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Are Climate Change Policies Politically Costly?

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Abstract

Are policies designed to avert climate change (Climate Change Policies, or CCPs) politically costly? Using data on governmental popular support and the OECD's Environmental Stringency Index, we find that CCPs are not necessarily politically costly: policy design matters. First, only market-based CCPs (such as emission taxes) generate negative effects on popular support. Second, the effects are muted in countries where non-green (dirty) energy is a relatively small input into production. Third, political costs are not significant when CCPs are implemented during periods of low oil prices, generous social insurance and low inequality.

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I. INTRODUCTION

There are few issues that have sparked more attention across the globe than how to avoid the environmental and human catastrophe that climate change is inflicting on our planet. In the period leading up to the pandemic, what started as a school strike by a sixteen-year old in Sweden has spurred unprecedented advocacy by an entire generation of policies to fundamentally change course in the pursuit of far-reaching climate changes policies (CCPs). But even in the wake of massive public protests and an ambitious agenda since the 2015 Paris Agreement, the hesitancy of politicians is remarkable. This is disturbing.

Why? First, because the evidence on the economics of climate change is very clear that the long-term costs of unmitigated global warming will outweigh the short-term adjustment costs from mitigation (Stern, 2006). Although economists expect that poor countries will be hit the most (Diffenbaugh and Burke, 2019; Dell et al., 2012), industrial countries are estimated to suffer losses of ½-2 percent of GDP if global temperatures rise 2-4 degrees Celsius by 2100 (Hsiang et al., 2017). Kahn et al. (2019) project that unmitigated climate change will reduce global real GDP per capita by more than 7 percent by the end of the century. Burke et al. (2015) shows that economic costs are increasing in time because greater accumulation of emissions in the atmosphere requires greater interventions to stabilize global temperatures in the future; moreover, unmitigated global warming would likely lead to global GDP declining by up to one quarter by 2100 (IMF, 2019).

Second, the politics of climate change has also become more favorable to CCPs. In the last decades, the number of green parties has increased in many advanced democracies (Dolezal, 2010). Because green parties (which possess issue ownership on CCPs) are a threat to many non-green parties, it seems likely that mainstream parties would adopt green issues in their manifestos (Meguid, 2005; Spoon et al., 2014; Green-Pedersen and Mortensen, 2010). In addition, with median voters greener over time, there is a re-positioning of mainstream parties in this same direction (Ezrow et al., 2010; Adams et al., 2006). Thus, from a political-party competition perspective, the growing salience of green parties in the landscape should make the adoption of CCPs more likely.

Third, public concerns about climate change have increased in many industrial countries (PEW Research Center, 2019). Among other reasons, this is driven by the increasing number of extreme weather phenomena and natural disasters in the last decade

(EMDAT, 2020) which make people more risk averse about climate change (Herrnstadt and Muehlegger, 2014; Bird et al., 2014; Welsch and Biermann, 2014). Social instability, global protests, rising environmental awareness and the rise of pro-environmental NGOs are also bringing green issues to the forefront of the political agenda (Hsiang and Burke, 2014; Liu et al., 2011; Steves and Teytelboym, 2013).

Despite these push factors, governments remain hesitant despite the availability of effective mitigation instruments (Weitzman, 1974; McKibbin and Wilcoxen, 2002; Aldy et al., 2003; Li and Lin, 2013; Carl and Fedor, 2016; Akerlof et al., 2019). While political costs of reform is a mainstay of the political-economy literature (Alesina and Drazen, 1991; Drazen and Grilli, 1993; Persson and Tabellini, 2000; Tsebelis, 2002; Alesina et al., 2006; Alesina et al. 2020a,b), the tie-in with climate policies has been less systematically investigated.[2] With the present paper, we aim to fill this void.

Our paper contributes to the literature on the political economy of CCPs in several aspects. First, we estimate the average effect of CCPs on support for the government implementing them (governmental support for short)—where CCPs are proxied by the OECD's Environmental Policy Stringency (EPS) indicators (Botta and Kozluk, 2014) and governmental support is proxied by the "Index of Popular Support" produced by the International Country Risk Guide (ICRG) from public opinion polls in a sample of 34 advanced and emerging market economies for the period 2001–2015. Our results show that CCPs reduce popular support on average, but this mainly reflects the impact of market-based policy measures, such as emission taxes, rather than non-market-based measures, such as emission limits. This result strikes us as central given that economists see Pigouvian taxation as the first-best corrective tool for carbon emissions (Jenkins, 2014; Goulder and Parry, 2008; Goulder et al., 2019; Stiglitz, 2019; IMF, 2019). Although often second-best options,

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[2] While some studies focus on the political economy of CCPs, empirical analyses testing for political costs are missing from the literature. For instance, Savocool et al. (2015) use a case study approach covering eight countries and discuss how economic, social and political trade-offs led to insufficient climate change adaptation efforts in the past. In a similar vein, Klenert et al. (2018a) review the economic and political science literature on the obstacles to carbon pricing and argue for revenue recycling mechanisms to reduce distributional concerns. These notions are in line with the earlier theoretical model in Klenert et al. (20158b), in which the authors formally show how public revenues – raised through emission taxation – can be re-directed to adversely-affected households and firms and in this way are able to reduce the efficiency-equity trade-off.

nonmarket-based measures are still efficient tools to combat climate change (IMF, 2019a) while being more politically viable (IMF, 2019).

Second, we examine whether the political consequences of CCPs vary depending on a number of mediating factors. We show that political costs are larger: (i) in times of high fuel prices; (ii) for countries with a high input share of dirty energy in mining and production; (iii) when CCPs are adopted when inequality is high; and (iv) when social expenditures do not compensate the losers from CCPs. Remarkably, in the counterfactual to all of these situations, the political cost of CCPs is not statistically different from zero.

Third, we propose an identification strategy that enables us to confirm our baseline estimates and draw causal conclusions. Beyond omitted variable bias, reverse causality is the key methodological problem in attempting to link policy change with popular support outcomes. On the one hand, the adoption of CCPs requires political support (Alesina et al. 2020) which would imply a positive effect of our dependent variable on our explanatory variable.[3] In this case, the OLS estimates would under-estimate the causal effect. On the other hand, an unpopular government may implement CCPs because it has nothing to lose. Additionally, events such as natural disasters may shift policy preferences and heighten incentives for governments to adopt stricter CCPs. In both cases, OLS estimates would overstate the political costs of CCPs.[4]

To address these problems, we use an Instrumental Variable (IV) approach that exploits cross-sectional variation in a country's likelihood to implement CCPs. To do this, we use an interaction between a time-varying global and a country-based geographical term. With the global term, we exploit the time variation in implementing CCPs through a variable that proxies for the "policy pressure" that climate change induces at the global level. Here, we use the occurrence and damage of global extreme weather events, which can be assumed to be exogenous with respect to country's levels of government support. The country-specific term consists of geographical indicators which determine the vulnerability of a country towards climate change. For instance, we use the length of the coastline as a proxy for vulnerability to

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[3] Alternatively, governments with weak popular support may delay the implementation of structural policies because of the fear of losing elections.

[4] There can also be measurement error on both sides which would lead to standard attenuation bias.

rising sea levels or the urban extent in coastal zones. It is reasonable to assume that such geographical indicators are randomly distributed and do not affect general levels of government support. The analysis also controls for country and time fixed effects and can therefore be seen as a *differences-in-differences* approach. Overall, our results show that endogeneity in the OLS estimator indeed is likely to result in underestimation of the true causal effect. These results are robust to various instrument and validity tests. Since establishing causality is a challenging undertaking in macroeconomic setups (Nakamura and Steinson, 2018), our IV approach can be a promising avenue for future analysis studying the effect of CCPs.

The paper is structured as follows. In the next section, we review the literature on CCPs, paying attention to the allocation of costs and benefits from stricter environmental regulation, the challenges of international agreements, and strategies on the mitigation of political costs. In section III, we present the data, outline our empirical approach and discuss the baseline results. Section IV presents the Instrumental Variable approach and results. Section V concludes with some implications for policymakers in advanced economies.

II. THE POLITICAL ECONOMY OF CLIMATE CHANGE POLICIES

There are parallels between the literature on the political economy of structural reforms and the political economy of environmental policies. Both types of reform are subject to resistance and a lack of political will despite the fact that they are welfare maximizing in the long run. In seminal contributions by Alesina and Drazen (1991), Drazen and Grilli (1993), Persson and Tabellini (2000) and Tsebellis (2002), the reform process is modelled as a function of veto players who are required to distribute the benefits and costs in a way that satisfies all of them.[5]

This approach applies also to environmental policies. On the one hand, short-term adjustments costs for the economy from CCPs are significant, while significant benefits materialize only in the very long run.[6] What constrains the political will even further is that the benefits of CCPs are not directly observable because they materialize in the absence of

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[5] Alesina et al. (2020a) find that domestic (product and labor market) and external (trade and capital) structural reforms tend to reduce the vote share of the incumbent's party, especially when implemented close to elections.

[6] In the shorter term, there could be some benefits from carbon pricing including carbon tax revenues and reductions in air pollution and associated mortality.

environmental damage. On the other hand, CCPs put a large burden of the adjustment on a few stakeholders, while the benefits from climate protection are distributed widely (Stockes, 2015; Tvinnereim and Iversflaten, 2016). In addition, the regressive nature of CCPs in many countries heighten inequality concerns (Goulder et al., 2019; Stiglitz, 2019; Rojas-Vallejos and Lastuka, 2020). This is especially so for market-based instruments such as emission taxes. Concentration of costs creates strong interest groups against change, while benefit dispersion diffuses support.

Beyond national considerations, environmental protection has always been hostage to the political agenda at the international level. Enforceable solutions are hampered by a dearth of sanctioning mechanisms, and CCP agreements are exposed to ex-post deviations by free-riding countries. A free-riding country gains competitive advantage against a complying country. For instance, the adoption of stricter climate change policies can incentivize multinational firms to relocate their production processes to regions with lower levels of environmental (Koch and Basse Mama, 2019). This constellation of national interests can induce a race-to-the-bottom. In addition, climate change is expected to elevate international economic inequalities further. The Center for Global Development (2015) estimates that 79 percent of global carbon emission from 1980 to 2011 have been generated by developed countries, but these countries are not the ones that struggle most from climate change today (Eckstein et al., 2020). Thus, those most responsible for climate change are precisely those who have not borne much of the cost. International agreements have to contend with achieving a political equilibrium by re-ordering the incentive structures of countries in a way that makes pollution economically unattractive.

In general, CCPs and other structural reforms—by changing the *status quo*—are likely to be accompanied by distributional consequences—winners and losers (Ostry et al., 2019; Ostry et al., 2020; Markkanen and Anger-Kraavi, 2019). In the realm of CCPs, such policies will make certain products more costly—either directly through taxation or indirectly through increasing production costs (Rao, 2013). Although price increases for oil and gas prices matter to all economic agents, they tighten the budget constraints especially for the poor. This regressivity is a common argument against CCPs and a source of political opposition (Metcalf, 2009; Habla and Roeder, 2017; Goulder et al., 2019). While certain instruments can be designed in a progressive manner, this is not applicable to all types of measures. Thus, when this is not possible, targeted redistributive policies or timing CCPs to periods when energy costs are low, as now, can help to relax pressures on the budget constraints of the poor, and thereby mitigate political costs.

CCPs have negative side effects for certain industries (Steves and Teytelboym, 2013; Fankhaeser et al., 2008). Although new employment opportunities created by CCPs could mitigate job displacements to some degree (IMF, 2019; IMF, 2020), the adjustment costs would nevertheless be salient in the short run. This is especially the case for the old, and for employees with specialized industry-specific capital or low educational levels (OECD, 2012; Guivarch et al., 2011). Given these considerations, it is not surprising that employees in adversely affected sectors have been vocal against stricter CCPs (Tvinnereim and Iversflaten, 2016). One way to reduce such opposition is for governments to increase targeted assistance programs for certain risk groups (i.e. the unemployed, the elderly), and to devote more resources to training and active labor market programs in sectors that are hurt by CCPs (dirty sectors).

III. EMPIRICAL ANALYSIS

This section examines empirically whether governments ought to fear the adoption of CCPs and, if so, whether the political consequences differ across types of CCP. In a further step, we investigate how political costs can be militated by judicious choice of timing and by complementary measures to mitigate associated rises in inequality.

A. Data

We use the Environmental Policy Stringency (EPS) indicators constructed by the OECD (Botta and Kozluk, 2014). These data are the most comprehensive available source for policy measures across countries (28 OECD and six BRICS countries) and time (1990–2015)[7]. The dataset helpfully provides a breakdown by instrument type. The EPS data will specifically allow us to test the effect of different instruments – scaled from 0 (not stringent at all) to 6 (very stringent) – relative to an overall aggregate index consisting of both market-based and non-market-based measures. Here, market-based measures include instruments such as taxation on emissions, trading schemes and feed-in tariffs, while non-market-based indicators capture legislation on emission limits and R&D subsidies, among others.

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[7] One concern might be that environment legislation is adopted at the supranational level such as the European Union. This would be problematic for our empirical analysis as national governments may then not directly responsible for the stricter environmental regulation. Despite this potential concern, it is noteworthy that substantial cross-country variation exists within the EU and environmental policymaking takes place at the national level.

We use a measure of governmental popular support constructed by the International Country Risk Guide (ICRG, 2020) to gauge political cost. The data cover all countries of interest on an annual basis from 2001 onwards. The support measure is a sub-indicator of the overall government stability score, which assesses the government's ability to carry out its declared program(s), and its ability to stay in office. The indicator ranges between 0 (very high risk) and 4 (very low risk). The annual frequency of the data is key to correctly associate the implementation of CCPs with political outcomes. Finally, the indicator of popular support is strongly correlated with other measures of political strength, such as the vote share of the incumbent party at election time and the probability that the leader of the party gets re-elected (Alesina et al. 2020a). Because of the wide coverage comprising 31 countries over the period 2001–15, the dataset affords the opportunity to assess countries that are, in sum, responsible for 43 percent of total global carbon emissions (Muntean et al., 2018). Tables 1 and 2 present an overview and summary statistics for all variables.

B.   Baseline Model

In the baseline specification, we regress popular support of the incumbent on the change in the overall EPS indicator. We use policy changes (instead of levels) based on our prior about the well-documented status-quo bias of voters and path dependency obstacles towards institutional change (Samuelson and Zeckhauser, 1988; Fernandez and Rodrik, 1991; Pierson, 1996). We test how the resistance of voters towards policy change is a threat to the popularity of the government.

$$y_{i,t} = \beta \Delta EPS_{i,t} + \theta X_{i,t} + \alpha_i + \gamma_t + \epsilon_{i,t} \qquad\qquad (1)$$

where y is the popular support for the incumbent and ΔEPS the change in the EPS indicator. Beyond including country (α_i) and year fixed effects (γ_t) to control for country- and time-specific unobserved factors and limit omitted variable bias, the specification includes various economic, political and demographic indicators X. Specifically, we control for fiscal deficit as percentage of GDP (OECD), the share of the elderly (65+) as percentage of population (OECD), real GDP growth rate (World Bank), unemployment rate (World Bank), consumer

price index (OECD), lagged government's vote share (Alesina et al., 2020a), tax revenue as percentage of GDP (World Bank), and numerous economic reform indicators by Alesina et al. (2020a) covering financial, labor and product market policy domains.

The inclusion of these variables is based on well-established streams in the economic literature related to economic voting, political capabilities and structural reforms. For instance, by controlling for tax revenue, we take account of the empirical finding that voters oppose taxes in general[8] (Alesina et al. 2020b). Beyond these baseline controls, we have tested how the inclusion of further covariates related to political systems changes the effect of our main variables of interest. Our results are robust towards the inclusion of the following variables: ideology of the incumbent, government fractionalization, legislative strength, type of the political system (presidential/assembly/parliamentary), number of years left in current term, margin of majority, Herfindal index of the opposition, democratic accountability index, bureaucratic quality index, civil disorder index, income tax wedge, net household income (level and growth rate) and share of green parties in parliament. While we do not report the single estimations due to space constraints, Table A1 provides model-averaging estimates obtained considering all these variables and using the Weighted Average Linear Squared (WALS) developed by Magnus, Powell, and Prufer (2010), and Bayesian Model Averaging (BMA) developed by Fernandez, Ley and Steel (2001). As the results show, the significant negative effect of CCP is robust to model uncertainty concerns. Since our results are robust to inclusion and exclusion of various sets of controls, we use the specification in Table 3 column 1 as our baseline for the subsequent estimations.

In line with our theoretical expectation, increasing environmental policy stringency has significantly negative and sizeable effects on the popular support for the government (Table 3). A major change in CCPs—equivalent to an increase in the EPS indicator from the 25th to the 75th percentile of the EPS distribution—is found to reduce popular support by about 10 percent. The statistical significance of the results is robust to alterative sets of controls and the magnitude of the coefficients does not change with model specification. Moreover, the cost of implementing stricter CCPs also holds when we use de facto measures of government popularity such the vote share of the incumbent government coalition at the

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[8] The results remain unchanged if we use the change in tax revenue as control variable instead of the levels.

elections as dependent variable (Table A2). In particular, the results, based on much smaller sample of 108 observations, suggest that a major change in CCPs—equivalent to an increase in the EPS indicator from the 25th to the 75th percentile of the EPS distribution—reduces the vote share of the government coalition by more than 7 percentage points.

In line with previous findings of the literature (see Alesina et al. 2020 and references therein), we also find that better economic conditions, higher deficits and lower inflation are associated with higher political support. We find that the level of popular support its typically larger in countries with an older share of the population as age tends to be positively associated with voting for the incumbent at the individual level (De Vries and Giger, 2014).

To understand whether different policy instruments have heterogenous effects on government popular support, we re-estimate the baseline specification for each sub-component of the EPS indicator. The literature portrays taxes on emissions as the most effective tool to reduce global greenhouse gases (Jenkins, 2014; Goulder and Parry, 2008; Goulder et al., 2019; Stiglitz, 2019; IMF, 2019). At the same time, carbon taxes play only a limited role in national environmental legislation (Beiser-McGrath and Bernauer, 2019; Carattini et al., 2018). Possibly different political consequences across types of CCPs thus seems to be of the essence.

The results suggest important differences between market-based and nonmarket-based instruments. Market-based measures, and especially taxes on emissions, are broadly consistent with the baseline results pointing to negative consequences on governmental support (Table 4). In contrast, non-market-based measures are typically not statistically significant (Table 5). One interpretation is that price mark-ups are more visible to consumers than supply limits. The fact that households are constantly confronted by fuel prices—for instance through household energy bills or at the gasoline station—makes it easy for them to trace the price mark-up back to tax rises. In contrast, non-market measures either do not spark nearly the same level of attention or limits translate into price changes only with a lag. Thus, using non-market-based measures—which overall are still efficient ways to reduce carbon emissions (IMF, 2019)—seem to stand a reasonable chance of escaping political blame (Weaver, 1986; Pierson, 1994) and thus of overcoming the political cost of CCPs.

An exception is the EPS on diesel. While diesel tax changes do not have a significant effect on popular support (Table 4), nonmarket-based measures on diesel do (Table 5). This

could be because such measures are visible inasmuch as they prevent diesel vehicles from entering inner-city areas. A price mark-up on diesel could be less politically costly owing to lower price elasticity for this fuel relative to others (Dahl, 2012; Winebrake et al., 2015).

C. Mediating factors

What factors might mediate political costs? In this section, we test four key channels that have widely been discussed in the literature. First, global energy prices (FRED, 2020a; FRED, 2020b) might play an essential role in how citizens perceive stricter climate change legislation. This is because environmental policy measures create economic costs through price or quantity changes which voters have to absorb. Second, because CCPs impact especially emission-intensive sectors (Fankhaeser et al., 2008), we would expect that countries with a large input share of dirty-energy would face higher political obstacles than others. Here, we proxy industrial vulnerability to CCPs by the value-added share of dirty-energy mining industries (OECD, 2018). Third, we look at whether CCPs are politically costly particularly at times of high income inequality, using the Gini and income share data as measures (OECD, 2017; WID, 2020). Fourth, we examine whether social insurance to the losers from CCP tightening helps mitigate political costs, proxying such compensation by the generosity of the social protection system measured as social expenditure relative to GDP (Adema et al., 2011). Our specification for examining these channels is:

$$y_{i,t} = F(z_{i,t}) * \left[\beta^L \Delta EPS_{i,t} + \theta^L X_{i,t} \right] + (1 - F(z_{i,t})) * \left[\beta^H \Delta EPS_{i,t} + \theta^H X_{i,t} \right] + \alpha_i + \gamma_t + \epsilon_{i,t}$$

with $F(z_{i,t}) = \dfrac{e^{-1.5*z_{i,t}}}{1+ e^{-1.5*z_{i,t}}}$ $\qquad\qquad\qquad\qquad\qquad$ (2)

where z is the z-score (normalized to have zero mean and unit variance) of the following variables (*M*): (i) oil and gas price; (ii) value added share of dirty-energy mining industries (mining and oil); (iii) income shares and inequality measures; and (iv) social expenditure in

percent of GDP. $z_{i,t} = \left(\frac{\bar{M}_{it} - \bar{M}_i}{\sigma_{Mi}}\right)$.[9] $F(z_{it})$ is the smooth transition function of the variable z. The coefficients β^L and β^H capture the impact of changes in EPS in cases of low M ($F(z_{it}) \approx 1$ when z goes to minus infinity) and high M (($1 - F(z_{it}) \approx 1$ when z goes to plus infinity), respectively.[10]

The use of the smooth transition function is equivalent to the smooth transition autoregressive (STAR) model developed by Granger and Teravirta (1993) to assess non-linear effects above/below a given threshold or regime. The main advantage of this approach relative to estimating SVARs for each regime is that it uses a larger number of observations to compute the effects, improving the stability and precision of the estimates. In addition, this estimation strategy can also handle the potential correlation of the standard errors within countries more easily by clustering at the country level.

Starting with the results for fuel prices, we find that the effect of CCPs on governmental support depends on gasoline and oil market conditions, consistent with both sources of energy being the main sources of household energy consumption in OECD countries (Eurostat, 2017). In times of high oil and gasoline prices, the political damage from CCPs is statistically significantly negative—with coefficients 1.5 to 2 times as large as the direct effect in our baseline (Table 6, column 1-2). In contrast, the effect is not statistically different from zero when EPS rises in times of low global fuel prices. This implies that timing plays a substantial role from a political standpoint: governments can mitigate the political costs to a large extend by passing new legislation when fuel market conditions are favorable. This said, our estimation strategy only allows us to draw conclusions about fuel prices in the year of CCP implementation. Thus, it could be that governments are penalized in subsequent years should fuel prices rise again.

Overall, we interpret these results in two ways. First, adopting new measures when fuel prices are low gives an offsetting effect from more stringent environmental regulation on domestic budget constraints. In such cases, CCPs may fly under the political radar. Second, there

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[9] Thus, we look at within-country variation over time. Since the importance of the dirty energy mining industry and the initial levels of EPS is influenced by CCPs, we address this endogeneity issue by looking at *cross-country* variation—defining $z_{i,t} = \left(\frac{M_{i_{first}} - \bar{M}}{\sigma_M}\right)$. In other words, we use the first value that is available in our dataset.

[10] $F(z_{it})$=0.5 is the cut-off between the weak and strong regimes. The approach is similar to considering a dummy variable that takes value 1 when the variable is about the country-specific mean—that is, $F(z_{it})$>=0.5, and zero otherwise. The difference is that instead of considering two discrete values (0 and 1), the smooth transition approaches allow the regimes to continuously vary between 0 and 1.

is a well-established literature on an inverted U-shaped relation between incomes and support for CCPs—a kind of environmental Kuznets curve (Dinda, 2004): environmental damage increases initially with rising levels of income per capita, but then diminishes. There is also micro evidence that support for CCPs rises with incomes, and so a greater willingness-to-pay for green policies (Carlsson and Johansson-Stenman, 2000; Franzen, 2003; Kotchen et al., 2013; see Torgler et al., 2007). So, it is plausible that relaxing budget constraints through real income gains in times of low fuel prices could increase support for CCPs among voters.

CCPs have differential effects across sectors. This is because environmental policy legislation adversely affects dirty-energy-producing industries more than other sectors (IMF, 2019). Our results show that in countries with a large share of dirty-energy sectors such as Norway and Indonesia, more stringent CCPs garner larger political costs (Table 6, column 3). The effect is not significant for countries where dirty-energy sectors have a very low weight in national input shares such as in Slovenia and Ireland. Employees in dirty-energy sectors will be adversely affected by more stringent EPS and would tend to oppose such policies owing to risks of worker displacement and lost wages. While CCPs would also create new employment opportunities, transition costs could be substantial, and matching skills in declining sectors to new job opportunities difficult. A labor force with lower educational levels, higher age structure and extensive skill-specificity could be difficult to re-allocate to new positions in clean-energy sectors. In addition, we have also tested whether left-wing parties – which have historically a strong commitment to industrial workers (Kono, 2020) – face greater political costs when adopting stricter CCP. Our results provide evidence for this hypothesis (Table A5). Finally, while we cannot draw conclusions about the lobbying power of such industries, it is possible that interest groups favoring emission-intensive sectors could substantially constrain environmental-friendly governments (Jenkins, 2014; Cadoret and Padovano, 2016). In addition to the industrial composition, we also find that the political costs of CCP are higher in countries which started from stricter levels of environmental legislation in the early 2000s. This finding points towards potential non-linear effects in the cost function in a cross-sectional perspective: voters from environmental-friendly countries may penalize the government for stricter CCP when they feel that their economy is losing out against others which do not undertake CCPs.

The political costs of CCPs are larger when changes in EPS are implemented when income inequality is high – measured through market- and net-based Gini coefficients (Table 7).

But the effects are not statistically different from zero when changes in EPS are adopted in times of low inequality. Likewise, when CCPs are implemented in times of high top-income shares, the political damage is statistically significant (Table 7). The effect disappears when top income shares is low. This pattern also holds for bottom income shares, i.e., when these are low, the political cost of CCPs is evident, but not vice-versa. A slight wrinkle is that the bottom 1 percent-income share does not affect political cost, consistent with the limited political representation and participation of the very poor (Lijphart, 1997).

All these results point to the conclusion that the level of inequality at times of stricter environmental policy adoption matter for the size of the political costs. One reason why inequality is an important driver of the political cost of EPS is that CCPs impact prices of basic goods which by their nature are not easily substitutable and have a low-price elasticity of demand (Eitches and Crain, 2016). This means that while all households are affected by CCPs, a price jump for energy products hits households at the lower end of the income distribution more than the upper income deciles (Rao, 2013); there are implications for preferences for more stringent EPS across income levels (Carlsson and Johansson-Stenman, 2000; Franzen, 2003; Kotchen et al., 2013; see Torgler et al., 2007).

In order to test whether direct compensation can reduce political costs, we examine the role of social expenditure (Table 7). We find that lower social expenditure makes EPS rises particular damaging politically, but that the coefficient on the interaction term loses significance when generous social expenditure compensates for the financial costs from CCPs. Our results show that unemployment benefits and active labor market policies are particularly effective tools in this respect, as are to some extent other components of social spending (except social expenditure on housing—see Table A3).

Although we have tested further transmission channels, they do not turn out to mediate the effect of CCPs on popular support. With respect to aggregate economic performance, we do not find a significant difference between low and high growth periods. Using the difference between market and net-based Gini coefficient as a measure for redistribution, we do not find an effect on political cost. Finally, neither the vote share of green parties, fractionalization of the opposition, nor the democratic accountability measure seem to drive the variation in political costs.

IV. INSTRUMENTAL VARIABLE ESTIMATION

Three considerations drive us to move beyond OLS estimation. The first is omitted variable bias (OVB). While our model accounts for unobserved cross-country and -period heterogeneity and controls for numerous determinants of government support in line with the political economy literature, we cannot exclude the possibility of OVB. A second issue is reverse causality, where the bias can actually go in either direction. On the one hand, governments might require political capital—proxied by popular support—to implement reforms. This selection bias implies a positive effect of the dependent variable on our policy variable and biases the OLS estimate towards zero. On the other hand, a government might implement CCPs because its unpopularity implies it has little to lose from implementing reforms. This would imply a negative effect of popular support on CCPs and thus the possibility that the OLS estimate could overestimate the true effect. Third, we cannot exclude measurement error, especially when using policy reform indicators.[11] If the measurement error is uncorrelated with the error term, we lose precision, but if there is correlation, attenuation bias leads to distorted coefficient estimates.

A. IV Estimation of the Baseline Model

Following Nunn and Quian (2014), out IV approach consists of interacting a time-varying global term and a constant country-specific term. The global term we consider approximates the "policy pressure" that climate change induces at the supra-national level. Here, we use indicators on the occurrence and consequences of extreme weather events in a given year. In particular, we look at the number of flood events, the number of major hurricanes in the North Atlantic, the number of people affected by earthquakes and the number of wildfires around the globe per annum. This choice of instrument is consistent with previous evidence that preferences toward CCPs changes after major natural disasters (Bird et al., 2014; Welsch and Biermann, 2014; Latré et al., 2017), with such changes in policy toward more sustainability (Meyer and Schoen, 2017). In addition,

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[11] There are numerous challenges in the measurement of reform indicators (Clinton, 2017. Reducing the dimensionality of legislation to a few numerical indicators is one issue (Benoit and Laver, 2012). Second, the scoring classification is ultimately based on arbitrary ex-ante defined schemes, and the integer-based nature of the indicators reduces substantial variation in the policy levels. In our context, the EPS indicators provide only an ordinal ranking and it is merely assumed that the gaps between the different levels are equal in size; the size of the measurement uncertainty is also unknown given the lack of standard errors around the estimates.

it is reasonable to assume that such global weather events are exogenous and do not drive popular support in a single country.

The country term we consider captures the vulnerability towards climate change which makes the adoption of CCPs more likely. For this purpose, geographical characteristics seem suitable measures since they can reasonably be assumed to be randomly distributed across countries and thus should not drive government support. We consider country-specific measures such as the length of the coastline, the minimum distance of a country's centroid to the coast, the share of urban population in coastal area, and agricultural land (in km^2) per capita.

The theoretical rationale for our pressure-vulnerability instrument is based on the "war of attrition" model proposed by Alesina and Drazen (1991). According to this model, reforms to correct unstainable long-term trends (such as persistent increases in debt or in emissions in our case) are often delayed when they have distributional implications. Reforms occur only when a given group concedes and is "forced" to bear the adjustment. In this model, a crisis (natural disaster) may induce reform "*because the relative cost of fighting the war tilts in favor of concession*" (Alesina et al., 2006, p.5). Overall, we have constructed five instruments consisting of interactions between a time-varying global term and a constant country term.[12] Our IV estimation reads as follows:

$$y_{i,t} = \beta \Delta \widehat{EPS}_{i,t} + \theta X_{i,t} + \alpha_i + \gamma_t + \epsilon_{i,t}$$

with $$\Delta \widehat{EPS}_{i,t} = \vartheta S_{i,t} + \varphi X_{i,t} + \alpha_i + \gamma_t + \upsilon_{i,t}. \tag{3}$$

where *S* is the instrument. The analysis also controls for country and time fixed effects and can therefore be seen as a *differences-in-differences* approach. As the baseline instrument, we use the interaction between the number of floods in the world per annum and a country's coastline length—this instrument exhibited the strongest F-test in the first stage regressions. For all five instruments, the first stage estimates suggest that the instrument is "strong" and statistically

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[12] The instruments are the interaction between: (i) the global number of flood events and the length of a country's coastline (WRI); (ii) the global number of flood events and the length of a country's coastline (FB); (iii) the frequency of major hurricanes in the North Atlantic and distance from the centroid of a country to the nearest coast; (iv) the global number of people affected by earthquakes and the (log) share of urban extent in coastal zone affected by sea level elevation; and (v) the global number of wildfires and a country's agricultural land per capita.

significant. The Kleibergen–Paap rk Wald F statistic—which is equivalent to the F-effective statistic for non-homoskedastic error in case of one endogenous variable and one instrument (Andrews et al., 2019)—is higher than the associated Stock-Yogo critical value.

Our IV results support the findings obtained with OLS (Table 8): the baseline IV result indicates a significant negative effect of EPS changes for all five instruments with similar-sized coefficients. However, the magnitude of the IV coefficient estimates is approximately three times larger than the OLS estimate, which suggests that OLS estimates are biased towards zero. This is an informative outcome given that the direction of bias was ambiguous *ex ante*. We also test the effect of EPS changes with the vote share of the incumbent on the left-hand-side. Although only one out of five instruments is sufficiently strong to draw conclusions, the IV results of this model are in line with the popular support estimation (Table A4).

To test the validity of our instruments, we run several checks. First, we test whether the instruments have a direct effect on popular support by including them stepwise as additional controls in the baseline model (Table 3 column 1). If the coefficients turn out to be significant, one can argue that the instruments are part of the error term and thus do not satisfy the exclusion restriction. As the results in Table 9 show, this is not the case, since all five instruments turn out to be insignificant. Second, instead of regressing popular support on the instruments, we also test directly the association of the baseline residuals and the instrument. Again, the relationship is indistinguishable from zero (Table 10), which supports the validity of our instruments.

Third, it might well be that our global term is associated with other global factors that affect political support. For instance, climate change pressure might be associated with global trends in political conflict or aversion towards globalization. These factors could affect popular support via country-specific factors spuriously related to countries' vulnerability to climate change. To test for this concern, we augment the baseline IV specification by including the interaction between length of the coastline with global factors such as: the overall globalization index from the KOF dataset (see Gygli et al., 2019), which captures an aggregate of economic, social and political dimensions of globalization; the global average of the weighted conflict index from the ICRG dataset; and the global average of the riot index from the ICRG dataset (Table 11 columns I-III).

Another concern is that our country's specific term is associated with country characteristics (such as country size) that may affect political support through factors spuriously

related to global pressure for climate change (such as globalization). For example, globalization may be more relevant for smaller countries and the popular support for globalization policies for smaller economies may be different than for larger economies. To address this concern, we further control for the interaction between the number of flood events and country size and the interaction the global globalization index and country size. The results are similar to, and not statistically different from, the baseline IV results (Table 11 column IV and V, respectively).

B.   IV Estimation of the Interaction Models

We now extend the IV estimation to include the mediating factors described previously. The setup is similar to the regime-dependent IV approach of Ramey and Zubairy (2018):

$$y_{i,t} = \beta \Delta \widehat{EPS^D}_{i,t} + \theta X_{i,t} + \alpha_i + \gamma_t + \epsilon_{i,t}$$

$$\text{with } \Delta \widehat{EPS^D}_{i,t} = \vartheta \left[F^D(Z) * S_{i,t} \right] + \varphi X_{i,t} + \alpha_i + \gamma_t + \upsilon_{i,t} \quad (4)$$

while $D \in \{H, L\}$.

The IV results in the extended model support our previous OLS estimates (Table 12). Adopting CCPs in times of high oil and gas prices raises the political costs (Table 12-1). Likewise, a higher dirty-energy input share raises political costs, which are three times as large as in countries with low dirty energy input shares (Table 12-2). We also find backing for the level effects in the IV estimations: the political costs of CCPs are higher in countries which had already higher levels of environmental protection in the early 2000s. In contrast, we do not find significant coefficients for low oil and gas prices as well as EPS levels, the respective instruments in these models are too weak to draw further conclusions.

The OLS results also hold for the mediating effects of inequality: CCPs in times of high (low) top (bottom) income shares lead to significantly larger political costs (Table 12-3 and Table 12-4). The results for the Gini are in line with this conclusion (Table 12-5), and the instruments are sufficiently strong. Finally, the findings on the mediating effect of social expenditure are also in line with the baseline estimates (Table 12-6): for each of the social

expenditure components (overall, ALMP-related and unemployment benefits), the conclusion is that CCPs are not politically costly when generous social benefits counteract regressivity of stricter environmental legislation.

V. POLICY IMPLICATIONS AND CONCLUSIONS

While our results confirm that CCPs on average may undercut the popularity of governments on average, we show that there are effective mitigating strategies that limit or even remove the detrimental political consequences.

First, the type of CCP matters. While market-based measures (i.e. emission taxes) lower government popularity substantially, the political cost is not significant for non-market-based instruments (e.g., emission limits). Since economists consider market-based instruments to be the most efficient, it is important to internalize when rendering policy advice that such measures might also be relatively costly politically. Since non-market-based measures remain viable instruments to reduce carbon emissions (IMF, 2019), such second-best options from an economic efficiency standpoint should be on the table, lest the perfect become the enemy of the good (Pearce, 2005; Stiglitz, 2019).

Second, timing matters. The consequences of tightening environmental regulation seem to be more visible when the changes adjust energy prices – especially energy and fuel prices for households. Thus, adopting CCPs when world energy prices are low can provide an effective avenue for overcoming political-economy challenges. There is a message for exploiting today's environment of very low world prices in an opportunistic fashion to move ahead with greener policies.

Third, inequality matters for the feasibility of CCPs. The economic burden from CCPs seem likely to be concentrated among certain groups, and especially those with weaker initial conditions and less resilience. Our results show that when CCPs are adopted in times of high inequality, political costs are magnified. Redistributive instruments targeted at the losers, and trampoline policies to allow workers to more easily migrate from losing sectors to growing ones, are a tried and, in our view based on the data, true recipe for overcoming the political fallout from CCPs.

CCPs will be more of a challenge when economies are more reliant on dirty industries (like mining) since the labor force in such sectors is likely to face extensive job risks. Since

these employees may not be easily able to migrate to job opportunities in emerging sectors, permanent unemployment is a concern. Thus, governments where these dirty sectors loom large are likely to be punished at the ballot box. In order to reduce such political costs, policies to enhance sectoral and regional diversification ex-ante are likely to pay dividends.

Climate change will be on the global policy agenda for years to come. As with all policies that generate winners and losers, CCPs require political support to be viable. Rational governments will continue to hesitate and delay because the political damage is palpable. Overcoming this bad equilibrium of inaction is urgent. Our hope is that the evidence and strategies identified in this paper may provide some guidance on ways forward.

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TABLE 1. VARIABLE DESCRIPTION

Variable	Source	Description
D_Popularsupport	ICRG	level of support for the government
D.EPS	OECD	Environmental Policy Stringency Index (aggregated) (first diff)
D.EPS_MKT	OECD	marked-based environmental policy stringency index (first diff)
D.EPS_NMKT	OECD	non-marked-based environmental policy stringency index (first diff)
D.TAXES	OECD	aggregated environmental taxation index (first diff)
D.taxPCA	OECD	PCA (first) of taxation sub-indices (first diff)
D.TAXCO2	OECD	taxation on CO_2 index (first diff)
D.TAXDIESEL	OECD	taxation on Diesel index (first diff)
D.TAXNOX	OECD	taxation on NOx index (first diff)
D.TAXSOX	OECD	taxation on SOx index (first diff)
D.STD	OECD	aggregated limits index (first diff)
D.limitPCA	OECD	first principale component of all limits sub-indices (first diff)
D.ELV_DIESELSO	OECD	emission limits on Diesel index (first diff)
D.ELV_NOX	OECD	emission limits on NOx index (first diff)
D.ELV_PM	OECD	emission limits on PM index (first diff)
D.ELV_SOX	OECD	emission limits on SOx index (first diff)
D.FIT	OECD	aggregated feed-in-tariffs index (first diff)
D.FIT_SOLAR	OECD	feed-in-tariffs for solar index (first diff)
D.FIT_WIND	OECD	feed-in-tariffs for wind index (first diff)
D.TRADESCH	OECD	aggregated trading scheme index (first diff)
D.TRADESCH_CO2	OECD	trading scheme for CO_2 index (first diff)
D.TRADESCH_EEF	OECD	trading scheme for Energy Efficiency Certificate index (first diff)
D.TRADESCH_REC	OECD	trading scheme for Renewable Energy Certificate index (first diff)
D.RD_RE	OECD	R&D subsidies for renewable energy index (first diff)
D.RD_SUB	OECD	aggregated R&D subsidies index (first diff)
deficit	OECD	general government deficit (as % of GDP)
elderly	OECD	share of the elderly (65+) (as % of total population)
gdpr	WB	growth rate of GDP at market prices (LCU)
unemprate	WB	unemployment, total (% of total labour force)
cpi	OECD	consumer price index
voteshare	Alesina	vote share of the main governing party (or coalition)
taxrevue	WB	tax revenue % GDP
finref_index	Alesina	domestic financial liberalization indicator
product_index	Alesina	product market liberalization indicator
current_index	Alesina	current account reform indicator

capital_index	Alesina	capital account reform indicator
labor_index	Alesina	labor market liberalization indicator
oilPrice	FRED	global price for brent crude
gasPrice	FRED	global price of natural gas
top1share	WID	share of pre-tax national income of top 1%
top10share	WID	share of pre-tax national income of top 10%
top20share	WID	share of pre-tax national income of top 20%
bot1share	WID	share of pre-tax national income of bottom 1%
bot10share	WID	share of pre-tax national income of bottom 10%
bot20share	WID	share of pre-tax national income of bottom 20%
giniMRKT	OECD	market-based Gini coefficient
giniNET	OECD	net Gini coefficient
socexp	OECD	total public social expenditure as percentage of GDP
almpExp	OECD	public social expenditure for ALPM as % of GDP
unempExp	OECD	public social expenditure for unemployment as % of GDP
famExp	OECD	public social expenditure for family as % of GDP
healthExp	OECD	public social expenditure for health as % of GDP
houseExp	OECD	public social expenditure for housing as % of GDP
incapExp	OECD	public social expenditure for incapatity as % of GDP
oldExp	OECD	public social expenditure for old age as % of GDP
energyShare	OECD	Share of value added of dirty energy producing/mining industries ("mining and extraction of energy producing products" (D05T06) and "Coke and refined petroleum products" (D19))
cenDist	PSU	distance from centroid of country to nearest coast (km)
coastLengthWRI	WRI	length of coastline
coastLengthFB	Factbook	length of coastline
eventsFlood	EMDAT/OWID	number of floods (global)
majorHurric	NOAA/OWID	frequency of major hurricanes in the North Atlantic
peopleAffectedEQ	EMDAT/OWID	number of people affected by earthquakes (global)
urbanCoastLOG	CIESIN	(log) share of urban extent in coastal zone (km2) affected by sea level elevation
eventsWildfire	EMDAT/OWID	number of wildfires (global)
agriLandPC	Nationmaster	agricultural land (km2) per capita
globalization	Gygli et al.	index of globalization (total) (global average)
riot	ICRG	index of riot (global average)
conflict	ICRG	index of total conflict (global average)
logpop	WB	(log) population (in million)
left	Cruz et al.	Dummy variable for left-wing chief of executive
center	Cruz et al.	Dummy variable for center-wing chief of executive
right	Cruz et al.	Dummy variable for right-wing chief of executive

TABLE 2. SUMMARY STATISTICS

Variable	Mean	Min	Max	SD
D_Popularsupport	2.223	0.500	3.792	0.579
D.EPS	0.106	-1.400	1.112	0.317
D.EPS_MKT	0.076	-2.000	2.083	0.425
D.EPS_NMKT	0.136	-2.000	2.000	0.449
D.TAXES	0.013	-1.000	1.750	0.280
D.taxPCA	0.054	-1.006	2.722	0.335
D.TAXCO2	0.031	-1.000	6.000	0.383
D.TAXDIESEL	-0.071	-4.000	3.000	0.666
D.TAXNOX	0.058	-1.000	6.000	0.513
D.TAXSOX	0.031	-2.000	6.000	0.413
D.STD	0.190	0.000	3.500	0.569
D.limitPCA	0.200	0.000	3.689	0.601
D.ELV_DIESELSO	0.147	0.000	2.000	0.387
D.ELV_NOX	0.220	0.000	5.000	0.906
D.ELV_PM	0.150	0.000	5.000	0.632
D.ELV_SOX	0.242	0.000	6.000	0.951
D.FIT	0.105	-6.000	6.000	1.057
D.FIT_SOLAR	0.118	-6.000	6.000	1.168
D.FIT_WIND	0.092	-6.000	6.000	1.175
D.TRADESCH	0.104	-2.400	2.800	0.780
D.TRADESCH_CO2	0.137	-6.000	6.000	1.878
D.TRADESCH_EEF	0.039	-3.000	4.000	0.317
D.TRADESCH_REC	0.104	-4.000	4.000	0.536
D.RD_RE	0.082	-4.000	4.000	0.708
D.RD_SUB	0.082	-4.000	4.000	0.708
deficit	-2.582	-32.063	18.633	4.613
elderly	14.367	4.099	26.648	4.715
gdpr	2.230	-9.132	25.163	2.977
unemprate	8.181	2.119	33.473	5.311
cpi	0.996	0.476	2.616	0.227
voteshare	-1.854	-29.860	27.230	8.839
taxrevue	18.946	7.919	36.500	6.403
finref_index	0.890	0.500	1.000	0.116
product_index	0.746	0.182	1.000	0.183
current_index	0.934	0.250	1.000	0.149
capital_index	0.897	0.500	1.000	0.165
labor_index	0.622	0.298	1.000	0.143
oilPrice	68.758	24.412	111.960	30.665
gasPrice	5.067	2.614	8.895	1.952
top1share	0.117	0.045	0.296	0.049

top10share	0.354	0.229	0.651	0.084
top20share	0.500	0.377	0.792	0.080
bot1share	0.000	-0.006	0.001	0.001
bot10share	0.010	-0.020	0.027	0.006
bot20share	0.042	0.012	0.078	0.014
giniMRKT	47.337	32.917	69.400	5.679
giniNET	31.890	22.093	59.400	7.952
socexp	20.742	4.793	32.213	5.698
almpExp	0.573	0.005	2.052	0.399
unempExp	0.923	0.000	3.555	0.767
famExp	2.034	0.134	3.997	0.949
healthExp	5.728	2.109	8.869	1.481
houseExp	0.338	0.000	1.745	0.347
incapExp	2.229	0.035	5.540	1.140
oldExp	8.456	0.853	17.126	3.509
energyShare	0.027	0.00043	0.265	0.044
cenDist	250.153	2.944	1712.840	377.420
coastLengthWRI	28511.000	0.000	265523.000	53217.010
coastLengthFB	16980.620	0.000	202080.000	39301.840
eventsFlood	165.867	128.000	226.000	27.711
majorHurric	3.200	0.000	7.000	1.835
peopleAffectedEQ	6370539	786413	46724390	11012820
urbanCoastLOG	0.007	0.000	0.082	0.017
eventsWildfire	10.600	4.000	22.000	4.791
agriLandPC	12.561	0.365	208.211	36.766
globaliz	58.290	53.319	61.753	2.899
riot	0.783	0.032	3.839	1.132
conflict	834.641	179.419	2891.194	822.479
logpop	3.268	0.688	7.152	1.470
left	0.431	0	1	0.496
center	0.093	0	1	0.29
right	0.476	0	1	0.5

TABLE 3. THE EFFECT OF EPS CHANGES ON POPULAR SUPPORT

	(I)	(II)	(III)	(IV)	(V)	(VI)
D.EPS	-0.231***	-0.246***	-0.247***	-0.266***	-0.295***	-0.284***
	(0.079)	(0.079)	(0.080)	(0.086)	(0.084)	(0.088)
deficit	0.039***	0.028*	0.027*	0.025	0.028*	0.029**
	(0.014)	(0.015)	(0.015)	(0.017)	(0.014)	(0.015)
elderly	0.163***	0.177***	0.194***	0.172**	0.196***	0.149**
	(0.051)	(0.055)	(0.063)	(0.069)	(0.074)	(0.066)
gdpr	0.063**	0.047	0.051	0.048	0.041	0.031
	(0.029)	(0.030)	(0.031)	(0.033)	(0.036)	(0.030)
unemprate		-0.035**	-0.033*	-0.028	-0.032	-0.034
		(0.017)	(0.018)	(0.022)	(0.022)	(0.025)
cpi			0.299	0.233	0.084	-0.807*
			(0.309)	(0.314)	(0.299)	(0.457)
voteshare				0.010*	0.009*	0.007
				(0.005)	(0.005)	(0.004)
taxrevenue					-0.006	-0.007
					(0.047)	(0.045)
finref_index						0.153
						(1.620)
product_index						-0.680
						(0.710)
current_index						1.987
						(1.316)
capital_index						0.706
						(1.170)
labor_index						2.260
						(1.998)
Constant	0.659	0.724	0.281	0.628	0.628	-1.750
	(0.624)	(0.673)	(0.924)	(1.010)	(1.508)	(1.731)
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.455	0.466	0.468	0.464	0.493	0.512
Observations	370	370	370	326	321	321

Note: The outcome variable is popular support. Estimated based on $y_{i,t} = \beta \Delta EPS_{i,t} + \theta X_{i,t} + \alpha_i + \gamma_t + \epsilon_{i,t}$. A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard deviations based on robust standard errors clustered at the country level in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

TABLE 4. THE EFFECT OF MARKET-BASED EPS CHANGES ON POPULAR SUPPORT

	(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)
deficit	0.038***	0.039***	0.040***	0.039***	0.039***	0.039***	0.039***
	(0.013)	(0.013)	(0.013)	(0.013)	(0.014)	(0.013)	(0.013)
elderly	0.166***	0.161***	0.162***	0.151***	0.159***	0.150***	0.153***
	(0.053)	(0.051)	(0.051)	(0.051)	(0.052)	(0.050)	(0.051)
gdpr	0.064**	0.058**	0.058**	0.062**	0.060**	0.059**	0.061**
	(0.028)	(0.027)	(0.027)	(0.027)	(0.027)	(0.027)	(0.027)
D.EPS_MKT	-0.153**						
	(0.074)						
D.TAXES		-0.199**					
		(0.092)					
D.taxPCA			-0.202***				
			(0.062)				
D.TAXCO2				-0.071**			
				(0.035)			
D.TAXDIESEL					0.019		
					(0.066)		
D.TAXNOX						-0.090**	
						(0.042)	
D.TAXSOX							-0.196***
							(0.052)
Constant	0.621	0.684	0.684	0.785	0.695	0.839	0.799
	(0.647)	(0.628)	(0.635)	(0.619)	(0.628)	(0.611)	(0.626)
Country FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.452	0.445	0.450	0.443	0.439	0.447	0.450
Observations	370	373	373	378	373	378	378

Note: The outcome variable is popular support. Estimated based on $y_{i,t} = \beta \Delta EPS_{i,t} + \theta X_{i,t} + \alpha_i + \gamma_t + \epsilon_{i,t}$. A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard deviations based on robust standard errors clustered at the country level in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

Table 4 (cont). The effect of market-based EPS changes on popular support

	(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)
deficit	0.039***	0.039***	0.039***	0.038***	0.038***	0.039***	0.039***
	(0.013)	(0.014)	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)
elderly	0.151***	0.147***	0.154***	0.159***	0.159***	0.148***	0.149***
	(0.051)	(0.052)	(0.051)	(0.050)	(0.051)	(0.050)	(0.051)
gdpr	0.061**	0.062**	0.061**	0.066**	0.065**	0.062**	0.061**
	(0.027)	(0.027)	(0.027)	(0.028)	(0.028)	(0.027)	(0.027)
D.FIT	-0.028						
	(0.022)						
D.FIT_SOLAR		-0.022					
		(0.018)					
D.FIT_WIND			-0.024				
			(0.020)				
D.TRADESCH				-0.084			
				(0.060)			
D.TRADESCH_CO2					-0.039		
					(0.026)		
D.TRADESCH_EEF						-0.133**	
						(0.058)	
D.TRADESCH_REC							0.021
							(0.038)
Constant	0.807	0.848	0.765	0.704	0.684	0.845	0.812
	(0.619)	(0.630)	(0.620)	(0.608)	(0.613)	(0.604)	(0.619)
Country FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.444	0.443	0.444	0.455	0.456	0.446	0.442
Observations	378	378	378	374	374	378	378

Note: The outcome variable is popular support. Estimated based on $y_{i,t} = \beta\Delta EPS_{i,t} + \theta X_{i,t} + \alpha_i + \gamma_t + \epsilon_{i,t}$. A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard deviations based on robust standard errors clustered at the country level in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

TABLE 5. THE EFFECT OF NON-MARKET-BASED EPS CHANGES ON POPULAR SUPPORT

	(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)	(VIII)
deficit	0.039***	0.039***	0.039***	0.039***	0.040***	0.039***	0.039***	0.039***
	(0.014)	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)	(0.014)	(0.014)
elderly	0.149***	0.149***	0.145***	0.148***	0.146***	0.150***	0.148***	0.148***
	(0.050)	(0.051)	(0.051)	(0.052)	(0.050)	(0.051)	(0.050)	(0.050)
gdpr	0.061**	0.061**	0.057**	0.061**	0.061**	0.062**	0.062**	0.062**
	(0.027)	(0.027)	(0.025)	(0.027)	(0.027)	(0.027)	(0.027)	(0.027)
D.EPS_NMKT	-0.066							
	(0.048)							
D.STD		-0.004						
		(0.065)						
D.ELV_DIESELSO			-0.206**					
			(0.103)					
D.ELV_NOX				0.014				
				(0.041)				
D.ELV_PM					-0.043			
					(0.041)			
D.ELV_SOX						0.020		
						(0.037)		
D.RD_RE							-0.050*	
							(0.027)	
D.RD_RE								-0.050*
								(0.027)
Constant	0.835	0.819	0.891	0.834	0.856	0.810	0.836	0.836
	(0.600)	(0.611)	(0.618)	(0.627)	(0.604)	(0.618)	(0.604)	(0.604)
Country FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.444	0.441	0.450	0.442	0.443	0.442	0.445	0.445
Observations	378	378	378	378	378	378	378	378

Note: The outcome variable is popular support. Estimated based on $y_{i,t} = \beta \Delta EPS_{i,t} + \theta X_{i,t} + \alpha_i + \gamma_t + \epsilon_{i,t}$. A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard deviations based on robust standard errors clustered at the country level in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

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TABLE 6. THE EFFECT OF EPS CHANGES ON POPULAR SUPPORT MEDIATED BY FUEL PRICES, SHARE OF DIRTY ENERGY INDUSTRIES AND INITIAL LEVEL OF EPS

	(I)	(II)	(II)	(IV)
deficit	0.039***	0.040***	0.040***	0.039***
	(0.014)	(0.013)	(0.009)	(0.009)
elderly	0.165***	0.165***	0.163***	0.160***
	(0.052)	(0.050)	(0.054)	(0.051)
gdpr	0.063**	0.061**	0.063***	0.065***
	(0.028)	(0.028)	(0.019)	(0.019)
D.EPS x low_oilPrice	-0.029			
	(0.272)			
D.EPS x high_oilPrice	-0.320**			
	(0.129)			
D.EPS x low_gasPrice		0.222		
		(0.301)		
D.EPS x high_gasPrice		-0.493***		
		(0.170)		
D.EPS x low_energyShare			-0.200	
			(0.119)	
D.EPS x high_energyShare			-0.264**	
			(0.119)	
D.EPS x low_initialEPS				-0.145
				(0.113)
D.EPS x high_initialEPS				-0.321***
				(0.076)
Constant	0.634	0.628	-0.681	0.000
	(0.633)	(0.614)	(0.896)	(0.000)
Country FE	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes
R²	0.456	0.460	0.337	0.338
Observations	370	370	370	366

Note: The outcome variable is popular support. Estimated based on $y_{i,t} = F(z_{i,t}) * \left[\beta^L \Delta EPS_{i,t} + \theta^L X_{i,t}\right] + (1 - F(z_{i,t})) * \left[\beta^H \Delta EPS_{i,t} + \theta^H X_{i,t}\right] + \alpha_i + \gamma_t + \epsilon_{i,t}$ with $F(z_{i,t}) = \frac{e^{-1.5*z_{i,t}}}{1 + e^{-1.5*z_{i,t}}}$. A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard deviations based on robust standard errors clustered at the country level in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

TABLE 7. THE EFFECT OF EPS CHANGES ON POPULAR SUPPORT MEDIATED BY INCOME SHARES

	(I)	(II)	(III)	(IV)	(V)	(VI)
deficit	0.040***	0.040***	0.040***	0.034***	0.039***	0.040***
	(0.014)	(0.014)	(0.014)	(0.013)	(0.014)	(0.014)
elderly	0.142**	0.144**	0.145**	0.126**	0.143**	0.147**
	(0.058)	(0.059)	(0.059)	(0.059)	(0.058)	(0.058)
gdpr	0.071**	0.071**	0.071**	0.108***	0.070**	0.070**
	(0.035)	(0.035)	(0.035)	(0.034)	(0.035)	(0.035)
D.EPS x low_top1share	0.023					
	(0.192)					
D.EPS x high_top1share	-0.300*					
	(0.156)					
D.EPS x low_top10share		0.033				
		(0.182)				
D.EPS x high_top10share		-0.320**				
		(0.155)				
D.EPS x low_top20share			0.022			
			(0.151)			
D.EPS x high_top20share			-0.320**			
			(0.146)			
D.EPS x low_bot1share				0.010		
				(0.264)		
D.EPS x high_bot1share				-0.254		
				(0.314)		
D.EPS x low_bot10share					-0.304*	
					(0.156)	
D.EPS x high_bot10share					0.021	
					(0.210)	
D.EPS x low_bot20share						-0.352**
						(0.150)
D.EPS x high_bot20share						0.082
						(0.198)
Constant	0.278	0.252	0.227	0.442	0.269	0.208
	(0.907)	(0.912)	(0.918)	(0.928)	(0.912)	(0.907)
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.513	0.513	0.513	0.524	0.513	0.514
Observations	301	301	301	259	301	301

Note: The outcome variable is popular support. Estimated based on $y_{i,t} = F(z_{i,t}) * [\beta^L \Delta EPS_{i,t} + \theta^L X_{i,t}] + (1 - F(z_{i,t})) * [\beta^H \Delta EPS_{i,t} + \theta^H X_{i,t}] + \alpha_i + \gamma_t + \epsilon_{i,t}$ with $F(z_{i,t}) = \frac{e^{-1.5*z_{i,t}}}{1 + e^{-1.5*z_{i,t}}}$. A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard deviations based on robust standard errors clustered at the country level in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

Table 7 (cont.). The effect of EPS changes on popular support mediated by inequality and social expenditure measures

	(I)	(II)	(III)	(IV)	(V)
deficit	0.038***	0.039***	0.033**	0.033**	0.034***
	(0.014)	(0.014)	(0.013)	(0.014)	(0.013)
elderly	0.099	0.103	0.156**	0.156**	0.171**
	(0.094)	(0.093)	(0.062)	(0.062)	(0.070)
gdpr	0.057**	0.057**	0.048**	0.047**	0.050*
	(0.026)	(0.026)	(0.024)	(0.023)	(0.027)
D.EPS x low_giniMRKT	-0.066				
	(0.279)				
D.EPS x high_giniMRKT	-0.380**				
	(0.170)				
D.EPS x low_giniNET		-0.006			
		(0.264)			
D.EPS x high_giniNET		-0.402**			
		(0.162)			
D.EPS x low_socExp			-0.504***		
			(0.167)		
D.EPS x high_socExp			-0.055		
			(0.143)		
D.EPS x low_almpExp				-0.483***	
				(0.158)	
D.EPS x high_almpExp				0.084	
				(0.153)	
D.EPS x low_unempExp					-0.544***
					(0.170)
D.EPS x high_unempExp					0.155
					(0.218)
Constant	1.487	1.445	0.845	0.832	0.666
	(1.149)	(1.139)	(0.734)	(0.743)	(0.827)
Country FE	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes
R²	0.470	0.470	0.461	0.462	0.498
Observations	326	326	334	334	321

Note: The outcome variable is popular support. Estimated based on $y_{i,t} = F(z_{i,t}) * [\beta^L \Delta EPS_{i,t} + \theta^L X_{i,t}] + (1 - F(z_{i,t})) * [\beta^H \Delta EPS_{i,t} + \theta^H X_{i,t}] + \alpha_i + \gamma_t + \epsilon_{i,t}$ with $F(z_{i,t}) = \frac{e^{-1.5*z_{i,t}}}{1 + e^{-1.5*z_{i,t}}}$. A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard deviations based on robust standard errors clustered at the country level in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

TABLE 8. THE EFFECT OF EPS CHANGES ON POPULAR SUPPORT USING INSTRUMENTAL VARIABLE REGRESSIONS

	(I)	(II)	(II)	(IV)	(V)
D.EPS	-0.994***	-0.741**	-0.635*	-1.047**	-0.879**
	(0.386)	(0.361)	(0.341)	(0.533)	(0.377)
Country FE	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes
Instrument:					
- Global term	eventsFlood	eventsFlood	majorHurric	peopleAffectedEQ	eventsWildfire
- Country term	coastLengthWRI	coastLengthFB	cenDist	urbanCoastLOG	agriLandPC
1st-stage coef.	0.003***	0.003***	-7.552***	0.016***	9.58***
	(0.001)	(0.000)	(1.165)	(0.004)	(2.542)
1st-stage F-Stat	30.79	66.47	42.05	14.98	14.21
Observations	370	370	370	370	361

Note: The outcome variable is popular support. Estimated based on $y_{i,t} = \beta \Delta \overline{EPS}_{i,t} + \theta X_{i,t} + \alpha_i + \gamma_t + \epsilon_{i,t}$ with $\Delta \overline{EPS}_{i,t} = \vartheta S_{i,t} + \varphi X_{i,t} + \alpha_i + \gamma_t + \upsilon_{i,t}$. Controls from the baseline specification (table 3 column 1) are included. A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard deviations based on robust standard errors clustered at the country level in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

TABLE 9. VALIDITY OF IV (1). EFFECT OF INSTRUMENTS ON POPULAR SUPPORT

	(I)	(II)	(III)	(IV)	(V)
D.EPS	-0.221***	-0.215***	-0.207***	-0.226***	-0.240***
	(0.078)	(0.079)	(0.080)	(0.079)	(0.081)
deficit	0.040***	0.039***	0.040***	0.040***	0.040***
	(0.014)	(0.014)	(0.014)	(0.014)	(0.014)
elderly	0.160***	0.164***	0.159***	0.163***	0.150***
	(0.050)	(0.053)	(0.050)	(0.051)	(0.048)
gdpr	0.062**	0.063**	0.061**	0.063**	0.064**
	(0.028)	(0.029)	(0.028)	(0.029)	(0.029)
eventsFlood* coastLengthFB	-0.000				
	(0.000)				
majorHurric* cenDist		0.000			
		(0.000)			
eventsFlood* coastLengthWRI			-0.000		
			(0.000)		
peopleAffectEQ* urbanCoastLOG				-0.000	
				(0.000)	
eventsWildfire* agriLandPC					-0.000
					(0.000)
Constant	0.775	0.569	0.988	0.660	0.962*
	(0.601)	(0.636)	(0.603)	(0.627)	(0.566)
Country FE	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes
R^2	0.456	0.456	0.459	0.456	0.458
Observations	370	370	370	370	361

Note: The outcome variable is popular support. Estimated based on $y_{i,t} = \beta \Delta EPS_{i,t} + \theta X_{i,t} + \alpha_i + \gamma_t + \epsilon_{i,t}$. The estimation is the same as in Table 3 column 1 (baseline) with the respective instrument as additional control. A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard deviations based on robust standard errors clustered at the country level in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

TABLE 10. VALIDITY OF IV (2). CORRELATION OF INSTRUMENTS ON BASELINE RESIDUALS

	(I)	(II)	(III)	(IV)	(V)
deficit	0.001	-0.000	0.001	0.000	0.000
	(0.014)	(0.014)	(0.014)	(0.014)	(0.014)
elderly	-0.003	0.000	-0.005	-0.000	-0.013
	(0.050)	(0.053)	(0.050)	(0.051)	(0.048)
gdpr	-0.000	0.000	-0.002	0.000	0.001
	(0.028)	(0.029)	(0.028)	(0.029)	(0.029)
eventsFlood* coastLengthFB	-0.000				
	(0.000)				
majorHurric* cenDist		0.000			
		(0.000)			
eventsFlood* coastLengthWRI			-0.000		
			(0.000)		
peopleAffectEQ* urbanCoastLOG				-0.000	
				(0.000)	
eventsWildfire* agriLandPC					-0.000
					(0.000)
Constant	0.116	-0.090	0.329	0.002	0.303
	(0.601)	(0.636)	(0.603)	(0.627)	(0.566)
Country FE	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes
R^2	0.00201	0.00259	0.00718	0.00141	0.00239
Observations	370	370	370	370	361

Note: The outcome variable is the residuals from Table 3 column 1 based on $y_{i,t} = \beta \Delta EPS_{i,t} + \theta X_{i,t} + \alpha_i + \gamma_t + \epsilon_{i,t}$. We regress these residuals on the baseline controls together with the respect instruments. Standard deviations based on robust standard errors clustered at the country level in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

TABLE 11. VALIDITY OF IV (3). ROBUSTNESS OF IV ESTIMATION WITH ADDITIONAL CONTROLS

	(I)	(II)	(III)	(IV)	(V)
D.EPS	-0.976**	-1.555***	-1.388***	-0.963***	-0.966***
	(0.405)	(0.433)	(0.396)	(0.320)	(0.368)
Country FE	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes
global correlate	globaliz	conflict	riot		
country correlate				logpop	
Additional interaction					globaliz* logpop
1st-stage F-Stat	26.36	23.46	18.38	29.45	28.17
Observations	370	357	357	370	370

Note: The outcome variable is popular support. Estimated based on $y_{i,t} = \beta \Delta EPS_{i,t} + \theta X_{i,t} + \alpha_i + \gamma_t + \epsilon_{i,t}$ with $\Delta \widehat{EPS}_{i,t} = \vartheta S_{i,t} + \varphi X_{i,t} + \alpha_i + \gamma_t + \upsilon_{i,t}$. We use the baseline IV (eventsFlood*coastLengthWRI) for all estimations. In column I-III, an interaction term between coastLengthWRI (country term of baseline instrument) and a global correlate is included, namely the global average of the overall globalization index from KOF dataset (I), the global average of the total conflict index from the ICRG dataset (II), and the global average of the total riot index from the ICRG dataset (III). In column VI, we add an interaction term consisting of eventsFlood (global term of baseline instrument) and a country's population (log). In column V, we add an interaction term consisting of the global globalization index from the KOF dataset and a country's population (log). in int A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard deviations based on robust standard errors clustered at the country level in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

TABLE 12-1. THE EFFECT OF EPS CHANGES ON POPULAR SUPPORT MEDIATED BY FUEL PRICES (IV)

	(I)	(II)	(III)	(IV)
D.EPS_low_oilPrice	2.065 (8.226)			
D.EPS_high_oilPrice		-1.241** (0.608)		
D.EPS_low_gasPrice			-1.808 (4.388)	
D.EPS_high_gasPrice				-1.184* (0.630)
Country FE	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes
1st-stage F-Stat	0.02	12.161	2.52	12.82
Observations	370	370	370	370

Note: The outcome variable is popular support. Controls from the baseline specification (table 3 column 1) are included. The instrument is peopleAffectEQ_urbanCoastLOG. A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard deviations based on robust standard errors clustered at the country level in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

TABLE 12-2. THE EFFECT OF EPS CHANGES ON POPULAR SUPPORT MEDIATED BY SHARE OF DIRTY ENERGY INDUSTRIES AND INITIAL LEVEL OF EPS (IV)

	(I)	(II)	(III)	(IV)
D.EPS_low_energyShare	-1.374* (0.773)			
D.EPS_high_energyShare		-3.675** (1.769)		
D.EPS_low_initialEPS			-3.530 (3.976)	
D.EPS_high_initialEPS				-1.343*** (0.513)
Country FE	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes
1st-stage F-Stat	18.10	9.85	1.631	25.166
Observations	370	370	366	366

Note: The outcome variable is popular support. Controls from the baseline specification (table 3 column 1) are included. The instrument is peopleAffectEQ_urbanCoastLOG. A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard deviations based on robust standard errors clustered at the country level in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively

Table 12-3: The effect of EPS changes on popular support mediated by income shares (IV)

	(I)	(II)	(III)	(IV)	(V)	(VI)
D.EPS_low_top1share	13.098					
	(16.554)					
D.EPS_high_top1share		-3.495***				
		(0.561)				
D.EPS_low_top10share			11.496			
			(14.939)			
D.EPS_high_top10share				-4.482***		
				(0.862)		
D.EPS_low_top20share					7.378	
					(7.012)	
D.EPS_high_top20share						-5.500***
						(1.546)
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes	Yes
1st-stage F-Stat	0.52	73.36	0.43	30.61	0.71	17.65
Observations	301	301	301	301	301	301

Note: The outcome variable is popular support. Controls from the baseline specification (table 3 column 1) are included. A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard deviations based on robust standard errors clustered at the country level in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

TABLE 12-4. THE EFFECT OF EPS CHANGES ON POPULAR SUPPORT MEDIATED BY INCOME SHARES (IV)

	(I)	(II)	(III)	(IV)	(V)	(VI)
D.EPS_low_bot1share	-2.051**					
	(0.837)					
D.EPS_high_bot1share		-0.458				
		(0.577)				
D.EPS_low_bot10share			-2.643**			
			(1.197)			
D.EPS_high_bot10share				0.707		
				(2.046)		
D.EPS_low_bot20share					-4.291***	
					(1.112)	
D.EPS_high_bot20share						4.840
						(3.334)
Observations	259	259	301	301	301	301
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes	Yes
1st-stage F-Stat	8.42	29.84	25.02	1.14	24.82	1.02
Observations	259	259	301	301	301	301

Note: The outcome variable is popular support. Controls from the baseline specification (table 3 column 1) are included. A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard deviations based on robust standard errors clustered at the country level in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

TABLE 12-5. THE EFFECT OF EPS CHANGES ON POPULAR SUPPORT MEDIATED BY INEQUALITY (IV)

	(I)	(II)	(III)	(IV)
D.EPS_low_giniMRKT	1.454			
	(1.876)			
D.EPS_high_giniMRKT		-4.569***		
		(1.757)		
D.EPS_low_giniNET			4.748	
			(3.315)	
D.EPS_high_giniNET				-3.343***
				(0.852)
Country FE	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes
1st-stage F-Stat	9.77	6.93	8.50	29.20
Observations	326	326	326	326

Note: The outcome variable is popular support. Controls from the baseline specification (table 3 column 1) are included. A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard deviations based on robust standard errors clustered at the country level in parentheses . ***,**,* denote significance at 1, 5 and 10 percent, respectively.

TABLE 12-6. THE EFFECT OF EPS CHANGES ON POPULAR SUPPORT MEDIATED BY SOCIAL EXPENDITURE (IV)

	(I)	(II)	(III)	(IV)	(V)	(VI)
D.EPS_low_socexp	-1.492***					
	(0.523)					
D.EPS_high_socexp		-2.694				
		(4.886)				
D.EPS_low_almpExp			-1.863*			
			(1.037)			
D.EPS_high_almpExp				10.620		
				(21.066)		
D.EPS_low_unempExp					-1.860**	
					(0.834)	
D.EPS_high_unempExp						-15.330
						(12.759)
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes	Yes
1st-stage F-Stat	44.32	2.54	76.26	0.38	36.28	1.18
Observations	334	334	334	334	321	321

Note: The outcome variable is popular support. Controls from the baseline specification (table 3 column 1) are included. A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard deviations based on robust standard errors clustered at the country level in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

TABLE A1. THE EFFECT OF EPS CHANGES ON POPULAR SUPPORT – WALS AND BMA

	(I)	(II)	(III)	(IV)
D.EPS	-0.269***	-0.249**	-0.270***	-0.262**
	(0.088)	(0.105)	(0.088)	(0.101)
	[-3.07]	[-2.37]	[-3.05]	[-2.59]
deficit	0.022*	0.001	0.018	0.000
	(0.011)	(0.014)	(0.018)	(0.003)
	[1.96]	[0.07]	[0.97]	[0.09]
elderly	0.143**	0.160*	0.208***	0.126
	(0.060)	(0.095)	(0.078)	(0.130)
	[2.40]	[1.69]	[2.67]	[0.97]
gdpr	0.035**	0.023	0.033	0.002
	(0.016)	(0.020)	(0.028)	(0.008)
	[2.19]	[1.19]	[1.16]	[0.20]
unemprate	-0.027**	-0.050**	-0.031	-0.082***
	(0.012)	(0.019)	(0.024)	(0.017)
	[-2.20]	[-2.59]	[-1.30]	[-4.86]
cpi	-0.417	-0.128	-0.037	-0.012
	(0.370)	(1.441)	(0.203)	(0.243)
	[-1.13]	[-0.09]	[-0.18]	[-0.05]
voteshare		-0.003		-0.000
		(0.005)		(0.001)
		[-0.60]		[-0.11]
taxrevenue	0.004	0.022	0.001	0.000
	(0.026)	(0.041)	(0.009)	(0.007)
	[0.16]	[0.55]	[0.09]	[0.03]
constant	0.759	6.760	0.886	3.982
	(1.355)	(4.769)	(1.165)	(4.329)
	[0.56]	[1.42]	[0.76]	[0.92]
Model Type	WALS	WALS	BMA	BMA
Country FE	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes
Observations	349	232	349	232

Note: The outcome variable is popular support. The estimates are the results of WALS (Column 1 and 2) and BMA regressions (Column 3 and 4). Column 1 and 3 use the control variables from Table 3 Column 6 without vote share (vote share was dropped in order to make the sample as large as possible). Column 2 and 4 use the control variables from Table 3 Column 6 (including vote share) plus the following additional covariates: share of green parties in parliament, partisanship of chief executive of government, government fractionalization index, numbers of years left in current term , margin of majority, Herfindahl index of opposition parties, index of democratic accountability, index of bureaucratic quality, index of civil disorder, index of legislative strength, tax wedge from national accounts, average tax wedge for one-earner married couple, net household income for couple with two children (levels and changes). A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard errors in parentheses, t statistics in brackets. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

TABLE A2. THE EFFECT OF EPS CHANGES ON VOTE SHARE

	(I)	(II)	(III)	(IV)	(V)
D.EPS	-6.908*	-6.901*	-6.368*	-9.042***	-12.341***
	(3.556)	(3.495)	(3.658)	(2.350)	(2.833)
L.voteshare	-0.391***	-0.406***	-0.494***	-0.604***	-0.669***
	(0.086)	(0.084)	(0.074)	(0.074)	(0.082)
deficit	0.686	0.488	0.514	0.605	0.366
	(0.675)	(0.768)	(0.727)	(0.673)	(0.737)
elderly	-0.574	-0.745	0.555	-3.340**	-2.750
	(1.608)	(1.671)	(1.193)	(1.153)	(1.657)
gdpr	1.481*	1.440*	1.585**	1.553**	1.384*
	(0.827)	(0.803)	(0.709)	(0.620)	(0.732)
unemprate		-0.451	-0.210	-1.032**	-1.189**
		(0.429)	(0.384)	(0.392)	(0.523)
cpi			31.725***	19.024	27.858*
			(11.205)	(15.291)	(15.076)
taxrevenue				0.358	0.112
				(0.980)	(1.180)
finref_index					-20.690
					(26.580)
product_index					0.576
					(7.794)
current_index					26.211
					(19.451)
capital_index					-65.451**
					(22.230)
labor_index					-7.841
					(28.372)
Constant	15.837	0.956	-50.243*	0.000	0.000
	(29.826)	(27.186)	(27.857)	(0.000)	(0.000)
Country FE	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes
R2	0.421	0.428	0.498	0.660	0.701
Observations	109	109	109	80	80

Note: The outcome variable is vote share of the incumbent. Estimated based on $y_{i,t} = \beta \Delta EPS_{i,t} + \theta X_{i,t} + \alpha_i + \gamma_t + \epsilon_{i,t}$. Standard deviations based on Driscoll-Kraay standard errors with two lags in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

TABLE A3. THE EFFECT OF EPS CHANGES ON POPULAR SUPPORT MEDIATED BY SOCIAL EXPENDITURE TYPES

	(I)	(II)	(III)	(IV)	(V)
deficit	0.032**	0.034***	0.031**	0.032**	0.032**
	(0.014)	(0.013)	(0.013)	(0.014)	(0.013)
elderly	0.153**	0.155**	0.143**	0.156**	0.155**
	(0.061)	(0.061)	(0.058)	(0.062)	(0.062)
gdprb	0.048**	0.048**	0.050**	0.048**	0.049**
	(0.024)	(0.024)	(0.024)	(0.023)	(0.024)
D.EPS_low_famExp	-0.408**				
	(0.195)				
D.EPS_high_famExp	-0.139				
	(0.136)				
D.EPS_low_healthExp		-0.894***			
		(0.260)			
D.EPS_high_healthExp		0.078			
		(0.147)			
D.EPS_low_houseExp			-0.226		
			(0.169)		
D.EPS_high_houseExp			-0.285		
			(0.192)		
D.EPS_low_incapExp				-0.348***	
				(0.115)	
D.EPS_high_incapExp				-0.145	
				(0.147)	
D.EPS_low_oldExp					-0.380**
					(0.188)
D.EPS_high_oldExp					-0.124
					(0.167)
Constant	0.875	0.849	0.987	0.833	0.849
	(0.735)	(0.734)	(0.693)	(0.737)	(0.739)
Country FE	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes
R^2	0.458	0.471	0.459	0.458	0.458
Observations	334	334	325	334	334

Note: The outcome variable is popular support. Estimated based on $y_{i,t} = F(z_{i,t}) * [\beta^L \Delta EPS_{i,t} + \theta^L X_{i,t}] + (1 - F(z_{i,t})) * [\beta^H \Delta EPS_{i,t} + \theta^H X_{i,t}] + \alpha_i + \gamma_t + \epsilon_{i,t}$. A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard deviations based on robust standard errors clustered at the country level in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

TABLE A4. THE EFFECT OF EPS CHANGES ON VOTE SHARE USING INSTRUMENTAL VARIABLE REGRESSIONS

	(I)	(II)	(II)	(IV)	(V)
D.EPS	-6.944	-15.049*	-7.467	-14.614	-8.173
	(11.215)	(8.404)	(9.378)	(17.845)	(8.903)
Country FE	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes
Instrument:					
- Global term	eventsFlood	eventsFlood	majorHurric	peopleAffectedEQ	eventsWildfire
- Country term	coastLengthWRI	coastLengthFB	cenDist	urbanCoastLOG	agriLandPC
1st-stage coef.	0.0027**	-11.18***	0.0022**	-0.0184	25.943
	(0.001)	(1.950)	(0.001)	(0.034)	(17.215)
1st-stage F-Stat	5.416	32.865	6.997	0.3	2.271
Observations	81	81	81	81	79

Note: The outcome variable is vote share. Estimated based on $y_{i,t} = \beta \widehat{\Delta EPS}_{i,t} + \theta X_{i,t} + \alpha_i + \gamma_t + \epsilon_{i,t}$ with $\Delta \widehat{EPS}_{i,t} = \vartheta S_{i,t} + \varphi X_{i,t} + \alpha_i + \gamma_t + \upsilon_{i,t}$. Controls from the baseline specification (table 3 column 1) are included. Standard deviations based on Driscoll-Kraay standard errors with two lags in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.

TABLE A5. THE EFFECT OF EPS CHANGES ON POPULAR SUPPORT MEDIATED BY IDEOLOGY

	(I)	(II)	(III)	(IV)	(V)	(VI)
D.EPS*left	-0.420**	-0.437**	-0.440**	-0.468**	-0.498**	-0.473**
	(0.185)	(0.179)	(0.180)	(0.207)	(0.210)	(0.208)
D.EPS*center	-0.190	-0.231**	-0.239**	-0.243**	-0.265**	-0.247**
	(0.119)	(0.115)	(0.117)	(0.119)	(0.121)	(0.114)
D.EPS*right	0.175	0.169	0.168	0.140	0.060	0.048
	(0.190)	(0.170)	(0.169)	(0.177)	(0.157)	(0.148)
deficit	0.036**	0.021	0.019	0.017	0.026**	0.027**
	(0.016)	(0.016)	(0.016)	(0.017)	(0.013)	(0.013)
elderly	0.071	0.063	0.081	0.079	0.105	0.097
	(0.097)	(0.093)	(0.094)	(0.096)	(0.096)	(0.093)
gdpr	0.075**	0.056*	0.061*	0.057	0.047	0.037
	(0.031)	(0.033)	(0.035)	(0.037)	(0.040)	(0.034)
unemprate		-0.050***	-0.048***	-0.046**	-0.050***	-0.054**
		(0.016)	(0.018)	(0.019)	(0.018)	(0.024)
cpi			0.501	0.416	0.205	-0.652
			(0.309)	(0.337)	(0.314)	(0.558)
voteshare				0.007	0.006	0.004
				(0.006)	(0.006)	(0.005)
taxrevenue					-0.041	-0.042
					(0.047)	(0.047)
finref_index						0.169
						(1.851)
product_index						-0.542
						(0.719)
current_index						1.828
						(1.556)
capital_index						0.514
						(1.170)
labor_index						2.562
						(2.829)
Constant	1.856	2.270**	1.655	1.777	2.622	-0.278
	(1.184)	(1.124)	(1.210)	(1.230)	(1.713)	(2.527)
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.466	0.484	0.488	0.487	0.521	0.535
Observations	305	305	305	294	289	289

Note: The outcome variable is popular support. Estimated based on $y_{i,t} = \beta \Delta EPS_{i,t} + \theta X_{i,t} + \alpha_i + \gamma_t + \epsilon_{i,t}$. A coefficient of -0.2 is equivalent to a 10 percent decline in popular support from an increase in EPS from the 1st to the 3rd quartile of the EPS distribution. Standard deviations based on robust standard errors clustered at the country level in parentheses. ***,**,* denote significance at 1, 5 and 10 percent, respectively.