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On war and political radicalization: Evidence from forced conscription into the Wehrmacht

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ABSTRACT

This paper investigates when and how political preferences that were shaped by conflict express themselves into electoral outcomes. During World War II, the Third Reich annexed the French eastern borderlands and forcibly conscripted their inhabitants into the Wehrmacht. While conscription was introduced in both annexed regions, the administrators' independence gave them broad discretion on how to apply the policy. As a result, different birth cohorts were drafted in different regions. I show that individuals conscripted and their descendants display reduced levels of political trust. This attitude is not reflected into aggregate municipal electoral outcomes at once. It translates into voting only with the emergence of parties that are large enough and radical enough. In the absence of parties that fulfill both conditions, these preferences lead to higher abstention.

1. Introduction

In the last decade scholars have studied extensively how war violence affects individual political preferences and attitudes. While qualitatively the effects seem to depend on whether one is looking at between or within-group cooperation, there is now broad quantitative evidence indicating that war violence does affect political attitudes (Bauer et al., 2016; Rohner and Thoenig, 2020). Nonetheless, little is known on how and under which conditions these attitudes translate into observable equilibrium outcomes, such as electoral results or policy. This comes as no surprise; by focusing on preferences and attitudes of survivors of conflict, the literature has put its focus on the demand-side of electoral politics, and for good reason. Since these studies typically take place in developing countries, shortly after the conflict has occurred, the context is not appropriate to assess equilibrium outcomes which is what economists are ultimately interested in. This paper does a step in this direction.

In the Downsian tradition of rational choice models, individuals weight the benefits and costs of joining collective action (the so-called "calculus of voting"). Attitudes might thus not be observable in electoral results, or might not be observable at all times. In this paper I demonstrate *when* and *how* political attitudes that were altered by war express themselves into equilibrium (electoral) outcomes. I show that – in the context of this study – individuals conscripted and their descendants display reduced levels of political trust. This attitude is not reflected into aggregate municipal electoral outcomes at once. It translates into voting with the emergence of parties that "radically criticize the existing social and economic order", defined as radical parties by Backes (2009).[2] In the absence of parties that are both radical enough and large enough, this attitude leads to higher abstention.

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[1] This paper features an Online Appendix containing additional results and data description.

[2] Backes' definition is also close to that of Capoccia (2002) who defines anti-system parties as "*parties or groups that exert a radical form of opposition*". The blanket term "right-wing populist" is increasingly used nowadays in economics to refer to such parties.

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The setting studied is Alsace and Moselle in France which were annexed by the Third Reich during WWII and whose inhabitants were forcibly conscripted into the Wehrmacht.[3] While this setting is peculiar in that conscription does not simply pertain participating to war, it offers several advantages. Firstly, it allows to identify a plausibly exogenous source of variation in conscription. Second, the time scope of this study (elections over more than 50 years), and its geographical extend (the annexed regions only represent 5% of the electorate) allow to exploit variation in national political supply that is credibly exogenous to this WWII experience. Finally, the ambiguous conditions of the re-integration of these soldiers into the post-WWII French state are prone to the emergence of attitudes of distrust towards established political institutions.

In the early post-WWII elections, in which the radical right was either absent or marginal, voting in the annexed lands was characterized by higher levels of abstention. In the mid-1980s, the radical right with a pronounced anti-establishment discourse emerged in France. Since then, these regions have been strongholds of such candidates, systematically exceeding the national averages. Building on the extensive literature that links war to political preferences, and the success of the radical right in attracting distrustful voters, I argue that this voting behavior reflects reduced political trust due to WWII conscription.

The identifying variation emerges from the fact that while conscription into the Wehrmacht was introduced in both Alsace and Moselle, different cohorts were conscripted in each. In Alsace, which was integrated within the neighboring German Baden *Gau*, the relevant cohorts were born between 1908 to 1927; in Moselle, which was integrated within the neighboring Saar-Palatinate *Gau*, the relevant cohorts were born between 1914 and 1927. The reasons for the diverging conscription rules remain unclear even today; nonetheless, it is generally thought that the discretionary powers of the administrators of these two *Gaue* – already appointed in 1933 and 1935 – allowed them to choose different policies when it came to ideological and cultural assimilation.

I exploit this variation to estimate the impact of war on municipality-level electoral outcomes for the 1965–2017 period in a fuzzy regression discontinuity design. Identification is achieved using the number of eligible births by municipality as an instrument for conscription while simultaneously controlling for local birth rates and a wide range of covariates (including pre-war voting). In other words, identification fixes the compliance rate (through the first-stage estimation) while accounting for variation arising from fertility. Results confirm the descriptive evidence: in the earlier elections, when the radical right is a marginal force, abstention is higher in places more affected by conscription, while in more recent elections, there is increased support for radical right-wing candidates in these locations. Conscription also affects the demographic structure of municipalities, namely overall population, and the share and age of the male population. While some of the observed differences in voting can be attributed to changes in demography, this channel can only account for a minor fraction of the effect. Moreover, differences in demography persist, and are thus unlikely to explain the transition from abstention to radical support.

I then make use of content analytic discourse data to determine the extend to which presidential candidates are radical, regardless of party family. Following Backes (2009), I focus on a candidate's discourse that has a negative tone to construct the measure. Heterogeneity with respect to a candidate's degree of radicalism reveals a consistent pattern. While presidential candidates receive on average lower vote shares in municipalities where more men were conscripted, the effect is mitigated by a candidate's radicalism. In early election however, no candidate is radical enough to overcome the decrease in average vote shares, while in later ones, the most radical candidates (those of the radical right) are. This leads to a pattern of "activated history", where attitudes are initially not detectable, but translate into support with the rise of parties that can capture the resentment (Ochsner and Roesel, 2017; Cantoni et al., 2019). Other policy positions that are central to the radical right (such as nationalism and authority) are unable to explain this pattern.

To understand the transition from abstention to radical support, I develop a theoretical framework based on Glaeser et al. (2005) that introduces war mobilization as a negative shock to voters' levels of political trust. This twist to the model yields several testable predictions: Within an election, in municipalities where a higher proportion of men were conscripted (1a) political candidates with a more radical program should have larger shares of the vote as long as platforms are not *too* similar; (1b) moderate candidates should get smaller shares of the vote; (1c) abstention should be higher. Across elections (2a) both radical and moderate candidates should have, all else held equal, lower shares of the vote in elections where policy platforms are similar, and (2b) this effect should be larger in places where a higher proportion of men were conscripted. The predictions of the model are then empirically tested using variation in political discourse across candidates (in the same election), and across time (in different elections). The results indicate that conscription affects candidate vote shares asymmetrically, such that in highly polarized elections the effect of conscription is captured by more radical candidates, while when platforms are similar, both radical and moderate candidates are penalized, resulting in higher abstention.

Finally, to disentangle the underlying mechanism, I exploit survey data and implement a simple difference-in-differences strategy. I first confirm the aggregate findings by showing that men born between 1907 and 1913 in Alsace (drafted) are more likely to state proximity to the radical-right than their Moselle counterparts (not drafted), while there is no difference for the cohorts born from 1914 to 1927 (both drafted). I then explore political attitudes. The treated group (men born in 1908–1913 in Alsace) is less trustful of elected politicians than the control one (men born in 1908–1913 in Moselle), while, once again, there is no difference for the 1914–1927 cohorts. This is not the case for women born during the relevant periods in Alsace and Moselle, nor for men born in regions where conscription did not take place. This difference cannot be attributed to differential career prospect, organizational skills, or endogenous casualties, and is present in descendants of conscripts, pointing towards a slow changing attitude of political alienation that persists through transmission, vertical but also potentially oblique (Cavalli-Sforza and Feldman, 1981).

[3] Until 2016, Alsace was a region that consisted of two departments (Bas-Rhin, Haut-Rhin), while Moselle was a department in Lorraine. Both belong to the Grand-Est region since 2016. They are henceforth referred to as annexed departments, regions, and lands, interchangeably.

Literature review – The recent electoral success of parties that fall outside the traditional, bipolar Social-Democrat and Christian-Democrat divide that has dominated European politics through the second half of the 20th century has attracted considerable attention in public and academic discourse. This success has coincided with the economic downturn which has focused attention on short-run determinants, such as social and economic insecurity.[4] In contrast, the analysis presented here focuses on long-run determinants of support for radical politicians which have roots in deeper and slower-changing attributes of the electorate, such as political preferences.

While economists have taken preferences as given for a long time, a recent literature has emerged that shows that preferences can not only be influenced by history, but they can also be reactivated by particular events.[5] Fontana et al. (2018) and Koenig (2015) look at political support in the immediate aftermath of wars and Ochsner and Roesel (2020) and Schindler and Westcott (2015) focus on long-run effects. The effects of war and military service on political participation were also examined by Alacevich and Zejcirovic (2020) and Fize and Louis-Sidois (2020). My paper reconciles these findings by demonstrating that abstention and political support can be the expression of the same attitude depending on the political landscape. The idea that political campaigns can reactivate pre-existing preferences is present in both Ochsner and Roesel (2017) and Cantoni et al. (2019). The main contribution with respect to these two papers is the longer temporal scope and the use of quantitative, time-varying measures of political discourse.

This paper naturally builds on a literature that has explored the impact of war on political attitudes in developing countries using survey data. The evidence is mixed, depending on whether one is looking at between or within-group cooperation. Blattman (2009), Bellows and Miguel (2009), Voors et al. (2012), Gilligan et al. (2014), and Bauer et al. (2018) show that witnessing violence had a positive impact on political participation while Cassar et al. (2013), Rohner et al. (2013), and Grosjean (2014) find detrimental effects on trust. For the right or wrong reasons, WWII Europe has attracted very little attention.[6] The analysis presented here also goes one step further by evaluating when and how these attitudes express themselves in aggregate electoral outcomes.

The historical developments in Alsace and Moselle are also exploited in work by Dehdari and Gehring (2019) on regional identity. Whereas Dehdari and Gehring (2019) exploit variation between Alsace–Moselle and the rest of Lorraine, I compare Alsace to Moselle. The two papers are thus complementary in explaining differences in attitudes and voting between Alsace, Moselle, and the rest of France. Nazi institutions such as Economic Aryanization, the Professional Civil Service, and the Hitler Youth have already been investigated in Huber et al. (2021), Waldinger (2011), and Voigtländer and Voth (2015), respectively. As far as I am aware, Wehrmacht conscription has not been examined before. Finally, by looking at attitudes over such a long time period this analysis is also linked to the literature on intergenerational transmission of values and beliefs, as presented theoretically by Bisin and Verdier (2001) and Doepke and Zilibotti (2008) and empirically tested by Campante and Yanagizawa-Drott (2015), and Fernández and Fogli (2009).

The structure of the paper is as follows: Section 2 presents the background. Section 3 presents municipal-level evidence that establishes a link from conscription to abstention in early elections and support for radical candidates in late ones. Section 4 proposes a theoretical framework to explain this transition and tests its predictions. Section 5 presents individual-level evidence relating conscription to political attitudes and discusses potential mechanisms. Section 6 concludes.

2. Historical background and data

2.1. Historical background[7]

The Alsace and Moselle border – The regions of Alsace and Lorraine were created in 1790 and their border was stabilized in 1793, when the County of Saarwerden was annexed by the French Republic. Following the defeat of the French troops in the Franco-Prussian War of 1870–1871, Alsace and parts of the Moselle and Meurthe departments in Lorraine jointly formed the *Imperial Territory of Alsace-Lorraine*. They remained a part of the German Empire until the end World War I and the 1918 armistice, that resulted in the re-integration of Alsace and the newly re-founded Moselle department to the Third French Republic.

Alsace and Moselle were annexed once again by the Third Reich following the French capitulation in June 1940. Unlike the previous annexation, in this case the two departments were absorbed into the neighboring pre-existing German *Gaue* of Baden (in the case of Alsace) and Saar-Palatinate (in the case of Moselle), as shown in Fig. 1. Alsace and Moselle were administrated separately: the first by Robert Wagner, a soldier by profession, and the latter by Josef Bürckel, a teacher by profession.

Wehrmacht conscription[8] – In general *Gauleiters* were subject to the authority of the occupying army. However, in August 1940, Hitler issued a decree granting full civil control to Wagner and Bürckel. Therefore, the two administrators of Alsace and Moselle possessed virtually unrestricted civil powers, and essentially were responsible only to Hitler himself. The two administrators held similar positions but their personalities and methods differed significantly (Iung et al., 2012). Wagner, a WWI veteran, was of the view that the Wehrmacht and the party would be the means by which the local youth would complete their ideological and cultural

4 See among others, Couttenier et al. (2019), Drago et al. (2017) for crime and voting; Dustmann et al. (2018), Halla et al. (2017), Otto and Steinhardt (2014) for immigration; Algan et al. (2017), Autor et al. (2020), Colantone and Stanig (2018) for unemployment and trade exposure.

5 Several papers look at determinants of differences in preferences for redistribution or gender attitudes (Alesina and Fuchs-Schündeln, 2007; Alesina et al., 2013; Campa and Serafinelli, 2019).

6 The sample in Grosjean (2014) also comprises five European countries and mainly focuses on WWII. Homola et al. (2020) also look at Nazi concentration camps and individual attitudes today.

7 An extended version of the background can be found in Online Appendix Section D.1.

8 The conscription background is taken from Riedweg (1995) unless otherwise specified.

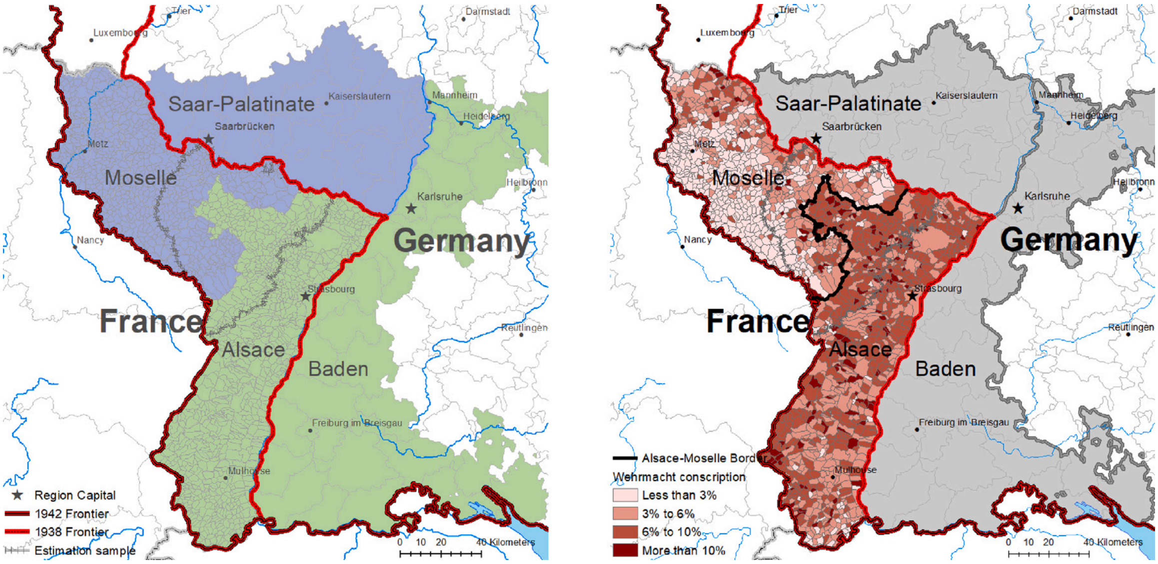

Fig. 1. WWII Annexation and Wehrmacht conscription.
Notes: Left-hand-side: Map of the western German border in 1938 and 1942. With the annexation of Alsace and Moselle by the Third Reich in 1940, Alsace consolidated the pre-existing Baden *Gau*; Moselle and Saar-Palatinate formed the Westmark *Gau*. The shaded area presents the sample used for the estimation. Right-hand-side: Wehrmacht conscription (% of 1936 population) by municipality.

assimilation. Bürckel, which was considered a "nazification" (*Gleichschaltung*) expert after being in charge of nazifying the Saar region, believed that assimilation through education was not possible and openly considered deporting part of the population and replacing them with German farmers. The course of the war on the Eastern Front favored Wagner's approach. The decision to grant German nationality to the populations of Alsace and Moselle was made on August 9, 1942. This made it possible to introduce military service in Alsace on August 25, 1942, and in Moselle on August 29, 1942. The administrators' independence gave them broad discretion in implementing the policy. Wagner who felt that assimilation could be accomplished through military service, mobilized 20 cohorts in Alsace (1908–1927), while Bürckel only mobilized 14 in Moselle (1914–1927).[9] In total 103,000 men from Alsace and 31,000 from Moselle were drafted into the Wehrmacht (MACVG, 1954). The process was otherwise identical in both departments (Iung et al., 2012). According to historians "*the responsibility falls entirely on Gauleiter Wagner who did everything in his hand so that a maximum of Alsatians are incorporated in the Wehrmacht*" (Riedweg, 1995, p.99), an explanation also put forward by the French National Statistical Institute after the War (INSEE, 1956, p.205).

Being in the Wehrmacht during the War – Conscription in the Wehrmacht was two years long and consisted of four stages: registration, first (medical) examination, drafting, and call-up (Iung et al., 2012). Police authorities and the ordinary local registration of the civilian population, were responsible for the registration of men liable for military service (Ambrose, 1990). Depending on the unit, instruction could last from two to several months.[10] Upon arrival in the barracks conscripts from the annexed lands were mixed with conscripts from all over the Reich.

The German military command distrust led to several special arrangements concerning these conscripts: Delays between the stages were shorted, they were not allowed on the occupied West Territories nor in certain units, there should not be more than 8% to 15% of them per ground army unit (5% for battle units), and they only served on the Eastern Front (Iung et al., 2012; Riedweg, 1995). Note however, that the German military allowances were larger than the French ones. They were capped at 85% of previous income and were valid for all employment types. As such, families of men incorporated were not in need during the annexation period (Riedweg, 1995).

Being in Moselle during the War – In the meanwhile, life in Moselle went slowly back to normal. Street were renamed, all administrative managers were replaced, and public administration was reorganized. An urbanization plan that intended to merge small municipalities into larger ones was introduced and public transport was intensified. Large manifestations were frequent

[9] In Luxembourg, *Gauleiter* Simon, who was against the introduction of conscription, only mobilized the 1920 to 1927 cohorts, without eliciting a reaction by the German High Command.

[10] Training in the Wehrmacht was primarily military, most of the indoctrination taking place in the "Empire Labor Service" for ages 17 to 25 and the "Hitler Youth" for ages 10 to 18.

during the period. Measures were introduced to overcome potential shortages such as training days for business managers and incentives to farmers (Sary, 1983). In parallel, Brückel proceeded to three waves of deportation of inhabitants not belonging to the "German people", recent movers from other regions of France and inhabitants of French-speaking municipalities.[11] Nonetheless, as summarized in Sary (1983), *"during fours years, the occupying authorities worked with determination to earn the sympathy of the population of Moselle. A situation and activity of quasi-normality was maintained in Metz, with an intense cultural life, several cultural events, and a supply of goods that was sufficient"*.

Returning after the War – In the immediate aftermath of the War, a strong anti-Germanic feeling dominated the French society. The figure of the patriot resistant monopolized the collective memory. Wehrmacht conscription was incompatible with the post-WWII impersonation of heroism in France (Bludszus, 2014). These soldiers had to justify their involvement to a French public opinion that was not moved by their fate. During their return through Paris, they were often insulted and spat on, Parisians not distinguishing between them and volunteers in the Nazi legion (Iung et al., 2012). The French state did not intervene, which contributed to the misconception of their role during the War; for much of the public, these men were traitors. As Paul Durant, a WWII veteran, recalled *"there was only bullying and persecution for the "Müss Preussen" soldiers that we were"*, while for Madeleine Lemoine, *"the climate [...] was as painful as the Hitler climate"* (Bludszus, 2014, p.141).

In their view, men conscripted were let down both by the Vichy government, that only protested mildly against their incorporation into the Wehrmacht, and by the re-founded French state for not taking a public stance to defend them. The ambiguous position of the French state is well illustrated by the 1953 "Bordeaux trial" that convicted conscripts involved in the 1944 Oradour-sur-Glane massacre. This trial, that stigmatized the *Malgré-nous* according to local newspapers, was experienced as an "intolerable humiliation" and led to an uproar of the local population, forcing the government to pass an amnesty law. The outcome was nonetheless interpreted as the expression of the misunderstanding of the annexation in the rest of the country (Iung et al., 2012).

The veterans went on to form "Against our will" associations (*Malgré-nous*) – a name that indicates the cynical view of their engagement – with the purpose of repatriating war veterans still in captivity, but also their recognition and compensation as war veterans to a full extend (Bludszus, 2014).[12] They did not form a political party, unlike veterans from Luxembourg that went on to form a single-issue anti-establishment party. In spite of that, this experience is still considered nowadays to have greatly affected the political and social life in the region (Vogler, 1995, p.274).

2.2. Post-WWII elections in France

The emergence of the radical right – The radical right resurfaced in the 1984 European parliament election when a list led by Jean-Marie Le Pen, who claimed that the policies of both left- and right-wing governments "betrayed popular trust", received 11% of the vote. Up to that point, the radical right-wing had campaigned with very limited success. In 1965 Jean-Louis Tixier-Vignancour received 5% of the votes on a platform to keep Algeria French, while in 1974 Le Pen only received 0.8% of the vote and did not run in 1981. The success of the radical right was confirmed in the 1988 Presidential election with 14% of the vote. In his campaign, Le Pen wondered out loud "why [mainstream candidates] would do tomorrow what they did not know how to do yesterday". The radical right has had a presence in every presidential election since.

Electoral participation was systematically lower in Alsace and Moselle than the rest of France until the 1980s. Since then, these regions have been strongholds of the radical right, where its share of the vote is well above the national average in every election. In 1988, Le Pen received 21% of the valid ballots in Alsace and Moselle, 47% more than the national average. In 1995, Le Pen received the highest share of votes in the annexed departments *out of the 101 departments in France* (66% more than the average). In the elections from 2002 to 2017, the radical right's vote share exceeded the national average by between 3 and 8 percentage points (22 to 41% above average).

Radical candidates and political distrust – Backes (2009) defines radical candidates as the ones that "*radically criticize the existing social and economic order*". Following this definition, I focus on a candidate's discourse that has a negative tone to construct a measure of candidate radicalism. Political discourse data comes from the *Comparative Manifesto Project* and the *Euromanifesto Project* (henceforth jointly CMP). This content-analytic data classifies political discourse by topic (e.g. military expenditures and tariffs), and tone (positive and negative mentions). Examples range from negative references to military power (usually a radical left topic), international co-operation (usually a radical right topic), but also the European Union (both radical right and radical left).

To assess whether candidates that use a negative language attract support from distrustful voters, I exploit the post-electoral surveys that have been taking place in France since the 1950s. These surveys ask questions both on voting in presidential elections and trust in politicians. Using this data, I estimate election-specific multinomial logit models of the effect of political distrust on the choice of presidential candidate.[13] The correlation between the measure of a candidate's radicalism and the relative likelihood

[11] Roughly 15% of the population was deported, as opposed to 3% in Alsace (INSEE, 1956).

[12] Books on conscription also evoke this feeling, such as "The great disgrace" in 1965, "The shameful soldier" in 1972, or "The night of the pariahs" in 1975 (Iung et al., 2012).

[13] The surveys used are the post-electoral surveys of 1969, 1978, 1995, 2002, 2007, 2012, and 2017. The political distrust question is "Do you think politicians on the whole care about what people like you think?". More details can be found in Online Appendix Section B.6.

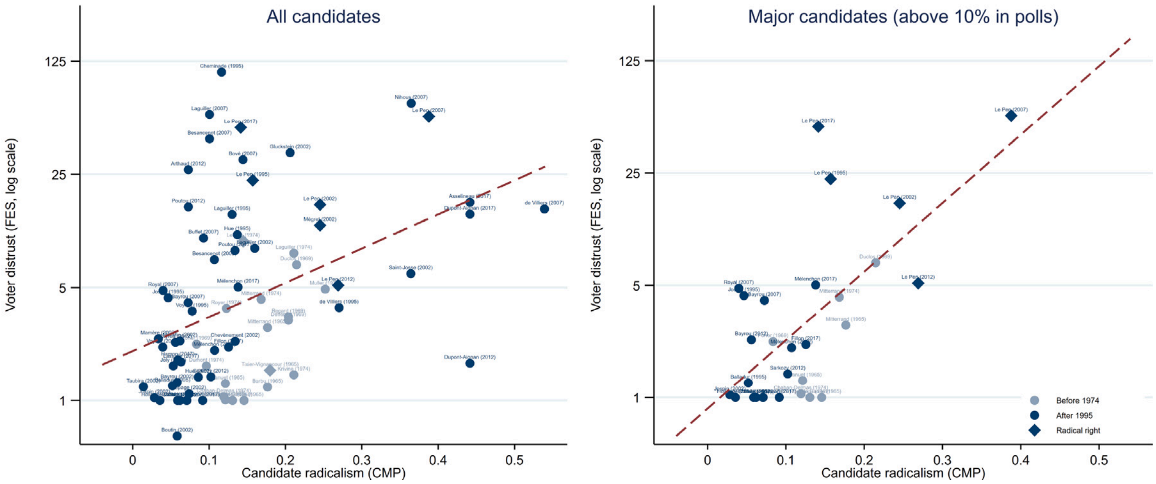

Fig. 2. Radical candidates and political distrust.
Notes: Use of political discourse with a negative tone and relative likelihood a distrustful voter chooses a candidate. Candidate radicalism is measured as the share of the discourse with a negative tone; the data comes from the *Comparative Manifesto Project* and the *Euromanifesto Project* (Volkens et al., 2018; Schmitt et al., 2018). Political distrust is the relative risk ratio of from a multinomial logit estimation of candidate choice on distrust in politicians; the data comes from the 1969, 1978, 1995, 2002, 2007, 2012, and 2017 *French Electoral Studies*. Candidates before 1974 in light blue; after 1995 in dark blue; radical right-wing candidates in diamonds. (For interpretation of the references to colour in this figure legend, the reader is referred to the web version of this article.)

a distrustful voter supports this candidate (the relative risk ratio) is presented in Fig. 2. Consistently with the anecdotal evidence, before the 1980s neither candidates' programs were very radical (in the Backes, 2009, sense), nor were there large differences between the electoral choices of trustful and distrustful voters.[14]

3. Abstention and support for radical candidates

I start by focusing on the relationship between aggregate voting outcomes and conscription. To do so, I combine municipality-level conscription data with electoral outcomes from 1965 to 2017 and candidate-level content-analytic data.

3.1. Data

Conscription proxy – Data on Wehrmacht conscription is taken from the *Index of French Nationals Compelled into German Armed Forces* (MACVG, 1945, 1946). It includes 44,527 individuals which were declared missing in an official census carried out in October 1945 with the sole purpose of repatriating French prisoners of war held in allied camps. The data was digitalized, and 44,034 individuals were matched to 1,435 contemporary municipalities, which means that 91% of all municipalities existing in 2018 have at least one man declared missing. In the case of an average municipality, 2.3% of the 1936 population was declared missing (28 men).

Two assumptions are necessary for the number of men figuring in the index to be a valid approximation of the number of men conscripted: (i) mortality and the imprisonment rate should be independent of municipality characteristics, and (ii) casualties should not be systematically communicated in some municipalities but not others. These assumption are supported by historical evidence indicating that men from Alsace and Moselle were scattered across the Eastern front, and that individual characteristics were not taken into account in this allocation. The key independent variable, namely the proxy of the fraction of a municipality's 1936 population conscripted into the Wehrmacht, is constructed by rescaling the number of men declared missing by $134,000/44{,}500$, where 134,000 (44,500) is the total number of individuals conscripted (declared missing).[15]

Pre-war births – Birth data is taken from the decennial civil status registers (*Tables décennales de l'état civil*), which are available on the websites of the Bas-Rhin (Alsace) and Moselle archives. The data was collected for three decennial censuses (1903–1932) for the 462 municipalities within 20 km of the Alsace–Moselle border. Using the censuses, I calculate the fraction of a municipality's 1936 population that was born during the 1908–1913 and 1914–1927 periods to get the number of eligible births. These are used as an instrument for conscription. Birth rates per decade (1903–1912, 1913–1922, and 1923–1932) are also constructed.

[14] The radicalism measure also correlates strongly with expert measures from the *Chapel Hill Expert Survey*. More details on its construction can be found in the Online Appendix Section B.5.

[15] Further details on the calculation of the proxy are presented in Online Appendix Section B.1.

Electoral outcomes – The 1962 French constitutional reform introduced universal suffrage in Presidential elections. Data for the 1965, 1969, and 1974 elections was obtained from the French National Archives and digitalized. Later electoral results (1995, 2002, 2007, 2012, and 2017) are available from the data platform of the French Republic. Surprisingly, the copying of the 1981 and 1988 municipality-level results is prohibited. The dependent variables, a candidate's gross share of the vote and abstention for these presidential elections, are constructed using this data.

Covariates – The 1936 parliamentary election results are taken from Lachapelle (1936), while other pre-war data come from the 1936 and 1946 census results in INSEE (1956), which are only available at the *canton* level.[16] The population of municipalities in 1936 is taken from the Cassini Database. Municipality-level religious affiliation (Roman Catholic, Protestant, Jewish) is approximated using church data. Municipality-level linguistic makeup (French- or German-speaking) is approximated using 1891 to 1940 family names at birth which are available in the *1891 to 1990 Family Name File* of the French National Statistics Institute. Data on contemporary municipality characteristics is taken from the 1968, 1975, 1982, 1990, 1999, 2009, and 2014 population censuses.

3.2. Estimation strategy

Conscription determinants – I first evaluate the importance of the draft rule in the conscription process. To do so, I collected data from the 1936 population census (*Listes nominatives*) for 134 Alsace municipalities that border Moselle.[17] This individual-level data includes characteristics such as date of birth, religion, nationality, and language. Individuals can thus be mapped from the census to the conscription data, since the latter is at the individual level as well, allowing to evaluate the importance of individual characteristics on conscription.

The individual-level regressions are presented in Table 1. Eligibility is the most important predictor (t-stat=6.2); "other religions" (primarily Jewish) correlates negatively with conscription; Protestants are not more likely to be conscripted than Roman Catholics (the omitted category). French-speaking individuals (as opposed to non-french speaking) are more likely to be conscripted. While eligibility alone can explain 10% of the total variation in conscription (see Column 1), adding individual characteristics and municipality fixed effects only increases the regression fit from 10 to 13% (Column 2). Columns (4) and (5) focus on the compliance decision by restricting the sample to individuals eligible for Wehrmacht conscription (born in 1908–1927). Despite their limited explanatory power, individual characteristics do matter jointly (as shown by the joint significance test), a fact that motivates the use of the draft rule as an exogenous source of variation in conscription. The last two columns of Table 1 focus on individuals that potentially volunteered (born in years not drafted). Individual characteristics have no explanatory power in this decision.

Identification strategy – The regression equation (omitting the time subscript) is

$$\pi_m = x'_m \gamma + \rho \cdot Conscription_m + \eta_m \tag{1}$$

where π_m is a voting outcome in municipality m, $Conscription_m$ is the share of m's population that was conscripted, and x_m is a column vector of municipality covariates. The coefficient of interest is ρ which captures the change in voting outcomes when conscription increases by 1 pp.

Variation in conscription originates from three sources: fertility, the draft rule, and differential compliance. To focus on the variation arising from the draft rule, I exploit the existence of a border discontinuity in the conscription rule and implement a fuzzy regression discontinuity design, in the spirit of Dell (2010). I hence estimate the following first-stage relationship

$$Conscription_m = x'_m \pi_{10} + \pi_{11} \cdot Eligible_m + \pi_{12} \cdot Fertility_m + f(lat_m, lon_m) + \xi_{1m} \tag{2}$$

where $Eligible_m$ is the eligible population in m, male births during 1908–1927 in Alsace and during 1914–1927 in Moselle, over pre-war population, and $f(lat_m, lon_m)$ is a polynomial in latitude and longitude of m. Variation in the eligible population also originates from differences in fertility, it is thus important to control for birth rates in m. The coefficient of interest in the first-stage ($\hat{\pi}_{11}$) captures average conditional compliance, in other words, the percent increase in conscription when eligible births increase by 1 percentage point, in nearby locations with similar fertility. Identification is therefore based on neighboring municipalities across the Alsace–Moselle border, that have comparable fertility rates but different fractions of men conscripted because of the different draft rules in the two regions. The instrumental variables specification is

$$\pi_m = x'_m \gamma + \rho \cdot \widehat{Conscription}_m + \delta \cdot Fertility_m + f(lat_m, lon_m) + \epsilon_m \tag{3}$$

The coefficient of interest in Eq. (3), namely ρ, is similar to a Wald estimator since it re-scales the effect of eligibility by average conditional compliance.

[16] *Cantons* are the second lowest administrative unit in France. In 1936 there were 93 cantons in Alsace and Moselle. An average canton comprised 18 municipalities and an area of 155 km^2.

[17] The 134 municipalities in the 1936 Saverne *arrondissement*, that correspond to 124 municipalities as of 2018. The arrondissement chosen has two advantages: (i) it borders on Moselle, and (ii) its municipalities are highly heterogeneous in terms of religious and linguistic composition. The Moselle census was destroyed in 1942. More details can be found in Online Appendix Section B.4.

Table 1
Estimation strategy: Conscription determinants.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dep. variable	Individual is conscripted						
Sample	Full sample (male)			Eligible (male)		Ineligible (male)	
Specification	No controls	Within municipality	Within year-of-birth	Within municipality	Within year-of-birth	Within municipality	Within year-of-birth
Born 1908–1927	0.148***	0.129***					
	(0.020)	(0.021)					
Age		−0.002		0.035		0.000	
		(0.001)		(0.032)		(0.000)	
Age × Age		0.000		−0.001		−0.000	
		(0.000)		(0.001)		(0.000)	
French national		0.020	0.018	0.048	0.058	0.002	0.004
		(0.015)	(0.015)	(0.052)	(0.051)	(0.002)	(0.003)
German national		0.106	0.109	0.139	0.131	0.108	0.112
		(0.082)	(0.084)	(0.138)	(0.148)	(0.100)	(0.102)
Protestant		0.004	0.004	0.002	0.006	0.001	0.001
		(0.009)	(0.009)	(0.029)	(0.028)	(0.002)	(0.002)
Other religions		−0.034**	−0.027*	−0.079**	−0.074**	−0.000	0.000
		(0.014)	(0.015)	(0.030)	(0.031)	(0.002)	(0.002)
French-speaking		0.024***	0.011**	0.039**	0.026	0.002	0.001
		(0.006)	(0.005)	(0.018)	(0.018)	(0.002)	(0.002)
Household head		0.007	0.002	0.041	0.025	−0.003	−0.002
		(0.011)	(0.006)	(0.030)	(0.031)	(0.003)	(0.002)
Household size		0.002	−0.000	0.000	0.000	−0.000	−0.000
		(0.001)	(0.001)	(0.004)	(0.004)	(0.000)	(0.000)
Municipality (1936) FE		Yes	Yes	Yes	Yes	Yes	Yes
Year-of-birth FE			Yes		Yes		Yes
Mean dep. variable	0.047	0.047	0.047	0.150	0.150	0.002	0.002
Observations	7,590	7,590	7,588	2,303	2,303	5,287	5,285
Clusters	93	93	91	20	20	73	71
R-squared	0.10	0.13	0.17	0.10	0.13	0.05	0.06
Controls F-test (p-value)		0.03	0.04	0.00	0.11	0.45	0.45

Notes: OLS estimates of the likelihood of appearing in the *Index of French Nationals Compelled into German Armed Forces* (MACVG, 1945, 1946). The unit of observation is a male individual. Standard errors clustered at the year-of-birth-level in parentheses. * significant at 10%; ** at 5%; *** at 1%.

Identifying assumptions – Variation in eligibility originates from fertility and the draft rule. It is therefore important to control for birth rates, which addresses the concern that fertility might simultaneously affect conscription rates and voting behavior, thus violating the exclusion restriction. It is also important to control for the share of the population that was deported (which was higher in Moselle). To further ensure cultural and geographical proximity, dialect fixed effects (that do not coincide with department borders) and border segment fixed effects are included.[18] The identifying assumption is thus that the administrators' conscription policies did not reflect particularities of the population which might also affect voting behavior and change discontinuously at the border. This assumption seems plausible, since Alsace and Moselle were integrated into neighboring, pre-existing German *Gaue*, whose administrators were already in place before the War.

The identifying assumption is tested using pre-War electoral outcomes from Lachapelle (1936), and measures of Nazi policies from INSEE (1956). This data is only available at the canton level. The results from regressing pre-War electoral outcomes and Nazi policies on an Alsace binary variable are presented in Table 2. Panel A presents unconditional differences; Panels B and C present differences conditional on geography and dialects. Columns (1) and (2) present the results using pre-War right-wing vote and abstention as outcomes, respectively. Neither support for right-wing parties, nor political participation were different before the War. Columns (3) to (5) present the results when using Nazi policies as the outcome. Unconditional estimates show differences in the policies chosen by the two administrators. Conditional estimates (that increase proximity and comparability) imply that the only policy that differed systematically is Wehrmacht conscription. Finally, Column (6) presents the municipality-level conscription proxy. Municipalities on the Alsace side of the border have a higher share of men conscripted.[19]

[18] Fertility decisions have been shown to be endogenous to cultural attitudes (Fernández et al., 2004; Alesina et al., 2011). Local dialects date back to the 5th century. Municipalities with the same dialect share a very long common history (Lévy, 1929).

[19] In Online Appendix Table C.1, I further evaluate whether municipalities on the two sides of the border are comparable with respect to their socio-professional composition. This analysis reveals no differences in occupation, population, and language, but does reveal differences in religion. The importance of this difference is thoroughly discussed in Online Appendix Section C.2.

Table 2
Estimation strategy: Pre-War voting and Nazi policies.

	(1)	(2)	(3)	(4)	(5)	(6)
	Pre-War voting (1936)		Nazi policies (INSEE, 1956)			Municipality-level
Dep. variable	Right-wing	Abstention	Deported	Incarcerated	Wehrmacht	Wehrmacht
			Panel A: Unconditional differences			
Alsace dummy	−1.543	−2.737	−20.169**	0.192***	3.548***	5.086***
	(5.461)	(3.310)	(7.638)	(0.040)	(0.588)	(0.502)
Mean dep. variable	61.73	18.71	10.58	0.36	5.66	7.05
Observations	92	92	92	92	92	1,579
Clusters	23	23	23	23	23	93
			Panel B: Conditional on geography and dialect			
Alsace dummy	−6.195	−0.651	−14.021*	0.075	2.509***	4.890***
	(6.099)	(3.401)	(7.705)	(0.072)	(0.484)	(0.486)
Mean dep. variable	61.73	18.71	10.58	0.36	5.66	7.05
Observations	92	92	92	92	92	1,579
Clusters	23	23	23	23	23	93
			Panel C: Conditional on geography and dialect (estimation sample)			
Alsace dummy	1.578	−4.594	−5.159	0.082	2.180***	3.678***
	(9.573)	(7.794)	(3.621)	(0.068)	(0.343)	(0.463)
Mean dep. variable	62.77	20.70	8.51	0.32	5.55	7.74
Observations	31	31	31	31	31	462
Clusters						32

Notes: Differences in pre-War voting, Nazi policies, and fertility between Alsace and Moselle. In Columns (1) to (5) the unit of observation is a 1936 canton; in Column (6) it is a municipality. Panel A: Unconditional differences; Panel B: Differences conditional on geography and dialects; Panel C: Differences conditional on geography and dialects (estimation sample). Column (1): Vote for right-wing parties in the 1936 parliament election. Column (2): Abstention in the 1936 parliament election. Column (3): Fraction of 1936 population that was deported. Column (4): that was incarcerated. Column (5): that was conscripted to the Wehrmacht (INSEE, 1956). Column (6): that was conscripted to the Wehrmacht (MACVG, 1945). Geography and dialects controls included in Panels B and C: access to waterways (binary), elevation (log mean, log std.dev.), distance to Germany (log km), 25 km border segment fixed-effects, historical dialect fixed effects, and a quadratic polynomial in latitude and longitude ($x+y+x^2+y^2+xy$). Standard errors clustered at the arrondissement level in parentheses in Columns (1) to (5) of Panels A and B; robust standard errors in Columns (1) to (5) of Panel C; standard errors clustered at the canton level in Column (6). * significant at 10%; ** at 5%; *** at 1%.

Table 3
Estimation strategy: First-stage estimates.

	(1)	(2)	(3)	(4)	(5)
Specification	No controls	Fertility controls	Lat-lon poly and dialect	Historical controls	Historical and contemporary
		First-stage estimates. Dep. variable: Conscription proxy (%)			
Eligible births (%)	0.372***	0.475***	0.357***	0.261***	0.240***
	(0.033)	(0.045)	(0.048)	(0.050)	(0.054)
Born 1903–1912 (%)		−0.019	−0.046	−0.011	−0.007
		(0.064)	(0.058)	(0.056)	(0.056)
Born 1913–1922 (%)		−0.299**	−0.198	−0.147	−0.126
		(0.146)	(0.169)	(0.165)	(0.160)
Born 1923–1932 (%)		−0.310***	−0.188	−0.114	−0.099
		(0.113)	(0.128)	(0.126)	(0.135)
First-stage F-statistic	129.04	110.64	55.53	27.37	19.95
Lat-lon polynomial			2nd	2nd	2nd
Border segment FE			Yes	Yes	Yes
Historical dialect FE			Yes	Yes	Yes
Historical controls				Yes	Yes
Contemporary controls (1965)					Yes
Mean dep. variable	7.74	7.74	7.74	7.74	7.74
Observations	462	462	462	462	462
Clusters	32	32	32	32	32

Notes: First-stage estimates of the effect of the draft rule on Wehrmacht conscription. The unit of observation is a municipality. Standard errors clustered at the canton level in parentheses. Columns (3) to (5) include a quadratic polynomial in latitude and longitude ($x + y + x^2 + y^2 + xy$). First-stage F-statistic is the Kleibergen–Paap rk Wald F-statistic. * significant at 10%; ** at 5%; *** at 1%.

First-stage results – I first estimate the first-stage relationship presented in Eq. (2). The outcome is the proxy of the share of a municipality's population that was conscripted. The exogenous instrument, eligible births, are male births during 1908–1927 in Alsace and during 1914–1927 in Moselle, over the 1936 population. The birth data was only collected for localities within 20 km from the Alsace–Moselle border, the first-stage is therefore estimated solely for these 462 municipalities. Since some pre-war characteristics to be included vary only across cantons, standard errors are clustered at this level to correct for potential serial correlation.

The estimation results are presented in Table 3, where control variables are included sequentially to assess the sensitivity of the first-stage. Column (1) presents the unconditional correlation between eligibility and conscription. An increase in the eligible population by 1 percentage point is associated to an increase in conscription by 0.37 pp. The relationship is highly statistically significant (p -value=11.2). In Column (2) I introduce a vector of variables to capture fertility differentials, namely the number of births during the 1903–1912, 1913–1922, and 1923–1932 periods over pre-war population. The effect of eligibility remains positive and precisely estimated. The coefficients on fertility have the expected signs and magnitudes. An increase in the 1903–1912 fertility increases conscription in Alsace (through the increase in 1908–1913 eligible cohorts) but not in Moselle, while an increase in the 1913–1922 and 1923–1932 births increases eligibility both in Alsace and Moselle. The total effect of an increase in the 1903–1912 births on conscription is therefore positive, while it is not statistically significant for the other two decades.[20]

Column (3) introduces a second degree polynomial in latitude and longitude ($x+y+x^2+y^2+xy$), 25 km border fixed effects, and fixed effects for the dialect that is historically spoken in the municipality. The introduction of the latitude–longitude polynomial makes this specification akin to a fuzzy regression discontinuity design where the eligibility probability jumps from low (Moselle) to high (Alsace) at the border. The coefficient of eligibility remains positive and highly statistically significant, with a t-stat of 7.45, or, a first-stage F-statistic of 55.

In the last two columns I introduce a vector of historical and contemporary characteristics. The historical covariates are intended to capture any observable differences at the Alsace–Moselle border, such as the difference in deportation policies between administrators. They consist of religion (% Protestant, Jewish presence), population in 1936 (log), French-sounding name births from 1891 to 1940 (%), dialect-speaking population in 1936 (%), foreign male population in 1936 (%), right-wing vote in 1936 (% registered), displaced population during WWII (%), access to waterways (binary), elevation (log mean, log std.dev.), and distance to Germany (log km).

In the last column I introduce contemporary determinants of voting. The inclusion of these controls should have no effect on conscription, but is potential important for the efficiency of the second-stage estimation.[21] The contemporary controls are the following: population (log total, % foreign), age (% of 6 groups), educational attainment (% high-school degree), males (%), employment (% blue collar workers, % unemployment), and income (log median, log std.dev.). Column (5) presents the results when introducing these controls for the earlier election in the data, that of 1965. As expected, the coefficient of eligibility remains unaltered, both in magnitude and in significance. Column (5) is the full first-stage estimation used in the subsequent sections. It will only change insofar as the contemporary controls (which are time-varying) change.

3.3. Abstention and the radical right-wing vote

Following the anecdotal and descriptive evidence presented in Fig. 2, I first focus on support for the radical right-wing. While reducing the radical right-wing to its protest dimension runs the risk of omitting other ideological aspects (Schwengler, 2003), such parties have nonetheless positioned themselves as the ultimate anti-system parties (Davies, 2002).

Reduced-form results – The reduced form relationship is an Intention-to-Treat (ITT) estimator obtained by substituting Eq. (2) into Eq. (1)

$$\pi_m = x'_m \pi_{20} + \pi_{21} Eligible_m + \pi_{22} Fertility_m + f(lat_m, lon_m) + \xi_{2m} \tag{4}$$

I estimate Eq. (4) for the 462 municipalities within 20 km of the Alsace–Moselle border for each election from 1965 to 2017 separately. π_m is the vote share of the radical right-wing candidates or abstention in municipality m.[22] Both outcomes are measured as fractions of registered voters. The vector x_m consists of the historical and contemporary controls presented in the first-stage estimation. All specification also include a quadratic polynomial in latitude–longitude, 25 km border fixed effects, controls for fertility, and historical dialect fixed effects.

[20] A joint significance test of the eligibility and birth decades coefficients yields the following estimates (p-values): 0.456 (0.00) for 1903–12, 0.176 (0.19) for 1913–22, and 0.165 (0.15) for 1923–32.

[21] The introduction of contemporary controls yields the threat of "bad controls", i.e. controls that are themselves affected by the treatment. To mitigate this threat, all post-war population controls have been calculated by excluding the 1900–1927 populations. A discussion on the potential effect of conscription via demography follows in the next sub-section. The sensitivity of the second-stage results to the inclusion of these controls is presented in Online Appendix Table C.5.

[22] Candidates considered are Tixier-Vignancour (1965), J.M. and M. Le Pen (1974, 1995 to 2017), and Mégret (2002). The sensitivity to this classification is tested in Online Appendix Table C.4.

Table 4
Abstention and the radical right-wing vote: Reduced-form estimates.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
		Earlier elections				Later elections		
Election year	1965	1969	1974	1995	2002	2007	2012	2017
	Panel A: Reduced-form estimates. Dep. variable: Radical right-wing (% of registered)							
Eligible births (%)	−0.016		0.003	0.268**	0.172**	0.012	0.053	0.021
	(0.028)		(0.004)	(0.100)	(0.084)	(0.050)	(0.077)	(0.086)
Mean dep. variable	1.21	0.00	0.31	22.18	22.25	14.89	23.24	27.27
	Panel B: Reduced-form estimates. Dep. variable: Abstention (% of registered)							
Eligible births (%)	0.248***	0.371***	0.201**	−0.036	−0.177**	0.101***	0.044	0.079
	(0.078)	(0.128)	(0.080)	(0.054)	(0.071)	(0.036)	(0.063)	(0.051)
Mean dep. variable	15.29	23.63	17.07	19.79	28.39	16.73	19.41	20.64
Lat-lon polynomial	2nd	2nd	2nd	2nd	2nd	2nd	2nd	2nd
Border segment FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Historical dialect FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Full set of controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	462	462	462	462	462	462	462	462
Clusters	32	32	32	32	32	32	32	32

Notes: Reduced-form estimates of the effect of conscription eligibility on support for radical right-wing candidates and on abstention. The unit of observation is a municipality. Standard errors clustered at the canton level in parentheses. Panel A: Reduced-form estimates with radical right-wing vote as the outcome; Panel B: Reduced-form estimates with abstention as the outcome. All specifications include a quadratic polynomial in latitude and longitude ($x + y + x^2 + y^2 + xy$). * significant at 10%; ** at 5%; *** at 1%.

The results are presented in Table 4, where each column represents a different election. For ease of interpretation, only the coefficient on eligibility is displayed. Recall that in 1969 there was no radical right-wing candidate. The results indicate that Wehrmacht draft eligibility is a powerful predictor of abstention in the 1965, 1969, and 1974 elections (Panel B). In later elections (after 1995) an increase in eligibility is associated to an increase in radical right-wing support (Panel A). The coefficients for 1995 and 2002 (Columns [4] and [5] of Panel A) are of similar magnitude as the early abstention coefficients. The effects in subsequent elections are positive, but fail to reach any conventional level of significance. The interplay between the effects is also noteworthy. When the coefficient on radical-right wing support is positive and significant, it is negative for abstention, while when the effect is not statistically significant, it is positive for abstention.

Second-stage results – The second-stage estimation results are presented in Table 5, where again each column represents a different election. Panel A presents the first-stage estimation, Panel B presents the 2SLS estimation with radical right vote share as the outcome, and in Panel C abstention is the outcome. In earlier elections (1965–1974), conscription has no effect on radical right-wing support, see Panel B. Conversely, abstention is higher in localities where more men were conscripted, as shown in Panel C. In later elections (1995–2017), conscription has a positive effect on the radical right vote. Note also that the second-stage results replicate the interplay pattern between abstention and radical right-wing vote presented in the reduced-form estimation.

The results imply that, an increase in conscription by 1 standard deviation (≈ 4.5 pp), increases abstention by 0.6 std.dev. in 1965, 0.75 std.dev. in 1969, and 0.6 std.dev. in 1974 (6 to 11 pp). An equivalent increase in conscription increases radical right-wing support by 1 and 0.6 std.dev. in 1995 and 2002 (8 and 6 pp), and between 0.05 and 0.17 std.dev. in 2007, 2012 and 2017 (0.4 to 1.6 pp, not significant). Extrapolated to municipalities further away from the border, this suggests a reduction of the radical right vote by 7.5 pp, 5.1 pp, 0.3 pp, 1.5 pp, and 0.6 pp, in the 1995–2017 elections, which is comparable to the difference between Alsace–Moselle and the rest of France.

Sensitivity analysis – In the Appendix, I perform the sensitivity analysis that is standard in regression discontinuity designs. Estimation results from higher order polynomials in latitude–longitude are presented in Table A.1. The sensitivity to the 20 km bandwidth is presented in Fig. A.1. I also present various sensitivity checks in the Online Appendix: testing for the radical candidate classification (Table C.4); the sensitivity to the inclusion of controls (Table C.5); excluding casualties (Table C.6); only instrumenting 1908–1913 conscription (Table C.7); polynomials in distance to the border (Table C.8). While in some cases the coefficients fluctuate, the results remain qualitatively the same.

Compositional effects – Before moving forward, a discussion on potential compositional effects of conscription is appropriate, since higher conscription is associated to higher casualties. The municipality-level results could thus be a combination of individual preference related effects and composition effects. To assess the importance and persistence of conscription-driven compositional changes, I estimate Eq. (3) using demographic variables and measures of human capital and income as outcomes. The results for 1965 are presented in Appendix Table A.2. As expected, municipalities where conscription was higher have a lower post-War population

Table 5
Abstention and the radical right-wing vote: Second-stage estimates.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
		Earlier elections				Later elections		
Election year	1965	1969	1974	1995	2002	2007	2012	2017
	Panel A: First-stage estimates. Dep. variable: Conscription proxy (%)							
Eligible births (%)	0.240***	0.251***	0.253***	0.247***	0.233***	0.247***	0.251***	0.247***
	(0.054)	(0.055)	(0.053)	(0.051)	(0.051)	(0.047)	(0.056)	(0.054)
First-stage F-statistic	19.95	21.10	22.52	23.41	20.75	27.87	20.41	20.69
Mean dep. variable	7.74	7.74	7.74	7.74	7.74	7.74	7.74	7.74
	Panel B: 2SLS estimates. Dep. variable: Radical right-wing vote (% of registered)							
Conscription proxy (%)	−0.068		0.012	1.085***	0.736**	0.048	0.210	0.085
	(0.125)		(0.017)	(0.378)	(0.322)	(0.199)	(0.287)	(0.349)
Mean dep. variable	1.21	0.00	0.31	22.18	22.25	14.89	23.24	27.27
	Panel C: 2SLS estimates. Dep. variable: Abstention (% of registered)							
Conscription proxy (%)	1.033**	1.472**	0.795**	−0.147	−0.761**	0.409**	0.177	0.320
	(0.412)	(0.624)	(0.330)	(0.214)	(0.296)	(0.190)	(0.258)	(0.223)
Mean dep. variable	15.29	23.63	17.07	19.79	28.39	16.73	19.41	20.64
Lat-lon polynomial	2nd	2nd	2nd	2nd	2nd	2nd	2nd	2nd
Border segment FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Historical dialect FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Full set of controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	462	462	462	462	462	462	462	462
Clusters	32	32	32	32	32	32	32	32

Notes: 2SLS estimates of the effect of conscription into the Wehrmacht on support for radical right-wing candidates and on abstention. The unit of observation is a municipality. Standard errors clustered at the canton level in parentheses. Panel A: First-stage estimates; Panels B and C: 2SLS estimates with radical right-wing vote and abstention as the outcome, respectively. Each column presents the estimation for a different election. All specifications include a quadratic polynomial in latitude and longitude ($x + y + x^2 + y^2 + xy$). First-stage F-statistic is the Kleibergen–Paap rk Wald F-statistic of the first-stage estimation. * significant at 10%; ** at 5%; *** at 1%.

and a lower post-War male population, see Panel A. Moreover, these differences persist. On the other hand, conscription has no effect on education, employment, and income, as shown in Panel B. This is potentially due to the age of the 1908–1913 cohorts at the time they were drafted.

To thoroughly assess the importance of these demographic changes on voting outcomes I use both individual survey data and municipality-level data from other regions in France where incorporation did not take place. The results imply that gender can account for at most a 0.4 pp increase in early abstention (out of a 6 to 11 pp difference), while age should reduce abstention. Compositional effects would lead to a reduction, if anything, in radical right-wing vote. The full results and procedure are presented in Online Appendix Section C.3.

3.4. Radical and moderate candidates

So far, I focused on support for the radical right under the assumption that its candidates are the only ones capturing political distrust. In this section, I relax this assumption and exploit a time-varying, continuous measure of radicalism for all candidates, regardless of party family.

Model specification – The specification closely follows the estimation strategy of the previous subsection. The main difference comes from the fact that now the conscription variable is interacted with the degree of radicalism of a candidate (measured as the share of sentences with a negative tone in the discourse). The instrumental variables specification thus becomes

$$\pi_{cm} = x'_m \cdot Radical_c \cdot \gamma + \rho_1 \cdot \widehat{Conscripted}_m + \rho_2 \cdot Radical_c \cdot \widehat{Conscripted}_m + \delta_1 \cdot Fertility_m + \delta_2 \cdot Radical_c \cdot Fertility_m + f_c(lat_m, lon_m) + \varepsilon_{cm} \quad (5)$$

where π_{cm} if the vote share of candidate c in municipality m, and $Radical_c$ is the degree of radicalization of candidate c. The specification is estimated by pooling all candidate vote shares by election. To preserve the identification strategy, fertility controls are introduced (both in levels and interacted with the degree of radicalism), control variables are interacted with the degree of radicalism, while the dialect fixed effects, border segment fixed effects, and latitude–longitude polynomial are candidate specific. Standard errors are clustered at the municipality level.

Table 6
Radical and moderate candidates: Reduced-form and second-stage estimates.

Election year	(1) 1965	(2) 1969	(3) 1974	(4) 1995	(5) 2002	(6) 2007	(7) 2012	(8) 2017
	Panel A: Reduced-form estimates. Dep. variable: Candidate vote (% of registered)							
Eligible births (%)	−0.259***	−0.449***	−0.404***	−0.042**	−0.009	−0.027**	−0.020	−0.034***
	(0.038)	(0.074)	(0.039)	(0.016)	(0.010)	(0.012)	(0.014)	(0.010)
Eligible births (%) × Radicalism [std]	0.099***	0.182***	0.162***	0.034**	0.011	0.011	0.012	0.017**
	(0.020)	(0.034)	(0.016)	(0.013)	(0.008)	(0.011)	(0.012)	(0.007)
	Panel B: 2SLS estimates. Dep. variable: Candidate vote (% of registered)							
Conscription proxy (%)	−0.726***	−1.258***	−1.133***	−0.117**	−0.024	−0.075**	−0.055	−0.095***
	(0.156)	(0.282)	(0.205)	(0.047)	(0.028)	(0.033)	(0.038)	(0.032)
Conscription (%) × Radicalism [std]	0.254***	0.489***	0.477***	0.112**	0.042	0.029	0.039	0.055**
	(0.073)	(0.120)	(0.092)	(0.046)	(0.028)	(0.036)	(0.042)	(0.024)
First-stage F-statistic	6.45	6.75	6.99	6.90	6.38	7.03	6.59	6.28
Underidentification F-statistic	12.54	13.13	12.96	13.36	12.31	13.34	12.70	11.98
Candidate-lat-lon polynomial	2nd	2nd	2nd	2nd	2nd	2nd	2nd	2nd
Candidate-border segment FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Candidate-dialect FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Full set of controls × Radicalism [std]	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Mean dependent variable	14.12	10.91	6.91	8.91	4.48	6.94	8.06	7.21
Observations	2,772	3,234	5,544	4,158	7,392	5,544	4,620	5,082
Clusters	462	462	462	462	462	462	462	462
Candidate radicalism distribution								
Median candidate (std.dev.)	1.05	0.87	1.52	0.19	0.51	0.23	0.71	0.80
Most radical candidate (std.dev.)	2.53	2.50	3.71	3.46	3.73	3.60	4.30	4.61

Notes: 2SLS estimates of the effect of conscription into the Wehrmacht on support for radical and moderate candidates. The unit of observation is a municipality × candidate. Standard errors clustered at the municipality level in parentheses. Panel A: Reduced-form estimates with respect to the degree of a candidate's radical discourse; Panel B: 2SLS estimates. Each column presents the estimation for a different election. The radicalism measure is constructed as the share of quasi-sentences with a negative tone from the Comparative Manifesto Project (Schmitt et al., 2018). The radicalism measure has been standardized and re-scaled to take a value of 0 for the most moderate candidate within an election. All specifications include a candidate-specific quadratic polynomial in latitude and longitude ($x + y + x^2 + y^2 + xy$). First-stage F-statistic is the Kleibergen–Paap rk Wald F-statistic of the first-stage estimation with eligible births as the exogenous instrument. * significant at 10%; ** at 5%; *** at 1%.

To facilitate interpretation and comparison across years, the radicalism measure has been standardized within election and re-scaled to take a value of 0 for the most moderate candidate. The coefficients of interest are the coefficient of conscription (ρ_1) and the interaction of conscription and a candidate's degree of radicalism (ρ_2). The coefficient ρ_1 is the effect of conscription on the vote share of the most moderate candidate; the coefficient ρ_2 captures the difference in the effect between the most moderate candidate in the election and a candidate that is 1 standard deviation more radical; the sum of the two coefficients thus captures the effect of conscription on the vote share of a candidate that is 1 standard deviation more radical than the most moderate candidate.

Results – Table 6 presents the results of the aforementioned estimation, where each column represents a different election. Panel B presents the 2SLS point estimates. Since the instrument in the first-stage estimation is weak, I also present reduced-form estimation results (Panel A). A consistent pattern emerges in that, in all elections, in localities where more men were conscripted, moderate candidates receive a lower share of the vote (as captured by the level effect of conscription). This negative effect is however mitigated by a candidate's degree of radicalism (the interaction-term effect). In other words, while the effect of conscription is negative for more moderate candidates, it can be positive for more radical ones.

There is also substantial heterogeneity across elections. In early elections there is a large premium to radicalism, but also a large penalty for moderate candidates. As a result, in the 1965 to 1974 elections, a candidate has to be roughly 3 std.dev. more radical than the most moderate one to benefit from conscription. No candidate is radical enough in the 1965 and 1969, while in 1974, only some minor candidates are beyond this threshold.[23] As such, until 1974 the effect of conscription on radical support remained latent. In later elections, the radicalism premium is smaller, as is the penalty for being moderate. However, a candidate only has to be 0.5 to 2.5 std.dev. more radical than the most moderate one to benefit from conscription, a threshold crossed by several candidates. This leads to a pattern of "activated history", where attitudes are initially not detectable, but translate into support with the rise of parties that can capture the resentment (Ochsner and Roesel, 2017; Cantoni et al., 2019).

[23] The threshold beyond which the conscription effect becomes positive is (year): 2.86 standard deviations (1965), 2.57 (1969), 2.37 (1974).

Other policy positions – Forced conscription into a foreign army may have led to a need to overcompensate in terms of national identity, thus increasing support for candidates with a strong nationalist discourse. Moreover, as a hierarchically organized authoritarian institution, the military has been shown to socialize its members into authoritarian modes of behavior (Jenning and Markus, 1977). The radical right has been characterized by a discourse that, apart its strong anti-establishment tone, is also nationalistic and authoritarian. Moreover, it is clearly located at the far right of the political spectrum. It could thus very well be that municipalities where more men where conscripted vote for such candidates because of these aspects of the program.

To evaluate whether this is the case, I repeat the analysis presented in Table 6 using other candidate positions. The results are presented in Table A.3. Panel A present the estimation results when focusing on the extend of nationalist discourse. In early elections, more nationalist candidates are penalized in municipalities where more men where conscripted. Panel B focuses on the extend to which a candidate's discourse is authoritarian; authoritarian candidates are penalized both in early and late elections. Panel C looks at whether this support is driven by the extend to which the discourse is right-wing. Again, evidence is against the hypothesis that support is driven by this aspect of candidates. Finally, Panel D looks at whether more extreme candidates (in the left–right dimension) are driving this support. While more extreme candidate receive higher shares of the vote in some elections, they are penalized in others. This indicates that these positions of the radical right-wing are unlikely to be driving support in municipalities with more conscripts.

Sensitivity analysis – In the Online Appendix, I perform two robustness exercises to assess the sensitivity of the results to the radicalism measure. The first exercise replicates the estimation strategy while using a direct measure of which candidates receive support from distrustful voters, already presented in Fig. 2. The results, presented in Table C.11, corroborate the findings. The second exercise consists of focusing on the aggregate vote share of most radical candidates and comparing it to the aggregate vote share of more moderate candidates. The results are presented in Table C.12. Conscription increases the radical-moderate difference, by decreasing aggregate vote share of more moderate candidates and increasing the aggregate vote share of most radical ones. The later exercise is also performed for the other candidate positions. The results, presented in Table C.13, confirm that aspects other than radicalism are unable to explain voting behavior.

4. From abstention to radical support

To explain the transition from abstention to radical support I build a theoretical framework that introduces conscription as a shock to political preferences, and yields testable predictions that are in line with the evidence in the previous sections. I then test the predictions of the model.

4.1. Theoretical framework

The framework is kept simple and builds on the expressive voting model in Glaeser et al. (2005). While some of the predictions are a direct consequence of the assumptions, the model is particularly useful in guiding the investigation of how political support is revealed when new actors enter the political scenery.

The candidates – Two candidates $C \in \{M, R\}$ run for office, where M stands for "Moderate" and R for "Radical". Policy platforms are uni-dimensional and refer to how critical candidates are of the political establishment, τ_C, where by definition $\tau_R \geq \tau_M$. Platforms are taken as exogenous since the analysis focuses on national elections in regions representing 5% of the electorate.

The voters – Each citizen i has a preferred policy τ^i. Military service during WWII acts as a constant (additive) shock to one's favorite policy

$$\tau^i = \bar{\tau} + \beta D^i + \varepsilon^i \tag{6}$$

where D^i is an indicator variable that takes the value 1 if individual i was conscripted (with probability α) and 0 otherwise, and ε^i is the idiosyncratic part of preferences that is assumed to follow a uniform distribution with a mean of 0 and a density of 1.[24]

The preferred policies τ^i are thus individual-specific, though they follow a uniform group-specific distribution τ^{iJ} on the support $\left[\bar{\tau}^J - \frac{1}{2}, \bar{\tau}^J + \frac{1}{2}\right]$, $J \in \{H, L\}$, where H stands for "High distrust" (conscripted) and L stands for "Low distrust" (not conscripted). The population shares of group H and L are α and $(1-\alpha)$ respectively. β is assumed positive, $\bar{\tau}^H > \bar{\tau}^L$, meaning group H is, on average, more distrustful of the political establishment relative to group L.

Individual i in group J gains utility from voting for candidate C which is equal to $W^{iJ}(\tau_C) = B - M(|\tau_C - \tau^{iJ}|)$, where B measures the psychological gain from expressing support for one's favorite policy τ^{iJ}. $M(\cdot)$ captures the fact that citizens derive less utility if they vote for a candidate whose policy proposal τ_C differs from their own bliss point τ^{iJ} and is assumed to be quadratic,

[24] The assumption that attitudes map one-to-one onto preferred policies can be relaxed by assuming that mobilization affects a characteristic θ^i and that τ^i is monotonically increasing in θ^i.

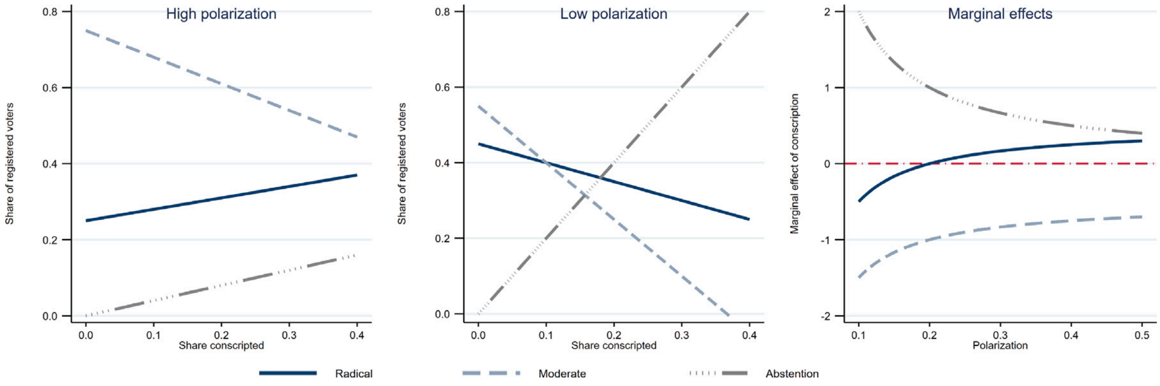

Fig. 3. Theoretical framework: Testable predictions.
Notes: Simulations of the testable predictions of the theoretical framework. Left-hand-side graph: vote shares in a highly polarized election; central graph: in an election with low polarization; right-hand-side graph: marginal effect of conscription with respect to electoral polarization. The parameters for the graphs are $\tau_M = 0$, $\tau_R = 0.1$ (low polarization), $\tau_R = 0.5$ (high polarization), $\bar{\tau}^H = 0.5$, $\bar{\tau}^L = 0$, $\bar{c}^H = 0.2$, $\bar{c}^L = 0$. Radical candidates in solid blue lines; moderate in light dashed lines, abstention in dashed gray lines.

$M(|\tau_C - \tau^{iJ}|) = \left(\tau_C - \tau^{iJ}\right)^2$. People also gain utility from voting against candidate C', which is equal to $-W^{iJ}(\tau_{C'})$. The benefit from voting is therefore

$$V\left(\tau^{iJ}, \tau_M, \tau_R\right) = \max \begin{cases} -\left(\tau_M - \tau^{iJ}\right)^2 + \left(\tau_R - \tau^{iJ}\right)^2, & \textit{Benefit from voting } M \\ -\left(\tau_R - \tau^{iJ}\right)^2 + \left(\tau_M - \tau^{iJ}\right)^2, & \textit{Benefit from voting } R \end{cases}$$

The act of voting involves a cost c^{iJ}. The cost c^{iJ} takes group-specific values $\bar{c}^J$, where $\bar{c}^H \geq \bar{c}^L$, i.e. the cost of voting is assumed higher for the group of individuals with low political trust.

Conditional on voting, people will support the candidate that is closer to their ideal platform and since the act of voting imposes a cost $\bar{c}^J$, people will vote as long as $V\left(\tau^{iJ}, \tau_M, \tau_R\right) \geq \bar{c}^J$. Voter heterogeneity is thus group-specific and expressed by the parameters $\bar{\tau}^J$ and $\bar{c}^J$.

Candidates' vote shares – It is straightforward to show that given the distributional assumptions on τ^{iJ} the overall vote share attained by candidate (π_C) and overall abstention (π_A) are

$$\pi_R = \underbrace{\frac{1}{2} - \frac{\tau_R + \tau_M}{2}}_{\beta_{0,R}} + \underbrace{\left(\bar{\tau}^H - \bar{\tau}^L\right)}_{\beta_{1,R} \geq 0} \alpha + \underbrace{\left(\frac{\bar{c}^L - \bar{c}^H}{2}\right)}_{\beta_{2,R} \leq 0} \frac{\alpha}{\Delta\tau_C} + \underbrace{\bar{\tau}^L - \frac{1}{2}\frac{\bar{c}^L}{\Delta\tau_C}}_{\varepsilon_R} \tag{7}$$

$$\pi_M = \underbrace{\frac{1}{2} + \frac{\tau_R + \tau_M}{2}}_{\beta_{0,M}} + \underbrace{\left(\bar{\tau}^L - \bar{\tau}^H\right)}_{\beta_{1,M} \leq 0} \alpha + \underbrace{\left(\frac{\bar{c}^L - \bar{c}^H}{2}\right)}_{\beta_{2,M} \leq 0} \frac{\alpha}{\Delta\tau_C} \underbrace{- \bar{\tau}^L - \frac{1}{2}\frac{\bar{c}^L}{\Delta\tau_C}}_{\varepsilon_M} \tag{8}$$

$$\pi_A = 1 - \sum_C \pi_C = \underbrace{\left(\bar{c}^H - \bar{c}^L\right)}_{\beta_{2,A} \geq 0} \frac{\alpha}{\Delta\tau_C} + \underbrace{\frac{\bar{c}^L}{\Delta\tau_C}}_{\varepsilon_A} \tag{9}$$

where $\Delta\tau_C \equiv \tau_R - \tau_M$ measures policy divergence. The coefficient signs are derived from the assumptions that $\bar{c}^H \geq \bar{c}^l$ and $\bar{\tau}^H \geq \bar{\tau}^L$.[25]

Testable predictions – This simple framework yields testable predictions both across candidates in the same election and within candidates across elections, presented in Fig. 3. Within an election (such that $\Delta\tau_C$ is constant), candidates with a more radical program (i.e. more critical of the establishment) would have, all else held equal, larger shares of the vote in municipalities were group H is larger ($\frac{\partial \pi_R}{\partial \alpha} \geq 0$) as long as platforms are not "too" similar (the effect becomes negative once $\Delta\tau_C \leq \frac{1}{2}\frac{\bar{c}^H - \bar{c}^L}{\bar{\tau}^H - \bar{\tau}^L}$). Moderate candidates should achieve lower shares of the vote ($\frac{\partial \pi_M}{\partial \alpha} \leq 0$) and abstention should be higher ($\frac{\partial \pi_A}{\partial \alpha} \geq 0$) in localities where group H is larger. Across elections, both radical and moderate candidates would have, all else held equal, lower shares of the vote in

[25] A voter i in group J whose ideal policy is τ^{iJ} votes for candidate R if $\tau^{iJ} \geq \bar{\tau}_C + \frac{\bar{c}^J}{2\Delta\tau_C}$ and for candidate M if $\tau^{iJ} \leq \bar{\tau}_C - \frac{\bar{c}^J}{2\Delta\tau_C}$; otherwise she abstains. The shares by group and candidate are $\pi_R^J = \frac{1}{2} - \bar{\tau}_C + \bar{\tau}^J - \frac{\bar{c}^J}{2\Delta\tau_C}$, and $\pi_M^J = \frac{1}{2} + \bar{\tau}_C - \bar{\tau}^J - \frac{\bar{c}^J}{2\Delta\tau_C}$. Group abstention is $\pi_A^J = \frac{\bar{c}^J}{\Delta\tau_C}$.

elections where policy platforms converge ($\frac{\partial \pi_C}{\partial \Delta \tau_C} \leq 0$), resulting in higher abstention ($\frac{\partial \pi_A}{\partial \Delta \tau_C} \geq 0$). This effect should be larger (in absolute terms) in places where group H is larger ($\frac{\partial^2 \pi_C}{\partial \Delta \tau_C \partial \alpha} \geq 0$).

4.2. Within election and municipality

To test these predictions, I combine municipality conscription, the 1965 to 2017 voting outcomes, and the measure of a candidate's degree of radicalism. Policy divergence is measured using the polarization index of Dalton (2008). The index weights the divergence between candidate positions and the election weighted average using the expected importance of each candidate as weights. It is election specific since it aggregates all candidate positions in a single election.[26]

Model specification – I use the radicalism measure, the divergence index, and the full spectrum of electoral results (i.e. all presidential candidates from 1965 to 2017) to estimate the impact political discourse has on the transition from abstention to radical support. Under the guidance of Eqs. (7) and (8), the instrumental variables regression equation becomes

$$\pi_{cmt} = x'_{mt}\gamma + \rho_1 \cdot Radical_{ct} \cdot \widehat{Conscripted}_m + \rho_2 \cdot \frac{\widehat{Conscripted}_m}{\Delta Radical_t} + \lambda_{ct} + \lambda_{mt} + \varepsilon_{cmt} \tag{10}$$

where π_{cmt} is the vote share of candidate c in municipality m in election year t and $\widehat{Conscripted}_m$ is the predicted conscription rate from the first-stage estimation(s) (which follow(s) in spirit Equation (2)). $Radical_{ct}$ is the measure of candidate c's degree of radicalism in election t; $\Delta Radical_t$ is the measure of policy divergence that only varies across elections.

Candidate × border segment fixed-effects (λ_{ct}), are introduced to capture candidate-specific characteristics and λ_{mt} are municipality × year (or municipality × party) fixed effects. The introduction of municipality × year fixed effects results in an estimation across candidates in the same election (within municipality); municipality × party fixed effects imply that candidates of the same party are compared (within municipality-party across elections). Eq. (10) is thus estimated both across political space, exploiting variation across candidates within an election, and across time, exploiting variation in the positions of a party across elections. To account for the fact that several things might be changing in time, some of the control variables (population, male fraction, and fraction of population displaced) are interacted with the discourse variables as well.

Results – The results from estimating Eq. (10) by 2SLS are presented in Table 7. Columns (1) and (2) present the results within a municipality × election (across candidates); Columns (3) to (6) present the results within a municipality × party (across time). The outcome in Columns (1) to (4) is a candidate's vote share; in Columns (5) and (6) it is abstention. Odd columns present the results when only looking at how radical a candidate is; pair columns also take into account other aspects that are central in the radical right-wing discourse (nationalism, authority).

In line with the predictions of the theoretical framework, conscription has a positive effect on the vote share of radical candidates, see Columns (1) to (4). This effect is mitigated by policy convergence, as shown in Columns (3) and (4). Moreover, as predicted, policy convergence leads to greater abstention in localities where more men where conscripted, see Columns (5) and (6).

The results indicate that conscription affects candidates asymmetrically in the presence of radical candidates. This is illustrated through the marginal effects of Column (4), that are presented in Fig. 4. In the left graph, the upper line presents the effect in a more polarized election ($\Delta Radical_t = 0.10 \approx \Delta Radical_{2002}$), while the lower one represents the effect in a less polarized one ($\Delta Radical_t = 0.02 \approx \Delta Radical_{1965}$). In highly polarized elections, conscription positively affects radical candidates and negatively affects moderate ones, however when polarization is low it affects both negatively. This graph highlights the main prediction of the model, namely that in elections with policy convergence *all parties are penalized*, which is expressed in higher abstention. In elections with policy divergence, the moderate candidates' loss of votes is partially captured by radical candidates. The right graph illustrates this asymmetry whereby the marginal effect of conscription on abstention by election is positive in all elections but much larger in the earlier ones which were less polarized.

Additional results – The reduced-form and first stage estimations are presented in Online Appendix Tables C.14 and C.15. Online Appendix Table C.16 also presents some evidence of intergenerational transmission by accounting for the population affected directly, and indirectly through parents and grandparents. The results indicate that, the effects are present for conscripts and their children, but no longer significant for grandchildren, implying that the effects dissipate in time.

5. Potential mechanisms

Following a large strand of literature, I turn to survey data to understand the underlying mechanism. Formally, the hypothesis to be tested in this section is $\beta \neq 0$ in Eq. (6).

[26] Dalton (2008)'s index is: $\Delta Radical_t = \left(\sum_C \left\{\hat{\pi}_{C,t} \times \left[Radical_{C,t} - \sum_C \hat{\pi}_{C,t} Radical_{C,t}\right]^2\right\}\right)^{\frac{1}{2}}$; it reveals little polarization in earlier elections compared to later ones (Online Appendix Figure B.5).

Table 7
From abstention to radical support: Second-stage estimates.

	(1)	(2)	(3)	(4)	(5)	(6)
Dep. variable	Candidate vote share				Abstention	
Specification	2SLS within election		2SLS across elections		2SLS across elections	
Conscription (%) × Radical	0.382***	0.399***	0.634***	1.293***		
	(0.112)	(0.114)	(0.234)	(0.369)		
Conscription (%) / ΔRadical			−0.016***	−0.014***	0.047***	0.050***
			(0.002)	(0.002)	(0.007)	(0.008)
Conscription (%) × Authoritarian		−0.945***		2.411***		
		(0.357)		(0.547)		
Conscription (%) / ΔAuthoritarian				−0.001		0.030***
				(0.003)		(0.010)
Conscription (%) × Nationalist		−0.544**		−1.441**		
		(0.220)		(0.650)		
Conscription (%) / ΔNationalist				−0.010**		0.036**
				(0.004)		(0.015)
Candidate-border segment FE	Yes	Yes	Yes	Yes	Yes	Yes
Election × Municipality FE	Yes	Yes				
Party × Municipality FE			Yes	Yes	Yes	Yes
Contemporary controls			Yes	Yes	Yes	Yes
Mean dependent variable	7.70	7.70	9.29	9.29	20.12	20.12
First-stage *F*-statistic	283.11	94.24	59.68	20.69	64.91	21.59
Underidentification *F*-statistic		219.18	116.92	121.65		49.87
Observations	38,346	38,346	30,030	30,030	3,696	3,696
Clusters	3,696	3,696	3,696	3,696	462	462

Notes: 2SLS estimates of the effect of conscription into the Wehrmacht on support for radical candidates and abstention across space and time (Eq. (10)). The unit of observation is a municipality × election × candidate. Standard errors clustered at the municipality × election in parentheses. Columns (1) and (2): results within a municipality × election (across candidates); Columns (3) to (6): results within a municipality × party (across time). Columns (1) and (2): 2SLS estimates for candidate vote shares within municipality and election (across candidates); Columns (3) and (4): 2SLS estimates for candidate vote shares within municipality and party (across time); Columns (5) and (6): 2SLS estimates for abstention and invalid ballots within municipality and party (across time). First-stage F-statistic is the Kleibergen–Paap rk Wald F-statistic of the first-stage estimation with Eligible births (%) × Radical and Eligible births (%)/ΔRadical as the exogenous instruments. * significant at 10%; ** at 5%; *** at 1%.

5.1. Data and estimation strategy

Survey data[27] – The *Interregional Survey of Political Phenomena* (henceforth ISPP) is an annual survey that took place from 1985 to 2004 in France. In total, roughly 250,000 individuals participated, of which almost 20,000 from the annexed lands. The survey contains information on the year-of-birth, gender, and department of residence of respondents, as well as several questions on political preferences and other individual characteristics.

The questions of interest are related to party preferences and trust in institutions. The question on party preferences is present in every wave and is formulated as follows: *"Here is a list of parties or political movements. Could you indicate which one you feel closer to, or less distant from?"*. The question on trust in institutions was asked in the 1987 and 1989 waves and is formulated as follows: *"Would you say you rather trust or not in [institution]?"*. The outcomes variables in the analysis, namely "proximity to [party]" and "trust in [institution]", are indicators that take the value 1 if the respondent answered that she feels closer to a specific party or that she rather trusts an institution.

Identification strategy – The identification strategy follows a large body of research that exploits draft rules to estimate causal impacts of conscription, as first proposed in Angrist (1990). To estimate the effect of Wehrmacht conscription on political preferences, I postulate the following difference-in-differences estimator for the sample of men born between 1908 and 1927

$$\tau_i = yob_i' + \rho_1 Alsace_i + \rho_2 \left[Alsace_i \times Born(1908-1913)_i\right] + x_i'\gamma + \eta_i \tag{11}$$

where τ_i is the outcome (party proximity or trust in institutions). $Alsace_i$ is an indicator that takes the value 1 if individual i was born in Alsace and $Born(1908-1913)_i$ is an indicator equal to 1 if i was born between 1908 and 1913. To focus on variation arising from the draft rule, all specification include year-of-birth fixed effects (yob_i), whose inclusion also captures birth period effects.

The coefficient of interest is ρ_2, which is a standard difference-in-differences coefficient. It captures the difference in political preferences between individuals born in 1908–1913 in Alsace and Moselle, while accounting for differences in preferences between Alsace and Moselle (estimated using the 1914–1927 cohorts), and differences in preferences between cohorts (captured by the year-of-birth fixed effects). The coefficient ρ_1 is of interest as well, since it captures Alsace–Moselle differences in political preferences for individuals born during the 1914–1927 period, therefore serving as an additional test for the validity of the setup.

[27] More details can be found in the Online Appendix Section B.8 .

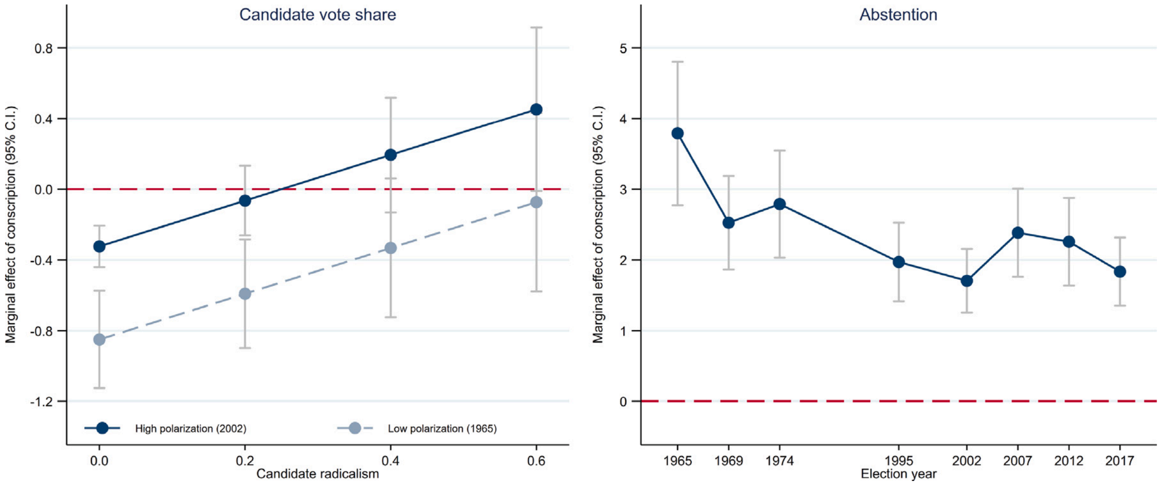

Fig. 4. From abstention to radical support: Marginal effects of conscription.
Notes: Marginal effects in Table 7. Left-hand-side graph, upper line: marginal effects at means of Column (4) for a highly polarized election ($\Delta Radical_t = 0.10 \approx \Delta Radical_{2002}$) by candidate degree of radicalism; left-hand-side graph, lower line: effects for an election with low polarization ($\Delta Radical_t = 0.02 \approx \Delta Radical_{1965}$). Right-hand-side graph: marginal effects at means of Column (6) by year.

Identifying assumptions – The central identifying assumption is a standard common trends assumption that can be summarized as follows: In the absence of Wehrmacht conscription, within year-of-birth Alsace–Moselle differences in political preferences for the 1908–1913 cohorts should not differ from differences for the 1914–1927 cohorts. The common trends assumption is tested by performing falsification exercises, namely using women in Alsace and Moselle, and men from regions where conscription did not take place. Since Eq. (11) is a reduced-form estimation (conscription is not observable), two additional assumption are necessary: (i) eligibility status should predict conscription, and (ii) eligibility should only affect political preferences via conscription when accounting for the year-of-birth and region. The former assumption was already tested in Table 1. Several tests are performed to assess the validity of the latter.

5.2. Party proximity and trust in institutions

Model specification – I start by estimating a Linear Probability Model of Eq. (11) using male respondents born during the 1908–1927 period in Alsace and Moselle. τ_i is either an indicator that takes the value 1 if i stated he feels closer to a specific party (party proximity), or an indicator variable that takes the value 1 if individual i answered "[he] rather trust[s]" a specific institution (trust in institutions). x_i is a vector of individual characteristics that includes year-of-birth fixed effects, religious affiliation, the origin of the father, how long the family has lived in the region, and survey year fixed effects. Standard errors are clustered at the year-of-birth × region level.

Results – I first investigate whether individuals born in 1908–1913 in Alsace are more likely to declare proximity to the radical right-wing. The results are presented in Panel A of Table 8. Consistently with the aggregate evidence from Section 3, the only parties for which men born in 1908–1913 in Alsace show more sympathy towards are those classified as belonging to the radical right. Men born in 1908–1913 in Alsace are 3.4 percentage points more likely to state they feel close to the radical right than their Moselle counterparts. Men born in 1914–1927 in Alsace are not more likely to state closeness to the radical right than men born during the same period in Moselle. Note moreover that, men born in Alsace in 1908–1913 are not more likely to state "None" as their preferred party. This is in line with the aggregate evidence indicating that by the 1990s the effects of conscription on voting behavior were already captured by the radical right.

Panel B focuses on trust in institutions. In Columns (1) to (5) I use measures of trust towards non-political institutions as the outcome. Men born in Alsace during the 1908–1913 period are not more likely to distrust these institutions. Columns (6) and (7) of Panel B focus on political institutions, namely political parties and elected politicians. While men born during the relevant cohorts in Alsace are not less trustful of political parties, they trust elected politicians significantly less.[28] The difference is sizable; the treated

[28] Constructing an "average political trust" variable indicates that men born in 1908–1913 in Alsace indeed show overall reduced political trust ($\hat{\rho}_2 = -0.689$, p-value=0.01).

Table 8
Party proximity and trust in institutions: Baseline results.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Panel A: Party proximity. *Which party do you feel closer to, or less distant from*						
Dep. variable	*Communist*	*Socialist*	*Ecologist*	*Liberal*	*Conservative*	*Radical right*	*None*
Alsace	−0.010	−0.024	0.005	0.039*	−0.066**	−0.003	0.057***
	(0.009)	(0.019)	(0.015)	(0.021)	(0.026)	(0.009)	(0.019)
Born 1908–1913 × Alsace	−0.020	−0.026	0.024	0.008	0.058	0.034**	−0.075
	(0.025)	(0.114)	(0.025)	(0.080)	(0.052)	(0.014)	(0.077)
Mean dep. variable	0.019	0.196	0.044	0.197	0.213	0.049	0.277
Observations	1,092	1,092	1,092	1,092	1,092	1,092	1,092
Clusters	40	40	40	40	40	40	40
	Panel B: Trust in institutions. *Would you say you rather trust or distrust in*						
Dep. variable	*Schooling system*	*Judicial system*	*Police*	*Church*	*Army*	*Political parties*	*Elected politicians*
Alsace	−0.059**	0.098	0.000	−0.048	−0.042	−0.082*	0.097
	(0.029)	(0.067)	(0.044)	(0.040)	(0.035)	(0.047)	(0.061)
Born 1908–1913 × Alsace	−0.004	0.069	−0.067	0.131	−0.205	−0.100	−0.549***
	(0.101)	(0.275)	(0.112)	(0.070)	(0.140)	(0.210)	(0.082)
Mean dep. variable	0.891	0.644	0.831	0.802	0.860	0.226	0.605
Observations	174	174	177	172	179	168	167
Clusters	36	35	36	36	36	34	34
Individual controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Survey year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year-of-birth FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Reduced-form estimates of the effect of Wehmacht eligibility on political preferences. Data comes from the Interregional survey of political phenomena (*Enquête interrégionale des phénomenes politiques*). The unit of observation is an individual. Standard errors clustered at year-of-birth × region level in parentheses. Panel A: Results with party proximity as the outcome; Panel B: Results with measures of trust in institutions are the outcome. Each column presents the estimation for a different party/institution. Year-of-birth × gender fixed effects are included in all specifications. * significant at 10%; ** at 5%; *** at 1%.

group is 55 percentage points less likely to trust elected politicians than its control counterpart, a magnitude close to the sample average. Once again, men born in 1914–1927 in Alsace are not less trustful of elected politicians than in Moselle.

Robustness – The first robustness exercise consists of estimating Eq. (11) using the sample of women born from 1908 to 1927. Since women in Alsace and Moselle were not conscripted, being born in 1908–1913 in Alsace should not make any difference in political preferences. The results, presented in Panels A and B of Table A.4, show that women born during the 1908–1913 period do not differ neither with respect to political trust, nor in their sympathy of the radical right.

The second falsification exercise focuses on departments in France in which Wehrmacht conscription did not take place. To perform this exercise I re-estimate Eq. (11) for any potential pair of departments, where treatment is allocated randomly within each pair. The Alsace–Moselle coefficient can then be compared to the distribution of the placebo coefficients. The results are presented in Panels C and D of Table A.4. While political trust in Alsace–Moselle is significantly different from the mean of the placebo distribution, proximity to the radical right is not (p-value=0.24).

5.3. Channels of persistence

The findings presented in Table 8 imply that individuals eligible for Wehrmacht conscription display lower political trust and are more likely to sympathize with the radical right. While I argue that the link between the two is direct, i.e. conscription reduced trust and increased support for the radical right, several channels could support these findings.

Transmission of attitudes – Previous research has demonstrated how beliefs can be transmitted from parents to children and persist in environments that are different than the ones they were developed in Guiso et al. (2006). To test whether this intergenerational transmission mechanism holds, I construct the likelihood that an individual's father or grandfather was born during the 1908–1913 and 1914–1927 periods, using the 1962 to 2011 censuses organized by INSEE (and available on IPUMS-I).[29] I then estimate Eq. (11) for the full sample of respondents. A separate coefficient is estimated for individuals affected directly, via their father, or via their grandfather(s).

The results are presented in Column (1) of Table 9. Two findings stand out: Firstly, both individuals whose father was eligible and those whose grandfather(s) was eligible display reduced political trust. Secondly, transmission from the first to the second generation, as captured by the ratio of the two coefficients, is very strong: The effect for the second generation (−0.6) is statistically

[29] More details on these variables can be found in Online Appendix Section B.7.

Table 9
Party proximity and trust in institutions: Channels of persistence.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Transmission	Service & casualties		Human capital & income			Organizational skills	
Dep. variable	Trust politicians	Trust politicians	Trust politicians	Higher education	Log income	Owns business	Party member	Member & Rad right
Alsace	0.064*	0.097	0.038	0.065**	0.032	−0.004	−0.032**	0.001
	(0.035)	(0.062)	(0.079)	(0.025)	(0.036)	(0.025)	(0.014)	(0.002)
Born 1908–1913 × Alsace	−0.536***	−0.538***		−0.014	−0.075	0.063	0.062*	−0.004
	(0.053)	(0.079)		(0.046)	(0.101)	(0.049)	(0.031)	(0.003)
B. 08-13 (Woman) × Als.	0.120							
	(0.187)							
B. 08-13 (Father) × Als.	−0.600***							
	(0.145)							
B. 08-13 (G.-father) × Als.	−0.218							
	(0.157)							
Born 1911–1913 × Alsace		−0.024						
		(0.162)						
B 08-13 (Husband) × Als.			0.012					
			(0.306)					
Individual controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Survey year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year-of-birth FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Mean dep. variable	0.535	0.605	0.669	0.110	8.847	0.145	0.019	0.002
Observations	1,935	167	166	1,086	816	981	525	525
Clusters	133	34	35	40	40	40	37	37

Notes: Reduced-form estimates of the potential mechanism. Data comes from the Interregional survey of political phenomena (*Enquête interrégionale des phénomenes politiques*). The unit of observation is an individual. Standard errors clustered at year-of-birth × region level in parentheses. Column (1): Intergenerational transmission of attitudes; Columns (2) and (3): Length of service and casualties; Columns (4) to (6): Human capital and earnings; Columns (7) and (8): Organizational skills. Year-of-birth × gender fixed effects are included in all specifications. * significant at 10%; ** at 5%; *** at 1%.

indistinguishable of the effect for the first (−0.536), while the effect for the third generation (−0.218) represents 35% of the effect for the second one (p-value=0.17, not statistically significant). These results point towards slow-moving attitudes that are inherited from past generations.[30]

Other potential mechanisms – The results presented could also be rationalized by several other mechanisms. If casualties are an increasing function of political trust, then the findings could reflect the pool of survey respondents. To address this hypothesis I exploit the fact that individuals born from 1911 to 1913 were drafted three months earlier than men born from 1908 to 1910. Since casualties are an increasing function of the length of service, if this hypothesis holds, the effect for the 1911–1913 cohorts should be larger than that of the 1908–1910 cohorts.[31] This does not appear to be the case, as shown in Column (2) of Table 9.

A different hypotheses is that it might not be disappointed men who become alienated per se, but their widows and orphans, whose father was taken from them by an alien power in a senseless last-minute bid. I thus construct the likelihood that a woman was married to a man born during the 1908–1913 period and estimate Eq. (11) using the sample of women born from 1908 to 1927. The results are presented in Column (3) of Table 9. Women married to men born during the 1908–1913 period in Alsace do not display reduced trust compared to their Moselle counterpart.

Conscription might also affect human capital and earnings potential of individuals (Blattman and Annan, 2010). While the plausibility of this channel is limited by the age of the affected 1908–1913 cohorts at conscription (from 31 to 36 years old), I nonetheless estimate Eq. (11) using: (i) an indicator for a university degree, (ii) self-declared income, and (iii) an indicator for business ownership. Conscription eligibility is not correlated with these measures of human capital and earnings, as shown in Columns (4) to (6) of Table 9.

Previous research has indicated that recruitment into the military is associated with organizational experience (Jha and Wilkinson, 2012). To test this mechanism I exploit information on party membership. The results, presented in Column (7) of Table 9, indicate that individuals born in 1908–1913 in Alsace are more likely to belong to a political party than in Moselle. Political support could thus be driven by local campaigning capacities if the radical right is more organized. To test this mechanism, I create an indicator that takes the value 1 if an individual is a party member and close to the radical right. The results, presented in Column (8) of Table 9, imply that this channel is unlikely to drive observed political attitudes.

30 Results that replicate this estimation for party proximity are presented in Table C.17 of the Online Appendix. The results indicate that, respondents affected directly are more likely to state proximity to the National Front, while their descendants to smaller radical right-wing parties.

31 The casualty rate for the 1911–1913 cohorts was of 20% and of 16% for the 1908–1910 ones.

Table 10
Party proximity and trust in institutions: Oblique transmission.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Trust in institutions					Radical right vote	
Dep. variable	Schooling system	Judicial system	Police	Army	Political parties	Ever voted for	Voted in 2002 R1
Alsace	0.117	−0.260	0.393	0.119	0.875***	−0.016	−0.082
	(0.321)	(0.314)	(0.316)	(0.264)	(0.286)	(0.097)	(0.140)
Conscription proxy (%)	−0.066**	−0.041	0.012	0.015	−0.036	0.033**	0.047***
	(0.031)	(0.036)	(0.032)	(0.030)	(0.032)	(0.014)	(0.014)
Born 1908-13 (Father) × Alsace	−0.457	0.432	−0.408	−1.290	−2.644***	0.098	0.287
	(1.165)	(1.478)	(1.121)	(0.961)	(0.849)	(0.422)	(0.548)
Born 1908-13 (G.-father) × Alsace	0.687	1.280	−0.845	−0.243	−2.178**	0.143	0.294
	(1.022)	(0.933)	(0.841)	(0.608)	(0.853)	(0.289)	(0.384)
Individual controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Survey wave FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year-of-birth FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Mean dep. variable	0.761	0.521	0.743	0.853	0.217	0.125	0.136
Observations	163	165	167	163	166	351	302
Clusters	68	69	69	67	69	99	92

Notes: Reduced-form estimates of the potential mechanism. Data comes from the 2002 and 2007 French Electoral Panel (*Panel électoral français*). The unit of observation is an individual. Standard errors clustered at year-of-birth × region level in parentheses. Columns (1) to (5): Results with measures of trust in institutions are the outcome; Columns (6) to (7): Results with past vote for the radical right as the outcome. Year-of-birth × gender fixed effects are included in all specifications. * significant at 10%; ** at 5%; *** at 1%.

Finally, different assimilation policies were implemented in Alsace and Moselle. Since the policies implemented in Alsace were potentially more assimilative than in Moselle (that mainly worked through deportations), this could affect political attitudes, as in Dehdari and Gehring (2019). As far as women experienced the same policies as men, Table A.4, that uses the female sample is informative on the potential effect of assimilation policies. The absence of any difference for females indicates that other policies are unlikely to be driving the differences in the male sample.

Oblique transmission – A last potential channel of persistence is transmission through individuals in the community, other than family members, also known as horizontal and oblique transmission (Cavalli-Sforza and Feldman, 1981). To identify such a channel, one should know the extend of conscription in the municipality respondents live. The *French Electoral Panel* (henceforth FES) of 2002 and 2007 contains information on the municipality of residence of respondents, as well as questions on trust in institutions and voting behavior.[32]

I use the FES information and re-estimate the regression of Column (1) of Table 9, while also including the conscription proxy for an individual's municipality of residence. The results are presented in Table 10.[33] Since municipality conscription is endogenous, one should be careful when causally interpreting the effects. The estimates imply that the environment in which one lives plays a role in the persistence of attitudes and voting behavior. The coefficient on the conscription proxy is significant for the two measures of past vote for the radical right-wing (Columns [6] and [7]), while it is negative, as expected, albeit not statistically significant, for trust in political parties, see Column (5).

6. Conclusions

This paper sheds light on the historical roots of political distrust and support for radical candidates. Identification exploits the fact that while conscription into the Wehrmacht took place in both Alsace and Moselle during their annexation to the Third Reich, different cohorts were drafted. I first show that WWII conscription results in increased abstention when policy platforms are similar, but increased support for radical candidates when there is polarization. I then provide survey evidence that conscription results in reduced political trust and that this attitude is transmitted from one generation to the next.

France's eastern borderlands have proven particularly fertile ground for radical candidates. These early forerunners have contributed to making such candidates relevant alternatives. By illustrating the political gains of campaigning on anti-establishment platforms, they also encouraged the formation of radical parties in other countries and altered the discourse of traditional parties. As such, this historical experience may have affected politics not only in France, but also in other European countries.

[32] More details can be found in Online Appendix Section B.9.

[33] Due to data scarcity it impossible to estimate the effect for individuals directly affected.

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Appendix A

A.1. Additional tables

See Tables A.1–A.4.

A.2. Additional figures

See Fig. A.1.

Table A.1
Abstention and the radical right-wing vote: Sensitivity to lat-lon polynomial.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Earlier elections			Later elections				
Election year	1965	1969	1974	1995	2002	2007	2012	2017
	Panel A: 2SLS estimates. Dep. variable: Radical right-wing vote (% of registered)							
Conscription proxy (%)	−0.054		0.010	0.952**	0.601**	−0.112	−0.001	−0.014
	(0.119)		(0.018)	(0.376)	(0.285)	(0.183)	(0.298)	(0.312)
Lat-lon polynomial degree	1st		1st	1st	1st	1st	1st	1st
First-stage F-statistic	28.00		31.88	32.79	29.83	47.72	38.11	35.61
Conscription proxy (%)	−0.068		0.012	1.085***	0.726**	0.045	0.203	0.081
	(0.125)		(0.016)	(0.380)	(0.318)	(0.191)	(0.316)	(0.390)
Lat-lon polynomial degree	3rd		3rd	3rd	3rd	3rd	3rd	3rd
First-stage F-statistic	29.16		28.09	34.19	27.44	39.77	22.94	24.56
	Panel B: 2SLS estimates. Dep. variable: Abstention (% of registered)							
Conscription proxy (%)	1.117***	1.488**	0.658**	−0.184	−0.855***	0.313**	0.073	0.252
	(0.385)	(0.564)	(0.310)	(0.200)	(0.294)	(0.145)	(0.228)	(0.206)
Lat-lon polynomial degree	1st	1st	1st	1st	1st	1st	1st	1st
First-stage F-statistic	28.00	31.17	31.88	32.79	29.83	47.72	38.11	35.61
Conscription proxy (%)	1.020***	1.466**	0.796**	−0.147	−0.779**	0.410**	0.169	0.346
	(0.356)	(0.581)	(0.330)	(0.226)	(0.298)	(0.185)	(0.285)	(0.245)
Lat-lon polynomial degree	3rd	3rd	3rd	3rd	3rd	3rd	3rd	3rd
First-stage F-statistic	29.16	31.04	28.09	34.19	27.44	39.77	22.94	24.56
Border segment FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Historical dialect FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Full set of controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	462	462	462	462	462	462	462	462
Clusters	32	32	32	32	32	32	32	32

Notes: Sensitivity of the results of Table 5 to different latitude–longitude polynomials. The unit of observation is a municipality. Standard errors clustered at the canton level in parentheses. Panel A: 2SLS estimates with radical right-wing vote as the outcome and a first order $(x + y)$ and third order $(x + y + x^2 + y^2 + xy + x^3 + y^3 + x^2y + xy^2)$ polynomial in latitude–longitude; Panel B: 2SLS estimates with abstention as the outcome. Each column presents the estimation for a different election. First-stage F-statistic is the Kleibergen–Paap rk Wald F-statistic of the first-stage estimation. * significant at 10%; ** at 5%; *** at 1%.

Table A.2
Abstention and the radical right-wing vote: Compositional effects.

	(1)	(2)	(3)	(4)	(5)	(6)
	Panel A: 2SLS estimates. Demographic structure (1965)					
Dep. variable	Population			Age structure		
	Log	Log Male	Male (%)	Average	Male	Female
Conscription proxy (%)	−0.038**	−0.041*	−0.263	0.417	0.680**	0.125
	(0.016)	(0.022)	(0.430)	(0.258)	(0.286)	(0.335)
Mean dep. variable	6.13	5.40	49.37	34.51	33.36	35.77
First-stage F-statistic	21.75	21.75	21.75	21.75	21.75	21.75
	Panel B: 2SLS estimates. Human capital (1965)					
Dep. variable	Education			Employment & income		
	Years	High school	University	Blue collar	Unemployment	Log income
Conscription proxy (%)	0.013	−0.018	−0.015	−0.080	0.125	−0.006
	(0.012)	(0.148)	(0.068)	(1.588)	(0.108)	(0.007)
Mean dep. variable	8.79	4.07	0.74	66.20	0.76	10.15
First-stage F-statistic	25.86	25.86	25.86	25.86	25.86	25.86
Lat-lon polynomial	2nd	2nd	2nd	2nd	2nd	2nd
Border segment FE	Yes	Yes	Yes	Yes	Yes	Yes
Historical dialect FE	Yes	Yes	Yes	Yes	Yes	Yes
Full set of controls	Yes	Yes	Yes	Yes	Yes	Yes
Observations	462	462	462	462	462	462
Clusters	32	32	32	32	32	32

Notes: 2SLS estimates of the effect of conscription into the Wehrmacht on population, age, education, and employment outcomes in 1965. The unit of observation is a municipality. Standard errors clustered at the canton level in parentheses. Panel A: 2SLS estimates for population and age; the contemporary controls vector does not include any demographic characteristics. Panel B: 2SLS estimates for education and employment; the contemporary controls vector does not include any socioeconomic characteristics. Each column presents the estimation for a different outcome. All specifications include a quadratic polynomial in latitude and longitude ($x + y + x^2 + y^2 + xy$). First-stage F-statistic is the Kleibergen–Paap rk Wald F-statistic of the first-stage estimation. * significant at 10%; ** at 5%; *** at 1%.

Table A.3
Radical and moderate candidates: Other policy positions.

Election year	(1) 1965	(2) 1969	(3) 1974	(4) 1995	(5) 2002	(6) 2007	(7) 2012	(8) 2017
	Panel A: 2SLS estimates. Nationalist discourse							
Conscription proxy (%)	0.431***	0.153***	0.144***	−0.215***	−0.107***	−0.061	−0.083**	−0.111*
	(0.096)	(0.051)	(0.043)	(0.057)	(0.036)	(0.041)	(0.037)	(0.060)
Conscription (%) × Nationalism [std]	−0.641***	−0.320**	−0.136***	0.384***	0.140***	0.010	0.178*	0.115
	(0.179)	(0.135)	(0.036)	(0.108)	(0.044)	(0.067)	(0.092)	(0.085)
First-stage F-statistic	6.45	6.75	6.99	6.90	6.38	7.03	6.59	6.28
Underidentification F-statistic	12.54	13.13	12.96	13.36	12.31	13.34	12.70	11.98
	Panel B: 2SLS estimates. Authoritarian discourse							
Conscription proxy (%)	0.529***	0.292***	−0.036	−0.169***	−0.011	0.040	−0.096*	−0.018
	(0.136)	(0.100)	(0.039)	(0.052)	(0.026)	(0.036)	(0.056)	(0.030)
Conscription (%) × Authoritarian [std]	−0.969***	−0.504**	−0.049	0.164***	0.056	−0.051	0.170**	0.014
	(0.289)	(0.202)	(0.037)	(0.051)	(0.035)	(0.044)	(0.075)	(0.051)
First-stage F-statistic	6.45	6.75	6.99	6.90	6.38	7.03	6.59	6.28
Underidentification F-statistic	12.54	13.13	12.96	13.36	12.31	13.34	12.70	11.98
	Panel C: 2SLS estimates. Right-wing discourse (Budge and Laver, 2016, scale)							
Conscription proxy (%)	0.328***	0.089**	0.185***	−0.203***	−0.088**	−0.106*	−0.170**	−0.019
	(0.090)	(0.045)	(0.050)	(0.063)	(0.036)	(0.061)	(0.081)	(0.043)
Conscription (%) × Right-wing [std]	−0.406***	−0.352**	−0.278***	0.202***	0.080***	0.049	0.135**	0.028
	(0.117)	(0.172)	(0.079)	(0.061)	(0.028)	(0.035)	(0.056)	(0.041)
First-stage F-statistic	6.45	6.75	6.99	6.90	6.38	7.03	6.59	6.28
Underidentification F-statistic	12.54	13.13	12.96	13.36	12.31	13.34	12.70	11.98
	Panel D: 2SLS estimates. Extremist discourse (absolute Budge and Laver, 2016, scale)							
Conscription proxy (%)	−0.711***	−1.299***	−1.223***	−0.250***	0.158***	−0.058	0.027	−0.182***
	(0.149)	(0.283)	(0.236)	(0.087)	(0.048)	(0.039)	(0.066)	(0.064)
Conscription (%) × Extremism [std]	0.306***	0.528***	0.789***	0.159***	−0.040***	0.014	−0.068*	0.065**
	(0.081)	(0.124)	(0.164)	(0.057)	(0.013)	(0.018)	(0.036)	(0.025)
First-stage F-statistic	6.45	6.75	6.99	6.90	6.38	7.03	6.59	6.28
Underidentification F-statistic	12.54	13.13	12.96	13.36	12.31	13.34	12.70	11.98
Candidate-lat-lon polynomial	2nd	2nd	2nd	2nd	2nd	2nd	2nd	2nd
Candidate-border segment FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Candidate-dialect FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Full set of controls × Position [std]	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Mean dependent variable	14.12	10.91	6.91	8.91	4.48	6.94	8.06	7.21
Observations	2,772	3,234	5,544	4,158	7,392	5,544	4,620	5,082
Clusters	462	462	462	462	462	462	462	462

Notes: 2SLS estimates of the effect of conscription into the Wehrmacht on support for radical and moderate candidates. The unit of observation is a municipality × candidate. Standard errors clustered at the municipality level in parentheses. Panel A: 2SLS estimates with respect to the degree of nationalist discourse; Panel B: 2SLS estimates with respect to the degree of authoritarian discourse; Panel C: 2SLS estimates with respect to the degree of right-wing discourse using the Budge and Laver (2016) index; Panel D: 2SLS estimates with respect to the degree of extremist discourse using the absolute value of the Budge and Laver (2016) index. Each column presents the estimation for a different election. All specifications include a candidate-specific quadratic polynomial in latitude and longitude ($x+y+x^2+y^2+xy$). First-stage F-statistic is the Kleibergen–Paap rk Wald F-statistic of the first-stage estimation with eligible births as the exogenous instrument. * significant at 10%; ** at 5%; *** at 1%.

Table A.4
Party proximity and trust in institutions: Falsifications exercises.

	Panel A: Gender falsifications. *Which party do you feel closer to, or less distant from*						
Dep. variable	*Communist*	*Socialist*	*Ecologist*	*Liberal*	*Conservative*	*Radical right*	*None*
Alsace	−0.020*	−0.026	0.022	0.047**	−0.011	0.012	−0.028
	(0.010)	(0.024)	(0.015)	(0.021)	(0.020)	(0.008)	(0.018)
Born 1908–1913 ×Alsace	−0.030	0.167***	−0.054**	−0.108	0.004	−0.006	0.031
	(0.022)	(0.044)	(0.025)	(0.083)	(0.067)	(0.009)	(0.046)
Mean dep. variable	0.014	0.169	0.071	0.157	0.219	0.027	0.341
Observations	1,095	1,095	1,095	1,095	1,095	1,095	1,095
Clusters	40	40	40	40	40	40	40
	Panel B: Gender falsifications. *Would you say you rather trust or distrust in*						
Dep. variable	*Schools*	*Justice*	*Police*	*Church*	*Army*	*Parties*	*Politicians*
Alsace	−0.085*	0.02 4	−0.02 1	−0.088**	−0.151***	−0.07	0.004
	(0.048)	(0.122)	(0.052)	(0.033)	(0.037)	(0.068)	(0.069)
Born 1908–1913 ×Alsace	0.179	0.052	0.284	0.127***	0.264	−0.204	0.221
	(0.118)	(0.220)	(0.198)	(0.034)	(0.188)	(0.170)	(0.204)
Mean dep. variable	0.917	0.640	0.860	0.899	0.872	0.272	0.669
Observations	169	172	171	179	164	151	166
Clusters	35	34	34	35	34	33	35
	Panel C: Region falsifications. *Which party do you feel closer to, or less distant from*						
Dep. variable	*Communist*	*Socialist*	*Ecologist*	*Liberal*	*Conservative*	*Radical right*	*None*
Born 1908–1913 × Alsace	−0.020	−0.026	0.024	0.008	0.058	0.034	−0.075
	(0.052)	(0.220)	(0.046)	(0.137)	(0.119)	(0.029)	(0.143)
Placebo tests (other departments)							
5th percentile	−0.098	−0.333	−0.074	−0.236	−0.279	−0.058	−0.280
95th percentile	0.123	0.355	0.078	0.219	0.238	0.077	0.243
Two-tailed test p -value	0.696	0.905	0.594	0.953	0.626	0.239	0.600
Mean coefficient (other departments)	0.011	0.020	0.004	−0.011	−0.013	0.006	−0.013
Observations (department pairs)	4,004	4,004	4,003	4,004	4,004	4,000	4,004
	Panel D: Region falsifications. *Would you say you rather trust or distrust in*						
Dep. variable	*Schooling system*	*Judicial system*	*Police*	*Church*	*Army*	*Political parties*	*Elected politicians*
Born 1908–1913 × Alsace	−0.004	0.069	−0.067	0.131	−0.205	−0.100	−0.549**
	(0.226)	(0.462)	(0.243)	(0.245)	(0.330)	(0.403)	(0.244)
Placebo tests (other departments)							
5th percentile	−0.428	−0.677	−0.469	−0.640	−0.602	−0.628	−0.641
95th percentile	0.369	0.688	0.434	0.744	0.545	0.678	0.674
Two-tailed test p -value	0.986	0.881	0.785	0.594	0.535	0.803	0.025
Mean coefficient (other departments)	−0.020	0.004	0.000	−0.002	−0.023	0.022	0.019
Observations (department pairs)	1,755	1,760	1,646	1,554	1,695	1,574	1,367

Notes: Falsifications exercises of the effect of Wehmacht eligibility on political preferences. Panel A: Gender falsifications with party proximity as the outcome; Panel B: Gender falsifications with measures of trust in institutions are the outcome. Panel C: Region falsifications with party proximity as the outcome; Panel D: Region falsifications with measures of trust in institutions are the outcome. Each column presents the estimation for a different party/institution. Year-of-birth × gender fixed effects are included in all specifications. The unit of observation in Panels A and B is a female individual; in Panels C and D it is a department pair. Standard errors clustered at year-of-birth × region level in Panels A and B. * significant at 10%; ** at 5%; *** at 1%.

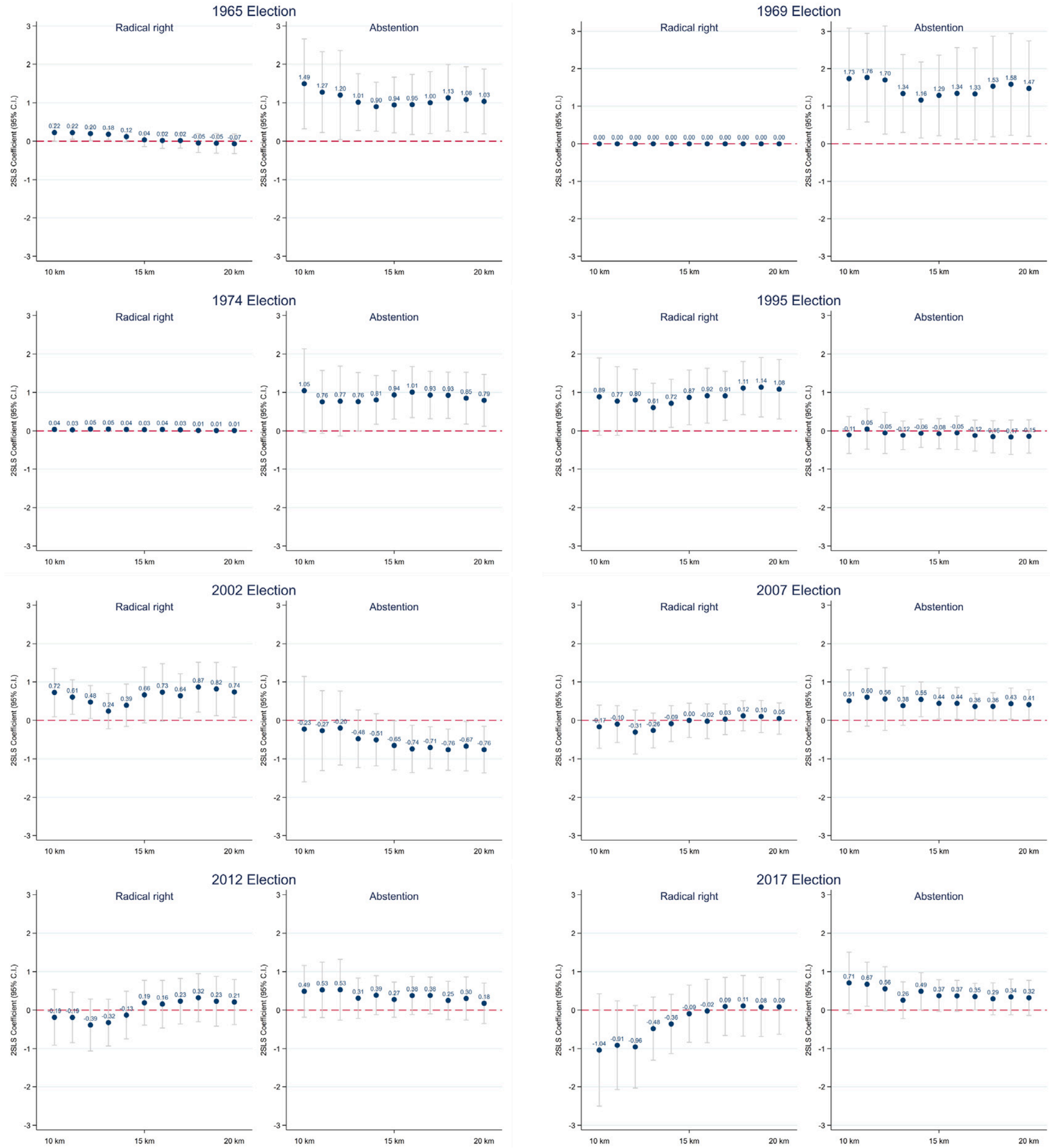

Fig. A.1. Abstention and the radical right-wing vote: Sensitivity to bandwidth selection.

Notes: Sensitivity of the results of Table 5 to different bandwidths (10 km to 20 km). Each graph presents the estimation for a different election. Left-hand-side sub-graph: 2SLS estimates with radical right-wing vote as the outcome; right-hand-side sub-graph: with abstention as the outcome. Standard errors clustered at the canton level; 95% confidence intervals presented in gray.

Appendix B. Supplementary data

Supplementary material related to this article can be found online at https://doi.org/10.1016/j.euroecorev.2022.104086.

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