Bicameralism and Policy Responsiveness to Public Opinion

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Abstract. Does the organization of the assembly affect whether governments deliver policy that reflects the changing preferences of the public? Cross-national analyses of public opinion and policy outputs, for welfare policies and immigration policies, show that governments respond to shifts in public opinion in systems with a dominant chamber but not where bicameralism is strong. Our theory's emphasis on the distribution of power *between* chambers further explains differences *within* bicameral systems: constraints on policy change mean that responsiveness is weaker where power is equally distributed between chambers but more robust where power is concentrated in the lower house. Evidence from institutional change in Belgium, where the fourth State Reform shifted power away from the senate and disproportionately toward the lower house, provides corroborating evidence that policy becomes more responsive when constitutions concentrate legislative power. Study findings have implications for our understanding of how bicameralism matters for government responsiveness to public opinion.

Verification Materials:

The data and materials required to verify the computational reproducibility of the results, procedures and analyses in this article are available on the *American Journal of Political Science* Dataverse within the Harvard Dataverse Network, at:
https://dataverse.harvard.edu/dataset.xhtml?persistentId=doi:10.7910/DVN/C7A5TN

Shortly after forming a new government in 2014, Italian Prime Minister Matteo Renzi embarked on a series of ambitious reforms. Unlike his predecessors, Renzi was keen to implement several constitutional reforms, the chief among them to abolish Italy's "perfect" form of bicameralism which grants commensurate powers to both houses of parliament in favor of a system which would shift power toward the Chamber of Deputies and away from the Senate. Supporters argued the reform would streamline Italy's sclerotic legislative process and enable the government to respond more quickly to social and economic challenges. In December 2016 the proposal to amend the constitution was put to voters. The referendum was unsuccessful, and Renzi resigned the premiership soon thereafter.

Was the public's distaste for reform justified? On the one hand, perhaps they were keen to limit representatives' power by dividing it more or less equally across two houses. The division of the legislature as an antidote for the concentration of power can be traced back to Madison's assertion in Federalist No. 62 that a second legislative chamber is essential to provide a salutary check on the government. On the other hand, horizontal diffusion of legislative powers, particularly when divided equally between chambers, may be a drag on legislative action. As such, Italy's perfect bicameralism may be counterproductive in situations requiring substantial reform.[1]

We examine the impact of legislative power-sharing on government responsiveness to the public. Responsiveness is a central feature of liberal democracy (Dahl 1956; Erikson, MacKuen,

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[1] It also could be argued that the referendum was less a vote against reform and more a tool for punishing the government. However, public opinion surveys suggest this was not the case: more than 60 percent of voters said they decided based on the content of the reform compared with less than 15% who used their vote to reward or punish Renzi's government (ITANES 2016). The lack of knowledge of legislative structures did not drive voting decisions.

and Stimson 2002; Hooghe, Dassonneville, and Oser 2019; Page and Shapiro 1983; Soroka and Wlezien 2010). The defining feature of political representation is "acting in the interest of the represented, in a manner responsive to them." (Pitkin 1967, 209). Responsiveness to public opinion, adds Powell (2004, 91), rates as "one of the justifications for democracy itself." To this point, many find that elected representatives and the governments they form are, indeed, responsive to public preferences (Budge et al. 2012; Canes-Wrone 2006; Kang and Powell 2010; Klüver and Spoon 2016; Soroka and Wlezien 2010; Spoon and Klüver 2014), though others are less sanguine (e.g., Achen and Bartels 2016; Giger, Rosset, and Bernauer 2012; Jacobs and Shapiro 2000).

But while theories of government responsiveness necessarily emphasize the influence of public sentiment, a range of factors contributes to policy outcomes. For instance, studies find economic downturns are associated with a decline in responsiveness (Clements, Nanou, and Real-Dato 2018; Ezrow, Hellwig, and Fenzl 2020) and with a widening gap in issue priorities of masses and elites (Traber, Giger, and Häusermann 2018). Others report that responsive government is shaped by existing levels of taxes, interest rates, social spending, pension reforms, and so on (e.g., Budge et al. 2012; Elsässer and Haffert 2021; Häusermann 2010; Kang and Powell 2010). Delving deeper, researchers identify the effects of institutional arrangements, including whether the opinion-policy link varies according to electoral rules, federalism, and multi-level governance (Soroka and Wlezien 2010; Peters 2016; Rasmussen, Reher, and Toshkov 2019).

Yet while political institutions figure prominently in current understandings of how representative democracy works, studies of opinion-policy linkages have paid little attention to the role of the assembly as the law-making apparatus. This omission is puzzling. The organization of assemblies deserves the attention of scholars of representation for several reasons. For one, as the setting where policies are proposed, assemblies are critical for crafting policy. Further, legislatures are the lynchpin connecting the voice of the people, who selected its members, to the government, from which it is formed. And perhaps more so than other institutional features, the organization of legislatures has attracted the attention of would-be reformers. We are aware of only one study that considers the organization of the legislature and its relationship to public opinion. In an impressive study on the effects of institutions, Rasmussen, Reher, and Toskhov (2019) find that countries with two legislative chambers have a lower likelihood of opinion-policy congruence than countries with only one.

Rasmussen and coauthors examine congruence, while our focus is on responsiveness. The distinction is not trivial. Congruence assesses factors that influence the "absolute ideological distance between the median citizen and the government" (Golder and Stramski 2010, 90), while responsiveness evaluates whether *changes* in citizen preferences are reflected by similar changes in preferences of elites or policy outputs *over time*.[2] The two concepts are not wholly independent: responsive policy tends to translate into congruence more than unresponsive policy. To the extent that policy is uniformly responsive to public opinion, congruence will not be sacrificed. However, arguably, unresponsive policy produces less congruent outcomes more times than not.

Perhaps for this reason, responsiveness of policymaker-agents to their citizen-principals occupies a privileged position for many theorists of democracy. Indeed, while theorists debate the direction and intentionality of voter-representative linkages, there is broad agreement that the linkage involves a *temporal* element. Dahl (1956) asserts that power is forward-looking such that the voter exerts control over the representative insomuch as the latter acts in response to the

[2] Golder and Stramski (2010) introduce several alternative conceptualizations of congruence which also account for the diversity of citizens' ideological preferences and parties' policy positions in parliament.

former. Nagel's (1975, 29) view of power as a "causal relation between the preferences of an actor regarding an outcome and the outcome itself" is also consistent with a temporal sequence. And Mansbridge (2003) offers a schema for understanding how temporal shifts relate to different forms of representation. These perspectives play out in empirical research. Government policy adjusts to changing preferences in the electorate with a lag (e.g., Kang and Powell 2010), whereas political parties may adjust more rapidly to shifting preferences during elections (e.g., Adams et al. 2004). Accordingly, we conceive of representation as a dynamic process encompassing short- and long-term effects (see also Erikson et al. 2002).

Just as the literature on responsiveness largely neglects the assembly, the range of studies on the implications of bicameralism have yet to turn their attention to elite responsiveness. This is true despite a wide-ranging literature on legislatures. Researchers have examined bicameralism's influence on many important political phenomena including intraparty bargaining (Bäck, Debus, and Klüver 2016), government formation and duration (Druckman and Thies 2002), budget deficits (Heller 1997), and party organization (VanDusky and Heller 2014). So while there are several studies of democratic responsiveness on the one hand, and bicameralism on the other, the direct link between them has yet to be made.

Does bicameralism matter for responsiveness, and if so, how? To the extent it enhances the attention paid to diverse groups, two chambers may be better than one. However, a more equitable distribution of power between chambers may constrain policy change, compared to a single chamber legislation process with less friction between citizens and policymakers. In this way, unicameral designs may be preferable for channeling responsive policies. After developing these competing claims, we present three tests of their implications. We first analyze public opinion and policy outcomes in two salient issue areas from fifteen developed democracies encompassing a wide range of cameral diversity. We show that while governments deliver policy in line with shifts in public preferences, responsiveness is weaker where bicameralism is stronger. Second, we examine differences within bicameral designs. We demonstrate that the basis of policymakers' reduced responsiveness under bicameralism is due to the balancing of formal powers across chambers. Other sources of diversity within bicameral systems, including congruence in method of selection, differences in partisan compatibility between chambers, and means by which agreements are reached across chambers, bear no effect. Third, we leverage the case of constitutional reform in Belgium to examine whether an increase in legislative power concentration enhanced policy responsiveness. Indeed, creating a weaker upper house paved the way for government social policies that were more responsive to public opinion. In short, we show that bicameralism diminishes policy responsiveness, that it is further diminished where the two chambers are equal in power, and that institutional change facilitates policy responsiveness.

Study findings are important for theoretical and policy reasons. With respect to theory, we test a heretofore overlooked prediction for how institutions affect the functioning of democracies. To some, bicameral institutions matter because the presence an upper chamber alters prospects for policy outcomes to diverge from the status quo (Druckman, Martin, and Thies 2005; Tsebelis and Money 1997). For others, the influence of institutions over policy outcomes stems from how authorized control over decision-making is organized and allocated (Eppner and Ganghof 2015; Lijphart 2012; Soroka and Wlezien 2010; Thies and Yanai 2014). Extending both perspectives, we show that the number of chambers matters for responsiveness and, furthermore, so too does the power distribution between them.

This study also carries policy implications. While constitutional features generally resist change in more established democracies, there are nevertheless many important post-war examples. Denmark and Sweden eliminated their upper houses in 1953 and 1970, respectively. In Belgium, constitutional reforms during the 1990s reallocated power vertically, by relocating powers to the regions, and horizontally, by shifting power to a more dominant lower chamber and away from the upper house. In Britain, reform of the House of Lords has been a point of discussion for well over a century by MPs and constitutional scholars seeking to modernize British democracy (Russell 2013). And while recent efforts in Italy were unsuccessful, discussions have continued about whether and how to reform the status quo. Furthermore, reforms to legislative institutions are frequently proposed in the service of making politics more inclusive of diverse communities or, alternatively, for more efficiency. Our results inform these discussions by spelling out how the organization of legislatures matters for the relationship between policy and the will of the public.

Unicameralism, Bicameralism, and Policy Responsiveness

According to a stylized chain of representation, citizens' interests are transmitted to and articulated by political parties which, after the election and upon the formation of governments, translate public demands into policy outcomes. In practice, however, institutions facilitate or impede the transmission of public preferences into policy outcomes. Opportunities for derailing responsive policy outputs are manifold. The electoral system is one such example. While proportional electoral rules mean that the median legislator will be closer to the median voter (Powell 2000), they also are more apt to produce multiparty governments, which blurs the lines of accountability and delays responsiveness (Hobolt, Tilley, and Banducci 2013). Consistent with this latter view, Ferland (2020) shows that responsiveness is inversely related to the number of parties in the cabinet, and that governments tend to be more responsive under majoritarian electoral designs. By awarding outsized influence to plurality winners, majoritarian rules enable governing parties to focus less on crafting concessions to junior partners and more on delivering for the electorate.

The vertical distribution of powers—through federalism, regional autonomy, or supranational governance—may constrain the capacity of national governments to respond to policy preferences. Peters (2016) finds federal systems to be more responsive than unitary systems, because local political elites are closer to the public and more sensitive to public opinion than are elites at the national level. The horizontal division of powers may also matter. Governments may be more responsive to changes in public opinion in "Madisonian" presidential systems than in systems where legislative and executive power are fused (Soroka and Wlezien 2010). The separation of powers grants proposer power to the legislature, relegating the president to react. In contrast, where the legislature and executive are fused, as in parliamentary systems, cabinet dominance leads to high levels of discretion in policymaking.[3] Building on these arguments, we concur that constitutional choices to concentrate or diffuse power matter for the connection between policy outcomes and public opinion. Our focus, however, is squarely on where policy is proposed: the legislature.

With respect to policy responsiveness, does it make a difference whether legislation requires one chamber or two? One possibility is that by creating more pathways for popular influence on agents, bicameralism facilitates responsiveness. This notion has a long history in liberal thought, from Blackstone's (1979[1785]) argument for a mixed constitution in England to Madison's articulation of the separation of powers in Federalist 51 (Hamilton et al. 2014[1788]). Shugart and Carey (1992) argue that presidential systems offer two agents for the electorate, the executive and the legislative, which increases opportunities for responsiveness. While their focus is on the two separate branches of government, the logic carries over to separation of powers within the assembly. As with arguments about the benefits of decentralization in general (Peters 2016), the horizontal division of legislative powers increases the opportunities for citizens to register their views with their representatives.

[3] Possession of dissolution powers in parliament grants prime ministers even more discretion in forming policy (Ecker, Schleiter, and Bäck 2019).

Similarly, a second chamber may provide voice to interests in the population that may have gone underrepresented in the lower chamber. Such is the case for legislatures in Germany, Switzerland, and the United States where legislators are allocated to the lower houses based on population and to the upper houses based on territory or region. Third, delaying and inter-cameral bargaining contribute a level of stability requisite for liberal democracy. Bicameralism limits the downside populist tendencies of majority rule (Riker 1992; cf Chiou and Rothenberg 2003). If political competition is in two or more dimensions, legislatures operating under majority rule permit out-of-equilibrium policies. Rules of inter-cameral bargaining, however, reduce disagreements among legislators to "one privileged dimension" (Tsebelis and Money 1997, 4), and responsiveness to the median voter is enhanced when issues are bundled into one dimension (Schofield 1985). Together, these features of bicameralism bring more representatives closer to the national publics' political preferences.

Yet a more straightforward expectation is that bicameralism reduces responsiveness. Bicameralism disperses decision-making authority (Powell 2000), requiring more elaborate rules and staging to pass legislation. While this results in outcomes more proximate to the status quo (or, more *stable* outcomes), policy decisions produced through multistage processes are ill-suited for efficient and responsive decision-making. Of course, if the public prefers stable policy outcomes over changes to the status quo, then dispersed authority may not be viewed as a disadvantage.

The argument that responsiveness is greater in systems with a single chamber follows directly from the original function of second chambers: as a brake on unfettered populist rule. The expansion of the franchise led to fears of reforms against the interests of the aristocracy, and an upper chamber served as a conservative safeguard against such excesses. While this initial function has become largely obsolete, upper chambers still carry a degree of *status quo bias*. By adding a chamber as an additional veto player within the legislative branch, the set of alternative bills that may beat the status quo shrinks, and with it the prospects for policy change (Tsebelis and Money 1997). Furthermore, a key consequence of reducing the range of plausible policy alternatives is that bicameralism subsequently reduces the agenda-setting role of its governments (Tsebelis 2017). This constraint weakens the capacity of governments to initiate responsive policies when compared to less constrained governments in unicameral systems.

The relationship between the design of the legislature and representation further requires a consideration that not all bicameral designs are alike. Upper chambers vary with respect to the selection of their members. In most, members are elected directly, but in others, such as the Dutch Eerste Kamer, selection is indirect or, in the case of the German Bundesrat, they are appointed. Other points of variation include whether similar or different political parties dominate in each chamber, and in the process for which agreements are reached between chambers.[4] This observation raises a second question, the answer to which has important implications for designing constitutions with the twin goals of stability and responsiveness in mind: Do differences among bicameral systems matter for policy responsiveness?

Since bicameral designs vary in the distribution of constitutionally endowed powers, we extend these arguments into theoretical expectations for government responsiveness *within* bicameral systems. In some cases, power is equally distributed across chambers, or what Lijphart (2012) calls "symmetric", and in others the lower house holds a preponderance of power ("asymmetric"). If policy is more responsive when legislative authority is concentrated, then responsiveness is weakest where the constitutional powers assigned to the second chamber are on par with the first, i.e., where it is symmetric. Likewise, when power is more concentrated

[4] We consider multiple points of variation within bicameral designs in the empirical analyses below.

in the lower chamber, the upper chamber is less likely to delay or prevent policy movement, and the government has more freedom to move policy in the direction favored by the public.

To summarize, bicameral systems may enhance coverage of political preferences in the population, but they also disperse authority which could inhibit responsive decision-making. These two competing sets of arguments support the following hypotheses:

Hypothesis 1a (Bicameral Responsiveness): Policy responsiveness to public opinion is stronger in systems with bicameral legislatures compared to those with unicameral legislatures.

Hypothesis 1b (Unicameral Responsiveness): Policy responsiveness to public opinion is stronger in systems with unicameral legislatures compared to those with bicameral legislatures.

A similar logic informs expectations about variations *within* bicameral systems. Power is most diffused when the power of the upper chamber is on par with the lower, and this dispersed authority should matter for responsiveness. This leads to a second pair of expectations:

Hypothesis 2a (Power Symmetry): Within bicameral systems, policy responsiveness to public opinion is stronger where powers are distributed symmetrically between chambers.

Hypothesis 2b (Power Asymmetry): Within bicameral systems, policy responsiveness to public opinion is stronger where powers are distributed asymmetrically between chambers.

We evaluate these hypotheses by examining the determinants of government policy in the short- and long-term. Our primary measure of policy pertains to welfare state generosity. In the Western democracies we analyze, social welfare is the most salient dimension of policy contestation.[5] We supplement these analyses by considering a second—and increasingly salient—issue divide in the form of immigration policy. For both expectations, we assume that social policy reacts to stimuli with a lag since it, as Kang and Powell (2010, 1017) note, "has a large inertial component, limiting programmatic change in a given year." Furthermore, to the extent short-term effects exist, the status quo bias arguments raised above suggest that unicameral systems are more likely to exhibit short-term responsiveness than bicameral systems.

Data and Measures

We measure social policy using the welfare state generosity index developed by Scruggs, Jahn, and Kuitto (2017). The index combines a range of social insurance benefits including employment insurance, sick pay insurance, and public pensions.[6] Our measure of public opinion

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[5] According to expert assessments of issue saliency (Benoit and Laver 2006, 176 and Supplementary Information (SI) Section D, pages 8-9), issues related to economic left-right (tax-spend and deregulation) rank as most salient in 11 of 15 cases we examine in depth below. In the remaining four, immigration (Denmark and the Netherlands) and the European Union (Portugal and the United Kingdom) ranked just ahead of left-right considerations. Responsiveness is expected to be weaker for less salient issues across political systems (Spoon and Klüver 2014).

[6] The index is the sum of the sub-indices for unemployment and sick pay insurance and pension generosity. Country-year values are standardized using on *z*-scores normalized on the cross-sectional mean and standard deviation in 1980 (Scruggs 2014). Proponents of this measure argue that it is a better measure of government policy than social welfare spending because the latter is influenced by unemployment rates and the population of pensioners that cause welfare

gauges the preferences of the median voter. For policy preferences we rely on the Left-Right scale from 1 (left) to 10 (right), using surveys from the Eurobarometer Trend File from 1978-2002 (Schmitt et al. 2008), all relevant Eurobarometer Surveys from 2003-10, and the Swiss Household Panel survey (Voorpostel et al. 2018) series to calculate the median positions for each country-year.[7] Previous studies report a strong relationship between redistributive attitudes and Left-Right self-placements, and they report strong estimates of this relationship (Benoit and Laver 2006; Hellwig 2015; Rohrschneider and Whitefield 2012).[8]

To assess the impact of cameral structure on the opinion-policy relationship, we classify national legislatures according to Lijphart's (2012) index of bicameralism. The index considers two criteria: i) similarity of the two houses' power over legislative outcomes, or "symmetry," and ii) similarity in the houses' method of selection, or "congruence." Combining these features

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spending to vary even if entitlement policies remain unchanged.

[7] If the distribution of citizen preferences is bimodal, this could have significant implications for the analysis. However, the median voter position and the *mean* voter position are expected to be similar because the distributions of respondents' self-placements are generally unimodal and symmetric (Adams and Somer-Topcu 2009, 682). Ward et al. (2011, fn. 50) report a correlation of 0.97 between estimates of the median and mean. Powell (2021) further shows that when the distribution of citizen self-placements is characterized as normal, there are only slight differences between the use of the interpolated median and the mean.

[8] Annual measures of public preferences for social welfare do not exist cross-nationally in the Eurobarometer or the other survey series. The SI Section D (pages 8-9) reports evidence from the European Social Survey (ESS) evaluating the relationship between preferences on redistribution and left-right positions that these preferences correlate at a relatively high level.

yields an index ranging from unicameralism to strong bicameralism (see SI Table S1, page 2).[9] To assess the robustness of our results, the analyses that follow employ both the index measure and a binary indicator coded 1 for systems which score three or higher on the index, and 0 otherwise.

Analysis

We begin by comparing changes in policy outputs. Does the capacity of two-chambered legislatures contribute to greater policy change? Or are legislatures with single chambers better able to depart from the status quo? To examine this question in this context, we compare absolute annual changes in *Welfare state generosity* for countries with unicameral and bicameral assemblies. We pool across all years available in the Comparative Welfare Entitlements Dataset (Scruggs et al. 2017).[10]

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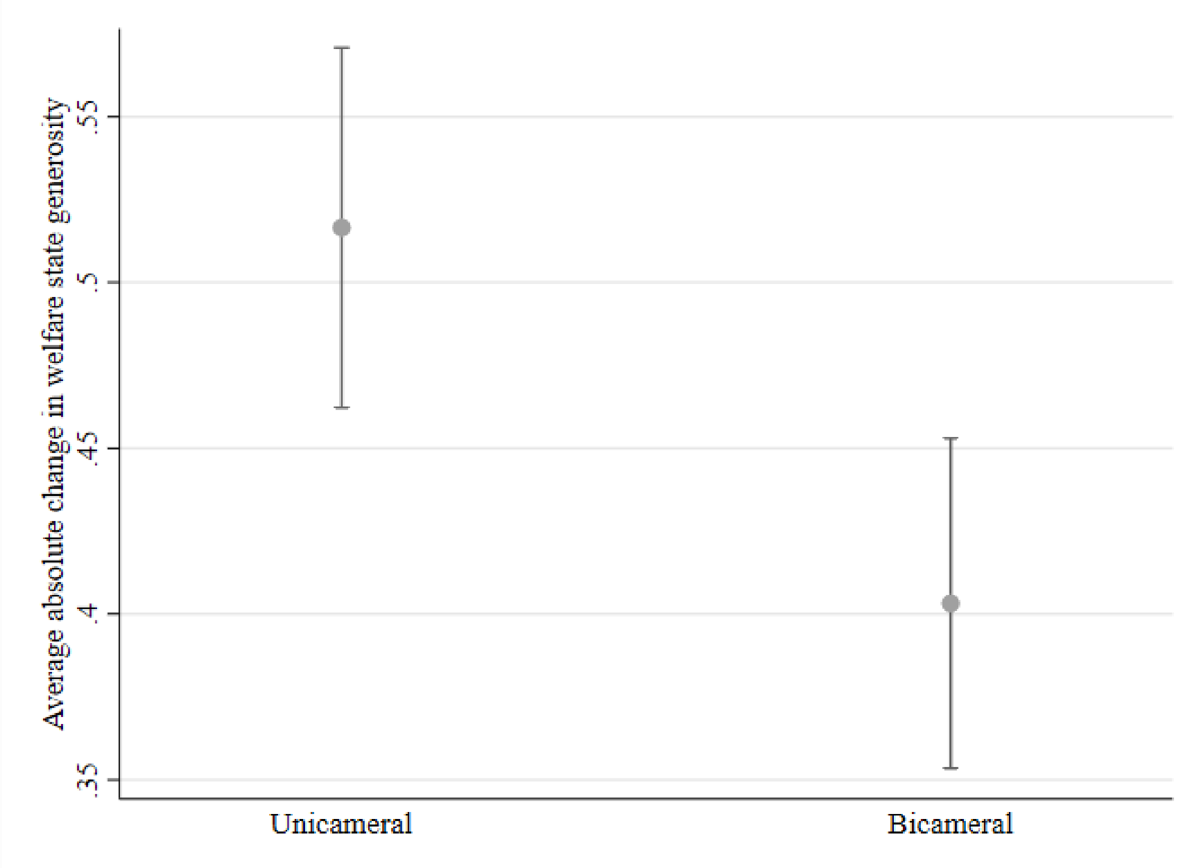
[9] The index yields values of 1, 2, 3, or 4. The only exception is the United Kingdom which Lijphart (2012, 200-201) "demotes" to 2.5 from a score of 3. The House of Lords is the product of a pre-democratic era which renders it less influential, according to Lijphart, compared to upper houses with stronger democratic underpinnings in other medium strength bicameral systems.

[10] Series run from 1971 to 2010 for Austria, Belgium (1975), Denmark, Finland (1976), France (1976), Germany, Greece (1981), Ireland (1973), Italy (1975), Netherlands (1975), Portugal (1986), Spain (1983), Sweden, Switzerland (1983), United Kingdom, Australia, Canada, Japan, New Zealand, Norway, South Korea (2008), and the United States. Note that the last six countries on the list are not in the cross-national analyses below that rely on Eurobarometer survey data.

Figure 1 shows that in unicameral systems, the estimated year-to-year change in *Welfare state generosity* is 0.52, and in bicameral systems the estimated change is 0.40. The difference in means is 0.11 ($p = 0.03$). Year-to-year changes to welfare state generosity policies are larger in unicameral political systems than in bicameral political systems. This finding is consistent with the Unicameral Hypothesis and, more generally, with claims of bicameralism's status quo bias.[11] We cannot infer from these estimates, however, that the larger policy changes reflect government responsiveness to changes in public opinion. On this basis we proceed.

[11] We provide further evidence of bicameralism's status quo bias in the SI Section C (pages 5-7).

Figure 1. Cameral Structure and Absolute Changes in Welfare State Generosity across 22 Countries



Notes. The figure shows the mean of the absolute changes in *Welfare state generosity* observed in unicameral and bicameral systems. To depict the difference between means with 95% confidence, vertical bars for each category report 84% confidence intervals (Julious 2004).

To capture responsiveness as a dynamic process, we test research hypotheses using dynamic models. We model the responsiveness of policy outputs to public opinion as a general error correction model (GECM). The model is of the form $\Delta Y_{it} = \alpha_0 + \alpha_1 Y_{it-1} + \beta_0 \Delta X_{it} + \beta_1 X_{it-1} + \varepsilon_{it}$, where Δ is the difference operator, t indexes time (in years), and i countries. Substantively, the model is flexible enough to uncover both the immediate and long-run impact of a shock to X on Y. Empirically, modelling shifts in policy outcomes rather than levels helps address potential issues of non-random error structures (Tromborg 2014). The contemporaneous impact of a shock to X is provided by β_0 while the cumulative impact is β_1/α_1 (De Boef and

Keele 2008). Models include country fixed effects to account for unobserved country-specific factors.[12]

We estimate the model on data from fifteen countries and report results in Table 1. Models 1 and 2 stratify the sample and report estimates separately for unicameral and bicameral systems, classified as in Figure 1 above. In these models, our chief interest is in the impact of *Median voter position* on *Welfare state generosity*. This coding convention means that a *negative* coefficient indicates responsiveness: if the median voter is moving rightward and increasing in value, this suggests that welfare state generosity should *decrease*. In Model 1 the coefficient on the long run effect for the lagged public opinion is negatively signed and precisely estimated, a finding consistent with previous research (Budge et al. 2012; Elkjær and Iversen 2020; Kang and Powell 2010; Soroka and Wlezien 2010). Estimates show that the transmission between public opinion and policy occurs over time rather than immediately. Model 2 reports estimates for bicameral systems. In this case, the contemporaneous effect of the median voter position is indistinguishable from zero while lagged series returns a positive-signed coefficient.[13]

Model 3 assesses whether these differences in the influence of public opinion on policy change is statistically significant by pooling all observations but interacting median voter positions with a dummy variable scored 1 if medium to strongly bicameral, and 0 otherwise. Recalling that a negative coefficient indicates more responsiveness, the positively signed

[12] Section E of the SI (page 10) discusses time series and stationarity considerations, and Table S5 (page 11) reports results of alternative estimators. Table S6 (page 13) provides a check of the robustness of these results to the inclusion of several control variables.

[13] This finding is not robust across specifications (see Figure 2), and analyses below suggest that it is driven by symmetrical bicameral systems (see footnote 19).

coefficients on the interaction terms implies that responsiveness to the median voter is weaker in systems with two chambers. For systems with a single chamber, the coefficient on the *Median voter position*$_{t-1}$ variable is negatively signed and precisely estimated (β = -0.95, s.e. = 0.14), but for bicameral systems it is no different from zero (β = 0.36, s.e. = 0.26).[14]

Figure 2 leverages the dynamics of the error correction model to chart the effects of public opinion shifts on policy change over time. We use Model 3 estimates to display a forecast of *Welfare state generosity* when the median voter position shifts one standard deviation to the left (see SI Section I, page 15, for details). As we would expect, a leftward opinion shift increases welfare state generosity steadily over time in unicameral systems. But in bicameral systems, positions of the median voter position produce almost no effect on policy.

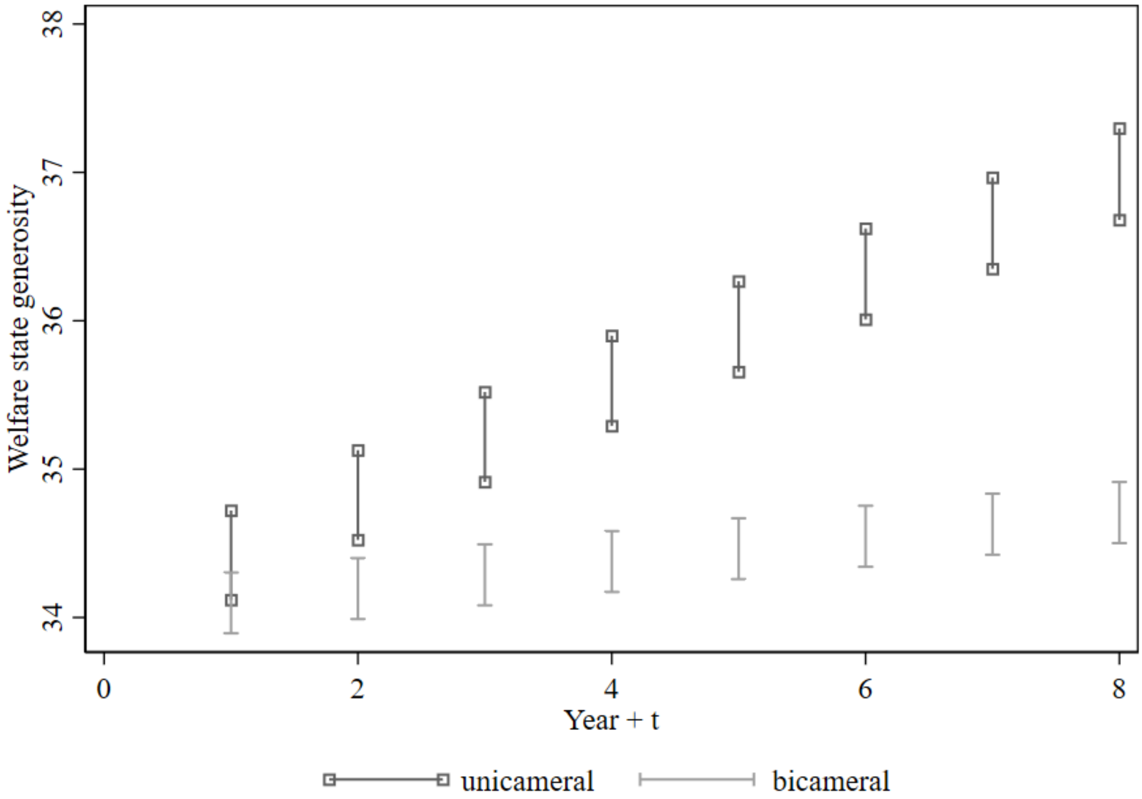
[14] In the SI Table S7 (page 14) we re-estimate Table 1 Model 3 dichotomizing the index of bicameralism at alternative points along Lijphart's four-point scale. The substantive findings are unchanged from those reported in Table 1. However, long-run effects notwithstanding, we further find that responsiveness is most rapid in the subset of strongly, or purely unicameral cases (that is, those scoring 1 on the index).

Table 1. Unicameralism, Bicameralism, and Policy Responsiveness to the Median Voter: Welfare State Generosity

	(1) Unicameral	(2) Bicameral	(3) All	(4) All
Welfare state generosity$_{t-1}$	-0.072**	-0.122**	-0.084**	-0.080**
	(0.025)	(0.027)	(0.018)	(0.018)
ΔMedian voter position$_t$	-0.437	0.245	-0.433	-0.998*
	(0.431)	(0.304)	(0.363)	(0.446)
Median voter position$_{t-1}$	-0.926**	0.449*	-0.949**	-1.271**
	(0.307)	(0.200)	(0.257)	(0.376)
Bicameral$_t$			-7.009**	
			(2.015)	
ΔMedian voter position$_t$ × Bicameral$_t$			0.616	
			(0.546)	
Median voter position$_{t-1}$ × Bicameral$_t$			1.313**	
			(0.370)	
Index of bicameralism$_t$				-9.608**
				(3.377)
ΔMedian voter position$_t$ × Index$_t$				1.945*
				(0.880)
Median voter position$_{t-1}$ × Index$_t$				1.839**
				(0.614)
Constant	7.364**	2.455	7.882**	9.449**
	(1.902)	(1.296)	(1.539)	(2.142)
Observations	202	184	386	386
R-squared	0.165	0.148	0.159	0.154

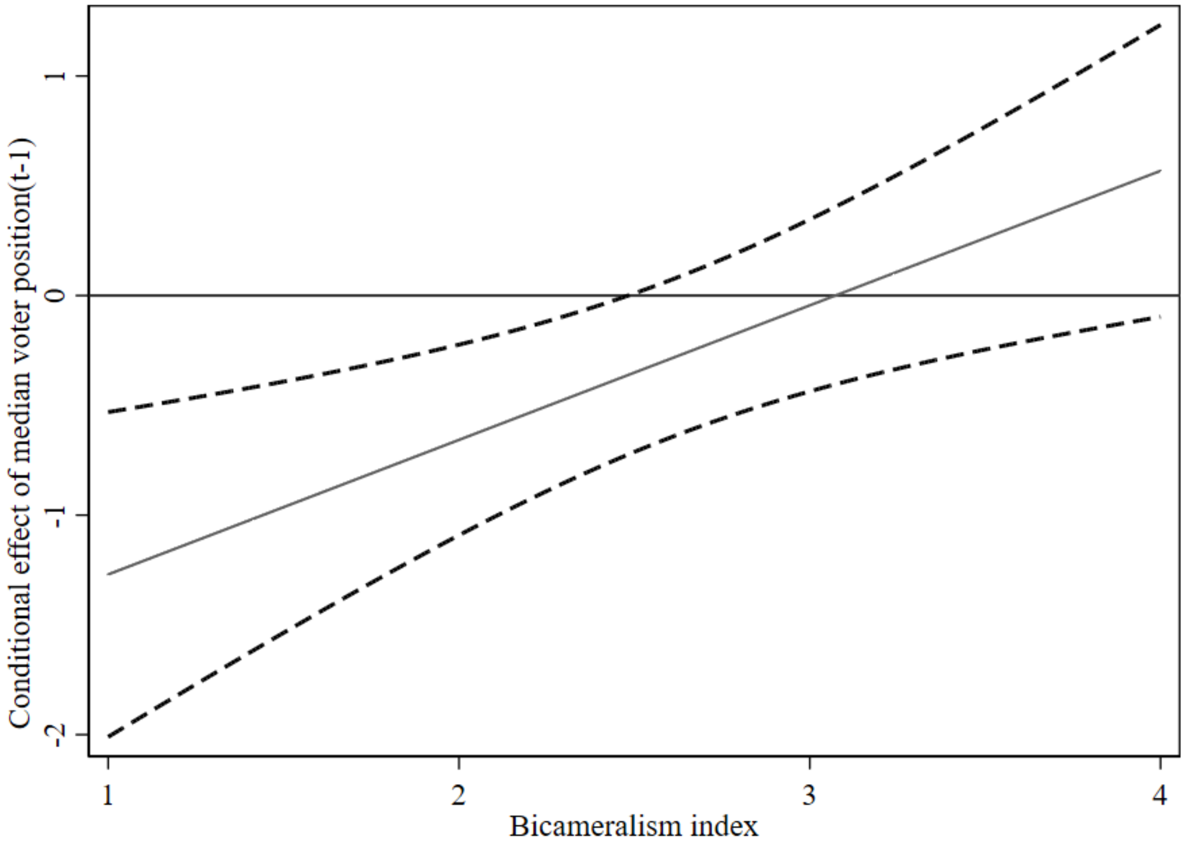
Notes. Cells report least squares coefficients with standard errors in parentheses. Models include country fixed effects. The dependent variable is Δ*Welfare state generosity$_t$*. Negative estimates indicate stronger government responsiveness to public opinion. * $p<0.05$, ** $p<0.01$, two tailed test.

Figure 2. Dynamic Simulation of Welfare State Generosity in Unicameral and Bicameral Systems



Notes. Figure displays expected values of *Welfare state generosity*$_t$ when the *Median voter position*$_{t-1}$ is set to one standard deviation to the left of its in-sample mean. Estimates are based on Table 1 Model 3. The initial value for *Welfare state generosity*$_t$ is set to 34. Vertical bars report 95% confidence intervals.

Figure 3. Marginal Effects of Median Voter Position on Welfare State Generosity across the Range of the Index of Bicameralism



Notes. Figure is produced with Table 1 Model 4 estimates, and it charts the coefficient on *Median voter position*$_{t-1}$ over values of the *Bicameralism index*. Negative estimates indicate stronger government responsiveness to public opinion. Dashed lines report 95% confidence intervals.

The last column in Table 1 treats the bicameralism as continuous.[15] We again find that policy responsiveness to public opinion declines as the strength of bicameralism increases. Further, treating cameral structure in terms of degrees of bicameralism, rather than its presence or not, has the effect of revealing a contemporaneous effect of public opinion change on policy change. The contemporaneous impact of the median voter position is also negatively signed and precisely estimated, both directly and conditioned by its interaction with the bicameralism index, a finding that is consistent with the Unicameral Hypothesis. Relying on the long-term

[15] We rescale the bicameralism index to 0-1 to facilitate comparability.

effects (*Median voter position$_{t-1}$*) we use these Model 3 estimates to display the marginal effects in Figure 3. Results show that responsiveness occurs only for systems with values less than 3 on the index.[16]

We show the organization of the assembly as a key determinant of responsiveness. Three considerations follow from this. The first pertains to whether this result is due to the explanation behind the Unicameral Hypothesis that policy change is less constrained in these systems. In the SI Section C (Table S3, page 6), we report analyses that greater policy changes are more likely to produce responsive outcomes than small policy changes. The second consideration is how other political institutions, such as federalism and electoral systems, may also influence responsiveness. To this end, we re-estimate Model 4 from Table 1 to consider a broad range of formal rules, and report results in Table S8 in the SI (pages 18-19) that control for vertical diffusion (i.e., federalism), horizontal diffusion (separation of legislative and executive powers), majoritarian electoral rules, assembly size, and Lijphart's executive-parties and federal-unitary dimensions of power-sharing. For all of the above, bicameralism's conditioning effect on government responsiveness remains intact. Finally, to evaluate whether the results extend to other types of legislation, we examine government responsiveness on immigration policy. Immigration is appropriate for our purpose because it has consistently ranked as one of the most important issues in Europe over the past decades (Böhmelt, Bove, and Nussio 2020),[17] and it

[16] This suggests, for example, that countries with symmetric (e.g., Belgium before 1995) and/or incongruent chambers (e.g., France) do not yield responsiveness policy making, while those without both of these features do exhibit responsiveness (e.g., Belgium after 1995). See SI Table S1 (page 2).

[17] For instance, in the United Kingdom between 2003 and 2017, nearly 30 percent of the population viewed immigration as one of the two most important issues that their country faces.

captures a dimension of issue contestation that is not coterminous with welfare state issues (Allen and Knight-Finley 2019). Results, reported in SI Table S9 (page 22), reveal a pattern is consistent with those for welfare state generosity.

Policy Responsiveness within Bicameralism

Our second pair of hypotheses examines responsiveness within bicameral designs, which further focus on the distribution of power between chambers. According to the Power Symmetry Hypothesis (H2a), equality between chambers facilitates responsiveness; the Power Asymmetry Hypothesis (H2b) posits the opposite expectation. To evaluate these contrasting views, we retain country cases with two chambers and remove those with only one chamber. Our measure of interest is the distribution of powers between the lower and upper houses. Following Lijphart (2012), we create a variable *Symmetrical* coded 1 for upper houses in Germany, Italy, the Netherlands, Switzerland, and, from 1978-1995, Belgium; and coded 0 otherwise. For asymmetric systems, the estimated yearly absolute change in *Welfare state generosity* is 0.53, and the estimated change is lower for symmetric systems (0.36). The difference in means is 0.17 ($p = 0.01$).

We interact *Median voter position* with *Symmetrical* and report estimates in Table 2 Model 1. The coefficient on *Median voter position*$_{t-1}$ is negatively-signed and precisely estimated, implying responsiveness when the lower house holds a preponderance of legislative power. However, for the symmetrical cases the effect of public opinion is positively signed, suggesting that governments shift policy in the direction opposite of public opinion on welfare state generosity.[18] Parity of power across the two chambers within bicameral systems diminishes estimates of responsiveness. This finding is consistent with the Power Symmetry Hypothesis.

The remaining columns in Table 2 assess the robustness of this result to additional sources of inter-cameral differences. First, we consider whether the method of selection to the chambers. Lijphart refers to "incongruent" legislatures in which the two chambers are selected by different methods. For example, in some systems selection to the upper house is designed to enhance regional representation (e.g., Germany). Such incongruent designs may be less adapted to responding to the median voter position. By contrast, "congruence" represents a similarity of selection methods (Lijphart 2012, 194), and these systems are potentially more responsive to the median. Accordingly, we use Lijphart's measure for chamber incongruence, which he assigns to France, Germany, Spain, Switzerland, and the United Kingdom. The conditional coefficients reported in Model 2 indicate that (in)congruence within bicameral systems does not affect responsiveness. We next consider differences in partisan control of the chambers. Regardless of power symmetry or incongruence in selection, partisan compatibility, or whether the same parties dominate across chambers, may vary over time. We therefore also consider variation in terms of the average absolute difference in party seat shares between the two houses.[19] Higher values of *Seat difference* indicate greater differences across the chambers with

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[18] The coefficient on *Median voter position*$_{t-1}$ conditional on *Symmetrical* taking a value of 1 is estimated as 0.43 with standard error 0.13. While not robust to all specifications, this finding may warrant further examination.

[19] Seat difference is calculated as $\left(\sum_{i=1}^{n}|UpperShare_i - LowerShare_i|\right)/n$, where $UpperShare_i$ is party i's seat share in the upper chamber, $LowerShare_i$ is party i's seat share in the lower chamber, and n is the number of parties. Data are from Druckman and Thies (2002), which we update for years since 2000.

respect to party dominance. To the extent that this would influence the status quo bias of the legislature, as our theory suggests, we would expect that seat difference diminishes government responsiveness. Model 3 shows that the estimate on *Median voter position*$_{t-1}$ × *Seat difference*$_{t-1}$ interaction variable is indeed positively-signed, suggesting that responsiveness is weaker where the differences in party seat shares are greater. This effect, however, is imprecise and, more germane to our efforts, the influence of power distribution is unchanged.

Lastly, we consider the power the second chamber extracts from the first due to systems of passing legislation. Nearly all bicameral assemblies, and all those in our dataset, employ a shuttle, or “navette,” system for passing legislation. It may be that the more shuttling required between chambers, the less likely that proposals result in legislation. Hence, more shuttling may yield less policy responsiveness. On the other hand, enhanced deliberation afforded by shuttling may well lead to more representative policy outputs. To gauge the influence of shuttling, we create a variable *Shuttles* to capture the number of rounds of exchange between chambers required to pass legislation.[20] Model 4 includes the interaction between *Shuttles* and *Median voter position*. Estimates on the interactive effects are imprecise. Yet again, for our focus, the conditioning influence of *Symmetry* is unchanged. In total, Table 2 results lend support to claims that an asymmetric distribution of formal powers between chambers increases government responsiveness to public opinion.

[20] *Shuttles* ranges from 1 to 3; and it is extended to a value of 4 for constitutions allowing for indefinite rounds. We rescale the measure from 0 to 1 before entering it in the model.

Table 2. Power Distribution between Chambers and Policy Responsiveness

	(1)	(2)	(3)	(4)
Welfare state generosity$_{t-1}$	-0.094**	-0.107**	-0.091**	-0.096**
	(0.021)	(0.022)	(0.023)	(0.021)
ΔMedian voter position$_t$	0.360	0.614	0.283	0.534
	(0.433)	(0.586)	(0.446)	(0.538)
Median voter position$_{t-1}$	-0.895**	-1.296**	-0.859*	-1.031**
	(0.277)	(0.350)	(0.343)	(0.337)
Symmetrical$_t$	-7.201**	-7.433**	-7.817**	-6.073*
	(2.221)	(2.224)	(2.580)	(2.709)
ΔMedian voter position$_t$ × Symmetrical$_t$	-0.119	-0.364	0.001	-0.174
	(0.677)	(0.732)	(0.689)	(0.702)
Median voter position$_{t-1}$ × Symmetrical$_t$	1.338**	1.409**	1.409**	1.149*
	(0.408)	(0.410)	(0.434)	(0.484)
Incongruent$_t$		-4.794*		
		(2.228)		
ΔMedian voter position$_t$ × Incongruent$_t$		-0.443		
		(0.745)		
Median voter position$_{t-1}$ × Incongruent$_t$		0.781		
		(0.415)		
ΔSeat difference$_t$			6.616	
			(4.290)	
Seat difference$_{t-1}$			-5.750	
			(22.17)	
ΔMedian voter position$_t$ × ΔSeat difference$_t$			18.59	
			(24.11)	
Median voter position$_{t-1}$ × Seat difference$_{t-1}$			-2.970	
			(4.609)	
Shuttles$_t$				-7.599
				(4.215)
ΔMedian voter position$_t$ × Shuttles$_t$				-0.510
				(1.023)
Median voter position$_{t-1}$ × Shuttles$_t$				0.532
				(0.728)
Constant	7.944**	10.45**	7.451**	8.730**
	(1.706)	(2.154)	(1.831)	(2.017)
Observations	280	280	280	280
R-squared	0.139	0.157	0.150	0.143

Notes. Sample excludes unicameral systems. Cells report least squares coefficients with standard errors in parentheses. Models include country fixed effects. The dependent variable is Δ*Welfare state generosity$_t$*. * p<0.05, ** p<0.01, two tailed test.

From Symmetric to Asymmetric Bicameralism: Policy Responsiveness in Belgium

Ample evidence above reveals that bicameral designs are associated with lower levels of government responsiveness to the median voter and, further, that a reduction in responsive policy is due mainly to cases where the powers of the upper house match those of the lower house, i.e., under symmetric bicameralism. Results support claims that bicameralism impedes movements from the status quo and toward changes in the preferences of the electorate. If we are right, *changes* in the design of legislative institutions should register in changes in policy responsiveness. As noted, many governments in recent years have contemplated such reforms. While instances of actual reform are rare, positive cases do exist. One such case is Belgium.

As part of a series of changes aimed at ameliorating tensions between its Flemish and Walloon communities, Belgium engaged in a series of constitutional changes in recent decades. Most fundamental of these was the fourth State Reform, passed in 1993. Along with consolidating earlier efforts to disperse power to the regions, the reform reallocated representation between the lower and upper houses of parliament. The Senate shrunk in size by 61%, its method of selection modified to reduce malapportionment and, most importantly given our findings above, many of its powers were stripped. Where the chambers shared powers in symmetrical fashion prior to the reform, afterwards, most legislation fell under so-called "unicameral matters" for which the Chamber of Representatives exercised sole power. The Senate retained competency on (pure or partial) "bicameral" matters, which include constitutional revisions, regional agreements, and international treaties (Alen and Peeters 1995).

Did the change in the distribution of power between chambers influence future policy responsiveness? In Table 3 we present results from multivariate analyses which estimate changes in welfare state generosity in Belgium. The first column shows that when we estimate our standard model for Belgium alone the coefficients on the *Median voter position* variables are statistically insignificant. Model 2 re-estimates the model and interacts the public opinion variables with a binary variable scored 1 for the post-reform period (1995-2010).[21] Estimates show that the difference in the coefficients on *Median voter position*$_{t-1}$ before and after the reform are statistically significant, and the coefficient on *Median voter position*$_{t-1}$ conditional on *Post Reform* (taking a value of 1) is -1.79 with standard error 0.77 (p = 0.01). These analyses suggest that when Belgium operated under symmetric bicameralism, policymakers were not systematically responsive to the median voter. When the balance of power had shifted towards the lower house, however, estimates of policy responsiveness were statistically significant.[22]

[21] The fourth state reforms took effect following the 1995 general election (Alen and Peeters 1995).

[22] There is little to suggest that the concomitant—but far more gradual in their roll-out—reforms to disperse power vertically to the regions contributed to the improvement in policy responsiveness. We also address this in relation to the synthetic control analysis.

Table 3. Constitutional Reform and Policy Responsiveness: The Case of Belgium

	(1) Belgium	(2) Belgium	(3) Synthetic Belgium
Welfare state generosity$_{t-1}$	-0.130	-0.117	-0.349**
	(0.070)	(0.075)	(0.116)
ΔMedian voter position$_t$	0.396	0.482	0.334
	(0.468)	(0.564)	(0.564)
Median voter position$_{t-1}$	-0.230	0.212	1.189*
	(0.351)	(0.670)	(0.574)
Post reform indicator		11.006*	-3.145
		(4.561)	(5.144)
ΔMedian voter position$_t$ × Post reform		-0.657	-0.239
		(1.014)	(1.078)
Median voter position$_{t-1}$ × Post reform		-2.105*	0.571
		(0.856)	(0.969)
Constant	6.613	3.655	6.336
	(4.449)	(6.275)	(3.80)
R-squared	0.18	0.35	0.31
N observations	33	33	33

Notes. Cells report least squares coefficients, with standard errors in parentheses. The dependent variable is Δ*Welfare state generosity$_t$*. Models 1 and 2 are estimated with data on Belgium; Model 3 estimates are produced using data for synthetic Belgium. * p<0.05, ** p<0.01, two tailed test.

These regression models leverage an intervention in a single country to show how responsiveness changes following a change in the locus of policy control. This test assumes that in the absence of the intervention we would not observe a move toward more responsiveness. While this assumption cannot be tested directly, we can approximate such a scenario by comparing policy changes in Belgium to similar cases. Accordingly, we utilize the methodology of synthetic controls (Abadie, Diamond, and Hainmueller 2015) to create a "synthetic" Belgium as a composite case that closely matches the attributes of Belgium before the shift in powers between chambers took effect. This allows for policy responsiveness to be estimated in Belgium, had the country not experienced constitutional reform.

The analysis begins with the selection of a donor pool of countries from which to create the synthetic Belgium. We are constrained in our choice of donors by those countries for which we have sufficient pre-treatment time periods.[23] We next identify cases on a set of matching variables that predict welfare state generosity levels in the donor pool of countries so that the levels match as closely as possible to Belgium before the reform. The algorithm assigns weights to the donor countries that minimize the differences between Belgium and synthetic Belgium during the pre-intervention period (Abadie, Diamond, and Hainmueller 2010). In this way, we can compare Belgium against a closely matched case as a test of the counterfactual.[24]

In the last column we re-estimate the interactive specification but, as a placebo test, for the synthetic Belgium policy series (in place of the Belgium series). Unlike Model 2, the estimates on the interaction variables in Model 3 do not achieve statistical significance. This suggests that the synthetic Belgium does not experience an increase in policy responsiveness post-treatment, from 1995 to the end of our time series. We conduct jack-knife analyses to check that these results are not being driven by one donor country in the construction of the synthetic control (SI Table S12, page 25).

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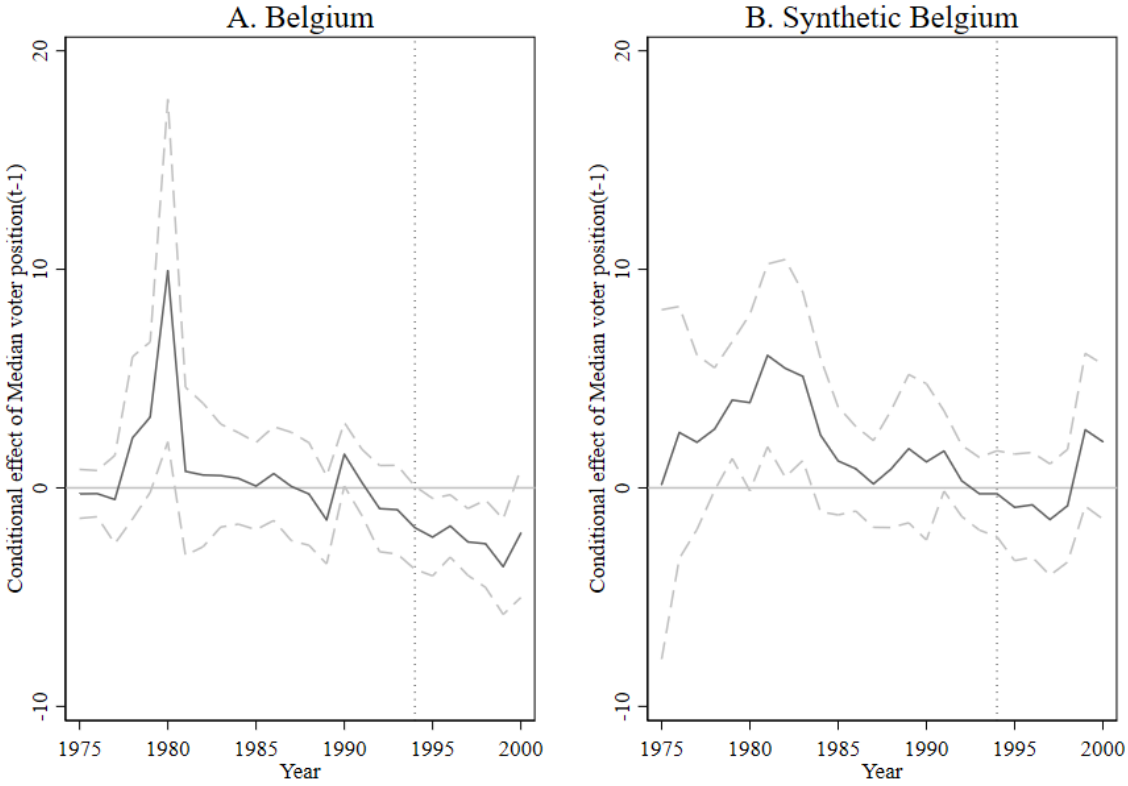
[23] Synthetic control models require panels to be balanced across units to estimate the counterfactuals. Our donor pool is comprised of 11 European countries (see SI Section L, pages 23-25) for which we have a sufficiently long series on *Welfare state generosity*.

[24] The choice of donors and weights is achieved by matching the predictors between the treated (Belgium) and non-treated (synthetic) units over the pre-intervention period. Matching variables are informed by research on social spending and include GDP growth, union density, government partisanship, electoral rules, trade openness, and percent of residents over age 65. Given the role of regional diffusion in Belgian politics in general and the series of State Reforms in particular, we also match on an indicator of vertical power diffusion. See SI Section L (pages 23-25) for details.

Finally, note that models specified in Table 3 assume that any change in policy responsiveness in Belgium would occur just after the fourth state reform's implementation, rather than before or after. We can relax this assumption by allowing the effect of the median voter's preference to vary over time rather than at the time of the intervention (treatment) alone. To do so, we re-estimate the specification in Model 1 using a rolling regression and the coefficients on *Median voter position*$_{t-1}$ are retained. Due to degrees of freedom, a 10-year window is required to estimate these time-varying coefficients. Figure 4A plots coefficient estimates on *Median voter position*$_{t-1}$ for Belgium between 1975 and 2000.[25] We see that for most of the period, the parameter estimates on variable are not statistically different from zero—a finding consistent the general lack of responsiveness in symmetric systems. But by 1995, estimates from the rolling regression veer negative. For comparison, Figure 4B displays the coefficients from a rolling regression using welfare state generosity from synthetic Belgium and this scenario provides no evidence of a shift toward responsiveness, neither in 1995 nor at any time. In total, the results corroborate that welfare state generosity in Belgium became systematically responsive to public opinion after the constitutional reform that weakened its upper house.

[25] The series ends in 2000 rather than 2010 due to the rolling window.

Figure 4. The Time-Varying Effect of Public Opinion on Policy: Belgium and Synthetic Belgium



Notes. The Figure shows estimates of rolling coefficients for the long-term effect of public opinion on changes in generosity, in Belgium (A) and Synthetic Belgium (B). The dashed lines show the 95% confidence intervals. The vertical dotted line separates the pre- and post-reform period. Estimates are based on a time window of 10 years. Negative estimates indicate stronger government responsiveness to public opinion.

Conclusion

How governments respond to public opinion is one of the most important measures by which we can evaluate a democracy. A consensus has emerged that the quality of democracy is influenced by political institutions (Ferland 2020; Hooghe, Dassonneville, and Oser 2019; Kang and Powell 2010; Soroka and Wlezien 2010). We have evaluated government responsiveness to public opinion as measured by how policies respond to changes in the median voter position, paying particular attention to legislature as the body responsible for channelling public preferences into policy proposals. Legislatures have long been identified as central institutions for translating voter preferences into policy decisions (e.g., Powell 2000). Yet which institutional designs for legislatures perform best in meeting that ideal? This paper has investigated an institutional feature foundational to legislatures: the number of chambers. When responsiveness under bicameralism and unicameralism is magnified, policy responsiveness to public opinion is observed to be stronger with one chamber than two, and within bicameral systems the more symmetrical the balance of power between chambers, the *weaker* the opinion-policy responsiveness.

In demonstrating how the structure of national legislatures matters for government responsiveness, this study lays a path for new research on how the views of the public are reflected in policy. Follow-up research should take up the question of responsiveness "to whom" by examining whether bicameralism means governments are more sensitive to the preferences of some subconstituencies over others (e.g., Elkjær and Iversen 2020; Gilens and Page 2014; Griffin and Newman 2005; Peters and Ensink 2015). It may be that governing parties may be responsive to their core supporters in bicameral systems. Another possibility is to test the bicameralism-responsiveness connection across more issues than welfare state generosity and immigration. Notable here is the finding by Rasmussen, Reher, and Toshkov (2019, 425-26; see also Elkjær forthcoming) that opinion-policy congruence varies more across issues than across countries. While these authors acknowledge that bicameralism reduces congruence, future studies should evaluate the relative importance of issue types and assembly types.

Yet another possible way forward follows innovative work by of Carey and Hix (2011) who propose an electoral "sweet spot" with respect to electoral system design that maximizes several features of representative democracy (e.g., responsiveness and stability). There may be a similar bicameral sweet spot that maximizes benefits of unicameral and bicameral systems. For example, although democratic theorists may view the policy responsiveness that we have attributed to unicameral systems as a normatively desirable feature of government, advocates of bicameralism could equally point to how upper chambers tend to protect the interests of regional or minority groups. Future work will examine these additional criteria for measuring the quality of democracy. Our analyses suggest that asymmetric bicameralism may be a case that maximizes advantages of both systems. Another promising extension would be to assess the connection between bicameralism and responsiveness in new democracies. Finally, we must ask to what extent ordinary citizens perceive the difference in policy responsiveness across systems. Corroborating evidence (see SI Section M, pages 26-29) suggests that the answer is very clearly "yes," and that individuals in countries with bicameral assemblies are less likely to believe that decisions taken by their national parliament are in the interest of people like themselves.

Caveats aside, the normative implications of our findings for democratic responsiveness, political institutions, and policy outcomes are striking. By slowing the process of law-making and tempering majority tyranny, bicameral constitutional structures have been argued to be a positive feature of liberal democracy (Riker 1992). These systems require more deliberation and (often) supermajorities that produce legislation that will more likely be favored by majorities. We uncover a potential trade-off of bicameralism's status quo bias: when legislation is difficult to pass, as is the case on most social welfare policies, it has also been *less* responsive to changing preferences in the electorate.

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Bicameralism and Policy Responsiveness to Public Opinion

Supplementary Information File

A. Chamber Classifications

Table S1. Chambers Classifications

Chambers symmetric	Chambers incongruent	Index of Bicam.	Countries
n/a	n/a	1	Denmark, Finland, Greece, *New Zealand*, *Norway (after 2009)**, Portugal, *South Korea*, Sweden
no	no	1.5	*Norway (before 2009)*
no	no	2	Austria, Belgium (after 1995)**, Ireland
no	yes	2.5	United Kingdom***
no	yes	3	*Canada*, France, Spain
yes	no	3	Belgium (before 1995)**, Italy, *Japan*, Netherlands,
yes	yes	4	*Australia*, Germany, Switzerland, *United States*

Notes: Countries listed are those included in Scruggs et al. (2017) and in Figure 1. Countries in italicized font are excluded from regression analyses reported in Tables 1 and 2.
* Until 2009 Norway elected legislators to one chamber, but after the election they divided themselves into two chambers.
** The 1993 revision to the Belgian constitution changed its upper house from a symmetric to an asymmetrically weaker legislative body vis-à-vis the lower house. The change took effect following the May 1995 federal election.
*** The British House of Lords is technically incongruent but its index value is "demoted" by 0.5 owing to its status as a pre-democratic relic (Lijphart 2012, 200-201).

B. Descriptive Statistics

Table S2a. Welfare State Generosity and Median Voter Position

Variable	Nr. of Obs	Mean	Std. Dev.	Min	Max
Welfare state generosity					
Δ	386	0.070	0.753	-2.900	7.600
t-1	386	33.503	4.947	23.100	44.900
Median voter position					
Δ	386	-0.008	0.145	-0.623	0.445
t-1	386	5.215	0.386	4.091	6.308

Table S2b. Bicameralism, Symmetry, and Congruence

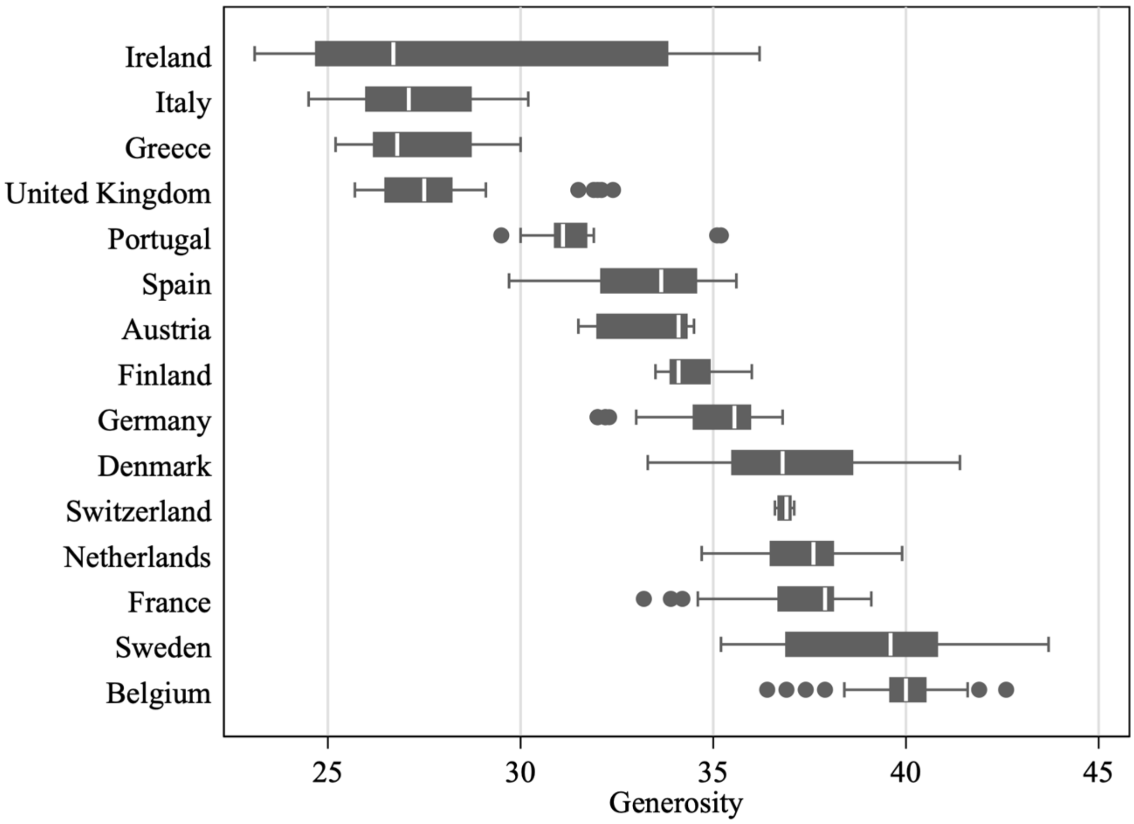
	Value	Frequency	%	%, Cum.
Bicameralism				
	0 (Unicameralism)	202	52.33	52.33
	1 (Bicameralism)	184	47.67	100.00
Index of bicameralism				
	1 (Unicameralism)	106	27.46	27.46
	2 (Weak bicameralism)	63	16.32	43.78
	2.5 (Med.-weak bicameralism)*	33	8.55	52.33
	3 (Medium bicameralism)	141	36.53	88.86
	4 (Strong bicameralism)	43	11.14	100.00
Symmetrical (bicameralism)				
	0 (Asymmetrical)	259	67.10	67.10
	1 (Symmetrical)	127	32.90	100.00
Incongruent (bicameralism)				
	0 (Congruent)	253	65.54	65.54
	1 (Incongruent)	133	34.46	100.00

* Following Lijphart (2012, 200-201), a value of 2.5 is assigned to the United Kingdom. While British bicameralism is technically incongruent, Lijphart "demotes" the score by half a point because the upper house is a relic of a pre-democratic era.

Table S2c. Seat Difference

Variable	Obs	Mean	Std. Dev.	Min	Max
Seat difference					
Δ	280	-0.001	0.012	-0.130	0.049
t-1	280	0.041	0.044	0.000	0.179

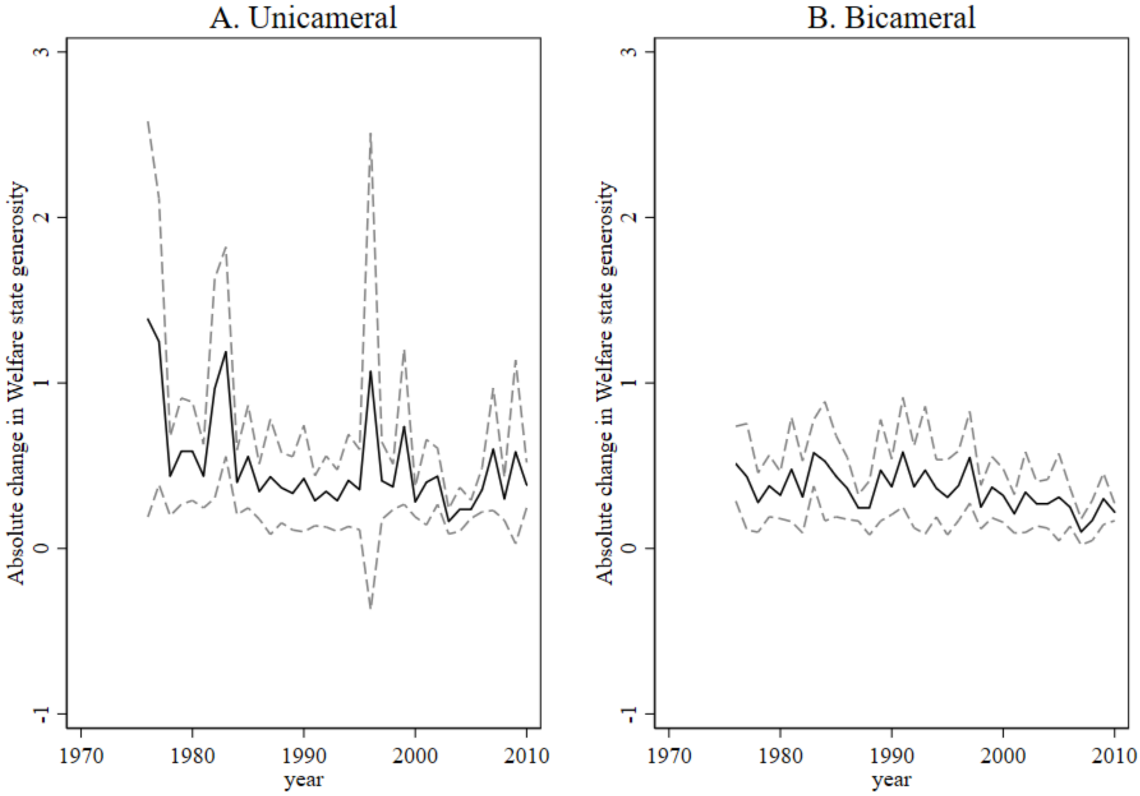
Figure S1. Boxplot: Welfare State Generosity, by Country



Notes. Figure shows the distribution of distributions of welfare generosity values over the sampled countries. The white bar in the boxes indicates median values. Then, all values between the 25th and 75th percentiles are represented by the dark grey box. Countries are sorted by the median value of the generosity index. Data on welfare generosity is from Scruggs *et al*. (2017). In addition, we report the mean annual change in welfare generosity for each country included: an average annual change of 0.20 in Austria; 0.19 in Belgium; -0.14 in Denmark; -0.14 in Finland; 0.15 in France; -0.11 in Germany; 0.14 in Greece; 0.38 in Ireland; 0.16 in Italy; 0.03 in the Netherlands; 0.43 in Portugal; 0.26 in Spain; -0.60 in Sweden; 0.04 in Switzerland; and -0.14 in the United Kingdom.

C. Bicameralism and Status Quo Bias

Figure S2. Absolute Change in Welfare State Generosity in Unicameral and Bicameral Systems across 22 countries



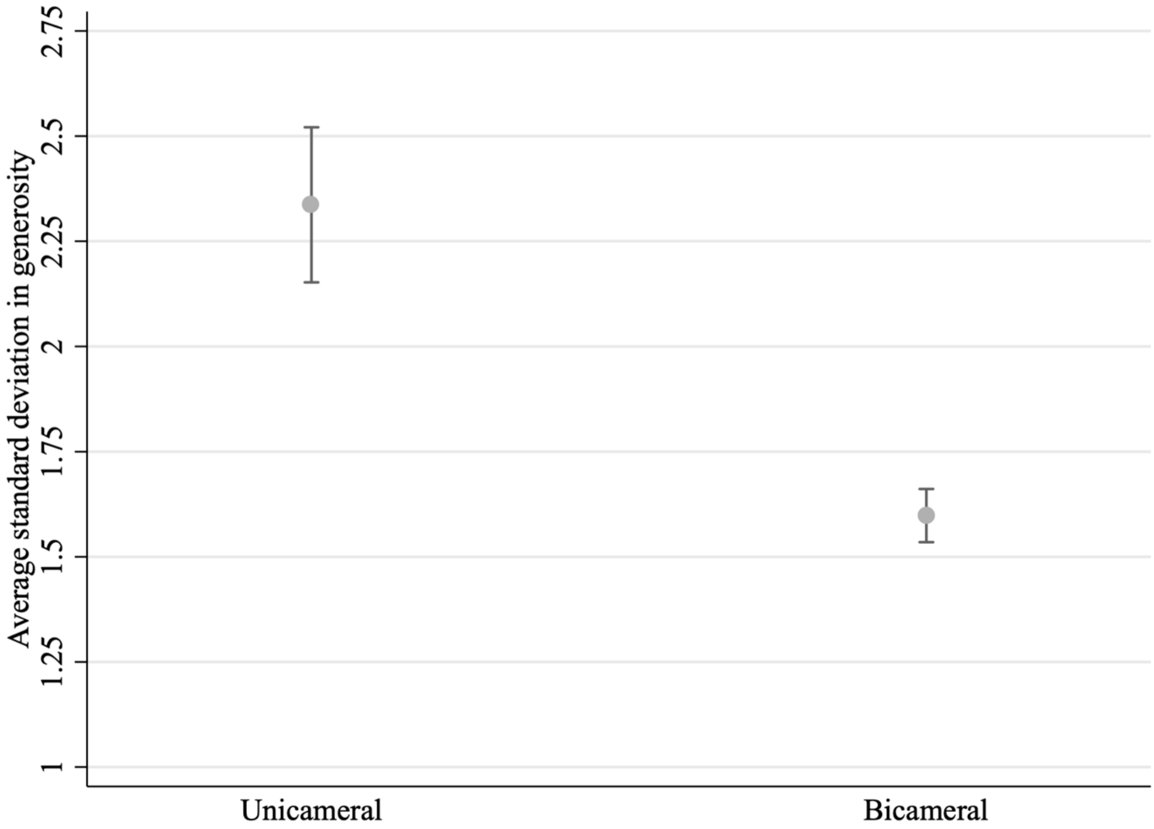
Note: Graphs display the absolute annual change in *Welfare state generosity* averaged over unicameral and bicameral systems. Solid lines report the country means with 95% confidence intervals in dashed lines. Cases are the same as reported in Figure 1 of the manuscript.

Table S3. The Effect of Absolute Policy Changes on Policy Responsiveness

	(1)
Welfare state generosity$_{t-1}$	-0.067**
	(0.017)
ΔMedian voter position$_t$	-0.266
	(0.340)
Median voter position$_{t-1}$	0.039
	(0.190)
Abs(ΔWelfare state generosity$_t$)	3.535**
	(0.890)
ΔMedian voter position$_t$ × Abs(ΔWelfare state generosity$_t$)	0.330
	(0.422)
Median voter position$_{t-1}$ × Abs(ΔWelfare state generosity$_t$)	-0.602**
	(0.168)
Constant	2.080
	(1.162)
Observations	386
R-squared	0.235

Notes. Cells report ordinary least squares estimates with robust standard errors in parentheses. Models include country fixed effects. The dependent variable is Δ*Welfare state generosity$_t$*. * p<0.05, ** p<0.01, two tailed test.

Figure S3. Cameral Structure and Absolute Changes in Welfare State Generosity across 22 Countries



Notes. The figure shows the mean of the standard deviations of *Welfare state generosity* observed in unicameral and bicameral systems. To depict the difference between means with 95% confidence, vertical bars for each category report 84% confidence intervals (Julious 2004).

D. The Relationship between Preferences for Redistribution and Left-Right Positions

In the paper we test claims about the effect of bicameralism on responsiveness by assessing the influence of changes in the median voter's position on the left-right continuum on changes in the direction of welfare state policy. This research design assumes that left-right positions serve as a proxy for preferences for welfare. In the main text we cite previous studies which find this to be the case; here we present our own findings which, again, demonstrate this connection.

To do so, we report findings from performing analyses on the European Social Survey (ESS) study series. While the ESS data do not enable us to go back as far in time, nor are the available at the same (annual) frequency as Eurobarometer, they do include items tapping preferences for redistribution. Specifically, the ESS asks respondents "To what extent do you agree or disagree with the statement: the government should take measures to reduce differences in income levels." This question has been asked consistently in Rounds 1-9 of the ESS. Using the cumulative data file (Rounds 1 to 9), in Model 1 we regressed respondents' left-right positions on preferences for distribution, controlling for the specific round (to account for possible time effects) and clustering standard errors by country. The results confirm a strong association between this item and the overall left-right placement of ESS respondents (Model 1 in Table S4).[26] Then, we evaluate how this association compared to other relevant dimensions of political competition. Round 8 of the ESS includes items on preferences for accepting refugees, taxing fossil fuels, and adoption for same sex couples (that is, items on environmental issues, immigration policy, and the cultural dimension) - along with the item on preferences towards redistribution. Importantly, all of these items were measured on the same 5-points scale, giving us the opportunity to better compare coefficient estimates. Model 2 estimates indicate that redistribution produce the largest coefficient, compared to the other issues attitudes, for explaining responses to the left-right self-placement item.

[26] The bivariate correlation between the items for Rounds 1-9 is 0.20 ($p < 0.000$).

Table S4. Modelling Left-Right Self-Placements, European Social Survey Data

Variables	(1) ESS Rounds 1-9 (2002-2018)	(2) ESS Round 8 (2016)
Opposing redistribution	0.391***	0.379***
	(0.044)	(0.038)
Opposing refugees		0.297***
		(0.052)
Opposing taxing fossil fuels		0.043
		(0.039)
Opposing adoption to same-sex couples		0.274***
		(0.029)
Constant	4.162***	2.438***
	(0.162)	(0.281)
Observations	198,512	20,585
R-squared	0.040	0.113

Notes. The dependent variable is Rightward positioning (measured on an 11-points scale). Cells report ordinary least squares estimates with robust standard errors in parentheses. Model 1 includes fixed effects for each ESS round, 1-9., * p<0.1, ** p<0.05, *** p<0.01, two-tailed test.
Source: European Social Survey

E. Alternative Estimators

In Table S5, we re-estimate our core model (Table 1 Model 4) using a series of alternative estimators. In all six cases, model estimates are consistent with those reported in Table 1 in the main text.

The first pair of models estimate fixed effects. Column 1 cross-country sources of heterogeneity by employing time fixed effects. In Model 2 we include both time and country fixed effects. The next two columns also address heterogeneity, but through the disturbance term rather than country or year intercepts. Model 3 employs Huber-White robust standard errors clustered by country, while Model 4 uses panel corrected standard errors (Beck and Katz 1995).

In the last two columns we filter the dependent variable, *Welfare state generosity*, prior to estimation. For Model 5 each country series is fitted with a linear time trend. For Model 6 we pre-whiten country series using an ARFIMA(p,d,q) process. The latter addresses the possibility of spurious effects owing to non-stationarity (Grant and Lebo 2016, 18). For each of the 15-country series we estimate d, the order of integration, using Sowell's (1992) procedure. In each case, d is between 0 and 0.5 ($0 < d < 0.5$), indicating series that are long-memoried but mean-reverting. We thus follow the recommendations of Helgason (2016) and of Esarey (2016) that advise to base inferences on the unfiltered data series (as in Table 1 Model 4). Helgason (2016) simulates fractionally integrated data and concludes that for models with short-run dynamics estimated on short series (less than 250 observations), GECM estimates produced by the original series outperform models using fractionally differenced data (see also Enns, Moehlecke, and Wlezien forthcoming). Similarly, Esarey (2016) concludes from his simulation analyses that in cases where data are fractionally integrated but stationary (that is, where $0 < d < 0.5$), then estimating a GECM or equivalent autoregressive distributed lag model is preferable to the ARFIMA model and, further, is less susceptible to over-fitting. We follow Helgason's and Esarey's recommendations for data with short time series and estimate GECMs.

Table S5. Bicameralism and Policy Responsiveness: Alternative Estimators

	(1) Time FE	(2) C & T FE	(3) Cluster SE	(4) PCSEs	(5) Detrended	(6) ARFIMA
Welfare state generosity$_{t-1}$	-0.026**	-0.077**	-0.027	-0.027**	-0.179**	-0.461**
	(0.008)	(0.019)	(0.014)	(0.005)	(0.030)	(0.034)
ΔMedian voter position$_t$	-0.945*	-0.982*	-0.914	-0.914**	-0.698	-1.176*
	(0.444)	(0.454)	(0.532)	(0.324)	(0.439)	(0.497)
Median voter position$_{t-1}$	-0.689**	-1.268**	-0.666**	-0.666**	-0.844*	-1.367**
	(0.247)	(0.375)	(0.193)	(0.182)	(0.365)	(0.408)
Index of bicameralism$_t$	-4.839*	-8.661*	-5.021*	-5.021**	-5.695	-10.214**
	(2.244)	(3.432)	(2.217)	(1.551)	(3.246)	(3.735)
ΔMedian voter$_t$ × Index$_t$	2.047*	2.089*	1.762*	1.762**	1.521	2.155*
	(0.897)	(0.911)	(0.774)	(0.592)	(0.861)	(0.966)
Median voter$_{t-1}$ × Index$_t$	0.925*	1.743**	0.964*	0.964**	1.239*	1.959**
	(0.424)	(0.627)	(0.417)	(0.289)	(0.592)	(0.682)
Constant	4.781**	9.218**	4.472**	4.472**	4.104*	22.580**
	(1.370)	(2.131)	(1.071)	(0.974)	(1.901)	(2.619)
Observations	386	386	386	386	386	371
R-squared	0.166	0.259	0.060	0.060	0.114	0.367

Notes. Cells report parameter estimates as described above. The dependent variable is Δ*Welfare state generosity$_t$*, calculated as the change in welfare state generosity in the current year *t* compared to welfare state generosity in the previous year. * p<0.05, ** p<0.01, two tailed test.

F. Description of Control Variables

Policy change may be driven by factors other than institutional arrangements. Governments, of course, can have clear preferences on policy that are distinct from those of the median voter. Right-of-center governments generally strive to reduce welfare effort while governments on the left increase it. To account for these factors, we include *Government ideology* from Seki and Williams (2014). The measure is based on party manifesto codes from the Comparative Manifesto Project (Volkens et al. 2017) and weighs the positions of governing parties based on cabinet shares. Welfare state generosity should be less under governments that score high on the left-right measure of government ideology. Second, economic performance may affect policy changes such that governments can afford to pursue more expansive policies during economic boom times and (often) are required to tighten the fiscal reins during downturns. Expanded models thus include *GDP growth* with the expectation that higher values will be associated with greater welfare effort. These data were retrieved from the Comparative Political Data Set (Armingeon et al. 2017, who in turn collected it from the OECD). From the same data set we also gathered data on *Unemployment* and *Trade openness* (imports and exports). (Higher) unemployment can create pressures on the welfare system because it creates (more) recipients of social programs. Over time this can, in turn, force the government to decrease generosity in order to sustain the higher number of recipients. Instead, trade can create demands in the electorate for compensation due to increased imports from abroad.

Results of these augmented models are reported below in Table S6. Consistent with expectations, higher growth leads to more generous welfare (because the economy becomes more able to sustain generous social programs). At least in some specifications, unemployment is also statistically significant and the effect is as expected: higher unemployment rates are bad news for welfare generosity. Surprisingly, instead, the coefficient estimate for government ideology doesn't reach statistical significance. The sign nonetheless goes in the expected direction and the lack of significance might be due to the contemporaneous presence of the median voter position. Being the two correlated, it is possible that the median voter simply absorbed the effect of government ideology. The effect of trade is more complex. In the short-run, trade dampens welfare regimes, but in the long-run compensatory forces prevail and more trade is associated with more generous welfare policies. However, these results are noisy across our sampled countries and the estimates do not achieve statistical significance. Pertaining instead to our research hypotheses, these effects do not bear on the estimates for the joint effects of the median voter and bicameralism described in the text.

G. Table S6: Bicameralism and Policy Responsiveness: Control Variables

	(1)	(2)	(3)
Welfare state generosity$_{t-1}$	-0.078**	-0.094**	-0.091**
	(0.018)	(0.020)	(0.020)
ΔMedian voter position$_t$	-0.941*	-0.903*	-0.862
	(0.449)	(0.453)	(0.456)
Median voter position$_{t-1}$	-1.140**	-0.901*	-0.795
	(0.388)	(0.397)	(0.406)
Index of bicameralism$_t$	-8.586*	-7.474*	-6.629
	(3.471)	(3.513)	(3.586)
ΔMedian voter position$_t$ × Index of bicameralism$_t$	1.821*	1.928*	1.838*
	(0.887)	(0.898)	(0.904)
Median voter position$_{t-1}$ × Index of bicameralism$_t$	1.664**	1.433*	1.294*
	(0.628)	(0.642)	(0.652)
ΔGovernment ideology$_t$	-0.006		-0.005
	(0.005)		(0.005)
Government ideology$_{t-1}$	-0.004		-0.004
	(0.003)		(0.003)
ΔGDP growth$_t$		0.049*	0.050*
		(0.024)	(0.024)
GDP growth$_{t-1}$		0.046	0.047
		(0.028)	(0.028)
ΔUnemployment$_t$		0.007	0.003
		(0.052)	(0.052)
Unemployment$_{t-1}$		-0.039*	-0.036
		(0.019)	(0.019)
ΔTrade openness$_t$		-0.014	-0.015
		(0.008)	(0.008)
Trade openness$_{t-1}$		0.004	0.004
		(0.004)	(0.004)
Constant	8.710**	7.756**	7.106**
	(2.207)	(2.242)	(2.301)
Observations	386	375	375
R-squared	0.158	0.190	0.193

Notes. Cells report least squares coefficients. Models include country fixed effects. The dependent variable is Δ*Welfare state generosity$_t$*. * p<0.05, ** p<0.01, two tailed test.

H. Table S7. Bicameralism and Policy Responsiveness: Alternative Cut-points for the Bicameralism Index

	(1)	(2)	(3)	(4)
Welfare state generosity$_{t-1}$	-0.071**	-0.092**	-0.084**	-0.076**
	(0.018)	(0.019)	(0.018)	(0.018)
ΔMedian voter position$_t$	-1.153*	-0.589	-0.433	-0.201
	(0.498)	(0.384)	(0.363)	(0.278)
Median voter position$_{t-1}$	-0.447	-1.125**	-0.949**	-0.449*
	(0.509)	(0.312)	(0.257)	(0.189)
Weak bicameralism$_t$	-0.832			
	(2.857)			
ΔMedian voter position$_t$ × Weak bicameralism$_t$	1.477*			
	(0.596)			
Median voter position$_{t-1}$ × Weak bicameralism$_t$	0.215			
	(0.542)			
Medium-weak bicameralism$_t$		-6.900**		
		(2.164)		
ΔMedian voter position$_t$ × MW bicameralism$_t$		0.784		
		(0.547)		
Median voter position$_{t-1}$ × MW bicameralism$_t$		1.306**		
		(0.402)		
Medium bicameralism$_t$			-7.009**	
			(2.015)	
ΔMedian voter position$_t$ × Medium bicameralism$_t$			0.616	
			(0.546)	
Median voter position$_{t-1}$ × Medium bicameralism$_t$			1.313**	
			(0.370)	
Strong bicameralism$_t$				-5.243*
				(2.609)
ΔMedian voter position$_t$ × Strong bicameralism$_t$				0.451
				(1.383)
Median voter position$_{t-1}$ × Strong bicameralism$_t$				1.025*
				(0.493)
Constant	4.596	9.050**	7.882**	5.041**
	(2.752)	(1.892)	(1.539)	(1.191)
Observations	386	386	386	386
R-squared	0.145	0.154	0.159	0.141

Notes. Cells report least squares coefficients. Models include country fixed effects. The dependent variable is Δ*Welfare state generosity$_t$*. Bicameral classifications are from Lijphart (2012). * p<0.05, ** p<0.01, two tailed test.

Table S7 replicates the model reported in Table 1 Model 3 for alternative cut-points on Lijphart's (2012) index of bicameralism such that *Weak bicameralism* is scored 1 for countries scoring 2 or above on the index if bicameralism and otherwise 0. *Medium-weak bicameralism* is scored 1 for countries scoring 2.5 or above on the index and otherwise 0. *Medium bicameralism* is scored 1 for countries scoring 3 or above on the index and otherwise 0. *Strong bicameralism* is scored 1 for countries scoring 4 on the index and otherwise 0.

I. Description of Dynamic Simulation in Figure 2

Figure 2 in the main text leverages the dynamic characteristics of our error correction model to forecast the effect of a given (one standard deviation) shift in the median voter's left-right position on government policy, as gauged by *Welfare state generosity*. To do so, we proceed as follows. First, following De Boef and Keele (2008), and given that our ECM model is estimated on stationary series, we re-specify our equation as follows:

$$\Delta Y_{it} = \alpha_0 + \alpha_1 Y_{(it-1)} + \beta_0 \Delta X_{it} + \beta_1 X_{(it-1)} + \varepsilon_{it} \quad \text{(s1)}$$

Which may be rewritten without difference operators as

$$Y_{it} - Y_{(it-1)} = \alpha_0 + \alpha_1 Y_{(it-1)} + \beta_0 X_{it} - \beta_0 X_{(it-1)} + \beta_1 X_{(it-1)} + \varepsilon_{it} \quad \text{(s2)}$$

Rearranging terms, we have

$$Y_{it} = \alpha_0 + (\alpha_1 - 1) Y_{(it-1)} + \beta_0 X_{it} + (\beta_1 - \beta_0) X_{(it-1)} + \varepsilon_{it} \quad \text{(s3)}$$

We can rewrite this as

$$Y_{it} = \alpha_0 + \alpha_1^{*} Y_{(it-1)} + \beta_0^{*} X_{it} + \beta_1^{*} X_{(it-1)} + \varepsilon_{it} \quad \text{(s4)}$$

where $\alpha_1^{*} = \alpha_1 - 1$, $\beta_0^{*} = \beta_0$, and $\beta_1^{*} = \beta_1 - \beta_0$.[27] We then use equation s4 to re-estimate our main models. We use these estimates to the long-run multiplier, which depicts the full effect of a change in an exogenous variable through all subsequent quarters in the series. We then simulate the over-time effects of *Median voter position* on the predicted levels of *Welfare state generosity* at *t* = 1, 2, 3,... Our forecasts depict two states of the world: one for unicameral systems and one for bicameral systems. *Welfare state generosity* is set to an initial value of 34 (the mean). Simulations are performed with STATA's dynsim package (Williams and Whitten 2011).

[27] De Boef and Keele (2008, 189-190) discuss the Bardsen transformation from the autoregressive distributed lag specification to the ECM specification; here, we perform the transformation in reverse.

J. Policy Responsiveness and Political Institutions

We re-estimate Table 1 Model 4 to consider a broad range of formal rules, and report results in Table S8 controls for vertical diffusion in Models 1-2 (i.e., federalism and regional authority from Hooghe *et al.* 2016), separation of legislative and executive powers (Model 3), judicial review (Model 4), majoritarian electoral rules (Model 5), assembly size (Model 6), and Lijphart's first and second dimensions of power-sharing institutions (Models 7-8).

The manuscript sets out arguments that more federal countries may influence policy responsiveness (see, e.g., Peters 2016). To address this possibility, we control for *Federalism* in Table S8 Models 1-2 that measure it relying on Huber et al. (2004; as in Armingeon et al. 2017) estimates for Model 1, and the regional authority index (*RAI*) developed by Hooghe et al. (2016) in Model 2. Since policy may be more responsive to the public in regimes which concentrate executive power, we also control for constitutional structures which have separate selection of executives in Model 3 (*Separate Executive*). Model 4 includes *Judicial Review* which is the presence of an independent body to review the legislature's decisions (based on Huber *et al.* 2004 and retrieved from Armigeon et al. 2017). If more powerful courts constrain policy change, we expect that governments in political systems that employ judicial review will be less responsive to the median voter position. Model 5 evaluates how majoritarian electoral rules may incentivize responsiveness to the public at large (Ferland 2020). If more disproportional rules foster opinion-policy linkages, then the interaction of *Disproportionality* and the median voter position should have a negatively-signed coefficient. *Disproportionality* is measured through the Gallagher (1991) index as a function of the difference between parties' seat shares from their vote shares, through the formula $\sqrt{1/2 \sum_{i=1}^{m} (v_i - s_i)^2}$, where *v* indicates vote share for party *i*, and *s* is its subsequent seat share in parliament. Data for disproportionality was retrieved from the Comparative Political Data Set (Armigeon et al. 2017). Model 6 includes *Assembly Size* which potentially matters for responsiveness, because larger assemblies render more proportional electoral outcomes with more parties to represent the preferences in the electorate (Hobolt and Klemmensen 2008; Taagepera and Shugart 1989; see also Wong 2018). Thus, we control for the size of the assembly relative to the country's population directly and interacted with the median voter position. *Assembly size* is the ratio between assembly size and population, and is the min-max feature rescaled with the formula $x_{scaled} = (x_i - x_{min}) / (x_{max} - x_{min})$. Data on assembly sizes is collected from *The World Factbook* [*The Wayback Machine* < https://tinyurl.com/9whw6rpt], and the Central Intelligence Agency [https://tinyurl.com/d77fpuye]. The data were updated for assembly size in 2014.

Finally, we include two broad indices of power-sharing arrangements developed and reported in Lijphart (2012: 304-309) for 1981-2010. It is possible that by focusing on bicameralism, the analyses overlook combinations of other power-sharing institutions that are potentially more powerful at explaining policy responsiveness. The *First "executive-parties" dimension* is based on Lijphart's factor analysis of five dimensions, including the effective number of parties; concentration of executive power in single party majority cabinets; electoral system disproportionality; interest group pluralism; and executive dominance (over the legislature).

The *Second "federal-unitary" dimension* is based on federal and decentralized government; bicameralism; constitutional flexibility; central banks that are independent of the executive; and levels of judicial review. We re-performed the factor analysis for this dimension of Lijphart (2012, 242), *omitting bicameralism*.

Estimates in Model 7-8 suggest that bicameralism – rather than the broader sets of political institutions, as measured by the executive-parties and federal-unitary dimensions -- conditions policy responsiveness to the median voter.

For all of the models described above, bicameralism's conditioning effect on government responsiveness remains intact.

Table S8. Policy Responsiveness and Political Institutions

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Welfare state generosity$_{t-1}$	-0.079** (0.018)	-0.107** (0.021)	-0.090** (0.020)	-0.076** (0.019)	-0.077** (0.019)	-0.098** (0.022)	-0.078** (0.018)	-0.079** (0.018)
ΔMedian voter position$_t$	-1.017* (0.451)	-0.691 (0.478)	-1.066* (0.501)	-0.852 (0.763)	-1.123* (0.470)	-1.958** (0.734)	-0.960* (0.456)	-1.186* (0.547)
Median voter position$_{t-1}$	-1.329** (0.405)	-1.251** (0.380)	-0.866 (0.465)	-1.335* (0.519)	-1.433** (0.440)	-0.034 (0.603)	-1.270** (0.377)	-1.067* (0.515)
Index of bicameralism$_t$	-11.360* (4.406)	-3.093 (4.959)	-7.917* (3.532)	-9.611** (3.393)	-10.260** (3.520)	-3.799 (4.180)	-9.360** (3.399)	-7.661 (4.608)
ΔMedian voter position$_t$ × Index$_t$	2.280* (0.962)	2.614* (1.179)	1.822* (0.884)	1.816* (0.911)	2.144* (0.927)	2.773** (1.006)	1.930* (0.883)	2.316* (1.049)
Median voter position$_{t-1}$ × Index$_t$	2.027* (0.810)	0.477 (0.984)	1.459* (0.659)	1.752** (0.621)	1.975** (0.643)	0.662 (0.786)	1.761** (0.624)	1.457 (0.876)
Federalism$_{t-1}$	1.084 (3.169)							
ΔMedian voter position$_t$ × Federalism$_{t-1}$	-0.695 (0.772)							
Median voter position$_{t-1}$ × Federalism$_{t-1}$	-0.265 (0.576)							
Regional authority index$_{t-1}$		-0.188 (0.119)						
ΔMedian voter position$_t$ × RAI$_{t-1}$		-0.033 (0.033)						
Median voter position$_{t-1}$ × RAI$_{t-1}$		0.043 (0.023)						
Separate Executive$_{t-1}$			4.101 (2.451)					
ΔMedian voter position$_t$ × SE$_{t-1}$			0.444 (0.590)					
Median voter position$_{t-1}$ × SE$_{t-1}$			-0.661 (0.465)					
Judicial review$_{t-1}$				-0.919 (2.759)				
ΔMedian voter position$_t$ × JR$_{t-1}$				-0.153 (0.688)				
Median voter position$_{t-1}$ × JR$_{t-1}$				0.093 (0.476)				
ΔDisproportionality$_t$					-0.053* (0.026)			
Disproportionality$_{t-1}$					-0.071 (0.136)			
ΔMedian voter position$_t$ × ΔDisprop.$_t$					-0.023 (0.131)			
Median voter position$_{t-1}$ × Disprop.$_{t-1}$					0.012 (0.027)			

Assembly size$_{t-1}$						10.780* (5.248)		
ΔMedian voter position$_t$ × Assembly size$_{t-1}$						2.507 (1.279)		
Median voter position$_{t-1}$ × Assembly size$_{t-1}$						-1.921* (0.820)		
First (executive-parties) dimension$_{t-1}$							0.432 (1.239)	
ΔMedian voter position$_t$ × First dim.$_{t-1}$							0.0826 (0.322)	
Median voter position$_{t-1}$ × First dim.$_{t-1}$							-0.149 (0.225)	
Second (federal-unitary) dim.$_{t-1}$								-0.423 (1.400)
ΔMedian voter position$_t$ × Second dim.$_{t-1}$								-0.388 (0.482)
Median voter position$_{t-1}$ × Second dim.$_{t-1}$								0.132 (0.264)
Constant	10.25** (2.433)	9.599** (2.135)	7.103** (2.421)	10.24** (2.986)	10.19** (2.493)	3.482 (3.482)	9.298** (2.143)	8.027** (2.802)
Observations	386	386	386	386	369	386	386	386
R-squared	0.156	0.180	0.163	0.157	0.160	0.187	0.155	0.157

Notes. Cells report least squares coefficients with standard errors in parentheses. Models include country fixed effects. The dependent variable is Δ*Welfare state generosity$_t$*. * $p<0.05$, ** $p<0.01$, two tailed test.

K. Immigration Policy

To get a sense of whether the generality of the results on welfare state generosity extend to other types of legislation, we examine government responsiveness on immigration policy. Immigration is appropriate for our purpose both because it is salient—in recent years immigration is consistently viewed as one of the most important issues in Europe (Böhmelt et al. 2020)[28]—and because it captures a dimension of issue contestation that is not coterminous with welfare state, left-right issues (Allen and Knight-Finley 2019).

Measures. As our dependent variable, *Immigration policy,* is a measure of national immigration policies, provided by the Immigration Policies in Comparison (IMPIC) database (Helbling et al. 2017).[29] Higher values correspond with more restrictive policies. An examination of government responsiveness across different political contexts requires that we have public opinion data over time. Although the Eurobarometer series furnishes such measures for left-right positions, it does not consistently collect policy opinions on immigration or other issue areas. Our analyses of immigration instead employ data from the European Social Survey series (ESS). We use the question, "Is [country] made a worse or a better place to live by people coming to live here from other countries?" Responses appear on a 0-10 scale, which we invert so that higher values represent less favorable attitudes and use the country-year mean as our measure of public opinion.[30] Measures for bicameralism are those used in Table 1. Lastly, models include a pair of controls which may influence some components of the outcome variable, depending on country. *Migrant and Refugee Population* is the total population size (or stock) of international migrants and refugees in a country, and *Population* is the country's midyear total population which counts all residents regardless of legal status or citizenship. For each of the measures, we enter the series in natural logs. Both series are from the World Development Indicators. With these measures, we construct a dataset of 150 country-year observations.[31]

Results. Table S9 reports results in four models using the same specification as above. We regress the policy measure, *Immigration policy*, on the mean voter's immigration preference for having more or fewer "people coming to live here from other countries." As with Table 1 above, the first two columns estimate models for unicameral and bicameral systems. In Model 1 the coefficient on the long-run effect for the lagged public opinion, *Median voter position*$_{t-1}$, is positively signed and precisely estimated, indicating that immigration policy responds in kind to annual shifts in public opinion.[32] Model 2 shows that for bicameral systems alone public opinion has no distinguishable impact on government policy.

[28] For instance, for the United Kingdom between 2003 and 2017, nearly 30 percent of the population viewed immigration as one of the two most important issues their country faces.
[29] Our choice of measures is informed by Böhmelt (2021).
[30] Note that the European Social Surveys are fielded in different years in different countries.
[31] Countries and years included are Belgium 2003-2012, 2015-2017; Denmark 2003-2011, 2014-2015; Finland 2009-2010, 2013-2017; France 2004-2011, 2014-2017; Germany 2003-2017; Ireland 2003, 2006-2007, 2010-2017; Israel 2003-20069, 2012-2013, 2016-2017; Netherlands 2003-2017; Norway 2003-2006, 2009-2014; Spain 2003-2009; Sweden 2005-2017; Switzerland 2003-2017; United Kingdom 2003-2017.
[32] Our findings again show that the transmission between public opinion and policy occurs over time rather than immediately.

The last pair of models report interactive specifications, with Model 3 interacting public opinion with the dichotomous *Bicameral* measure and Model 4 substituting in the ordinal *Index*. In both cases, the models return a positive estimate on *Median voter position*$_{t-1}$, but the negatively-signed interactions with bicameralism measures suggest this effect is weaker as the existence of (Model 3) or strength of (Model 4) the second chamber are taken into account. Given less extensive coverage, these analyses of responsiveness to immigration must be taken as suggestive rather than definitive. Nonetheless, the pattern is consistent with those we report here for social welfare policies: for immigration, policy government responsiveness to public opinion is stronger in unicameral systems than bicameral ones.

Table S9. Unicameralism, Bicameralism, and Policy Responsiveness to the Median Voter: Immigration Policy

	(1) Unicameral	(2) Bicameral	(3) All	(4) All
Immigration policy$_{t-1}$	0.067	-0.094	0.044	0.040
	(0.036)	(0.058)	(0.028)	(0.024)
ΔMedian voter position$_t$	0.014	-0.004	0.013	0.016*
	(0.007)	(0.008)	(0.007)	(0.007)
Median voter position$_{t-1}$	0.011*	-0.002	0.011*	0.017**
	(0.004)	(0.006)	(0.004)	(0.005)
Bicameral$_t$			0.026	
			(0.031)	
ΔMedian voter position$_t$ × Bicameral$_t$			-0.015	
			(0.009)	
Median voter position$_{t-1}$ × Bicameral$_t$			-0.007	
			(0.006)	
Index of bicameralism$_t$				0.072
				(0.037)
ΔMedian voter position$_t$ × Index$_t$				-0.023
				(0.014)
Median voter position$_{t-1}$ × Index$_t$				-0.020*
				(0.008)
Δ(ln)Migrant and Refugee Population$_t$	0.449*	-0.115	0.132	0.164
	(0.142)	(0.079)	(0.124)	(0.131)
(ln)Migrant and Refugee Population$_{t-1}$	0.020*	-0.036	0.013	0.019*
	(0.006)	(0.019)	(0.006)	(0.007)
Δ(ln)Population size$_t$	-0.647	0.205	-0.396	-0.687
	(0.461)	(0.420)	(0.369)	(0.441)
(ln)Population size$_{t-1}$	-0.019*	0.024	-0.012*	-0.016*
	(0.007)	(0.014)	(0.005)	(0.006)
Constant	-0.068	0.187	-0.058	-0.102*
	(0.055)	(0.101)	(0.040)	(0.040)
R-squared	0.261	0.090	0.170	0.195
N observations	86	64	150	150

Notes. Cells report least squares coefficients; parentheses report standard errors clustered by country. The dependent variable is Δ*Immigration policy$_t$*, calculated as the change in policy in the current year *t* compared to the previous year such that higher values are more restrictive (Helbling et al. 2017).
* p<0.05, ** p<0.01, two tailed test.

L. Synthetic controls analysis

In the manuscript we report results from a synthetic controls analysis. The objective is to construct a synthetic Belgium based on values from a set of matching variables in the pre-intervention period from a donor pool, and then to compare the effect of public opinion on changes in *Welfare State Generosity* in that synthetic Belgium (the control case, without institutional reform) to actual Belgium (the treated case, with institutional reform). To construct the synthetic case, we first identify a set of donor countries. Donor units must i) contain measures on all time-periods (i.e., balanced panel), ii) broadly share the characteristics of the treated case, and iii) not experience the treatment. Given these requirements, our donor pool is comprised of Austria, Denmark, Finland, France, Germany, Ireland, Italy, the Netherlands, Norway, Sweden, and the United Kingdom.[33]

The synthetic control method (Abadie, Diamond, and Hainmueller 2015) weighs contributions from each donor based on a set of matching variables. Accordingly, we select a set of covariates as informed by research on the predictors of welfare state size and social spending. Variables are as follows:

- GDP growth: Nominal GDP growth, percent change from previous year. Source: OECD Statistics
- Union density: Net union membership as percent of employees. Source: Armingeon *et al.* 2020.
- Government partisanship: Schmidt index of cabinet composition, such that 1 = hegemony of right/center parties, 2 = dominance of right/center parties, 3 = balance of power between right and left, 4 = dominance of left parties, 5 = hegemony of left parties. Source: Armingeon *et al.* 2020.
- Proportional representation: 1 = seats to lower house of legislature selected using proportional electoral rules, 0 = otherwise
- Percent elderly: percentage of population aged above 65 years. Source: Armingeon *et al.* 2020.
- Trade openness: KOF index of economic globalization, trade *de facto* measure. Source: Gygli *et al.* 2019.
- Self-rule index: The authority exercised by a regional government over those who live in the region. Source: Hooghe *et al.* 2016.

Based on these matching variables, we assign weights to donors such that the resulting synthetic Belgium minimizes the root mean square prediction error (RMSPE) over the pre-intervention period of 1978-1994. The weights for each country in the resulting synthetic Belgium are displayed in Table S10. In Table S11 we report the pre-reform attributes of Belgium, synthetic Belgium, and the mean of a 16-country OECD sample.[34] Note that for each of the matching

[33] The size of the pool is mainly constrained by the need to have enough pre-intervention time-periods for *Welfare state generosity*. Note that none of these cases experienced a reform to the organization of their national legislatures, i.e., none of the other cases received the treatment. We performed jackknife analyses as a check that findings are not driven by inclusion of a particular case.

[34] Countries are Australia, Austria, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, the Netherlands, New Zealand, Norway, Sweden, the United Kingdom, and the United States.

variables the weighted synthetic control better approximates the value for Belgium than does the16-country mean value.

Table 3 in the main text reports the difference in results between a model of *Welfare State Generosity* using values for Belgium and one using values for synthetic Belgium. The key result is that responsiveness improved during the post-reform period for Belgium but not for synthetic Belgium (the rolling regression results in Figure 4 provide corroborating evidence). It is possible, however, that these differences are driven by a specific donor. To explore this possibility, jackknife analyses are reported which drop one donor country from the synthetic control pool at a time, and then re-estimates parameters for Table 3 Model 3 in the manuscript. Table S12 reports results. The models' estimates do not support the finding that the relationship between mean voter position and welfare state generosity changes before and after the year of the Belgian constitutional reform had taken effect (1994). The analyses provide confidence that the differences observed in Table 3 are not due to inclusion of a specific donor.

Table S10: Synthetic Weights for Belgium

Country	Weight
Austria	0
Denmark	0.39
Finland	0
France	0
Germany	0.36
Ireland	0
Italy	0
Netherlands	0.23
Norway	0.02
Sweden	0
United Kingdom	0

Note: Weights are country weights assigned by the synthetic control method.

Table S11: Welfare State Generosity Predictor Means before the Reform

	Belgium	Synthetic Belgium	OECD Sample Mean
GDP growth	6.66	6.73	9.01
Union density	50.35	50.26	46.39
Government partisanship	2.16	2.25	2.36
Proportional representation	1.00	1.00	0.56
Percent elderly	14.44	14.38	13.09
Trade openness	75.11	48.44	42.52
Self-rule index	17.87	16.75	13.37

Note: Values are means from 1976-1994.

Table S12. Jackknife Analyses Re-estimating Table 4 Model 3, Omitting One Country at a Time from the Donor Pool

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
Donor country excluded in construction of synthetic control:	AUS	DNK	FIN	FRA	DEU	IRL	ITA	NLD	NOR	SWE	GBR
Welfare state generosity$_{t-1}$	-0.36**	-0.23*	-0.38**	-0.35**	-0.38**	-0.38**	-0.34**	-0.53**	-0.34**	-0.36**	-0.35**
	(0.12)	(0.09)	(0.12)	(0.12)	(0.12)	(0.12)	(0.12)	(0.13)	(0.12)	(0.12)	(0.12)
ΔMedian voter position$_t$	0.34	-0.30	0.26	0.36	0.12	0.25	0.35	0.63	0.35	0.28	0.35
	(0.58)	(0.38)	(0.50)	(0.55)	(0.49)	(0.52)	(0.57)	(0.59)	(0.57)	(0.56)	(0.54)
Median voter position$_{t-1}$	1.37*	0.12	1.24*	0.97	1.60**	1.28*	1.17	0.74	1.14	1.35*	0.98
	(0.61)	(0.38)	(0.51)	(0.54)	(0.47)	(0.53)	(0.59)	(0.54)	(0.58)	(0.58)	(0.54)
Post reform indicator	-1.38	-9.74	0.11	-6.49	7.42	0.82	-3.89	-9.62	-4.28	-0.29	-6.34
	(5.05)	(5.91)	(4.37)	(5.36)	(4.35)	(4.46)	(5.32)	(5.31)	(5.36)	(4.79)	(5.35)
ΔMedian voter position$_t$ × Post reform	-0.41	0.53	-0.39	-0.12	-0.46	-0.43	-0.20	-0.63	-0.18	-0.42	-0.12
	(1.09)	(0.80)	(0.95)	(1.06)	(0.93)	(0.98)	(1.10)	(1.15)	(1.10)	(1.04)	(1.06)
Median voter position$_{t-1}$ × Post reform	0.25	1.78	-0.03	1.21	-1.38	-0.16	0.71	1.94	0.78	0.05	1.18
	(0.95)	(1.10)	(0.82)	(1.01)	(0.82)	(0.84)	(1.00)	(1.01)	(1.01)	(0.90)	(1.01)
Constant	5.55	8.02	7.14	7.24	5.30	6.98	6.23	13.78**	6.31	6.01	7.20
	(3.71)	(4.83)	(3.81)	(3.95)	(3.81)	(3.84)	(3.82)	(4.88)	(3.83)	(3.73)	(3.95)
R-squared	0.30	0.48	0.33	0.31	0.40	0.33	0.30	0.43	0.30	0.32	0.31
N observations	33	33	33	33	33	33	33	33	33	33	33

Notes. Cells report least squares coefficients, with standard errors in parentheses. The dependent variable is Δ*Welfare state generosity$_t$* for synthetic Belgium, whereby synthetic Belgium is constructed by iteratively removing one country case from the donor pool. * $p<0.05$, ** $p<0.01$, two tailed test.

M. Bicameralism and Perceptions of Representation through Legislatures

We show in the main text that bicameralism matters for policymaker responsiveness. But are individual citizens attuned with how cameral structure matters for how their interests are translated into policy? What if citizens care little about policy but instead evaluate political outcomes in partisan or, alternatively, "valence" terms? A consideration of public opinion after the Italian referendum in 2016 suggests that the lack of knowledge of legislative structures did not drive voting decisions as over 60 percent of voters said they based their decision on the content of the reform, compared with less than 15% who voted to reward or punish the government (ITANES 2016). Do citizens, all else equal, feel less represented in bicameral systems?

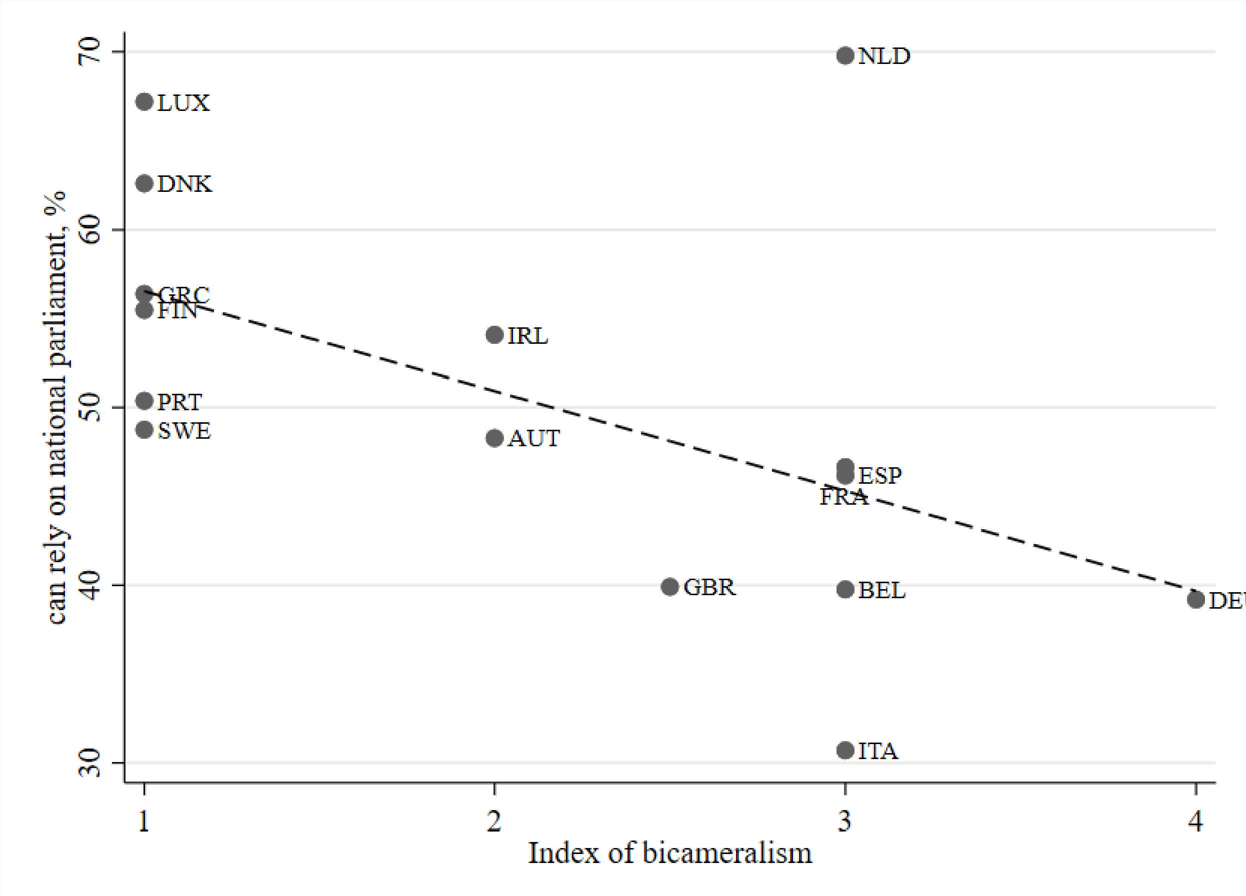
This section assesses how bicameralism affects mass perceptions of representation through legislatures. Although cross-national surveys do not regularly ask respondents their perceptions of legislative responsiveness, fortunately during the mid-1990s the Eurobarometer series posed such a question during the mid-1990s, a time period squarely within the time frame of our analysis. Respondents in the 15 Member States were asked whether they believed their national parliament could be relied on to make decisions "in the interest of people like yourself."[35] Consequently, we can examine whether – beyond actual responsiveness – bicameralism also influences representativeness perceptions. Figure S4 plots the percent in each country affirming that they can rely on the parliament to represent their interests (on the y-axis), against the Index of bicameralism (x-axis). The dotted line shows the bivariate regression coefficient, and the negative slope of which is statistically significant at $p = .039$.[36] Consistent with the Unicameral Hypothesis, citizens in countries that score *higher* on bicameralism report *lower* ratings of their national parliaments.

Table S13 reports the parameter estimates of individual-level logistic regression models with intercepts allowed to vary randomly by country-survey. The first model reports estimates from a binary specification; Model 2 also includes a measure of *Policy distance to government*, measured as the absolute distance between the respondent's position on the left-right scale and the *Government ideology* from Seki and Williams (2014). As we would expect, individuals who are located further (closer) from their governments on in terms of left-right positions are less (more) likely to feel their interests are represented. The effect of bicameralism, however, remains unchanged from the binary specification in Model 1. The final column adds a battery of individual controls for income, education, gender, age, and rural residence.[37] Again, *Index of bicameralism*

[35] The survey item reads in full: "To what extent do you feel you can rely on each of the following institutions to make sure that the decisions taken by this institution are in the interest of people like yourself? – the National Parliament." Countries included are Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, the Netherlands, Portugal, Spain, Sweden, and the United Kingdom. Data are from Clark and Rohrschneider (2019) drawing on Eurobarometers 44.1 (Nov-Dec 1995), 44.2bis (Jan-Mar 1996), and 47.2 (Apr-Jun 1997).
[36] The regression line is $Y = 62.2 - 5.6*$Index of bicameralism, $R^2=.29$, N=15. With Netherlands removed, it is $Y = 63.8 - 7.3*$Index of bicameralism, $R^2=.60$, N=14.
[37] Income is a dummy variable scored 1 for those in the upper two quintiles, education is scored 1 for those with some tertiary-level education or later, gender is scored 1 for female and 0 for male, age is age in years divided by ten, and rural is a dummy variable scored 1 for those self-identifying as living in rural areas.

exerts a negative impact on perceptions of interest representation. The supplementary information file reports additional specifications controlling for federalism, electoral system disproportionality, and semipresidentialism, and results remain unchanged. The picture at the individual level is consistent with that shown in the aggregate: individuals in countries with bicameral assemblies are less likely to believe that decisions taking by their national parliament are in the interest of people like themselves.

Figure S4. Bicameralism and Perceptions of Interest Representation through Legislatures, Country Percentages



Source: Clark and Rohrschneider (2019).

Table S13. Modelling Citizen Perceptions of Interest Representation of National Assemblies

	(1)	(2)	(3)	(4)	(5)	(6)
Index of bicameralism	-0.511**	-0.504**	-0.479**	-0.444***	-0.605***	-0.663***
	(0.165)	(0.165)	(0.165)	(0.096)	(0.096)	(0.101)
Policy distance government		-0.024**	-0.032**	0.004	-0.032***	-0.031***
		(0.005)	(0.005)	(0.008)	(0.005)	(0.005)
High income			0.189**	0.188***	0.133***	0.187***
			(0.014)	(0.014)	(0.023)	(0.014)
University education			0.148***	0.148***	0.147***	0.147***
			(0.014	(0.014)	(0.014)	(0.014)
Female			-0.206**	-0.205***	-0.206***	-0.206***
			(0.012)	(0.012)	(0.012)	(0.012)
Age			0.043***	0.042***	0.043***	0.034***
			(0.003	(0.003)	(0.003)	(0.006)
Rural			0.026	0.025	0.026	0.027
			(0.015	(0.015)	(0.015)	(0.015)
Index × Policy distance				-0.078***		
				(0.013)		
Index × High income					0.116**	
					(0.039)	
Index × Age						0.020*
						(0.010)
Constant	0.315**	0.350**	0.180*	0.145	0.211**	0.236**
	(0.085)	(0.085)	(0.087)	(0.078)	(0.078)	(0.080)
Random intercept	0.221**	0.220**	0.221**	0.183***	0.181***	0.183***
	(0.038)	(0.038)	(0.038)	(0.033)	(0.033)	(0.033)
Random slope				0.251***	0.255***	0.245***
				(0.051)	(0.053)	(0.050)
Log-likelihood	-81848.81	-81834.89	-81453.32	-81426.81	-81439.89	-81442.05
N individuals	122376	122376	122370	122370	122370	122370
N country-surveys	72	72	72	72	72	72

Notes: Cells report logit coefficients with random intercepts by country-survey and standard errors in parentheses. * p<0.05, ** p<0.01, two tailed test. Source: Clark and Rohrschneider (2019).

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