✓ | | ✓ | |
| State \times year dummies | ✓ | ✓ | ✓ | ✓ |
| No. obs. | 2025 | | | |

Note: Clustered standard errors are reported in parentheses. ** and * indicate statistical significance at the 1% and 5% level, respectively.

lum seekers equals 0.012 in the FE model and 0.011 in the LDV model. These values fall in the top one percent of the distribution of *asylum seekers*, suggesting that the rate at which the effect weakens is gradual but nevertheless becomes marginally negative within the range of the data. Figure 5 presents the associated APEs from the models of the far right over the different deciles of unemployment. The estimates are statistically significant over all deciles and are nearly double the magnitude of those from the model in which *asylum seekers* enters linearly (Figure 3). On average, the estimates suggest that a one percent point increase in the density of asylum seekers increases the vote for the far right by 1.2 percent points, with the point estimate reaching 1.7 in the ninth decile of the FE model. This range overlaps substantially with that reported by Dustmann et al. (2019), who find effect sizes ranging between 1.34 and 2.32 in all but the most urban Danish municipalities.

In the model of the far left, only the LDV specification suggests evidence for a quadratic effect of asylum seekers, but one that does not substantially change the earlier findings. A plot of the associated APEs over unemployment deciles reveals a pattern of estimates that is nearly identical to that in Figure 3 (not presented). A similar conclusion is reached from a model in which *asylum seekers* is interacted with *foreigners*, presented in Table 5. In the models of the far left only, the interaction terms are positive and statistically significant, indicating that the positive effect of *asylum seekers* increases with increases in the density of foreigners. Ultimately, however, the inclusion of the interaction term has no bearing

Figure 5: Average partial effects for the far right based on the model with a quadratic specification

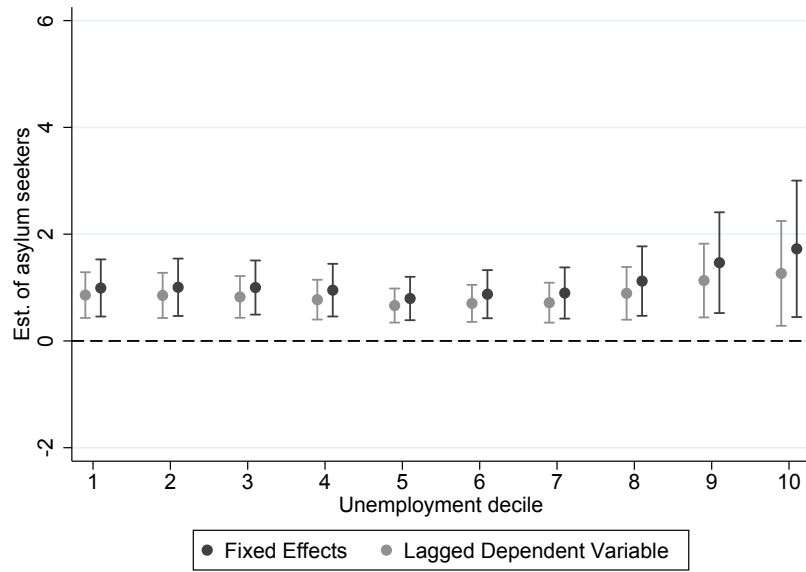


Table 5: Fractional probit coefficients with the asylum seekers \times foreigners interaction

| | Far right | | Far left | |
|--|---------------------|---------------------|------------------------|-----------------------|
| | FE | LDV | FE | LDV |
| Asylum seekers per cap | 3.370 (4.071) | 5.148 (3.321) | 5.587** (1.951) | 5.991** (1.734) |
| Foreigners per cap | 1.824** (0.666) | 0.045 (0.127) | 0.480 (0.560) | 0.311** (0.080) |
| Unemployment rate | 0.545* (0.274) | 0.802** (0.156) | -0.022 (0.154) | 0.233** (0.091) |
| Population density | -0.482** (0.119) | -0.006 (0.009) | 0.006 (0.068) | 0.003 (0.004) |
| Share Abitur | -0.163 (0.096) | -0.291** (0.047) | 0.004 (0.041) | 0.128** (0.025) |
| Share old | 0.679 (0.497) | 0.394** (0.141) | -1.228** (0.318) | 0.041 (0.104) |
| Lagged dependent variable | - (-) | 11.379** (0.498) | - (-) | 3.440** (0.092) |
| Unemployment \times asylum seekers | 1.442 (44.987) | -10.626 (38.106) | -130.095** (23.402) | -75.788** (22.901) |
| Foreigners \times asylum seekers | 35.975 (26.356) | 4.084 (17.454) | 52.035** (14.233) | 25.127* (11.490) |
| Constant | -2.147** (0.090) | -2.360** (0.045) | -1.457** (0.094) | -1.650** (0.027) |
| Coef. of \hat{v}_2 (2nd stage S.E.) | -0.964* (0.488) | 0.158 (0.167) | -0.326 (0.461) | -0.192 (0.120) |
| F-Stat of IV in 1st stage | 118.20 | 154.77 | 118.20 | 153.99 |
| County fixed effects | ✓ | | ✓ | |
| State \times year dummies | ✓ | ✓ | ✓ | ✓ |
| No. obs. | 2025 | | | |

Note: Clustered standard errors are reported in parentheses. ** and * indicate statistical significance at the 1% and 5% level, respectively.

on the conclusions reached from the more parsimonious specification of the model presented in Table 3: A plot of the APEs over unemployment (not presented) reveals essentially the same pattern as seen in Figure 4.

Additional robustness checks, documented in the Appendix, were undertaken to explore whether

the results are sensitive to the definition of the dependent variable, the model specification, and the estimation technique. Following Otto and Steinhardt (2014), we redefined the dependent variable in the model of the far left to focus exclusively on the share of votes going to the Green party. This redefinition decreases the magnitude of the estimates but retains the pattern seen in Figure 4. Recognizing the possibility of bad controls – variables that are themselves consequences of asylum seekers – we also estimated models that excluded the control variables, finding the estimates to be largely robust to their omission. Last, we employed a linear estimator, the standard method among studies that use regional data to investigate election outcomes. While this estimator produces results that are nearly identical to those of the non-linear model of the far left, it uncovers evidence for a positive and statistically significant interaction effect between asylum seekers and the unemployment rate in the model of the far right. As elaborated in the Appendix, we suspect that this divergence relative to the non-linear model, which detects no such interaction, owes to clustering of values of the far right share near zero, which the linear model is ill-equipped to handle.

5.3 To what extent did asylum seekers bear on the 2017 election?

The unprecedented influx of refugees to Germany in 2015 created a charged political atmosphere leading up to the 2017 election. Over this period, widespread media coverage of asylum seekers being welcomed with food and clothing as they arrived at the nation’s train stations was juxtaposed with images of mass anti-immigrant protests, particularly in the eastern cities. It has been a commonly held viewpoint in Germany that asylum seekers were a key determining factor in the outcome of the election (Bukow 2017, Korte 2017, Haas 2017). We explore this question by moving beyond a discussion of the model estimates to consider their implications for the outcome of the most recent election in 2017. Specifically, we want to consider the question of how parties on the far right and far left would have fared had no asylum seekers been present. To this end, we undertake a simple counterfactual exercise in which we use the coefficient estimates from the models to predict the vote share when the share of asylum seekers is set at its observed level in 2017, and subtract from this the model prediction when the share of asylum seekers is set to zero, holding all other variables fixed at their observed values.

Estimating such counterfactual quantities is, of course, an approximate undertaking for which caveats abound, one of the most important being the absence of general equilibrium effects. However, we would expect such effects to be modest given that asylum seekers comprise a small share of the population, well below one percent in most counties. Therefore, the results can serve as a broad gauge of the magnitude of the effects implied by the econometric estimates. Overall, we calculate that, on average, the absence of asylum seekers would have reduced the vote share for the far right by between 1.2 and 1.5 percentage points based on the predictions from the LDV and FE models that include the quadratic specification of *asylum seekers*. The corresponding predictions for the far left range between 0.4 and 0.5

percentage points. Although substantial, we would regard these estimates as conservative, since they reflect the effect of county-level variation in asylum seekers and do not fully incorporate the influence exerted through a broader awareness of asylum seekers in the country as a whole.

6 Conclusion

Three years following her pronouncement that Germany would manage the challenges of migrant inflows, Chancellor Angela Merkel's interior minister, Horst Seehofer, called immigration "the mother of all political problems," (Eddy 2018) and expressed understanding for the anger that fueled right wing protests in the town of Chemnitz in September of 2018. A month later, Merkel's Christian Democrats suffered heavy losses in local elections held in Bavaria, a state with the lowest unemployment rate in Germany, currently at 2.8%. Polling following the vote suggested that the refugee issue played a substantial part in an outcome that resulted in big gains for both the Greens and the AfD, with 28% of Bavarian voters citing refugees as the biggest problem at the state level (RTL/n-tv - Trendbarometer 2018).

Hence, from the partisan perspective of Germany's center right party, the CDU, the anecdotal and econometric evidence presented above suggest that the migration issue does indeed present a problem for the party's political prospects, one threatening its ability to maintain a governing coalition despite having presided over a prolonged period of economic stability pre-dating the financial crisis of 2008. Not only do we find that higher regional concentrations of asylum seekers have a direct and powerful effect of mobilizing electoral support for far right parties, but also that this support appears to emerge independently of economic conditions. Our preferred model shows that a one percent point increase in the share of asylum seekers is associated on average with a 1.2 percent point increase in the vote share for the far right. Moreover, we find only moderate and statistically insignificant variation in this effect size over different levels of unemployment. Following Golder's (2003) *ideational argument*, it may be that voters compartmentalize their responses to the unemployment rate and immigration, regarding the latter as a threat to national identity and culture irrespective of the state of the economy.

Our results fail to corroborate Steinmayr's (2016) analysis of the contact hypothesis, which predicts that interpersonal contact would dampen hostility toward asylum seekers and by extension dampen support for anti-immigrant far right parties. Steinmayr (2016) notes that in testing this hypothesis, it is conceptually important to distinguish between effects on the macro level, such as those transmitted via the media and political campaigns, and the micro level, as are transmitted through direct encounters between people. His analysis of election outcomes among rather small communities in Austria uncovers persuasive evidence that such direct encounters change voter perceptions about immigrants and draws them away from far right parties. Our use of a larger administrative unit, counties, dissuades us

from claiming a direct test of the contact hypothesis. As Golder (2003) suggests, differential effects of asylum seekers on voting outcomes across scales is not inherently inconsistent. Thus, our finding of a positive effect of asylum seekers on the far right vote share does not preclude the possibility that direct interactions with asylum seekers in one's immediate neighborhood could dampen support, as Steinmayr (2016) finds. Identifying the existence of such cross-scale differences and their underpinnings is a promising line of inquiry for future research.

On the left, we find that the presence of asylum seekers increases the support for far left parties, but only when the unemployment rate is moderate, with an effect size varying between 0.4 and 1.1. These results suggest that "postmaterial" considerations influence voters who share the humanitarian and egalitarian values championed by far left parties. As Inglehart's (1977) theory predicts, however, the prioritization of postmaterialist values might wane as socioeconomic conditions become more precarious, a notion our data supports. Our findings show that support for the left tapers as unemployment rises and eventually becomes negative, with the point estimate reaching a magnitude of -3.2.

In the coming years and decades, immigration to Europe will likely increase, either because refugees are fleeing civil wars or persecution or – as some have argued – because climate change will render neighboring regions less hospitable. How will this alter Europe's political landscape? For that matter, how will a global rise in refugees affect politics? Our study indicates that parties on the far right and far left will benefit at the expense of mainstream and centrist parties. Moreover, if an immigrant influx is accompanied by a rise in unemployment, then the far right may continue to benefit while support for the far left dampens. Given the confluence of high unemployment and high immigration, the picture in Germany would be of an electoral tilt to the right.

Appendix

A Instrumenting for foreigners and first stage IV-results

To account for the possible endogeneity of the variable *foreigners*, we construct an instrument that is based on the settlement pattern of foreigners distinguished by home country in the year 1999. This settlement pattern is calculated as the share τ of the number of foreigners from country g living in county i with respect to the aggregate number of foreigners from country g in Germany, i.e. $\tau_{git_0} = \text{foreigners}_{git_0} / \sum_g \text{foreigners}_{git_0}$. As a second step, we use this share to calculate a proxy for the number of foreigners from country g in county i for a later year t . To this end, we multiply the τ_{git_0} with the aggregate number of foreigners from country g in year t , i.e. $\tau_{git_0} \times \sum_i \text{foreigners}_{git} \approx \text{foreigners}_{git}$. As a last step, we take the sum of these proxies over all home-countries, which yields the shift share instrument for the aggregate number of foreigners in county i :

$$IV \text{foreigners}_{it} = \sum_g (\tau_{git_0} \sum_i \text{foreigners}_{git}). \quad (3)$$

The validity of this instrument primarily rests on the assumption that the past settlement pattern of foreigners has predictive power for the settlement pattern of newly arriving compatriots of these foreigners. This assumption, which is tested below, is widely supported by the empirical literature (Bartel 1989, Edin et al. 2003, Chiswick and Miller 2005, Glitz 2014). The second assumption, which cannot be empirically tested, is that this past settlement pattern is uncorrelated with confounders that jointly determine later settlement and voting outcomes. We deem this second assumption plausible given that we condition on all time-constant cofounders by the fixed-effects estimator. Moreover, it seems unlikely that past settlement is correlated with recent time-varying confounding variables.

Table A1 catalogues the results of the first stage regressions of the potentially endogenous variable *foreigners* on the shift share instrument *IV foreigners* and the full set of explanatory variables. Column 1 presents the results for the FE models, while columns 2 to 4 present the results for the LDV models, which differ according to the considered dependent variable in the second stage.

The estimates from Table A1 support the strength of the shift share variable as an instrument for the foreigner share. The coefficients of *IV foreigners per cap* are statistically significant at the 1% level or below and the associated F-statistics are uniformly higher than 100, thereby exceeding Staiger and Stock's (1997) rule of thumb according to which the F-statistic should exceed 10 to reject weak identification.

Table A1: First stage results with foreigners per cap as the dependent variable

| | FE - right, left, Greens | LDV - right | LDV - left | LDV - Greens |
|---------------------------|-----------------------------|---------------------|---------------------|---------------------|
| IV foreigners per cap | 0.187** (0.058) | 0.817** (0.066) | 0.816** (0.066) | 0.816** (0.066) |
| Unemployment rate | -0.069** (0.024) | 0.030 (0.025) | 0.027 (0.025) | 0.025 (0.026) |
| Population density | -0.025* (0.012) | 0.010** (0.002) | 0.011** (0.003) | 0.010** (0.003) |
| Share Abitur | -0.002 (0.007) | 0.006 (0.012) | 0.013 (0.011) | 0.007 (0.010) |
| Share old | -0.296** (0.061) | -0.149** (0.036) | -0.157** (0.036) | -0.149** (0.037) |
| Lagged dependent variable | – | -0.060 (0.068) | -0.031 (0.024) | -0.003 (0.003) |
| Constant | 0.013 (0.007) | 0.237** (0.006) | 0.025** (0.001) | 0.023** (0.007) |
| F-Stat of IV in 1st stage | 118.20 | 154.77 | 153.99 | 152.57 |
| County fixed effects | ✓ | | | |
| State × year dummies | ✓ | ✓ | ✓ | ✓ |
| No. obs. | | 2025 | | |

Note: Clustered standard errors are reported in parentheses. ** and * indicate statistical significance at the 1% and 5% level, respectively.

B Marginal effects and interactions

We calculate the average partial effects (APE) by differentiating Models 1 and 2 with respect to the variable *asylum seekers*. To illustrate, we rewrite Models 1 and 2 by subsuming all covariates, including the state × year dummies, into the vector \mathbf{X}_{it} . Only our variables of interest, *asylum seekers*, *unemployment* and their interaction are stated explicitly:

$$E(\text{vote}_{it} | \text{asyl}_{it}, \text{unemp}_{it}, \mathbf{X}_{it}) = \Phi(\alpha_1 \text{asyl}_{it} + \alpha_2 \text{unemp}_{it} + \alpha_{1,2} \text{asyl}_{it} \times \text{unemp}_{it} + \boldsymbol{\beta}^T \mathbf{X}_{it}).$$

To calculate the marginal effect, we differentiate this equation with respect to *asylum seekers*, where ϕ represents the density function of the standard normal distribution:

$$\begin{aligned} \frac{\partial E(\text{vote}_{it} | \text{asyl}_{it}, \text{unemp}_{it}, \mathbf{X}_{it})}{\partial \text{asyl}_{it}} &= \phi(\alpha_1 \text{asyl}_{it} + \alpha_2 \text{unemp}_{it} + \alpha_{1,2} \text{asyl}_{it} \times \text{unemp}_{it} + \boldsymbol{\beta}^T \mathbf{X}_{it}) \\ &\quad \times (\alpha_1 + \alpha_2 \text{unemp}_{it}). \end{aligned}$$

It is evident that the marginal effect of *asyl_{it}* on *vote_{it}* depends on *unemp_{it}* in two ways, firstly through its inclusion in the density function in the first part of the right hand side of the equation. This is a typical feature of non-linear regression models that even occurs if no interaction term is included, a source of variation referred to by Berry et al. (2010) as the *compression effect*. Second, the remaining part of the right hand side of the equation also depends on *unemp_{it}*. This is due to the explicit formulation of a multiplicative interaction term.

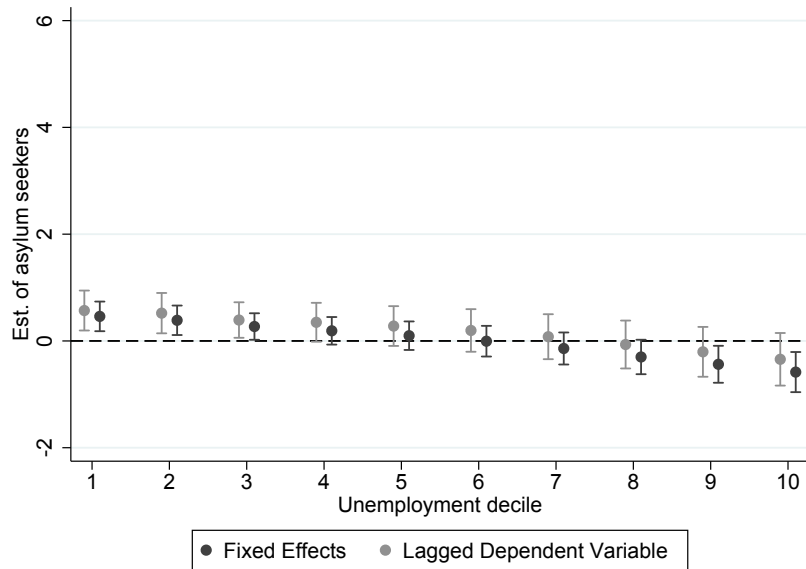
C Greens as the dependent variable

In Table A2 we follow Otto and Steinhardt (2014) by redefining the dependent variable in the model of the far left to focus exclusively on the share of votes going to the Green party. Compared to the baseline estimates in Table 3, the estimated effects of the asylum seekers are lower and no longer statistically significant at the 5% level. However, the introduction of an interaction effect between asylum seekers and the unemployment rate again leads to significant estimates of the interaction term and the coefficient for the asylum seekers. The graphic representation of the corresponding average partial effects in Figure A1 reveals a similar pattern to that of average partial effects in the baseline estimate (Figure 4), though the effects are somewhat smaller in magnitude.

Table A2: Fractional probit coefficients for the Greens

| | FE | | LDV | | FE | LDV |
|--|---------------------|---------------------|---------------------|---------------------|-----------------------|----------------------|
| | Coef. | APE | Coef. | APE | Coef. | Coef. |
| Asylum seekers per cap | 1.433 (0.800) | 0.207 (0.118) | 2.093 (1.148) | 0.301 (0.156) | 5.611** (1.253) | 5.742** (1.701) |
| Foreigners per cap | -0.086 (0.563) | -0.012 (0.084) | 0.173 (0.134) | 0.025 (0.018) | 0.051 (0.588) | 0.183 (0.134) |
| Unemployment rate | -0.401** (0.142) | -0.058** (0.021) | -0.902** (0.166) | -0.130** (0.023) | -0.345* (0.138) | -0.810** (0.173) |
| Population density | -0.059 (0.052) | -0.009 (0.007) | 0.019* (0.008) | 0.003* (0.001) | -0.055 (0.052) | 0.019* (0.008) |
| Share Abitur | -0.042 (0.047) | -0.006 (0.007) | 0.144** (0.046) | 0.021** (0.007) | -0.043 (0.046) | 0.143** (0.046) |
| Share old | -0.582 (0.392) | -0.084 (0.053) | 2.276 (0.146) | 0.040 (0.021) | -0.605 (0.394) | 0.270 (0.144) |
| Lagged dependent variable | — | — | 5.149** (0.318) | 0.741** (0.044) | — | 5.144** (0.318) |
| Unemployment × asylum seekers | — | — | — | — | -69.692** (17.239) | -55.576* (22.427) |
| Constant | -1.270** (0.092) | — | -1.672** (0.032) | — | -1.292** (0.103) | -1.676** (0.031) |
| Coef. of $\hat{\nu}_2$ (2nd stage S.E.) | 0.147 (0.517) | — | 0.065 (0.161) | — | 0.076 (0.538) | 0.041 (0.163) |
| F-Stat of IV in 1st stage | 118.20 | | 152.57 | | 118.20 | 152.57 |
| County fixed effects | ✓ | | | | ✓ | |
| State × year dummies | ✓ | | ✓ | | ✓ | ✓ |
| No. obs. | | | | 2025 | | |

Note: Clustered standard errors are reported in parentheses. ** and * indicate statistical significance at the 1% and 5% level, respectively.

Figure A1: Average partial effects for the Greens

D Estimates without control variables

We estimated parsimonious models that only included the variable *asylum seekers* to assess the stability of the results in the absence of potentially bad controls. This omission results in slightly smaller estimates in the FE model of the hard right share (Table A3) relative to the baseline estimates in Table 2. This downward tendency is more pronounced in the LDV model, rendering the estimates statistically insignificant. Conversely, removing the control variables from the model for the far left, depicted in Table A4, increases the effect of asylum seekers on votes shares compared to the baseline estimates. Again, this increase is small in the FE case and larger in the LDV case, with both sets of estimates retaining statistical significance. We conclude that the FE models are robust to the exclusion of the control variables. As the LDV models without fixed effects are generally more dependent on control variables in order to account for potential bias due to unobserved heterogeneity, it stands to reason that the estimates are less stable.

Table A3: Fractional probit coefficients for the far right without control variables

| | FE | | LDV | |
|---------------------------|---------------------|-------------------|---------------------|--------------------|
| | Coef. | APE | Coef. | APE |
| Asylum seekers per cap | 5.891* (2.706) | 0.485* (0.223) | 3.287 (1.901) | 0.270 (0.156) |
| Lagged dependent variable | — | — | 12.925** (0.479) | 1.060** (0.039) |
| Constant | -2.018** (0.036) | — | -2.294** (0.033) | — |
| County fixed effects | ✓ | | | |
| State × year dummies | ✓ | | ✓ | |
| No. obs. | 2025 | | | |

Note: Clustered standard errors are reported in parentheses. ** and * indicate statistical significance at the 1% and 5% level, respectively.

Table A4: Fractional probit coefficients for the far left without control variables

| | FE | | LDV | |
|---------------------------|---------------------|--------------------|---------------------|--------------------|
| | Coef. | APE | Coef. | APE |
| Asylum seekers per cap | 3.415** (0.953) | 0.853** (0.238) | 4.176** (0.937) | 1.035** (0.232) |
| Lagged dependent variable | – – | – – | 3.853** (0.078) | 0.955** (0.019) |
| Constant | -1.339** (0.044) | – – | -1.622** (0.022) | – – |
| County fixed effects | ✓ | | | |
| State × year dummies | ✓ | | ✓ | |
| No. obs. | | 2025 | | |

Note: Clustered standard errors are reported in parentheses. ** and * indicate statistical significance at the 1% and 5% level, respectively.

E Linear estimates

Table A6 presents the coefficients from a linear estimator, the approach employed in the majority of studies of voting outcomes using regional data. For the model of the far left vote share, we find a tight correspondence with the estimates from the non-linear model, with the pattern of APEs (Figure A3) nearly identical to that presented in Figure 4. The estimates from the linear model of the far right in Table A5, by contrast, diverges from the non-linear model, revealing an effect of asylum seekers that is roughly zero at a low unemployment rate but that increases markedly with higher unemployment. By the tenth unemployment decile, the estimated effect on the far right vote exceeds two in both the FE and LDV models, over a four-fold increase relative to the corresponding estimates from the non-linear model.

This weak correspondence may owe in part to the fact that a sizable share – roughly 10% – of the observations on the far right share cluster near zero. The resulting build-up of observations near the lower bound may render the linear model’s assumption of a constant marginal effect of the explanatory variables unrealistic. One possible manifestation associated with this pattern is predictions that fall above one or below zero. Eleven such cases of predictions below zero are found for the model of the far right (and none for the model of the far left). Owing to the potential inconsistency of the linear estimator under this circumstance (Horrace and Oaxaca 2006), coupled with a pronounced pattern of APEs whose magnitude far exceeds estimates found elsewhere in the literature, we would regard the non-linear model as the preferred estimator in this case.

Figure A2: Linear marginal effects for the far right

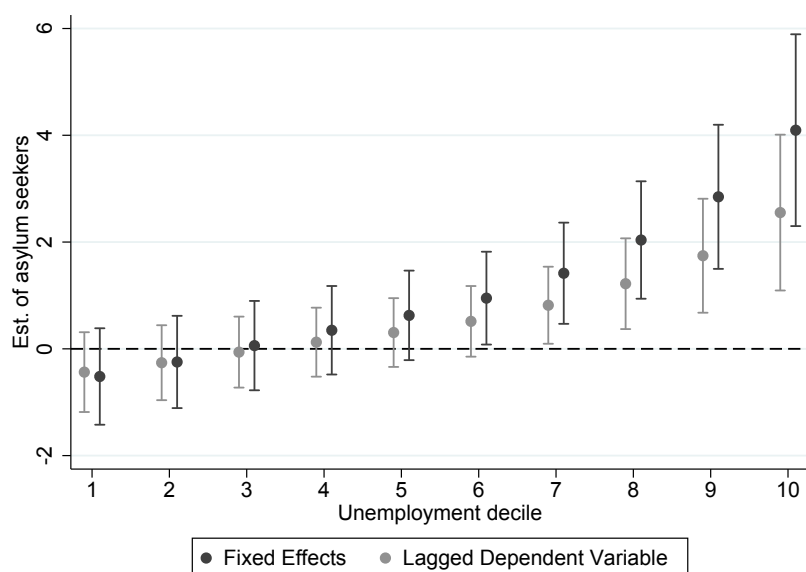


Table A5: OLS coefficients for the far right

| | FE | LDV | FE | LDV |
|--|---------------------|---------------------|---------------------|---------------------|
| Asylum seekers per cap | 0.264 (0.421) | 0.191 (0.335) | -1.886** (0.639) | -1.322* (0.557) |
| Foreigners per cap | 0.131* (0.064) | 0.008 (0.014) | 0.101 (0.063) | 0.004 (0.014) |
| Unemployment rate | -0.041 (0.034) | 0.065** (0.013) | -0.075* (0.034) | 0.029* (0.013) |
| Population density | -0.085** (0.020) | -0.001 (0.001) | -0.085** (0.019) | -0.001 (0.001) |
| Share Abitur | -0.030* (0.012) | -0.027** (0.004) | -0.030* (0.012) | -0.026** (0.004) |
| Share old | 0.101 (0.067) | 0.056** (0.018) | 0.125 (0.067) | 0.057** (0.018) |
| Lagged dependent variable | – | 0.900** (0.041) | – | 0.891** (0.041) |
| Unemployment × asylum seekers | – | – | 34.795** (7.516) | 22.539** (6.509) |
| Constant | 0.221** (0.010) | 0.200** (0.008) | 0.219** (0.010) | 0.202** (0.008) |
| Coef. of $\hat{\nu}_2$ (2nd stage S.E.) | -0.058 (0.046) | 0.038 (0.034) | -0.061 (0.047) | 0.047 (0.033) |
| F-Stat of IV in 1st stage | 118.20 | 154.77 | 118.20 | 154.77 |
| County fixed effects | ✓ | | ✓ | |
| State × year dummies | ✓ | ✓ | ✓ | ✓ |
| No. obs. | 2025 | | | |

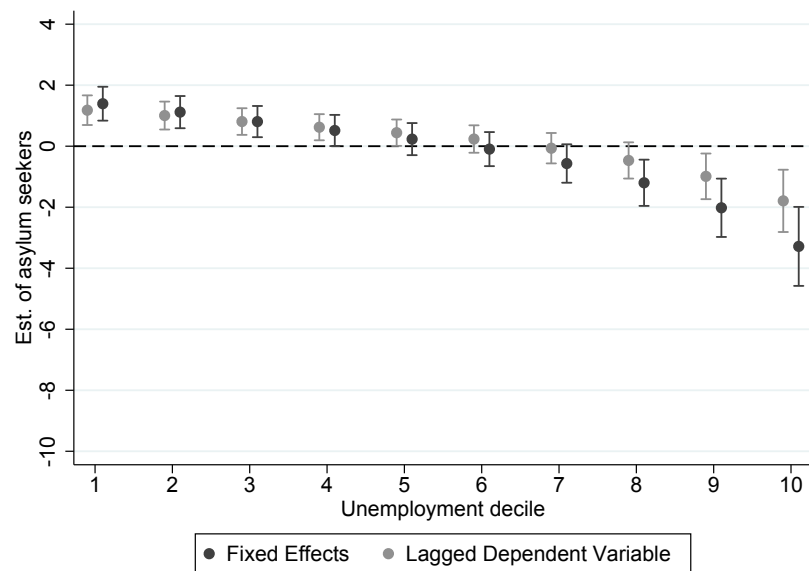
Note: Clustered standard errors are reported in parentheses. ** and * indicate statistical significance at the 1% and 5% level, respectively.

Table A6: OLS coefficients for the far left

| | FE | LDV | FE | LDV |
|--|---------------------|---------------------|----------------------|----------------------|
| Asylum seekers per cap | 0.601* (0.238) | 0.559** (0.200) | 2.780** (0.417) | 2.062** (0.365) |
| Foreigners per cap | 0.107 (0.114) | 0.020 (0.011) | 0.143 (0.122) | 0.024* (0.011) |
| Unemployment rate | 0.078 (0.040) | 0.067** (0.012) | 0.112** (0.039) | 0.104** (0.014) |
| Population density | 0.037* (0.015) | 0.002** (0.001) | 0.038* (0.015) | 0.002** (0.001) |
| Share Abitur | 0.017 (0.011) | 0.035** (0.005) | 0.018 (0.011) | 0.034** (0.005) |
| Share old | -0.497** (0.077) | -0.072** (0.015) | -0.519** (0.078) | -0.073** (0.015) |
| Lagged dependent variable | – | 0.925** (0.012) | – | 0.925** (0.012) |
| Unemployment × asylum seekers | – | – | -35.274** (5.428) | -22.416** (4.451) |
| Constant | 0.201** (0.022) | -0.071** (0.006) | 0.201** (0.024) | -0.073** (0.012) |
| Coef. of $\hat{\nu}_2$ (2nd stage S.E.) | -0.086 (0.114) | -0.015 (0.021) | -0.089 (0.121) | -0.025 (0.021) |
| F-Stat of IV in 1st stage | 118.20 | 153.99 | 118.20 | 153.99 |
| County fixed effects | ✓ | | ✓ | |
| State × year dummies | ✓ | ✓ | ✓ | ✓ |
| No. obs. | 2025 | | | |

Note: Clustered standard errors are reported in parentheses. ** and * indicate statistical significance at the 1% and 5% level, respectively.

Figure A3: Linear marginal effects for the far left



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