Trade shocks and relative consumption : why the European middle class is turning (far) right

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Benedicta Marzinotto

European University Institute
Robert Schuman Centre for Advanced Studies

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Abstract

This paper relates far-right political preferences to changes in relative consumption stemming from trade exposure to low-quality producers such as China. The availability of more affordable varieties benefits the poor, while low-to-middle income groups stand to lose from such import shocks. Relative consumption deprivation awakens their perception of losing out relative to other once marginalised groups in the same society. Resentment for status loss explains the recent rightward drift in politics that is then channelled into support for the far-right especially during the main part of the China shock from 2000 to 2006. I empirically explore this hypothesis by relating measures of relative consumption deprivation to survey-based data from the European Social Survey (ESS) on a sample of 18 European countries over 2002- 2014.

JEL Code

E21; F10; F62; P16

Keywords

Political radicalism; political economy; China shock; consumption; European Union

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1. Introduction

This paper aims to explain the rightward shift in political preferences up to the fringes of the political spectrum that has been recorded in Europe over the last three decades, mirroring a rather global trend. The recent growth of empirical scholarship on the origins of political radicalism is impressive but one argument that remains central to much of the available literature is the idea that trade globalization in the form of a "China shock or syndrome" prompted affected workers to hide behind the protectionist electoral platforms of far-right political parties (Algan et al 2017; Autor et al 2016a, 2016b; Colantone and Stanig 2018a; Colantone and Stanig 2018b; Rodrik 2018, 2021; Dippel et al. 2022; Backes and Müller 2024). At the same time, it is generally recognized that trade globalization and in particular the emergence of cheap producers have increased low-income agents' purchasing power. While this benefit is unlikely to redress the loss in income from the displacement of low-skilled jobs, it may well trigger knock-on effects elsewhere, a mechanism that is able to reconcile the fact that the pool of far-right voters is much larger than that the size of trade-displaced manufacturing workers.

I start off by conceptualising trade globalization as a demand- rather than a supply-side shock. The theoretical framing of the paper follows influential work on the pro-poor bias in international trade through its impact on expenditures. While low-income consumers are the main beneficiaries because they spend a significant share of their income on cheap tradables, middle-income consumers stand to lose from the process of trade liberalization (Fajgelbaum and Khandelwal 2016; Heins 2023). Seen this way, trade exposure especially in the form of a China shock is relevant through a mechanism that is complementary to standard supply-side effects from import competition yet at the same time more cross-cutting. The relative consumption loss of middle-income agents becomes politically consequential because they are generally more concerned with social status than any other income group, as acknowledged in the standard literature on social comparison and relative deprivation (Smith et al 2012; Kim et al 2018).[1] Far-right political preferences should thus mainly concern low-to-middle income groups, who see their consumption behaviour approximating that of lower income households, while departing from that of high-income agents, thereby feeling threatened by societal decline. This is in line with existing research showing that support for conservative views and parties is concentrated amongst low-to-middle income groups and indeed driven by a perception of relative economic decline amongst politically powerful groups (Kurer and Palier 2019; Kurer 2020; Engler and Weisstanner 2021). In a nutshell, building on a well-defined trade model, I aim to show that relative consumption misfortunes change the middle class' perceived social status and spur (median-voter) support in favour of far-right politics.

The argument largely resonates with that of Burgoon et al (2019), who show that radical preferences in Europe are informed by a sense of status loss driven by positional deprivation, where the latter is measured by the extent to which the growth in income of individuals is outpaced by that of others in the same society. I depart from them by accounting for relative losses in consumption rather than income. Building on the notion of subjective social status as used in Gidron and Hall (2017, 2020), I hypothesise that one's positioning in the social ladder is accurately and better informed by relative changes in consumption expenditures because these are socially more visible than relative changes in income, where the latter normally correlate with subjective social status.

1 Richer households are likely to lose in consumption terms relatively to the low-income as well because their consumption behavior is unaltered by trade exposure to a low-quality producer; yet, they should not perceive a threat of imminent social-status loss nor generally oppose globalization as they might well benefit from the financial dimensions of globalization in the form of enhanced returns on capital (Pastor and Veronesi 2021).

To account for political preferences, I use pooled European Social Survey (ESS) data registering individual-level attitudes in up to 18 European countries from 2002 (ESS1) to 2014 (ESS7).[2] I consider, first, whether respondents self-identify with the right and, second, whether they "feel closest" to parties that are classified in the available political-science literature as radical (or populist) configurations (Mudde 2009; Roodujin and Burgoon 2018; Burgoon et al 2019). This way, I am able to provide an explanation for the long-standing rightward drift in politics (Gethin et al 2022) starting from the mid-1990s as well as the more recent radicalization of right-wing voters. Because none of the ESS rounds contains information on individuals' consumption behaviour, I measure changes in consumption in the five years before each survey round using data from the Household Budget Survey (HBS), which include mean consumption expenditures by income quintile in a given country-time, and then match these aggregate measures to the ESS data.[3] I define relative consumption deprivation as a situation in which one's consumption growth (i.e. the consumption of respondents belonging to each income quintile) is outpaced by that of other groups in the same society (i.e. the average gain in consumption across all the remaining income quintiles).

My claim is that relative consumption during the main part of the China shock is an important driver of radical far-right preferences beyond and on top of income effects, whether trade-induced or not. Crucially, the theoretical framing of the paper implies that consumption gains from trade are not evenly distributed and, most importantly, they are distributed differently from trade-induced income effects. Low-income agents loose from the supply-side effects of trade globalization as their jobs are displaced but gain in consumption terms. Working out the net effect is a complex exercise and is likely to vary as a function of multiple individual and group-level factors. Moreover, there is no guarantee that the net effect becomes politically consequential, not least because it is ambiguous in the eyes of the individual herself. There is one category for which the net effect is more clear-cut. The middle-class is untouched by the supply-side effects of trade globalization but would unambiguously suffer from relative consumption deprivation; concerns for social status do the rest.[4]

Figure 1 correlates percentage changes in consumption by income quintile retrieved from the HBS with the change in the share of imports from China over total imports from the early 2000s to 2014. It provides the empirical motivation for the paper. The association between trade with China and consumption gains is more clearly positive for the first quintile and loses significance when moving up to higher income groups.[5] Moreover, almost independently of the degree of exposure to Chinese imports, it is generally true that the first quintile is gaining more in consumption terms than the fifth (or even lower) quintile. Differently put, countries do not need to be large importers of Chinese products, the distributive effect will always be there even in the presence of modest import penetration from China. Overall, the evidence delivers a very intuitive message and one that is fully consisted with the sophisticated trade models in Fajgelbaum and Khandelwal (2016) and Heins (2023): China has been flooding international product markets with products that cater to low-income groups, whose consumption grew by more than that of other groups in the same society. It remains to be explained why such distributive effects fuel resentment, then channelled into support for far-right parties, for which there is arguably a sociological explanation.

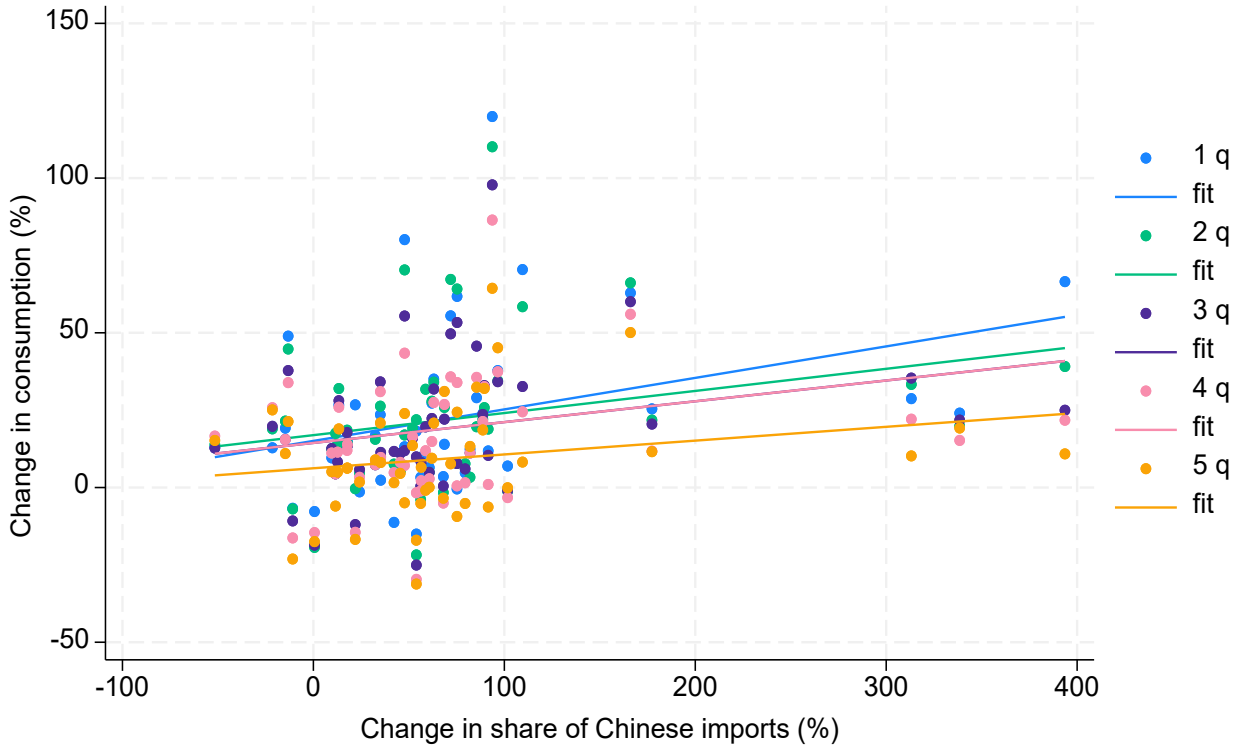
2 The countries included in the baseline analysis are: Austria, Belgium, Czech Republic, Denmark, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Netherlands, Poland, Slovenia, Slovakia, Spain, Sweden, and the United Kingdom.

3 The repeated cross-section in the HBS circumvent the problem of individuals that change income group from one survey round to the other.

4 The data suggest a quantitative effect but there might be also a quality effect. Products not made in China have become relatively expensive such that the mean real or perceived quality content of the consumption of the middle class has fallen relatively to richer households, who remain the only ones that can afford top-of-the-line products.

5 The correlation between the change in the share of Chinese imports and the consumption growth of the first income quintile is 0.35 ($p<0.05$). In the case of all other income quintiles, the correlation coefficient is not statistically significant.

Figure 1: China-induced shifts in consumption by income group



Key: Own elaboration based on Eurostat and Household Budget Survey (HBS).

Treating trade globalization as a demand-side shock comes with two important implications that frame this paper's analytical set-up. First, the distributive effects from trade exposure do not vary from industry to industry but from one income group to the other against the premise that consumer preferences are non-homothetic and that different types of households consume cheap tradables at different intensities. Conceptually, consumption inequalities might emerge even at constant income if there is an exogenous source of variation in the cost of living that stems, for example, from the large supply of cheap Chinese products. This is crucial to the identification of a consumption channel that is not necessarily driven by changes in income, as I do here. Second, gains from trade also vary from country to country as a function of relative technology endowment (Heins 2023), which is an additional source of variation I exploit in the data. Trade exposure might be politically consequential in the rich Western European economies of the euro area (EA) mainly because trade increases the availability of more affordable products rather than of more sophisticated varieties as it happens in the less affluent non-EA Eastern European countries.

I obtain results that are largely supportive of my hypotheses. Firstly, I find that relative consumption is an important driver of right self-identification, net of other individual-level determinants of political preferences (i.e. age, gender, education, employment status, foreign-born status, urban status, religiosity, subjective income). My findings underscore the importance of distinguishing between consumption and income as those that suffer from consumption deprivation might still be wealthier than those that gain in consumption terms, which is an assumption that is both plausible and at the core of the reference trade model. The intuition is confirmed: in countries where middle-income groups have seen their consumption growth outpaced by that of low-income agents, the rightward drift in politics is more evident and is further channelled into political support for far-right parties especially during the main part of the China shock (2000-2006).

The rest of the paper is structured as follows. Section 2 provides a critical overview of the available literature. Section 3 describes the data and the empirical strategy. Section 4 discusses the main results and provides additional robustness tests and extensions. Section 5 concludes.

2. Review of the literature

The literature on the origins and drivers of political radicalism is buoyant. I focus here on the different forms of economic distress that have been identified in the available studies and the mechanisms that turn economic misfortunes into resentment, then channelled into support for radical political ideas and parties. Given that most far-right parties rely on electoral platforms that oppose globalization and advocate a return to more protectionist policies under the label of "nationalistic autarky" (Burgoon 2009; Colantone and Stanig 2018), the bulk of the literature has been concerned with the distributive effects of globalization.[6] A common explanation that falls under the name of "China shock or syndrome" alludes to the fact that import competition from China has threatened jobs in manufacturing sectors and contributed to the (marginal) electoral success of far-right parties in the (manufacturing-intensive) regions that were mostly exposed to trade with China both in the U.S. (Autor, Dorn, and Hanson 2013, 2016; Autor, Dorn, Hanson, and Majlesi 2016; Autor 2017) and in Europe (Colantone and Stanig 2018a; Colantone and Stanig 2018b; Dippel et al 2022; Backes and Müller 2024).

A further implication from this literature is that globalization increases income inequality. Resentment may arise not only through concentrated wage and displacement effects, but because of either the electorate's (more or less diffuse) dislike of inequalities (e.g. Han 2016) or, similarly, their tendency to develop socio-tropical preferences (Colantone and Stanig 2018a). It is probably not all about trade. Pastor and Veronesi (2018) build a model showing that income inequality driven by globalization is a compelling explanation for the success of populism on a global scale because financial globalization has provided investment opportunities for wealthy groups, who have benefited from high returns to assets, whilst not impacting others.

The implicit trade models that underpin the globalization-related literature help to identify groups of losers and provide guidance for the operationalization of the independent variables that should drive support for populist parties, but often lack sophistication. So, for example, the principle that employment losses from trade integration persist even if agents consume entirely the imported good is valid only against a number of unrealistic assumptions, one of which is perfect competition. Second, there is variation within the manufacturing sector itself; some manufacturers are either well integrated into global value chains or produce high-value-added products for which China is hardly a true competitor (e.g. Germany); China may in fact become an export destination as product quality becomes appealing to a rising Chinese middle class.[7] Third and central to this paper, trade does not only affect wages and income on the supply-side, but comes with unequal effects on welfare when relaxing the standard assumption that consumer preferences are homothetic. Both Fajgelbaum and Khandelwal (2016) and Heins (2023) provide substantial empirical evidence that international trade is benefiting low-income consumers, while negatively affecting low-to-middle income groups. Moreover, the "pro-poor bias" of trade is stronger in rich Western European economies, which import cheap varieties, than in Eastern European economies where the price of imported goods is often higher than that of domestically produced goods, implying that trade increases the availability of more sophisticated varieties that possibly cater to higher-income groups.

6 A later literature has addressed forms of economic disruption other than globalization including immigration (e.g. Halla et al 2017; Mayda et al 2022) and crisis-related economic insecurity (e.g. Guiso et al 2024).

7 Relatedly, a general finding in the available literature is that European production systems suffered more from an export than an import China shock (e.g. Flückiger and Ludwig 2015).

There is a more recent body of literature that has the merit of merging the economic and the sociological explanations of political radicalism and that is more explicit about the transmission mechanism going from material distress to political support for one or the other party configuration. Burgoon et al (2019) use individual data on political preferences to test whether feelings of positional deprivation impact on the choice in favour of non-mainstream parties independently of the income group agents belong to, i.e. whether their starting position is that of a low- or high-income household. They find that positional deprivation measuring the loss in real disposable income relatively to other income groups in the same society is a superior explanation for radicalism than a loss in absolute income. In looking for a reason why some respondents support far-right and others far-left radicalism, they find that agents whose income fell (or grew less) relatively to lower deciles of the country's income distribution display far-right preferences, whereas those losing out relatively to higher deciles support far-left radicalism. Their data however do not allow identifying whether agents belonging to one or the other income group are more prone to develop resentment as their positional deprivation intensifies. Gidron and Hall (2017, 2020) employ the concept of social marginalization to capture the idea that a worsening of agents' subjective social status - defined as the perception of where they stand relatively to others in society - taps into discontent and might lead to radicalism in voting.

In my approach, differently from Burgoon et al (2019), I account for the "income identity" of respondents to account for the possibility that the same level of deprivation fuels resentment amongst certain income groups more than amongst others. Furthermore, I leverage relative perceived distress borrowing from the notion of subjective social status in Gidron and Hall (2017, 2020), but contend that one's positioning in the social ladder is accurately informed by consumption expenditures, which deserve special attention above and beyond income for at least two reasons. First, inequalities in consumption are more visible than income inequalities and hence more likely to trigger social comparisons. This is inspired by an older literature on positional goods, whose utility stems not only from absolute, but also relative levels of consumption (Duesenberry 1949; Frank 1985; Hopkins and Kornienko 2004; Carlsson et al 2007).[8] The scant empirical evidence suggests that this kind of interpersonal competition is there independently of age, gender, and income (Solnick and Hemenway 2005). Such an approach possibly contributes to explaining why subjective income may not fully coincide with actual income (Gidron and Hall 2017, 2020) as long as agents judge how well they fare in economic terms based on their consumption behaviour rather than on either income flows or wealth.[9] Second, trade-induced income inequality is not necessarily accompanied by an increase in consumption inequality if trade benefits disproportionally low-income agents via the expenditure channel. The focus on consumption is thus useful to explain why resentment may arise amongst richer groups in spite of the fact that their income is preserved, for example, because they are employed in sectors that are not exposed to trade. This seems a more convincing explanation of why service-sector workers and public employees develop radical political preferences than the general principle that their response is socio-tropical (Colantone and Stanig 2018a), which is in fact at odds with the recent literature on rising polarization along both class and cultural identities (Bonomi, Gennaioli and Tabellini 2021).

The centrality of the relative loss of socially powerful groups results also from previous analyses. Kurer and Palier (2019) and Kurer (2020) leverage the redistributive effects of broad technological change and automation respectively and show that those that are most exposed to it (i.e. politically powerful routine workers) tend to support far-right populist parties. Similarly, Engler and Weisstanner (2021) demonstrate that it is higher income groups that support the far-right as income inequality increases the probability that powerful groups undergo social decline. From this standpoint, the demand for radicalism is an inter-class and/or a median-voter phenomenon.

8 In the words of Galbraith "emulation [operates] mainly on behalf of privately produced goods and services" (1958). Houses, cars and potentially household appliances are typical positional goods.

9 The alternative explanation is that agents suffer from misperceptions that are fully irrational and unrelated to their actual economic situation, however measured.

3. The data

3.1. Measuring political preferences

To account for political preferences, I rely on various rounds of the European Social Survey (ESS). The surveys include repeated questions on values and attitudes over a number of waves starting 2002 together with granular individual-level characteristics for a large sample of European countries.[10] I consider two survey questions. One is whether respondents think of themselves as right-wing voters (i.e. the "right self-identification" question). The other is whether they "feel closest" to parties that are normally classified in the political-science literature as radical populists. For the latter, I follow closely the classification used in Burgoon et al (2019) and related literature (Mudde 2009; Roodujin and Burgoon 2018) to be able to compare my results to theirs.

Right self-identification tends to be a "pure" political preference because it is not affected by the decision to either participate in elections or not; nor by the appearance and stance of national political parties; nor by electoral politics altogether. I find it useful to start off from this aspect for two related reasons. First, right self-identification allows me to account for the long-run rightward shift in politics (Gethin et al 2022), of which radicalism might be just one manifestation. Second, by comparing statistical models that predict political support for the far-right with and without this additional independent variable, I am able to quantify, by approximation, the additional impact of economic shocks on political attitudes in the form of a "radicalization effect" that is plausibly purging my results from the possibility that the overall demand for populism is driven by long-standing changes in cultural frameworks. This way, I address some of the concerns in Margalit (2019), who argues that the globalization-related literature accounts for why the demand for populism increases (i.e. the marginal effect), but is unable to gauge the determinants of the general level of support for populist platforms (i.e. the average effect).

3.2. Consumption behaviour

Unfortunately, the ESS datasets do not contain any information about individuals' consumption behaviour and preferences. For that, I rely on the Household Budget Survey (HBS) to generate consumption growth variables at the country-year-quintile level and match these aggregate measures to the ESS data. Mean consumption expenditures of private households are collected through a combination of one or more interviews and diaries maintained by households approximately every five years. I use HBS data for the years 1994, 1999, 2005, 2010 and 2015 and match them to the relevant (i.e. after an average of five years) ESS rounds from 2002 up to 2014, hence limiting the analysis to ESS rounds 1, 3, 5 and 7.[11] A significant time lag between surveyed political attitudes and the evolution of consumption is warranted for two reasons. First, agents might need time to realise that their economic situation has changed both over time and relatively to others in the same society. Second, I am interested in enduring changes in consumption behaviour rather than in consumption volatility over the cycle because the latter is less likely to feed into perceptions of one's social status than the former.[12] To allow for cross-country comparability, consumption expenditures are expressed in purchasing power standard (PPS).

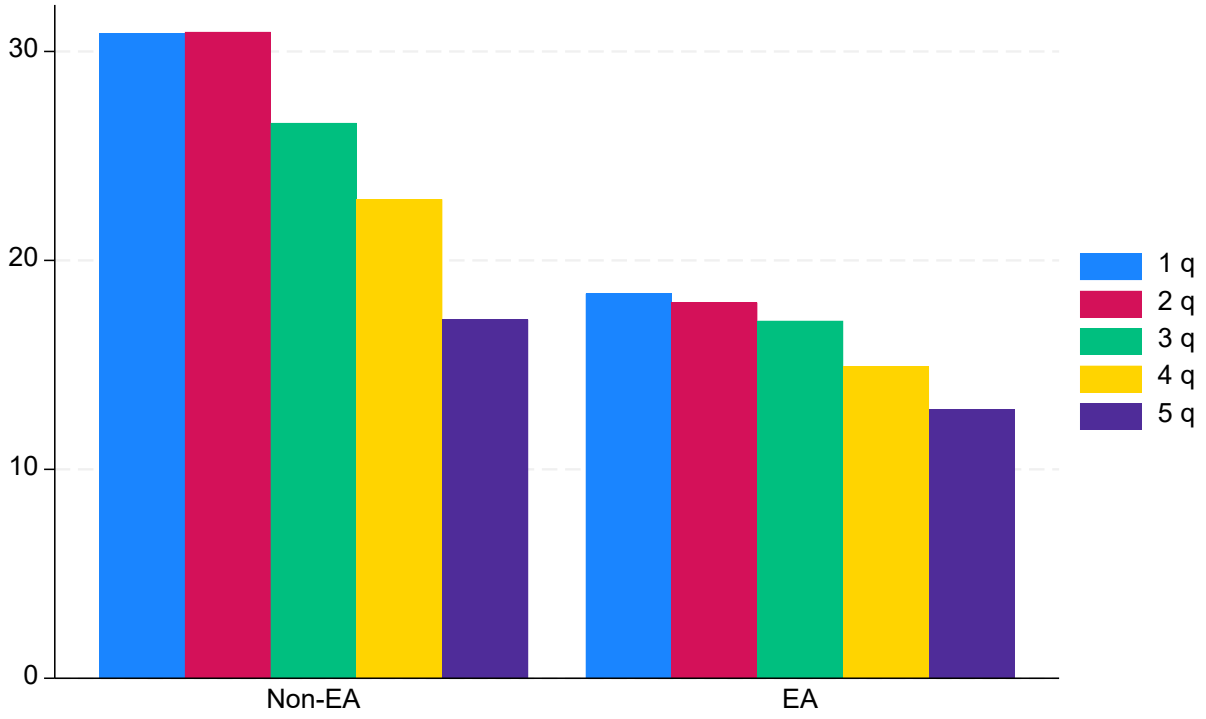
10 The size of the sample changes from one wave to the other. See Table A2 in the Appendix for details on data availability.

11 Using all (overlapping) ESS rounds would deliver biased results possibly stemming from an over-estimation of the impact of consumption on political preferences. To clarify, political attitudes from the ESS wave of 2002 (round 1), 2006 (round 3), 2010 (round 5) and 2014 (round 7) are linked to changes in consumption retrieved from the HBS over 1994-1999, 1999-2005, 2005-2010, and 2010-2015, respectively.

12 A third additional reason is that the five-year period is the average duration of an electoral cycle. Yet, because I am not looking at actual voting behavior, this is second-order.

Figure 2 shows the "consumption incidence curve" across sampled countries over the period 1994-2015, including the distinction between EA (i.e. Western Europe) and non-EA (i.e. Eastern Europe) countries.[13] On average, consumption gains fall linearly as income increases.[14] This aggregate result may seem at odds with the received wisdom of a marked steady rise in income inequalities (e.g. Lakner and Milanovic 2016). Yet, it is not. The reason is that the propensity to consume is differently distributed across income groups. The consumption of the richest households has increased less even if their income has arguably risen by more than that of other households over the same time span because they tend have a low propensity to consume. Symmetrically, even if the income of low-skilled workers in countries exposed to trade might have fallen through a wage or displacement effect, the same agents may have nonetheless gained in purchasing power by more than higher income agents just because they consume cheaper tradables at higher intensity.[15]

Figure 2: The consumption incidence curve by country grouping (1994-2015)



Key: Own elaboration based on Household Budget Survey (HBS).

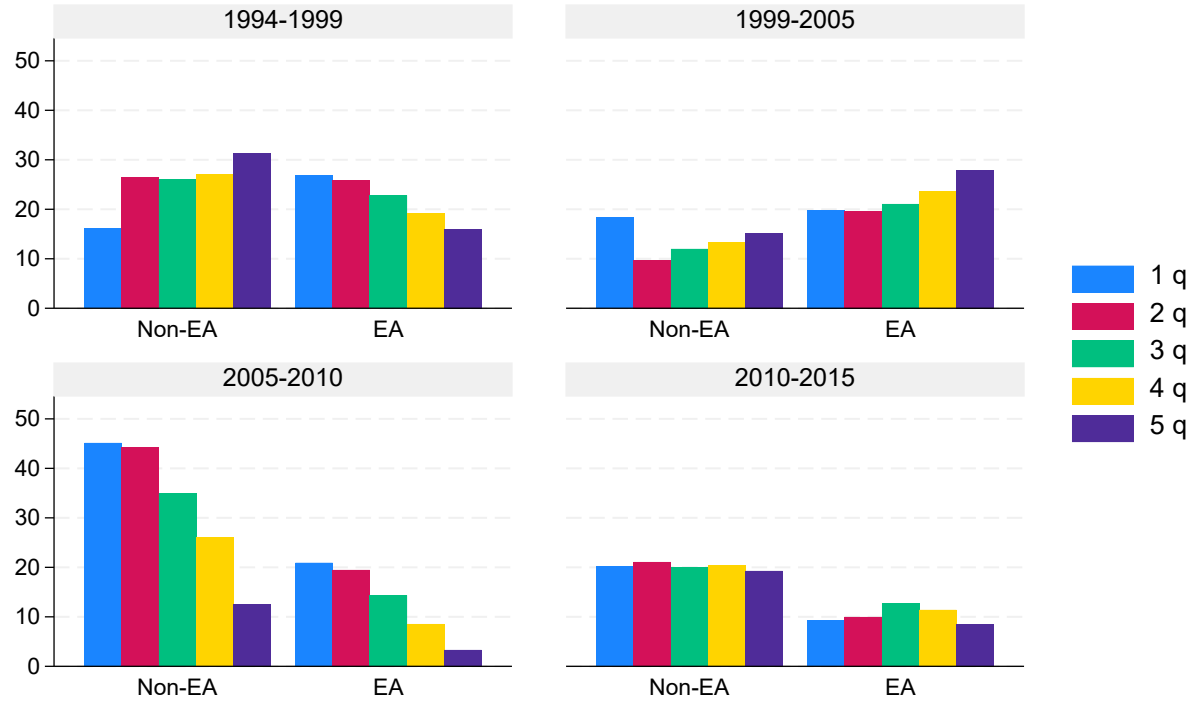
Notably, Figure 2 reveals differences in the level of consumption between EA and non-EA economies consistent with the idea of a catching-up process, but not in the shape of the consumption incidence curve. Rather, differential patterns emerge when isolating distinct time periods, with EA and non-EA countries being similar specifically over 2005-2010 but not at other times (Figure 3). This variation across countries and time is key for my identification.

13 The sample used here is larger than the available ESS sample and includes 28 European countries (i.e. Austria, Belgium, Bulgaria, Croatia, Cyprus, Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Latvia, Lithuania, Luxembourg, Malta, Netherlands, Poland, Portugal, Romania, Slovenia, Slovakia, Spain, Sweden, and the United Kingdom).

14 This is a simple average across all countries and times excluding all cases where there are missing data. See Table A2 in the Appendix for details on data availability.

15 Obviously, there are additional factors that might explain the difference between income and consumption incidence curves, one of which being access to finance.

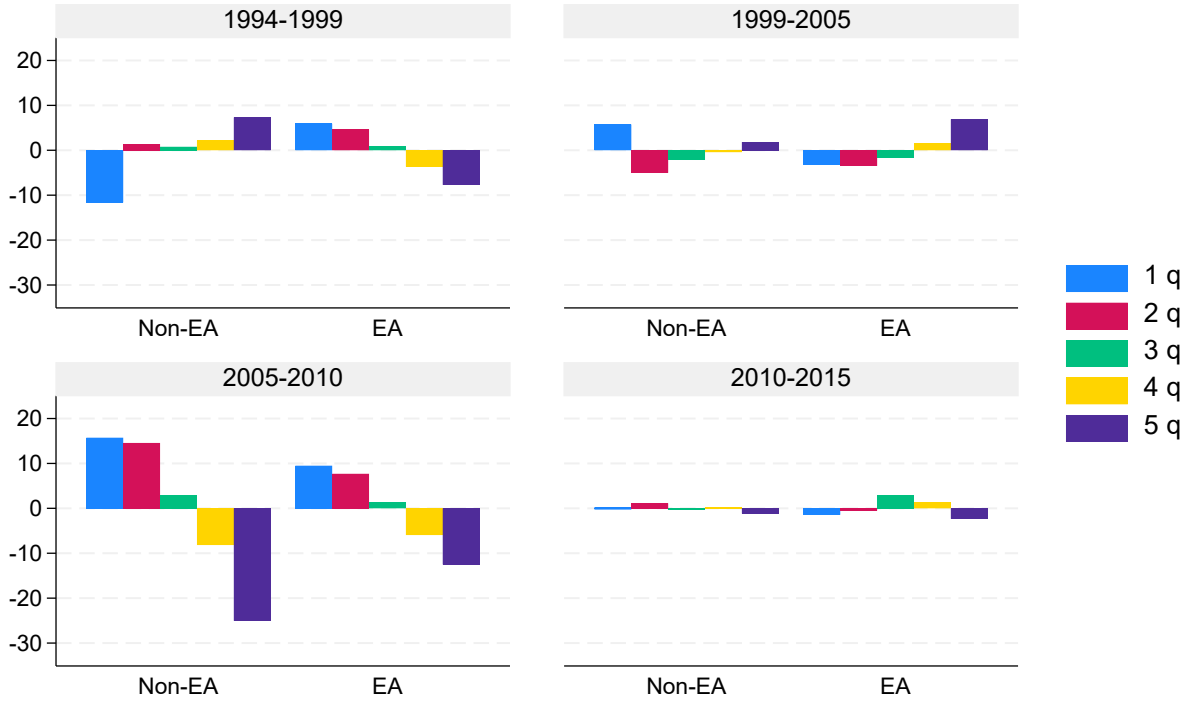
Figure 3: The consumption incidence curve by time and country grouping



Key: Own elaboration based on Household Budget Survey (HBS).

Figure 4 further displays relative as opposed to absolute consumption growth for each income quintile by time period (1994-1999, 1999-2005, 2005-2010 and 2010-2015) across EA and non-EA members. Relative consumption is calculated as the change in consumption for each income quintile minus the average change of all remaining quintiles. The average pattern coincides with the results from the counterfactual analysis in Fajgelbaum and Khandelwal (2016) on the unequal gains from trade along the income distribution. By and large, the first income quintile gains and the fifth quintile loses; the third quintile is stuck in between with relative gains at around zero. When taking the behaviour of the rich out of the picture - whose performance in fact varies depending on whether consumption of services is included or not (Fajgelbaum and Khandelwal 2016), the data unequivocally reveal a significant loss for the third quintile, which thus identifies the group of losers. This aspect is further developed in the next section where I propose a proxy to capture a “politically salient trade-induced consumption shock”.

Figure 4: Relative consumption by time period and country grouping



Key: Own elaboration based on Household Budget Survey (HBS).

3.3. The empirical strategy

Data on consumption growth by income quintile from the HBS described above are matched to the quintile each respondent in the respective ESS wave belongs to. So as to corroborate my argument, I test for the impact of both absolute and relative changes in consumption over time. Absolute consumption deprivation is simply the over-time change in mean consumption expenditures for the income group each respondent belongs to. It coincides with the data used to construct the "consumption incidence curve", yet with an inverted sign such that an increase in the value of the variable signals an increase in absolute economic distress. Relative consumption deprivation is here defined as the average change in consumption across (remaining) income groups minus the change registered by the income group the respondent belongs to. Higher values indicate that one's consumption is outpaced by that of the other members of the same society. Like in Burgoon et al (2019), I end up relying on "anonymous" measures of absolute and relative deprivation based on the respondent's income category. With the lack of accuracy comes but the fact that, first, the measure is certainly exogenous; moreover, the repeated cross-section in the HBS circumvent the problem of individuals that change income group from one survey round to the other, while also allowing me to rely on a large sample size that is less exposed to a violation of the assumption of random sampling.

My main claim is that the dynamic and relative evolution of consumption has a bearing on political attitudes after controlling for standard socio-economic determinants at the individual level. In my initial exercise, I focus on right self-identification because it can be considered by and large independent of the structure of the national party system and then move onto explaining whether this is channelled into electoral support in favour of far-right political parties. For that purpose, I estimate binary logistic regression models of the following type:

$$v_{ict} = \beta_0 + \beta_1 Cons_{ict} + \beta_2 X_{ict} + \beta_3 z_{it} + \beta_4 t_i + \varepsilon_{ict} \quad [1]$$

where v_{ict} is the odds that respondent *i* either self-identify with right-wing ideologies or supports far-right political formations in country *c* and ESS round *t*; *X* are individual-level characteristics that are expected to influence political preferences; *Cons* is my measure of consumption deprivation; *z* are country fixed effects and *t* are ESS wave fixed effects. My fixed effects imply that the impact of consumption deprivation is identified only out of variations across income groups within the same country and year (i.e. ESS round). I cluster standard errors at the country-time-quintile level because there might be autocorrelation stemming from the fact that I proxy individual-level consumption by changes in mean consumption expenditures in a respondent's income quintile. All models include design and population weights.

Individual-level controls consist of: a respondent's *level of education* (1 = less than lower secondary education, 2 = lower secondary education, 3 = upper secondary education, 4 = post-secondary non-tertiary education, 5 = tertiary education); *age* (in years); *sex* (1 = female); whether the respondent is *unemployed* (1 = unemployed); *subjective religiosity* (on a scale from 0 to 11, with 0 = not at all religious and 10 = very religious); *foreign-born status* (0 = native born, including parents); *urban living status* (rural = 0, urban = 1). As I test for the role of consumption patterns above and beyond income dynamics, I require some measure of income that does not coincide with the income quintiles that are used to calculate my consumption variable. For that purpose, I use *subjective income* (from "finding it very difficult on present income" = 1 to "living comfortably on present income" = 4). To the extent that available studies show that this co-moves with actual income yet not one to one (Hall and Gidron 1997; Burgoon et al 2019), there should not be auto-correlation with my consumption variable but the measure is at the same time a valid proxy for actual income.

As such, the results from equation [1] would not allow to establish whether differential consumption patterns drive political preferences conditional on one's income category. For that purpose, I construct relative consumption deprivation for specific groups at the country-year level. I account for the evolution in consumption of each quintile relative to the average and, based on results, consider how certain quintiles relate to either lower or higher quintiles of the income distribution. I re-estimate [1] using these alternative measures. Importantly, I further explore whether right self-identification is channelled into support for a radical far-right party in a given country and ESS round using the question about whether respondents "feel closest" to a particular party configuration. I classify parties as either far-right or far-left populists using existing categorizations (Mudde 2009; Roodujin and Burgoon 2018; Burgoon et al 2019).[16] Building on my own evidence, I focus in particular on the consumption deprivation of the third income quintile relative to the first, which I use to capture what I refer to as a "politically salient trade-induced consumption shock".[17] Trade especially with a low-quality producer such as China comes with unequal effects that leave low-to-middle income groups relatively worse off, a politically relevant result for a social group that is generally found to be sensitive to perceptions of societal status, as shown in the literature on social comparisons.

16 Table A1 in the Appendix provides an overview of the classification used.

17 Interestingly, this is exactly the income group that is found to lose the most in the counterfactual exercise developed in Fajgelbaum and Khandelwal (2016), which is itself a test of the robustness of my results.

4. Main results and robustness tests

4.1. Baseline results

Table 1 shows the results from equation [1]. While absolute consumption deprivation is insignificant, relative deprivation increases the probability that respondents self-identify with the right, net of other key socio-economic individual-level indicators, including subjective income. In particular, my results indicate that consumption and income differentially structure political preferences, which underscores the choice of treating them separately. Higher subjective income increases the probability of right self-placement; in other words, respondents that feel financially strong align with right-wing political ideas, but absolute levels of income are not telling of each respondents' capacity to spend over others, which is instead what relative consumption deprivation would capture. Taken together, my results indicate that right self-identification is probably a feature of well-off socio-economic groups that, for reasons exogenous to income (e.g. trade-induced changes in spending), feel outpaced by others in the same society. Most of the individual-level controls are significant and rightly signed. Consistent with the available literature, I find that young low-educated males living in rural areas tend to self-identify with the right, as do agents that believe to be religious and are non-foreign born.

Table 1: Consumption deprivation and right self-identification

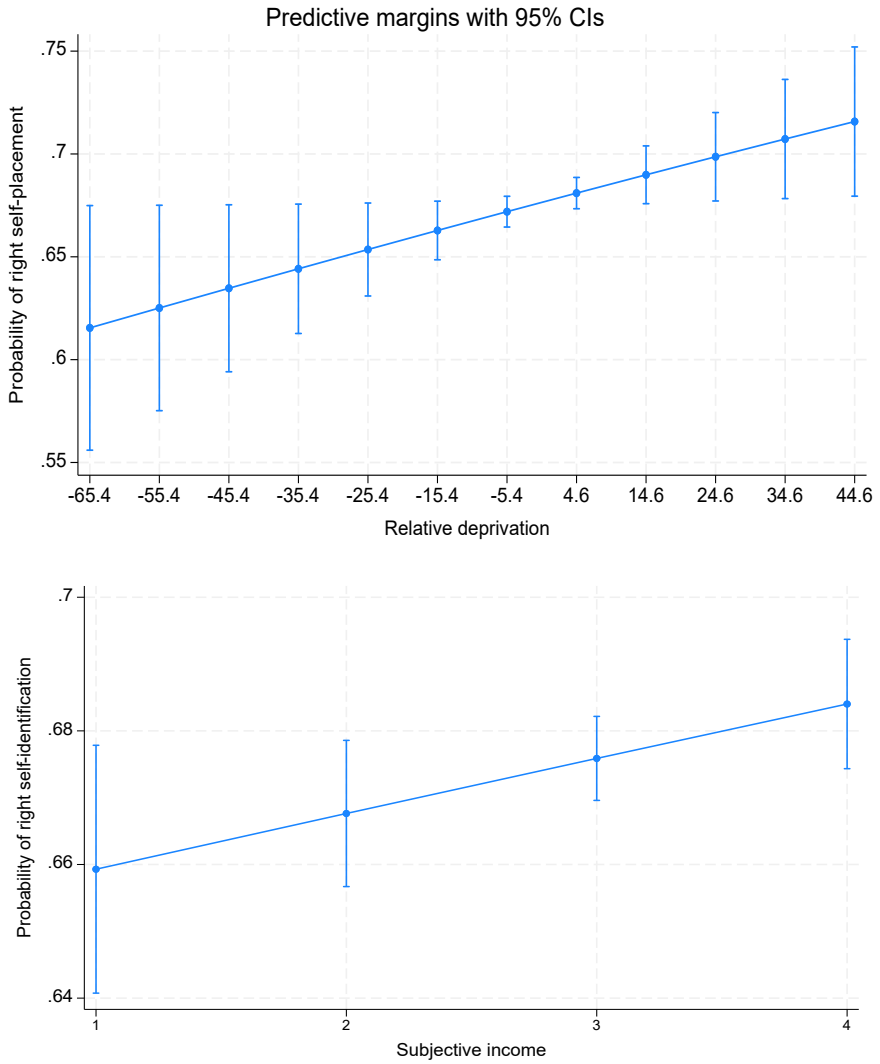
	(1)	(2)	(3)	(4)
Chinese penetration	0.000990*			
	[1.735]			
Absolute deprivation		0.000759		-4.96e-05
		[0.667]		[-0.0389]
Relative deprivation			0.00432**	0.00436**
			[2.397]	[2.085]
Subjective income	0.0746***	0.0453**	0.0396**	0.0396**
	[3.242]	[2.239]	[1.963]	[1.963]
Age (years)	-0.00279***	-0.00210**	-0.00201**	-0.00201**
	[-2.673]	[-2.494]	[-2.408]	[-2.396]
Education	-0.141***	-0.146***	-0.149***	-0.149***
	[-10.87]	[-13.31]	[-13.44]	[-13.44]
Female	-0.0864***	-0.101***	-0.0989***	-0.0989***
	[-2.795]	[-3.912]	[-3.824]	[-3.837]
Unemployed	-0.0995	-0.00608	-0.00169	-0.00167
	[-1.430]	[-0.101]	[-0.0283]	[-0.0280]
Religious	0.104***	0.106***	0.106***	0.106***
	[16.72]	[19.72]	[19.76]	[19.67]
Foreign born	-0.109**	-0.117***	-0.117***	-0.117***
	[-2.552]	[-2.994]	[-2.977]	[-2.978]
Urban	-0.187***	-0.195***	-0.195***	-0.194***
	[-5.769]	[-6.879]	[-6.876]	[-6.871]
Country dummies	yes	yes	yes	yes
Survey time dummies	yes	yes	yes	yes
Pseudo R-squared	0.043	0.044	0.044	0.044

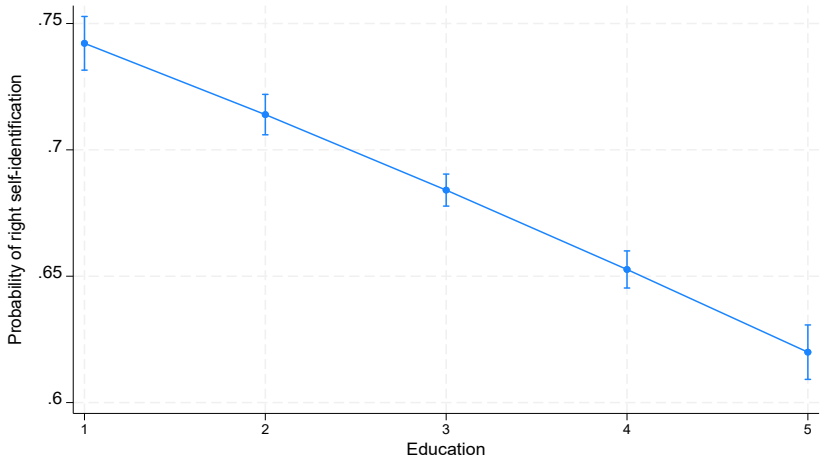
Cluster level	c-t-d	c-t-d	c-t-d	c-t-d
Observations	54,288	73,491	73,491	73,491

Key: Logistic regression estimates (binary dependent variable, 0-1). All models with fixed effects for countries and survey rounds (not shown) and robust-cluster standard errors. Robust z-statistics in brackets. *** p<0.01, ** p<0.05, * p<0.1. All models include design and population weights.

To account for the size of the effect of relative consumption deprivation, I compute the predicted probability of right self-identification across the full distribution. The results reported in Figure 5 show that relative consumption explain more variation in one's positioning than subjective income, albeit less than the level of education, which remains a key determinant of respondents' political affiliation.

Figure 5: Probability of right self-identification





Key: Predicted probability of support for government redistribution based on Table 1 (3) across the full distribution of relative deprivation, subjective income and education (predictive margins estimated from *Bars r*epresent 95% confidence intervals

I re-estimate equation [1] using country-year measures of consumption deprivation that are explicit about the income group of origin so as to test whether the income group a respondent belongs to is important. These country-year measures are calculated as the consumption deprivation of each quintile relatively to the average across the remaining four quintiles. There is obviously need for caution when interpreting longitudinal effects out of cross-national variation, but differences in the coefficients would be sufficient to prove that the income identity of the respondents in each country and time is a key factor in the mechanism that is described. To make sure my results are not driven by country-level income inequalities but truly capture the consumption dimension, I include the Gini coefficient. Additionally, I account also for perceived ethnic threat to be able to control for cultural drivers of self-identification and indeed gauge the additional effect from consumption-related distress. Table 2 shows the estimated coefficients. While the consumption-related distress of the first and the second quintile decreases self-identification with the right (Columns 1 and 2), the correlation turns positive and highly significant for the third quintile (Columns 3), consistent with the idea that negative economic shocks become politically consequential when they hit in particular the middle class. The estimated coefficient indicates that the odds of a rightward drift in politics increase by over 50% for each unit increase in the consumption deprivation of the middle-income relative to the poorest households.[18] To further corroborate the view that the middle class' perception of social status has driven the rightward drift in politics over the last two decades, I explore whether it is true, as in my assumption, that middle-income groups fear being outpaced by once marginalised groups in the same society rather than by the higher ends of the income distribution. My results show that when the third quintile is deprived in consumption terms relative to the first, a country's political orientation shifts to the right (Column 6), whereas "spending distress" relative to wealthier members of society is not politically consequential (Column 7).

18 My interpretation of the overall results is not affected by the positive yet weaker link between the (large) deprivation of the richest households and right-wing political preferences (Column 5 in Table 2). In fact, the general point I want to make is that different income groups are differently sensitive to changes in relative consumption and that the same or smaller relative loss in consumption triggers a shift to the right only amongst those that are concerned with social status and/or are at risk of societal decline, which is unlikely to be the case for the richest households.

Table 2: Consumption deprivation and right self-identification conditional on income identity

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	1st q/av.	2nd q/av.	3rd q/av.	4tht q/av.	5th q/av.	3rd/1st q	3rd/5th q
Relative deprivation	-0.00211*	-0.00604*	0.00801**	0.00623	0.00355*	0.00232*	-0.00280
	[-1.795]	[-1.818]	[1.997]	[1.614]	[1.893]	[1.917]	[-1.135]
Subjective income	0.124***	0.124***	0.124***	0.124***	0.124***	0.124***	0.124***
	[5.516]	[5.502]	[5.519]	[5.519]	[5.503]	[5.516]	[5.509]
Gini coefficient	-0.0442***	-0.0383**	-0.0427***	-0.0465***	-0.0416**	-0.0445***	-0.0395**
	[-2.684]	[-2.376]	[-2.660]	[-2.663]	[-2.559]	[-2.713]	[-2.418]
Age (years)	-0.00196**	-0.00195**	-0.00197**	-0.00196**	-0.00195**	-0.00196**	-0.00195**
	[-2.370]	[-2.360]	[-2.384]	[-2.376]	[-2.355]	[-2.374]	[-2.353]
Education	-0.0770***	-0.0768***	-0.0771***	-0.0770***	-0.0769***	-0.0771***	-0.0768***
	[-6.007]	[-5.972]	[-6.029]	[-5.994]	[-5.986]	[-6.014]	[-5.984]
Female	-0.0770***	-0.0764***	-0.0770***	-0.0768***	-0.0768***	-0.0770***	-0.0766***
	[-3.051]	[-3.034]	[-3.053]	[-3.044]	[-3.044]	[-3.052]	[-3.041]
Unemployed	-0.0871	-0.0863	-0.0887	-0.0875	-0.0858	-0.0875	-0.0857
	[-1.594]	[-1.578]	[-1.626]	[-1.603]	[-1.567]	[-1.603]	[-1.562]
Religious	0.119***	0.119***	0.119***	0.119***	0.119***	0.119***	0.119***
	[18.79]	[18.78]	[18.77]	[18.79]	[18.79]	[18.79]	[18.77]
Foreign born	0.0331	0.0328	0.0333	0.0331	0.0329	0.0331	0.0328
	[0.914]	[0.908]	[0.918]	[0.913]	[0.910]	[0.915]	[0.908]
Urban	-0.137***	-0.136***	-0.137***	-0.137***	-0.136***	-0.137***	-0.136***
	[-4.154]	[-4.141]	[-4.164]	[-4.169]	[-4.136]	[-4.158]	[-4.130]
Ethnic threat	0.184***	0.185***	0.184***	0.184***	0.185***	0.184***	0.185***
	[17.00]	[17.00]	[17.00]	[16.99]	[17.00]	[16.99]	[16.99]
Country dummies	yes	yes	yes	yes	yes	yes	yes
Survey time dummies	yes	yes	yes	yes	yes	yes	yes
Pseudo R-squared	0.061	0.061	0.061	0.061	0.061	0.061	0.061
Cluster level	c-d	c-d	c-d	c-d	c-d	c-d	c-d
Observations	88,422	88,422	88,422	88,422	88,422	88,422	88,422

Key: Logistic regression estimates (binary dependent variable, 0-1). All models with fixed effects for countries and survey rounds (not shown) and robust-cluster standard errors. Robust z-statistics in brackets. *** p<0.01, ** p<0.05, * p<0.1. All models include design and population weights.

The consumption deprivation of the third relatively to the first quintile is labelled as "politically salient trade-induced consumption shock". The reasoning is as follows: available trade models as well as my descriptive statistics identify the third quintile as the group of relative losers from trade consumption shocks, which happens also to be the one segment in society that is likely to be more sensitive to relative deprivation and social positioning. I explore whether this specific consumption shock spurs support in favour of the far right. Table 3 shows results of a logistic regression that estimates respondents' closeness to the radical right-wing in countries where the middle class is outpaced in consumption terms by the poor. [19] The pooled ESS results do not deliver statistically significant results (Column 1). When differentiating by ESS wave, I obtain that the deprivation of the third relatively to the first quintile increases support for the far-right both in 2002 and 2006, being especially strong in 2006 when, according to the available literature (Colantone and Stanig 2018a; Colantone and Stanig 2018b), Chinese import penetration reached its peak (Columns 2 and 3). During this time, the odds of far-right radicalism increase by over 90% for each unit increase in the consumption deprivation of the middle class relative to the poorest households. When controlling for the fact that part of the population might be already right-leaning, the relative consumption loss of the middle class is found to increase the odds in favour of the radical right by an additional 20%. This is to say, the unequal effects of trade via the expenditure channel generate a statistically significant and sizeable "radicalization effect" above and beyond long-standing changes in cultural frameworks (Table 4).[20]

Table 3: The China consumption shock and support for radical right

	(1)	(2)	(3)	(4)	(5)
	2002-2014	2002	2006	2010	2014
China shock (3rd/1st q)	-0.00899	0.0126**	0.488***	0.00289	-0.237***
	[-1.318]	[2.435]	[3.680]	[0.308]	[-8.414]
Gini coefficient	0.370***	0.0128	-0.315***	-0.140	0.792***
	[4.331]	[0.463]	[-4.977]	[-0.440]	[12.51]
Age (years)	-0.0161***	-0.0185***	-0.0315***	-0.00816	-0.0133***
	[-7.033]	[-3.918]	[-5.758]	[-1.512]	[-3.665]
Education	-0.135***	-0.0771	-0.152**	-0.0822	-0.190**
	[-3.375]	[-1.017]	[-2.413]	[-1.260]	[-2.436]
Female	-0.438***	-0.627***	-0.589***	-0.318**	-0.375***
	[-6.167]	[-4.942]	[-3.011]	[-2.312]	[-3.035]
Unemployed	0.297	0.126	1.085***	0.234	0.120
	[1.524]	[0.552]	[3.015]	[0.618]	[0.319]
Religious	-0.00560	-0.00426	-0.0407	0.0215	-0.00314
	[-0.306]	[-0.122]	[-1.258]	[0.706]	[-0.107]
Foreign born	-0.180*	-0.590**	0.244	-0.178	-0.115
	[-1.694]	[-1.960]	[1.094]	[-0.746]	[-0.799]
Urban	0.0508	0.514***	0.209	-0.142	-0.270**
	[0.502]	[2.796]	[1.106]	[-0.766]	[-2.331]

19 Because of data availability, the sample of countries is slightly smaller than the one used for Tables 1 and 2 and includes following countries: Austria, Belgium, Bulgaria, Croatia, Cyprus, Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Ireland, Italy, Lithuania, Luxembourg, Netherlands, Poland, Portugal, Slovakia, Slovenia, Spain, and United Kingdom.

20 The estimate whose coefficients are reported in Table 4 includes right self-placement but excludes perceived ethnic threat as the two are colliders.

Ethnic threat	0.403***	0.370***	0.470***	0.226***	0.483***
	[13.90]	[4.674]	[11.30]	[3.375]	[12.95]
Country dummies	yes	yes	yes	yes	yes
Survey time dummies	yes	no	no	no	no
Pseudo R-squared	0.24	0.19	0.19	0.13	0.32
Cluster level	c-d	c-d	c-d	c-d	c-d
Observations	36,081	7,278	8,370	8,381	7,521

Key: Logistic regression estimates (binary dependent variable, 0-1). All models with fixed effects for countries and survey rounds (not shown) and robust-cluster standard errors. Robust z-statistics in brackets. *** p<0.01, ** p<0.05, * p<0.1. All models include design and population weights.

Table 4: The radicalization effect of the China consumption shock

	(1) 2002	(2) 2006	(3) 2010	(4) 2014
China shock (3rd/1st q)	0.0111**	0.624***	0.00433	-0.138***
	[2.155]	[4.664]	[0.446]	[-5.089]
Gini coefficient	0.0383	-0.370***	-0.266	0.576***
	[1.420]	[-5.636]	[-0.921]	[9.471]
Right self-placement	3.451***	1.747***	0.398	1.326***
	[10.58]	[5.547]	[1.538]	[8.435]
Age (years)	-0.0166***	-0.0255***	-0.00498	-0.0103***
	[-3.977]	[-4.607]	[-0.975]	[-2.946]
Education	-0.163*	-0.320***	-0.176***	-0.344***
	[-1.926]	[-5.422]	[-3.168]	[-5.310]
Female	-0.569***	-0.424**	-0.368***	-0.104
	[-3.914]	[-2.510]	[-2.729]	[-1.025]
Unemployed	0.435**	1.190***	0.294	0.175
	[2.108]	[3.130]	[0.800]	[0.382]
Religious	-0.0984***	-0.100***	0.0153	-0.0428
	[-3.844]	[-3.192]	[0.519]	[-1.430]
Foreign born	-0.701**	0.297	-0.169	-0.365***
	[-2.566]	[1.026]	[-0.644]	[-2.619]
Urban	0.516***	0.249*	-0.201	-0.221**
	[2.735]	[1.665]	[-1.200]	[-2.137]
Country dummies	yes	yes	yes	yes
Survey time dummies	no	no	no	no
Pseudo R-squared	0.26	0.14	0.12	0.25
Cluster level	c-d	c-d	c-d	c-d
Observations	7,788	8,699	8,802	7,816

Key: Logistic regression estimates (binary dependent variable, 0-1). All models with fixed effects for countries and survey rounds (not shown) and robust-cluster standard errors. Robust z-statistics in brackets. *** p<0.01, ** p<0.05, * p<0.1. All models include design and population weights.

The sign of the coefficient turns negative just for the 2014 wave (Column 5 in Table 3). Possible explanations are that the financial crisis has introduced additional confounding factors or that a China-induced consumption shock is naturally receding or that the heterogeneity of the sampled countries is hiding significant differences. Indeed, when restricting the sample to the EA, the estimated coefficient is positive across all periods, with the probability of supporting right-wing radicalism diminishing over time consistent with a phasing out of the China shock (Table 5).

Table 5: EA - The China consumption shock and support for radical right

	(1) 2002	(2) 2006	(3) 2010	(4) 2014
China shock (3rd/1st q)	0.0126**	0.289***	0.00279	0.0407**
	[2.424]	[3.567]	[0.301]	[2.147]
Gini coefficient	0.0125	4.134***	-0.153	-0.227***
	[0.454]	[3.766]	[-0.479]	[-2.691]
Age (years)	-0.0190***	-0.0304***	-0.00825	-0.0202***
	[-3.866]	[-4.848]	[-1.212]	[-4.310]
Education	-0.0741	-0.175*	-0.102	-0.287**
	[-0.957]	[-2.212]	[-1.286]	[-2.491]
Female	-0.616***	-0.681***	-0.277*	-0.280
	[-4.749]	[-3.173]	[-1.761]	[-1.573]
Unemployed	0.138	0.916**	0.503	0.474
	[0.602]	[2.286]	[1.289]	[1.536]
Religious	-0.00322	-0.0276	0.0236	-0.0656*
	[-0.0914]	[-0.758]	[0.647]	[-1.893]
Foreign born	-0.605**	0.384*	-0.336	-0.390
	[-1.972]	[1.683]	[-1.249]	[-1.556]
Urban	0.524***	0.254	-0.0103	-0.119
	[2.801]	[1.175]	[-0.0483]	[-0.763]
Ethnic threat	0.367***	0.457***	0.190**	0.548***
	[4.449]	[9.079]	[2.316]	[15.93]
Country dummies	yes	yes	yes	yes
Survey time dummies	no	no	no	no
Pseudo R-square	0.19	0.18	0.13	0.25
Cluster level	c-d	c-d	c-d	c-d
Observations	6,407	6,284	6,889	5,465

Key: Logistic regression estimates (binary dependent variable, 0-1). All models with fixed effects for countries and survey rounds (not shown) and robust-cluster standard errors. Robust z-statistics in brackets. *** p<0.01, ** p<0.05, * p<0.1. All models include design and population weights.

Differently from Guiso et al (2019), for whom far-right radicalism is generally more pronounced in the EA because the economic effects from the financial crisis are exacerbated by the inability to use the devaluation option, I argue that the different performance across the two country groupings owes to the fact that the EA is differently exposed to the unequal effects of trade compared with non-EA countries, having a more pronounced pro-poor bias because it imports varieties of goods that are unavailable domestically, as shown in Heins (2023). To test for this proposition, I include in the estimation the average pro-poor bias as calculated by Heins (2023) for each country in my sample.[21] More specifically, I use the absolute consumption growth of the middle class (i.e. the third income quintile) in interaction with Heins' (2023) country-level pro-poor bias and further split radical preferences into support for the far-right and the far-left.[22] As shown in Table 6, the higher the consumption of the third quintile, the lower the support for radicalism, consistent with previous results for which it is deprivation that spurs far-right support. But when adjusting the consumption of the middle class for how such gains fare relatively to those of the poorest households, the results indicate that demand for radicalism increases when international trade is benefiting low income groups at the detriment of the middle class (Column 1) and, interestingly enough, only to the right of the political spectrum (Columns 2 and 3).

Table 6: The trade-induced consumption shock and support for radical parties

	(1)	(2)	(3)
	Radical	Radical right	Radical left
Pro-poor bias (Heins 2023)	-0.432**	-1.171***	-1.140
	[-2.553]	[-3.575]	[-1.098]
Absolute consumption (3rd)	-0.0185	-0.0631**	-0.0265*
	[-1.442]	[-2.429]	[-1.755]
Pro-poor bias *x* absolute consumption	0.00374	0.0160**	0.00602
	[1.119]	[2.455]	[1.142]
Gini coefficient	0.168***	0.315***	0.164***
	[2.967]	[3.338]	[2.780]
Age (years)	-0.00903***	-0.0160***	-0.00429*
	[-5.527]	[-6.958]	[-1.864]
Education	-0.0610**	-0.133***	-0.0176
	[-2.414]	[-3.341]	[-0.537]
Female	-0.221***	-0.437***	-0.0782
	[-4.818]	[-6.096]	[-1.224]
Unemployed	0.304***	0.305	0.339***
	[3.011]	[1.560]	[3.406]
Religious	-0.151***	-0.00683	-0.233***
	[-9.027]	[-0.375]	[-14.03]
Foreign born	-0.0395	-0.178*	0.0273

21 Heins (2023) develops a structural multisector model of international trade with i) non-homothetic preferences, ii) heterogeneous households, iii) cross-country variation in production technologies and endogenous vertical differentiation augmented with a log-logit preference structure. Against this framework, he predicts the counterfactual welfare of each consumer group under autarky and compares it to that in the current trade equilibrium. The bias that is quantified under this set-up shows the gain of a household at the 20th percentile of the income distribution relative to one at the 80th percentile.

22 I use a measure of absolute rather than relative changes in the consumption of the relevant income quintile because the pro-poor-bias variable contains already the distributive dimension of international trade.

	[-0.541]	[-1.668]	[0.284]
Urban	0.163***	0.0576	0.207***
	[3.210]	[0.577]	[3.285]
Ethnic threat	0.140***	0.403***	-0.0626**
	[5.960]	[13.96]	[-2.501]
Country dummies	yes	yes	yes
Survey round dummies	yes	yes	yes
Pseudo R-squared	0.12	0.24	0.09
Cluster level	c-d	c-d	c-d
Observations	45,362	36,081	33,147

Key: Logistic regression estimates (binary dependent variable, 0-1). All models with fixed effects for countries and survey rounds (not shown) and robust-cluster standard errors. Robust z-statistics in brackets. *** p<0.01, ** p<0.05, * p<0.1. All models include design and population weights.

How do these results speak to the literature that has engaged with the different drivers of right versus left radicalism? In brief, there are two categories of explanations for why individuals ally with either the far-right or the far-left. The first treats radical preferences as a form of protest against the incumbent government. Considering social democratic parties have been in government in most European countries over the last two decades, right-wing radicalism is largely an anti-incumbent vote. The other explanation is that the radical right and the radical left speak to different forms of positional deprivation. Those that feel to be losing out relatively to the poor are more likely to prefer rightist radicalism, while those lagging behind relatively to high-income agents are generally supportive of leftist radicalism, which often comes with a strong pro-redistribution agenda (Burgoon et al 2009). A possible interpretation of my results is that the middle class is fascinated by far-right electoral platforms because they tend to be associated with the protection of acquired rights against the rise of once marginalised groups at the lower end of the end of the income distribution. By contrast, leftist platforms normally promote the acquisition of rights for those that do not have them as yet, which is exactly the dimension that is animating powerful groups' concerns for social-status loss.

4.2 Substantive platforms

It may well be that countries move to the far right for reasons other than the fact that the middle in each country and time resents for having been consumption-deprived relatively to the poor. The results I obtain may be just random and unrelated to the fact that once privileged agents revolt against globalization. To control for this possibility, I relate the "politically salient trade-induced consumption shock" to the content of electoral platforms, where my working assumption is that the consumption deprivation of the middle class relatively to the poorest members of society would spur support in favour of party configurations that engage with anti-globalization sentiments in the form of an agenda of nationalism and autarchy. As dependent variable, I use the now standard net nationalistic autarky score (in logs), which is a composite indicator calculated from the contents of electoral platforms using information from the Comparative Manifesto Project as in Burgoon et al (2019)[23]. The score is then attributed to the political party each respondent feels closest to. I include right self-identification as a control for each respondent's general orientation. The estimation uses OLS, where the variable of interest is consumption deprivation of each quintile relatively to the other four at the country-year level. Table 8 shows that the consumption-related distress of the second and third quintile increases support for electoral platforms that revolve around anti-globalization sentiments (Columns 2 and 3). The same is not true when relative economic distress affects the first quintile (Column 1). In

23 For details on the construction of the indicator, see Burgoon et al (2019).

particular, in countries where the consumption of the third quintile is outpaced by that of the first quintile, a majority of respondents is supporting of nationalist electoral platforms (Column 6). The latter exercise allows me to show that, even when accounting for substantive aspects of the supply of radicalism, my results are largely preserved.

Table 7: Consumption deprivation and net nationalistic autarky conditional on income identity

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	1st q/av.	2nd q/av.	3rd q/av.	4th q/av.	5th q/av.	3rd/1st q	3rd/5th q
Relative deprivation	-0.00955+	0.0180*	0.0380**	0.0123	0.00633	0.0102+	0.000317
	[-1.730]	[1.977]	[3.065]	[0.783]	[0.764]	[1.934]	[0.0274]
Subjective income	0.0495**	0.0487**	0.0509**	0.0493**	0.0480**	0.0499**	0.0483**
	[3.775]	[3.603]	[3.911]	[3.694]	[3.623]	[3.815]	[3.638]
Gini coefficient	-0.00145	0.0336	-0.0154	0.0146	0.0266	-0.00679	0.0323
	[-0.0259]	[0.520]	[-0.255]	[0.237]	[0.467]	[-0.121]	[0.500]
Age (years)	0.00194*	0.00200*	0.00182*	0.00193*	0.00200*	0.00191*	0.00198*
	[2.456]	[2.460]	[2.332]	[2.442]	[2.501]	[2.427]	[2.480]
Education	0.000985	0.00341	0.000895	0.00232	0.00227	0.000823	0.00306
	[0.0856]	[0.305]	[0.0771]	[0.201]	[0.199]	[0.0714]	[0.270]
Female	-0.124**	-0.122**	-0.125**	-0.122**	-0.122**	-0.124**	-0.122**
	[-7.020]	[-6.956]	[-7.067]	[-6.936]	[-6.947]	[-7.033]	[-6.919]
Unemployed	-0.0314	-0.0333	-0.0435	-0.0325	-0.0280	-0.0340	-0.0315
	[-0.657]	[-0.679]	[-0.905]	[-0.670]	[-0.584]	[-0.711]	[-0.653]
Religious	0.0410**	0.0404**	0.0409**	0.0407**	0.0407**	0.0410**	0.0405**
	[6.693]	[6.506]	[6.692]	[6.597]	[6.638]	[6.698]	[6.592]
Foreign born	-0.0275	-0.0260	-0.0274	-0.0270	-0.0276	-0.0275	-0.0270
	[-1.024]	[-0.945]	[-1.018]	[-0.986]	[-1.011]	[-1.026]	[-0.980]
Urban	-0.0195	-0.0169	-0.0206	-0.0185	-0.0178	-0.0199	-0.0173
	[-1.069]	[-0.924]	[-1.135]	[-0.997]	[-0.966]	[-1.092]	[-0.943]
Ethnic threat	0.104**	0.105**	0.104**	0.105**	0.105**	0.104**	0.105**
	[15.53]	[15.84]	[15.48]	[15.75]	[15.53]	[15.53]	[15.54]
Country dummies	yes	yes	yes	yes	yes	yes	yes
Survey time dummies	yes	yes	yes	yes	yes	yes	yes
Pseudo R-squared	0.38	0.38	0.38	0.38	0.38	0.38	0.38
Cluster level	c-d	c-d	c-d	c-d	c-d	c-d	c-d
Observations	30,522	30,522	30,522	30,522	30,522	30,522	30,522
R-squared	0.385	0.383	0.386	0.382	0.382	0.386	0.381

Key: OLS with fixed effects for countries and survey rounds (not shown) and robust-cluster standard errors. Robust z-statistics in brackets. *** p<0.01, ** p<0.05, * p<0.1. All models include design and population weights.

4.3 Robustness tests

To confirm the interpretation of my estimates, I test whether deprivation of the third quintile relative to the fifth delivers different results. Moreover, I explore whether the pattern of support for the radical right changes over time when focusing specifically on the within-country variation in the level of positional deprivation experienced by the first quintile, namely those that should gain the most from the China consumption shock but be hardest it by the China supply shock and the Global Financial crisis. I find that, in countries where the third quintile loses relatively to the richest households, respondents are in fact less likely to support right-wing radicalism. This is true across the board including in the aftermath of the financial crisis (see Table A3 in the Appendix). These results should be interpreted as evidence that the middle class is concerned with deprivation relatively to once marginalised groups rather than in comparison with richest higher-status households. When accounting for the impact of consumption deprivation of the poorest households relatively to everyone else, I find that the consumption-deprived poor have been generally less likely to support the far right in the heyday of globalization, hence during the main part of the China shock. Instead, relative deprivation in the immediate aftermath of financial crisis is related to rising support for the far-right (Table A4 in the Appendix), in line with existing analyses looking at the role of economic insecurity (Guiso et al 2024).[24]

I perform a similar exercise to the one reproduced in Table 6 substituting the pro-poor bias with a crude proxy for the China shock, which is calculated as the percentage increase of Chinese imports over total imports in interaction with the consumption growth of the third quintile. Chinese penetration should largely equate with the size of the pro-poor bias in EA, whose members tend to vertically differentiate production from within, becoming progressively more specialised in the supply of high-quality products, whilst importing cheap tradables. By contrast, the simple share of Chinese over total imports in the case of non-EA does not necessarily imply that low-income agents stand to gain the most because, as shown elsewhere (Heins 2023), these countries import goods whose average price is higher than that of domestically produced goods. If my argument around the differentiation between EA and non-EA countries is correct, Chinese penetration should lead to strong demand for the far-right in EA and be either non-consequential in non-EA or associated with opposite political attitudes. The results largely confirm the hypothesis. It is found indeed that Chinese penetration reduces the distance between the middle class and radical right-wing attitudes and, at the same time, reduces the odds that countries lean towards the left only in the EA (Table A5 in the Appendix), while leading to opposite results when accounting for the overall sample including non-EA members (Table A6 in the Appendix).

Conclusions

A thorough understanding of the political economy of political radicalism requires first an economic model to identify the losers - whether current or prospective - and secondly a sociological mechanism that explains how economic distress manifests itself into political preferences and behaviour. This paper has taken issues with the trade model that has been used thus far to explain the globalization backlash. My starting point is that trade does not only cause labour market shifts, as in the standard Heckscher-Ohlin trade model, but also unequal changes in the cost of living due to non-homothetic preferences. More to the point, in their role of consumers, agents receive benefits from trade that are distributed in a fashion that is different if not opposed to the distribution of trade-induced income effects on the supply side. Low-income groups might lose income but they gain unambiguously in purchasing power. Middle-income groups are potentially unaffected by the China supply-side shock but lose in consumption terms relatively to low-income agents. My analysis further reveals that their relative disadvantage becomes politically consequential and drives a general rightward drift

24 Given that the financial crisis was associated with a steep rise in unemployment amongst the low-skilled, it is likely that the consumption deprivation of the lowest quintiles during this period is endogenous to changes in income.

in politics as well as its radicalization in the form of support for far-right parties, whose substantive platforms advocate nationalistic autarky. This is especially evident during the main part of the China shock from the early 2000s to 2006 and mostly in the EA, whose members tend to import low-quality products that are not supplied domestically. The mechanism that explains why the European middle class turns to the (far) right revolves around the feeling of being “threatened” by a rising group of consumers at the left hand of the income distribution, which comes with a perceived risk of societal decline. My results provide important insights. First, it is desirable to distinguish between consumption- and income-related misfortunes as they lead to differential preference formation both in ideological and substantive terms. Second, as alluded to earlier, it is the consumption deprivation of specific income categories that drives my overall results and, more generally, shapes political preferences at the country level especially if the relevant group coincides with the median voter. To wit, recent political shifts are part of the revolt against trade globalization, as now in the received wisdom, but my results would indicate that the backlash is coming from groups that have not been traditionally identified as the main losers from globalization. I believe this approach is able to explain various puzzles in the literature about trade and politics, opening up new fruitful avenues for future research.

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Appendix

Table A1: Classification of European political parties

Country	Radical left party	Radical right party
Austria		Freedom Party of Austria (FPÖ), Alliance for the Future of Austria (BZÖ)
Belgium		Flemish Interest (VB) National Front Belgium (FNb)
Czech Repub.	Commun. Party Boh.&Mor.(KSCM)	Dawn (Úsvit Tomia Okamury)
Denmark	Socialist People's Party (SF) Red-Green Alliance (EL)	Danish People's Party (DF)
Finland	Left Alliance (VAS)	Finns Party/True Fins (PS)
France	French Communist Party (PCF) Worker's Struggle (LO) Revolut. Commun. League (LCR)	National Front (FN) National Republican Movement (MNR)
Germany	The Left (Die Linke) Party of Democratic Socialism (PDS)	The Republicans (REP) National Democratic Party (NPD)
Greece	Communist Party of Greece (KKE) Coalition of the Left (SYN/Syriza)	Popular Orthodox Party (LAOS) Gold.Dawn Golden Dawn
Hungary	Workers Party (WP)	Jobbik (Movement for a Better Hungary)
Ireland	Sinn Fein (SF)	
Italy	Commun. Refoundation Party (PRC) Communists (Comunisti)	Northern League (LN) National Alliance (AN) Tricolor Flame (FT)
Netherlands	Socialist Party (SP)	List Pim Frotuyn (LPF) Party of Freedom (PVV)
Poland		League of Polish Families (LPR) Congress of the New Right (KNP)
Slovakia	Communist Party Slovakia (KSS)	Slovak National Party (SNS)
Slovenia	Združena levica (ZL)	Slovene National Party (SNS)
Spain	United Left (IU/Podemos)	
Sweden	Left (V)	Sweden Democrats (SD)
United Kingdom		British National Party (BNP) UK Independence Party (UKIP)

Key: The classification is taken from Burgoon et al (2019).

Table A2: ESS-HBS data availability

	HBS 1994-1999	HBS 1994-2000	HBS 2005-2010	HBS 2010-2015	ESS 1 (2002)	ESS 3 (2006)	ESS 5 (2010)	ESS 7 (2014)	ESS 1 (2002) - HBS 1994-1999	ESS 3 (2006) - HBS 1999-2005	ESS 5 (2010) - HBS 2005-2010	ESS 7 (2014) - HBS 2010-2015
AT	v	v	v	v	v	v	v	v	v	v	v	v
BE	v	v	v	v	v	v	v	v	v	v	v	v
BG	x	x	v	v	x	v	v	x	x	x	v	x
CY	x	x	v	v	x	v	v	x	x	x	v	x
CZ	x	x	v	v	v	x	v	v	x	x	v	v
DE	v	v	v	v	v	v	v	v	v	v	v	v
DK	v	v	v	x	v	v	v	v	v	v	v	x
EE	x	x	v	v	x	v	v	v	x	x	v	v
ES	v	v	v	v	v	v	v	v	v	v	v	v
FI	v	v	v	v	v	v	v	v	v	v	v	v
FR	v	v	v	x	v	v	v	v	v	v	x	v
GB	v	v	v	v	v	v	v	v	v	v	v	v
GR	v	v	v	v	v	x	v	x	v	x	v	x
HR	x	x	v	v	x	x	v	x	x	x	v	x
HU	x	x	v	v	v	v	v	v	x	x	v	v
IE	v	v	v	v	v	v	v	v	v	v	v	v
IT	v	v	x	x	v	x	x	x	v	x	x	x
LT	x	x	v	v	x	x	v	v	x	x	v	v
LU	v	v	x	x	v	x	x	x	v	v	x	x
LV	x	x	v	v	x	v	x	x	x	x	x	x
MT	x	x	v	v	x	x	x	x	x	x	x	x
NL	v	v	v	v	v	v	v	v	v	v	v	v
PL	x	x	v	v	v	v	v	v	v	v	v	v
PT	v	v	v	v	v	v	v	v	v	v	v	v
RO	x	x	v	v	x	v	x	x	x	x	x	x
SE	v	v	v	v	v	v	v	v	v	v	v	v
SI	x	x	v	v	v	v	v	v	x	x	v	v
SK	x	x	v	v	x	v	v	x	x	x	v	x

Table A3: Consumption deprivation (3rd q/5th q) and support for radical right

	(1)	(2)	(3)	(4)	(5)
	2002-2014	2002	2006	2010	2014
Consumption deprivation (3rd q/5th q)	-0.129***	-0.0283**	-0.208***	0.0208	-0.531***
	[-5.959]	[-2.435]	[-3.680]	[0.308]	[-8.414]
Gini coefficient	0.269***	-0.0474	-0.0210	0.109	0.236***
	[2.740]	[-0.948]	[-0.437]	[0.183]	[5.966]
Age (years)	-0.0158***	-0.0185***	-0.0315***	-0.00816	-0.0133***
	[-6.904]	[-3.918]	[-5.758]	[-1.512]	[-3.665]
Education	-0.135***	-0.0771	-0.152**	-0.0822	-0.190**
	[-3.414]	[-1.017]	[-2.413]	[-1.260]	[-2.436]
Female	-0.440***	-0.627***	-0.589***	-0.318**	-0.375***
	[-6.186]	[-4.942]	[-3.011]	[-2.312]	[-3.035]
Unemployed	0.332*	0.126	1.085***	0.234	0.120
	[1.656]	[0.552]	[3.015]	[0.618]	[0.319]
Religious	-0.00526	-0.00426	-0.0407	0.0215	-0.00314
	[-0.284]	[-0.122]	[-1.258]	[0.706]	[-0.107]
Foreign born	-0.180*	-0.590**	0.244	-0.178	-0.115
	[-1.700]	[-1.960]	[1.094]	[-0.746]	[-0.799]
Urban	0.0367	0.514***	0.209	-0.142	-0.270**
	[0.367]	[2.796]	[1.106]	[-0.766]	[-2.331]
Ethnic threat	0.406***	0.370***	0.470***	0.226***	0.483***
	[14.02]	[4.674]	[11.30]	[3.375]	[12.95]
Country dummies	yes	yes	yes	yes	yes
Survey time dummies	yes	no	no	no	no
Pseudo R-squared	0.25	0.19	0.19	0.13	0.32
Cluster level	c-d	c-d	c-d	c-d	c-d
Observations	36,081	7,278	8,370	8,381	7,521

Key: Logistic regression estimates (binary dependent variable, 0-1). All models with fixed effects for countries and survey rounds (not shown) and robust-cluster standard errors. Robust z-statistics in brackets. *** p<0.01, ** p<0.05, * p<0.1. All models include design and population weights.

Table A4: Consumption deprivation (1st q/av.) and support for radical right

	(1)	(2)	(3)	(4)	(5)
	2002-2014	2002	2006	2010	2014
Consumption deprivation (1st q/ av.)	-0.00235	-0.0110**	-0.308***	-0.00317	0.180***
	[-0.351]	[-2.435]	[-3.680]	[-0.308]	[8.414]
Gini coefficient	0.343***	0.00940	-0.0924**	-0.144	0.718***
	[3.582]	[0.327]	[-2.348]	[-0.438]	[12.73]
Age (years)	-0.0160***	-0.0185***	-0.0315***	-0.00816	-0.0133***
	[-7.032]	[-3.918]	[-5.758]	[-1.512]	[-3.665]
Education	-0.136***	-0.0771	-0.152**	-0.0822	-0.190**
	[-3.395]	[-1.017]	[-2.413]	[-1.260]	[-2.436]
Female	-0.437***	-0.627***	-0.589***	-0.318**	-0.375***
	[-6.147]	[-4.942]	[-3.011]	[-2.312]	[-3.035]
Unemployed	0.300	0.126	1.085***	0.234	0.120
	[1.536]	[0.552]	[3.015]	[0.618]	[0.319]
Religious	-0.00570	-0.00426	-0.0407	0.0215	-0.00314
	[-0.311]	[-0.122]	[-1.258]	[0.706]	[-0.107]
Foreign born	-0.179*	-0.590**	0.244	-0.178	-0.115
	[-1.684]	[-1.960]	[1.094]	[-0.746]	[-0.799]
Urban	0.0501	0.514***	0.209	-0.142	-0.270**
	[0.497]	[2.796]	[1.106]	[-0.766]	[-2.331]
Ethnic threat	0.402***	0.370***	0.470***	0.226***	0.483***
	[13.88]	[4.674]	[11.30]	[3.375]	[12.95]
Country dummies	yes	yes	yes	yes	yes
Survey time dummies	yes	no	no	no	no
Pseudo R-squared	0.24	0.19	0.19	0.13	0.32
Cluster level	c-d	c-d	c-d	c-d	c-d
Observations	36,081	7,278	8,370	8,381	7,521

Key: Logistic regression estimates (binary dependent variable, 0-1). All models with fixed effects for countries and survey rounds (not shown) and robust-cluster standard errors. Robust z-statistics in brackets. *** p<0.01, ** p<0.05, * p<0.1. All models include design and population weights.

Table A5: EA – Chinese penetration and support for radical parties

	(1) Radical	(2) Radical right	(3) Radical left
Chinese penetration	0.00384	-0.0820***	0.0157***
	[0.832]	[-2.650]	[3.993]
Absolute consumption (3rd q)	0.00473	-0.294*	0.0243+
	[0.290]	[-1.839]	[1.796]
Chinese penetration *x* absolute consumption	-0.000372	0.00586***	-0.00101***
	[-1.574]	[3.673]	[-4.925]
Gini coefficient	0.455**	0.926***	0.229
	[2.443]	[3.081]	[1.254]
Age (years)	-0.0120***	-0.0154***	-0.00973***
	[-5.141]	[-3.773]	[-3.400]
Education	-0.0290	-0.181***	0.0215
	[-0.873]	[-2.998]	[0.567]
Female	-0.261***	-0.264**	-0.243***
	[-4.665]	[-2.204]	[-3.658]
Unemployed	0.324***	0.523**	0.301**
	[2.851]	[2.311]	[2.367]
Religious	-0.185***	-0.0227	-0.235***
	[-9.393]	[-0.951]	[-13.16]
Foreign born	-0.0411	-0.311*	0.0460
	[-0.458]	[-1.664]	[0.377]
Urban	0.199**	-0.0414	0.282***
	[2.405]	[-0.307]	[2.858]
Ethnic threat	0.0721**	0.386***	-0.0403
	[2.375]	[8.082]	[-1.162]
Country dummies	yes	yes	yes
Survey time dummies	yes	yes	yes
Pseudo R-squared	0.11	0.18	0.13
Cluster level	c-d	c-d	c-d
Observations	17,269	12,743	13,953

Key: Logistic regression estimates (binary dependent variable, 0-1). All models with fixed effects for countries and survey rounds (not shown) and robust-cluster standard errors. Robust z-statistics in brackets. *** p<0.01, ** p<0.05, * p<0.1. All models include design and population weights.

Table A6: Chinese penetration and support for radical parties

	(1) Radical	(2) Radical right	(3) Radical left
Chinese penetration	0.0254***	0.0447***	0.00155
	[7.852]	[7.255]	[0.370]
Absolute consumption (3rd q)	0.0678***	0.427***	-0.0386***
	[5.777]	[8.661]	[-2.961]
Chinese penetration *x* absolute consumption	-0.000710***	-0.00274***	8.66e-05
	[-8.725]	[-9.964]	[0.792]
Gini coefficient	0.818***	0.0168	0.202
	[5.490]	[0.0836]	[1.112]
Age (years)	-0.00900***	-0.0119***	-0.00744***
	[-4.359]	[-4.026]	[-2.698]
Education	-0.0498*	-0.152***	0.0225
	[-1.685]	[-2.867]	[0.644]
Female	-0.252***	-0.342***	-0.186***
	[-4.883]	[-3.695]	[-2.878]
Unemployed	0.230*	0.165	0.323***
	[1.780]	[0.572]	[2.689]
Religious	-0.138***	0.00396	-0.226***
	[-6.906]	[0.181]	[-13.09]
Foreign born	-0.0146	-0.142	0.0745
	[-0.190]	[-1.123]	[0.651]
Urban	0.0783	-0.205**	0.259***
	[1.045]	[-1.985]	[2.809]
Ethnic threat	0.159***	0.411***	-0.0397
	[5.313]	[11.60]	[-1.232]
Country dummies	yes	yes	yes
Survey time dummies	yes	yes	yes
Pseudo R-squared	0.15	0.32	0.12
Cluster level	c-d	c-d	c-d
Observations	22,701	18,175	16,340

Key: Logistic regression estimates (binary dependent variable, 0-1). All models with fixed effects for countries and survey rounds (not shown) and robust-cluster standard errors. Robust z-statistics in brackets. *** p<0.01, ** p<0.05, * p<0.1. All models include design and population weights.

Authors

Benedicta Marzinotto

Robert Schuman Centre, European University Institute (EUI)

Johns Hopkins School of Advanced International Studies (SAIS)

Department of Economics and Statistics, University of Udine

benedicta.marzinotto@uniud.it