Radical moderation: recapturing power in two-party parliamentary systems*

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Abstract

We estimate the parameters of a reputational game of political competition using data from five two-party parliamentary systems. We find that latent party preferences (and party reputations) persist with high probability across election periods, with one exception: parties with extreme preferences who find themselves out of power switch to moderation with higher probability than the equivalent estimated likelihood for parties in government (extreme or moderate) or for moderate parties in opposition. We find evidence for the presence of significant country-specific differences. We subject the model to a battery of goodness-of-fit tests and contrast model predictions with survey and vote margin data not used for estimation. Finally, according to the estimated model parameters Australia is less than half as likely to experience extreme policies and Australian governments can expect to win more consecutive elections in the long-run as compared to their counterparts in Greece, Malta, New Zealand, and the United Kingdom.

Short title: Radical moderation

Keywords: Party Reputations; Policy Extremism; Two-Party Competition; Westminster Systems.

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1 Introduction

How fast can political parties shed a reputation for being out of line with the wishes of the pivotal segment of the electorate? In a two party system, the speed of this adjustment can largely determine the nature of political competition and the outcome of the policy process. If adjustment to a better reputation is uniformly slow, then political competition may remain lopsided for long periods of time once either party acquires a 'bad name' (intentionally or not). On the other hand, vigorous competition is possible if parties can instantaneously build a new reputation (e.g., through the ability to commit to a platform, as in stylized models of political competition a la Downs (1957)).

Casual empiricism suggests that reality lies somewhere in between the two extremes outlined above. For example, it took the British Labour party almost twenty years to effectively re-brand itself as the 'New Labour', a version of the left far more popular with "middle England" than the party that was defeated by Margaret Thatcher's Tories in 1979 or 1983. Similarly, largely as a consequence of the policy choices of the conservative Mitsotakis government of 1990-93, the Greek electorate remained unconvinced of the right-wing party's overt attempts to position itself as 'a party of the center' until the 2004 elections.

Despite its importance for the functioning and performance of two-party systems, it is quite hard to move beyond anecdotal evidence of the type above in order to quantify the persistence of party reputations. These reputations amount to the electorate's perceptions or beliefs about the parties' true policy preferences, and such beliefs are hard (or ex post impossible) to elicit systematically across a sufficiently long period of time.[1] Even if available, data on party reputations is not readily interpretable, as the evolution of these reputations is mediated by the strategic behavior of the parties that stand to benefit from successful manipulation of the electorate's perception of the party. Finally, beliefs about latent party preferences are derived from assessments of intangible or complex features of the competitive environment within parties, such as the relative power of various groups or coalitions of partisans, the rate at which senior party members retire or cease exercising influence within their party, etc. For that reason, we take the stance that the central question in this study can best be answered if empirical evidence is combined with the discipline imposed by theoretical equilibrium arguments. We pursue such a line of attack by structurally

[1] But, see Section 6 where we contrast the model's predictions of party reputations with survey data and electoral returns not used for estimation.

estimating a theoretical model of party reputations in this paper.

The theoretical model we use is due to Kalandrakis (2009) and features two parties that compete repeatedly and at any point in time may be controlled by either extremists or moderates. Because all party members benefit when the party has a reputation for being moderate, parties cannot perfectly convince the electorate that moderates hold the upper hand in the balance of power within the party and true party preferences are private information. The electorate, modeled as a pivotal voter, chooses between the two parties on the basis of their reputations, which evolve endogenously in equilibrium. In particular, voters dynamically update their beliefs about parties' extremism in light of past performance, the rate of persistence of latent party preferences when parties are in opposition or in government, and from the new information that arises from the policy process. The probabilities with which latent party preferences persist are exogenous in the model and constitute the central focus of our estimation efforts. In particular, party reputations cannot improve quickly unless latent party preferences shift at a sufficiently high rate. Thus estimates of these structural parameters, along with the equilibrium choices of governing parties, determine the overall speed with which party reputations adjust.

In ways more subtle than the intuition we outlined in the introductory paragraph, the (unique under relevant assumptions) equilibrium of the model exhibits radically different predictions regarding the pattern of alternation in office between the two parties, the expected duration of spells in office for the government, and the incidence of extreme policies. All of these forecasts depend on the values of the structural parameters that determine the persistence of latent party preferences. Importantly for our purposes, the equilibrium induces a likelihood over the observed sequence of victorious parties, which we use along with data from five countries with two-party parliamentary systems (Australia, Greece, Malta, New Zealand and the United Kingdom) in order to recover model parameters estimated with a Bayesian approach.

We find that latent party preferences (and reputations) are quite persistent across election periods: parties in government (extreme or moderate) and moderate parties in the opposition maintain their preferences across consecutive election periods with probabilities that exceed 80% (or with an annualized rate that exceeds 95%). Importantly, the exception to this pattern are extreme parties in opposition, which we estimate are almost half as likely to remain extreme and tend to be more likely to switch to moderate latent preferences than to remain extreme. These findings provide mixed support for the interpretation that political parties are able to strategically shift in order to achieve electoral success, as this incentive is triggered *only* when the parties are in opposition. Furthermore, in accordance with the anecdotal evidence we presented, the adjustment of opposition parties to moderation is neither instantaneous nor guaranteed: the hypothesis that the shift occurs with probability one can be rejected.

Despite the fact that these general patterns are shared by all countries in our sample, we find evidence for the presence of significant country-specific differences, notably, a distinct pattern of equilibrium electoral and policy dynamics in Australia. In the version of the model without electoral shocks, we find that a significant source for this difference can be traced to the smaller (estimated) value of government office in Australia compared to the remaining countries in our study. In versions of the model we estimate in which we allow for the possibility of electoral surprises occurring at a fixed rate across elections, we recover the Australian exceptionalism noted above, and obtain an alternative interpretation for its emergence; namely, we trace Australia's distinct pattern of electoral competition to the (estimated) tendency of Australian parties to remain/become moderate with higher probability while in government compared to their counterparts in the other countries.

We subject these models to a series of model checks comparing posterior predictive distributions with observed outcomes. Both when it comes to the predicted number of alternations in office and the balance of power between the two parties, p-values computed using posterior predictive distributions indicate a good fit. All versions of the estimated formal models outperform naive alternatives using a battery of goodness-of-fit statistics, with the model without electoral shocks having an edge over the models with electoral surprises.

We also discuss model predictions regarding quantities that are for our purposes unobserved, namely, the beliefs of the electorate about the true preferences within the two parties and the policies (moderate or extreme) implemented by the parties elected in government. With regard to implemented policies, all models predict relatively low probabilities of extreme policies with that probability increasing sharply toward the end of a party's spell in office. With regard to party reputations, we show that the predicted beliefs of the electorate about the true preferences within the two parties oscillate over time at least partly in line with what surveyed voters from the UK suggest (Section 6, Table 9). We also correlate differences in predicted posterior party reputations with the magnitude of the vote margin between the two parties across all five countries, and show that the model predictions consistently (although not uniformly) capture variation that was not originally used for the purposes of estimation (Section 6, Table 10).

Finally, we explore the predictions of the estimated models for the nature of political competition in the long-run. In all cases, the model predicts alternation of parties in government in the long-run, with Greece, Malta, New Zealand, and the United Kingdom experiencing a change of party in government once every two to three elections. In contrast, consistent with evidence pointing to significant country-specific differences, we estimate that in the long-term Australian governments can expect to stay in power for as many as five consecutive elections. We also estimate that in the long-run Australia is nearly half as likely to experience extreme policies compared to the remaining countries, while Greece and the UK provide the maximum long-run frequency with an extreme policy occuring in one of every seven or eight elections.

The connections of the theoretical component of our study with the formal theoretic literature of two-party competition are discussed by Kalandrakis (2009). Our study is also related to a large literature in comparative politics concerned with the logic by which political parties arrive at their policy positions (e.g., Budge, 1994), and with the measurement of these positions as in the Manifesto project (Budge, Klingemann, Volkens, Bara, and Tanenbaum, 2001). Besides our distinct methodological approach that marries observables with a game-theoretic model, our study is distinguished from this literature in that we are concerned with the persistence of political parties' real (but unobserved) policy preferences, as opposed to the *stated* preferences of political parties. Significantly, in our analysis the publicly announced party platforms or manifestos need not coincide with the reputation of that party. But, even though in the theoretical model government policies are the sole endogenously determined actions on the part of the parties that the electorate uses as 'hard evidence' to discern party preferences, our estimates of the rate of persistence of party reputations can be interpreted as a quantification of the real effect, if any, that party manifestos or other party actions that are omitted from the model may have on the electorate's perception about the prevailing preferences within the competing parties.

On the methodological front, the present paper follows in an established tradition of studies that combine a likelihood derived from the equilibrium of a game-theoretic model along with data for the purposes of estimation. Contributions in that vein using experimental data are ably reviewed in Palfrey (2006). Among studies that use observational data, this approach has been used in comparative politics (e.g., Diermeier, Eraslan, and Merlo, 2003; Merlo, 1997), international relations (e.g., Signorino, 1999; Smith, 1999), and American politics (e.g., Diermeier, Keane, and Merlo, 2005). Besides the distinct substantive focus, one difference of the present study from this literature is that the above models typically make use of action-specific preference shocks (as in McKelvey and Palfrey (1995)) in order to ensure that all actions have positive probability of occurring, whereas in the present study we do not rely on such shocks in order to rationalize the data. Also, among models that focus on some aspect of the political process in democracies the present model is, to our knowledge, the first in which both the electoral outcomes and policy-making are determined endogenously.

We organize our presentation as follows. First we briefly overview the theoretical model and state its equilibrium in Section 2. We then discuss how this equilibrium induces a likelihood over observed data, and we present our data and estimation strategy in Section 3. We report our main estimates in Section 4. In Section 5 we use the posterior predictive probabilities in order to evaluate the model's fit. We discuss predictions on unobservables in Section 6 and present estimated long-run dynamics in Section 7. We conclude in Section 8.

2 Model

Our estimation is based on a model of parliamentary systems of government in which two parties alternate in power controlling a legislative majority. In the model, two political parties, party L of the left and party R of the right, and an electorate interact over a sequence of periods indexed by $t = 1, 2, \ldots$. The electorate is modeled as a pivotal voter M. Each of the two parties contains individuals with two different ideological convictions, call them *moderates* and *extremists*. These two groups disagree as to the ideal government policy. In each period the balance of power within each party either tilts in favor of the moderates or in favor of the extremists. Thus, if extremists hold the upper hand in a party in some period, then we say that that party is an extreme type, e, in that period and it is a moderate type, m, otherwise. We call these types the parties' *latent* preferences.

We emphasize that these latent party preferences are not determined endogenously in the model and are not directly observable by the electorate. This is in contrast to models of political competition in which it is assumed that parties are able to commit to a future policy plan, for example, via the public announcement of a policy platform prior to the election. Of course, we do not rule out the possibility that parties can collectively take decisions that influence the power of moderates and extremists within their ranks. If such actions influence the probability that a particular party remains or becomes moderate, then that effect is captured by the estimated structural parameters of the model, as will become apparent in what follows. But any such actions, e.g., the announcement of a manifesto, or the election of a leader with a particular policy predisposition, are imperfect commitment instruments:[2] if extremists can prevail within the party after re-election there is nothing, except their own strategic calculation, that would prevent them from enforcing their will when it comes to government policy. Simply put, individuals within each party cannot choose their own preferences. Yet, as becomes clear in what follows, these latent party preferences do not determine the policy choices of the party: extremists can strategically choose which policy to implement if their party is elected in office.

In each period t the beliefs of the electorate about latent extremism within the two parties are summarized by a pair of probabilities $b^t = (b^t_L, b^t_R)$. For example, the electorate believes that party L is extreme in period t with probability b^t_L. Elections take place at the beginning of each period with voter choosing one of the two parties to be in government. We allow for the possibility that the outcome of the election is a surprise (i.e., different than the intended or preferred choice of voter M)[3] with probability $s < \frac{1}{2}$. Substantively, if elections are probabilistic, then governments with reputations for more policy extremism than oppositions might, in fact, win (and moderate governments might lose). The 'rally-round-the-flag' effect that returned the British Tories to power after the 1983 Falklands conflict is perhaps an example of the kind of electoral surprise that we wish to capture with this parameter. We denote the party elected to govern in period t by $P^t \in \{L, R\}$. Once elected, the party in government then chooses and implements a policy x^t that is publicly observed. There are three possible policies drawn from a set $X = \{x_L, x_M, x_R\}$. Moderate party types always pursue the moderate policy, x_M. If the government party P^t's type is extreme, it chooses between a moderate policy, x_M, or the extreme policy that corresponds to that party

[2] First, manifestos are promises about future policy choices; most electorates we are familiar with would not object to the association of politicians with broken promises. In fact, in recent work Adams, Ezrow, and Somer-Topcu (2011) find that electorates in a set of European democracies tend to discount or ignore political parties' platform announcements. Second, party leaders are ambitious individuals whose appetite for power may be the exact reason for their proffered (but imperfectly tested) moderate policy ideals.

[3] Such electoral surprises can be rationalized in the model as representing the unusual alignment of the choices of groups of voters different than the typically pivotal electorate captured by voter M.

$x_{P^t} \in X$. Thus, while parties cannot commit to a platform or choose their own preferences, they can strategically choose whether to pursue their ideal policy or follow a moderate policy in order to manipulate the party's reputation.

We now describe how party preferences change between periods. If a party is in government and is of type τ (extreme or moderate) in period t, then it is of the same type in period $t+1$ with probability π^g_τ. For example, if party L is in government in period t with extreme latent preferences, then that party is extreme in the next period with probability π^g_e and moderate with probability $1-\pi^g_e$.[4] We assume these changes are probabilistic due to the influx of a new generation of partisans with unknown preferences and the retirement of the 'old guard,' or due to vicissitudes in the level of influence that different individuals exercise within parties (but behind the scenes). Once more, we emphasize that these probabilistic shifts may also reflect the residual uncertainty in voter perceptions after parties have optimally taken any unmodeled actions (e.g., selection of suitable leaders, changes in party manifestos, etc.) that might affect extra-partisan players' perceptions. We denote the corresponding probabilities for a party in the opposition by π^o_τ. These probabilities, π^g_e, π^g_m for parties in government, and π^o_e, π^o_m for parties in the opposition, satisfy

$$\pi^g_e > \pi^o_e > 1 - \pi^o_m > 1 - \pi^g_m. \tag{1}$$

Note that $1-\pi^g_m$ and $1-\pi^o_m$ are the probabilities that a moderate party in government or the opposition, respectively, becomes extreme in the next period. Thus, in Inequality (1) we assume parties are extreme with higher probability (π^g_e or π^o_e) if the party was controlled by extremists in the previous period. The analogous statement is true for the case where the party was moderate in the previous period (e.g., $\pi^g_m > 1-\pi^o_e$). In addition to this serial correlation in latent party preferences, assumption (1) also states that parties are less likely to change type while in government compared to parties in the opposition ($\pi^g_\tau > \pi^o_\tau$). This is a natural assumption as parties in the opposition are more likely to undergo the kind of internal restructuring that results in an ideological shift within the party. On the other hand, the prevailing ideological group within a party in government is more likely to maintain control of the party.

All actors know that parties undergo these internal shifts in preferences according to the

[4] We preclude the singular case $\pi^g_e = 1$ that implies perpetual extremism and require that $\pi^g_e < 1$, so that there is some (possibly very small) probability that an extreme party in government switches to a moderate type.

above probabilities. Thus, given party reputations (b_L^t, b_R^t) and publicly observable information (P^t, x^t) for period t, the voter rationally updates beliefs about the probability that either party is extreme, and the parties do the same for the type of the opposition party, so that a new pair of reputation levels (b_L^{t+1}, b_R^{t+1}) corresponding to period $t+1$ is obtained. Then the voter elects a new party in government, P^{t+1}, given beliefs (b_L^{t+1}, b_R^{t+1}); this party then sets a new policy x^{t+1}, and so on. Thus, the interaction of the players in the game determines a sequence of publicly observable outcomes (P^1, x^1), (P^2, x^2), …, (P^t, x^t), …, as well as a sequence of beliefs or reputation pairs for the two parties b^1, b^2, …, b^t, ….

Extremists of each party receive a utility of one when they implement their favorite policy, a payoff of zero when the opposition party implements an extreme policy, and a payoff of one half when a moderate policy is implemented. In addition to these policy payoffs, extremists of each party receive an office benefit G when their party is elected in government, where

$$G \geq \frac{1 + s(\pi_e^o + \pi_e^g)}{2(1-2s)}. \tag{2}$$

Extreme parties care about the electoral and policy outcome in two periods, the current period t as well as period $t+1$, so that their payoff is the sum of the utilities accrued from these two periods. With condition (2) we restrict the analysis to the case partisans are sufficiently motivated by office. This is the interesting scenario, as when the value of winning office, G, is low, the game becomes strategically trivial in that extreme party types always pursue an extreme policy when in government. The voter strictly prefers the moderate policy over the two partisan policies, and is indifferent between the two extreme policies. The voter's payoff depends only on the policy outcome implemented immediately following the election, that is, partisans look further into the future, compared to the electorate, when making strategic decisions.

We focus on strategies in which both political parties and the voter choose among available actions by conditioning on the level of political competition as reflected by the reputations of the two parties. Note that party reputations are updated rationally given previous period' reputations as priors and given the publicly available information. Appendix A contains an exact statement of the equilibrium that satisfies these conditions on which we base our estimation. This equilibrium draws on the results in Kalandrakis (2009) where the interested reader can find a detailed discussion of equilibrium properties. To facilitate the substantive interpretation of our empirical findings we summarize key properties of the equilibrium in the following proposition:

Proposition 1 *The equilibrium is such that:*

1. *An extreme policy occurs with positive probability following an election with reputations b_L, b_R if either:*

   (a) *$|b_L - b_R|$ is sufficiently small and party reputations satisfy*

   $$b_L, b_R > b^* = \frac{\pi_m^g - \pi_m^o}{\pi_m^g - \pi_m^o + \pi_e^g - \pi_e^o}, \text{ or} \tag{3}$$

   (b) *An electoral surprise occurs (with probability s) electing a party with a sufficiently worst reputation than the opposition party.*

2. *Incumbent parties are re-elected with (weakly) higher probability when the value of office G is lower.*

3. *If inequality*

   $$\frac{1 - \pi_e^g}{1 - \pi_m^g} < \frac{1 - \pi_e^o}{1 - \pi_m^o}, \tag{4}$$

   holds, then government parties tend to lose any reputational advantage to the opposition party and extreme policies and alternation of the parties in office occur along the path of play even without electoral surprises.

4. *If inequality (4) is reversed, then the government party tends to enjoy a persistent reputational advantage to the opposition and extreme policies occur only after electoral surprises.*

According to part 1 of Proposition 1 an extreme policy occurs with positive probability either when a party with a reputational disadvantage is elected in government (after an electoral surprise has occurred with probability s), or when both parties have a reputation for being relatively extreme and neither party has a reputational disadvantage (whether an election surprise occurs or not). Part 2 concludes that incumbents survive with higher probability when office is less valuable. Finally, parts 3 and 4 state that, while the model induces a unique equilibrium for any single set of parameters, it produces two radically different long-term dynamics for different parameter values: one dynamic (when (4) holds) involves regular alternation in office and extreme policies, while the second implies an incumbency advantage and lower probability of extreme policies.

Given that different values of model parameters imply significantly distinct predictions, it is reasonable to ask whether we can use observed data in order to recover these parameters? To answer this question, in the following section, we show how the equilibrium of the model induces a likelihood over the observed data, as a function of the model parameters and describe an estimation method.

3 Data and Estimation Method

Consistent with the motivation for the model, we confine our empirical analysis to parliamentary systems of government in which two parties alternate in power controlling a majority of seats in parliament. Among two party parliamentary systems we focus on a subset of five developed countries,[5] namely Australia, Greece, Malta, New Zealand, and the UK. Excluded from the analysis are a large number of smaller former British colonies governed in the Westminster tradition.[6] Many of these countries feature too small a number of elections with a consolidated party system, while some among those with a larger number of potential observations, such as Jamaica and the Barbados, feature idiosyncratic forms of two party competition in which, for example, both of the two competing parties are Labour parties traditionally on the left of the policy spectrum.

Our data consist of a sequence of binary observations encoding the victorious party in the five countries included in the analysis. The available data series are listed in Table 1. The data series start immediately after WWII for all countries except Greece and Malta. In Greece, immediate post-WWII politics involved multiparty systems and political instability leading to a coup in 1967. Thus, the Greek data start with the 1977 election, the first election in which the first two vote receiving parties coincide with the two parties that governed the country following restoration of democracy in 1974. We include in the analysis all the Maltese elections starting with the first election following independence. With the exception of New Zealand, where the data series was interrupted in 1996 due to electoral reforms,[7] the remaining data extended to the present day.

[5]These uniquely meet the criterion of an advanced economy according to the IMF (2009) and high income countries according to the World Bank (2008).

[6]E.g., Bahamas, Barbados, Belize, Jamaica, St. Kitts & Nevis, St. Lucia, St. Vincent, etc.

[7]Both the UK and Greece experienced an apparent discontinuity in the data sequence in our sample because of an episode of inconclusive elections that occurred in each. In the UK this occurred in the first of the dual 1974 elections, while in Greece in the first two of the trio of elections in 1989-90. In these cases, we ignore the initial inconclusive elections that produced short inter-election periods, continuing the sequence with the first ensuing election that produced a single party majority in parliament. Similarly, we code the latest elections in the UK and Australia as

Thus, our data vary in length from a maximum of 26 periods (Australia) to a mere 10 elections in the Greek and Maltese cases. These data exhibit considerable alternation in power over time within each of the four countries. Note that competition between the two parties appears to be more balanced in Greece, Malta, and the UK, compared to New Zealand and Australia both of which seem to experience longer spells in power for the right party.

Australia		Greece		Malta		New Zealand		United Kingdom	
t	P^t	t	P^t	t	P^t	t	P^t	t	P^t
1946	L	1977	R	1966	R	1946	L	1945	L
1949	R	1981	L	1971	L	1949	R	1950	L
1951	R	1985	L	1976	L	1951	R	1951	R
1954	R	1990	R	1981	L	1954	R	1955	R
1955	R	1993	L	1987	L	1957	L	1959	R
1958	R	1996	L	1992	R	1960	R	1964	L
1961	R	2000	L	1996	L	1963	R	1966	L
1963	R	2004	R	1998	R	1966	R	1970	R
1966	R	2007	R	2003	R	1969	R	1974	L
1969	R	2010	L	2008	R	1972	L	1979	R
1972	L					1975	R	1983	R
1974	L					1978	R	1987	R
1975	R					1981	R	1992	R
1977	R					1984	L	1997	L
1980	R					1987	L	2001	L
1983	L					1990	R	2005	L
1984	L					1993	R	2010	R
1987	L								
1990	L								
1993	L								
1996	R								
1998	R								
2001	R								
2004	R								
2007	L								
2010	L								

Table 1: Election years and party in government.

Assuming observations from a particular country and from periods $t = 1, ..., T$, we denote the corresponding sequence of parties in government by a vector $\mathbf{P} = (P^1, P^2, \ldots, P^T)$. We also denote the (for our purposes unobserved) sequence of implemented policies by a vector $\mathbf{x} = (x^1, x^2, \ldots, x^T)$. We will first discuss the derivation of the likelihood from the model assuming the existence of the corresponding data on implemented policies $\mathbf{x}$, although for practical purposes we later treat these policies as unknowns to be included in the estimation. We use the notation $\mathbf{P}^t$ and $\mathbf{x}^t$, to indicate

victories of the Conservative and Labour parties, respectively.

a data sequence truncated up to period t, for example, $\mathbf{P}^t = (P^1, \ldots, P^t)$. We assume that these observations are generated from the equilibrium of the model that corresponds to a given set of (unknown) values for the exogenous parameters. For notational sanity, we compactly represent the exogenous parameters by (s, θ, b^1) where $\theta = (\pi_e^g, \pi_m^g, \pi_e^o, \pi_m^o, G)$. The parameters in θ are the primary focus of our estimation efforts. Parameters b^1 constitute the pair of the initial probabilities with which the two parties are drawn to be extreme (and also amount to the initial reputations of the two parties), and are estimated along with parameters θ.[8] Starting with period 1, we can compute the probability that party P^1 is elected given parameters s, θ, b^1. Also, conditional on P^1 being the elected party, we can similarly compute the probability that policy x^1 is chosen. Denote these probabilities $\mathsf{Pr}(P^1|s, \theta, b^1)$ and $\mathsf{Pr}(x^1|P^1, s, \theta, b^1)$, respectively. Now, given these choices it is possible to compute from the equilibrium the new party reputations that prevail in period 2, b^2. Using these reputations we can compute the conditional probability of observing election outcome P^2 and policy x^2 and, proceeding inductively in this fashion, we can compute the conditional probabilities $\mathsf{Pr}(P^t \mid \mathbf{P}^{t-1}, \mathbf{x}^{t-1}, s, \theta, b^1)$ and $\mathsf{Pr}(x^t \mid \mathbf{P}^t, \mathbf{x}^{t-1}, s, \theta, b^1)$ for the party and policy choice P^t and x^t in any period $t > 1$.

We can now combine the above probabilities in order to write the likelihood as a product of conditional probabilities as follows:

$$\begin{aligned}\mathcal{L}(s, \theta, b^1 \mid \mathbf{P}, \mathbf{x}) &= \mathsf{Pr}(P^1 \mid s, \theta, b^1)\mathsf{Pr}(x^1 \mid P^1, s, \theta, b^1) \times \\ &\prod_{t=2}^{T} \mathsf{Pr}(P^t \mid \mathbf{P}^{t-1}, \mathbf{x}^{t-1}, s, \theta, b^1)\mathsf{Pr}(x^t \mid \mathbf{P}^t, \mathbf{x}^{t-1}, s, \theta, b^1). \end{aligned} \tag{5}$$

As we explicitly show in Appendix B, all the equilibrium quantities that determine the probability with which parties choose extreme policies and the probability with which the voter elects either party to govern as well as the rule for updating party reputations are available in closed form as functions of the exogenous parameters. Thus, we can easily evaluate the likelihood in (5) for any complete data $\mathbf{P}, \mathbf{x}$ and any values of the parameters s, θ, b^1.

A major difficulty in our analysis is the fact that, unlike data on the sequence of the governing parties, it is much harder to objectively measure the nature of the policy choices, x^t, made by the governments in these countries. This is true even for the coarse binary distinction between

[8]Subsequent period reputations are not estimated but determined (rationally) in equilibrium as a function of parameters θ as specified in equation (10) in Appendix A.

extreme and moderate policies that we wish to make. Recall that a moderate policy in the model corresponds to the most preferred policy of a pivotal voter in the left-right policy dimension. Indeed, such moderate policies differ both over time and across countries, so that reasonable people may disagree about the classification (moderate or extreme) of even objectively identical policy bundles implemented by two different governments in two countries, or in different periods within the same country. Instead of pretending to be able to determine the nature of government policy choices across our diverse sample, we pursue an estimation strategy that consistently treats the implemented policies as unknown quantities generated from the equilibrium of the model.[9] In principle, there are various methods to implement this approach, but we opt for a data augmentation scheme (Tanner and Wong, 1987) which has a natural implementation in the context of the Markov Chain Monte Carlo (MCMC) Bayesian estimator we employ. The exact sampling scheme used is described in detail in Supplementary Appendix C. We discuss issues of identification of model parameters based on the observed data in Supplementary Appendix G.

Before we move to the next section, we note that the electoral shock parameter s is not among the parameters we estimate. Naturally, the parameter s can take different values, reflecting contrasting levels of stochastic fluctuation in the political environment. Yet, the fact that probabilistic elections introduce considerable noise in model predictions, does not permit us to estimate this parameter with the available data. An additional complication is that in order to ensure conditions for existence of the same equilibrium across different values of s, we must impose restriction (2) which affects the comparability of estimates of the value of office for different estimates of the parameter s, even if estimation of this parameter were obtained from our small sample. So in what follows we proceed by estimating the model for three fixed values for the probabilistic election parameter: $s = 0$, meaning no electoral shocks, $s = 0.1$, meaning that one election in ten is decided by factors other than electorate beliefs of relative extremism, and $s = 0.2$ which corresponds to the one in five elections case.

[9]This choice does not preclude the possibility of a more systematic attempt to obtain more information on policy choices in future research. Indeed, our estimation approach allows us to incorporate such evidence either directly or by use of appropriate prior specifications, where these priors may be elicited from country experts.

4 Results

Application of the estimation method outlined in the previous section allows us to draw several interesting conclusions. As a brief preview, on which we will expand substantially, we find that latent preferences within parties are quite stable over time, although extreme parties do tend to moderate after a spell out of office. Related to this observation, the value of government is generally high for partisans, though noticeably lower for Australian politicians in the model without electoral shocks ($s = 0$). We find evidence for significant country specific structural differences, and against pooled models in which groups of countries share structural parameters. Most notably in that regard, Australia emerges as a distinct case in our data. Finally, we find that on the basis of the available evidence the models without electoral shocks outperform their counterparts with electoral shocks (with the possible exception of Australia).

4.1 Persistence of latent party preferences and value of office

In Table 2 we report the parameter estimates for both the country-specific models (first five columns) and the models that pool observations across countries (last two columns).[10] Recall that π parameters are conditional probabilities of parties maintaining the same latent preferences; the superscript g or o refers to parties in government and opposition respectively, while m and e refer to their preferences/types—either moderate or extreme. Looking across the first two rows for all three levels of electoral shocks, we observe that the estimated probabilities for the persistence of preference of parties in government (π_e^g, π_m^g) are consistently at or above the 0.8 to 0.9 range, with the lowest value being $\pi_e^g = 0.69$ in Australia for the model with $s = 0.2$ (i.e., extreme parties in government tend to remain extreme with relatively lower probability in Australia according to the models with electoral shocks). These estimates are quite large,[11] implying parties in government maintain the same latent preferences with high probabilities across electoral periods.

Note that by assumption (1), parties in the opposition switch preference with higher probability than parties of the same type in government. We also estimate the probability that opposition parties remain moderate, π_m^o, to be consistently high. In the model without electoral shocks ($s = 0$)

[10] We report means of the relevant posterior. Medians of those posteriors yield essentially identical substantive conclusions, and they may be found in Supplementary Appendix D.

[11] For example, converting modal values of these per inter-election period probabilities to annual rates we obtain values between $0.95 - 0.98$.

		Australia	Greece	Malta	New Zealand	UK	Pooled	Pooled(4)
$s=0$	π_e^g	0.80	0.84	0.82	0.83	0.84	0.77	0.79
		(0.15)	(0.14)	(0.15)	(0.14)	(0.14)	(0.17)	(0.16)
	π_m^g	0.87	0.82	0.85	0.85	0.80	0.92	0.90
		(0.09)	(0.12)	(0.12)	(0.10)	(0.12)	(0.06)	(0.08)
	π_e^o	0.45	0.47	0.44	0.45	0.53	0.27	0.32
		(0.18)	(0.18)	(0.18)	(0.18)	(0.18)	(0.13)	(0.14)
	π_m^o	0.75	0.73	0.75	0.75	0.71	0.88	0.85
		(0.14)	(0.15)	(0.15)	(0.14)	(0.15)	(0.08)	(0.10)
	G	0.61	1.19	0.98	0.87	0.93	0.67	0.79
		(0.09)	(0.62)	(0.45)	(0.31)	(0.39)	(0.11)	(0.19)
$s=0.1$	π_e^g	0.73	0.85	0.83	0.84	0.85	0.80	0.81
		(0.18)	(0.13)	(0.15)	(0.14)	(0.13)	(0.16)	(0.15)
	π_m^g	0.87	0.79	0.82	0.82	0.79	0.90	0.87
		(0.11)	(0.14)	(0.13)	(0.12)	(0.13)	(0.08)	(0.09)
	π_e^o	0.52	0.51	0.54	0.53	0.55	0.36	0.40
		(0.21)	(0.19)	(0.21)	(0.20)	(0.19)	(0.17)	(0.18)
	π_m^o	0.65	0.68	0.69	0.70	0.68	0.83	0.80
		(0.19)	(0.17)	(0.18)	(0.17)	(0.17)	(0.11)	(0.12)
	G	1.42	1.58	1.46	1.24	1.26	0.91	1.10
		(0.89)	(0.81)	(0.79)	(0.51)	(0.53)	(0.15)	(0.29)
$s=0.2$	π_e^g	0.69	0.83	0.80	0.82	0.83	0.81	0.83
		(0.18)	(0.15)	(0.16)	(0.15)	(0.14)	(0.15)	(0.14)
	π_m^g	0.87	0.77	0.83	0.83	0.80	0.88	0.84
		(0.12)	(0.15)	(0.14)	(0.13)	(0.14)	(0.09)	(0.11)
	π_e^o	0.53	0.54	0.56	0.56	0.56	0.43	0.46
		(0.20)	(0.20)	(0.21)	(0.21)	(0.19)	(0.20)	(0.19)
	π_m^o	0.62	0.65	0.65	0.67	0.66	0.78	0.75
		(0.20)	(0.18)	(0.19)	(0.18)	(0.18)	(0.14)	(0.14)
	G	2.00	2.04	1.96	1.82	1.79	1.31	1.59
		(1.00)	(0.94)	(0.90)	(0.80)	(0.74)	(0.30)	(0.47)

Table 2: Parameter estimates (models with $s = 0$, $s = 0.1$, $s = 0.2$)

Point estimates are posterior means, with posterior standard deviations in parenthesis. The first five columns correspond to country-specific estimates. The 'Pooled' model pools observations across the five countries, and 'Pooled(4)' pools observations across the four countries, excluding Australia.

the minimum estimated value for that probability is 0.71 for the UK, and estimates rise as high as 0.88 in the case of the pooled model. Estimates for models with electoral shocks tend to be lower in the range between 0.6 and 0.7. On the other hand, the estimates reported in the third row of Table 2 corresponding to the probability that opposition parties remain extreme, π_e^o, are consistently lower. In particular, certainly for the pooled models and the model without electoral shocks, extreme parties in opposition are more likely to switch preferences to moderation (with probability $1-\pi_e^o$) than to remain extreme, with the UK being the only exception ($1-\pi_e^o = 0.47 < 0.53 = \pi_e^o$). Furthermore, the estimates of persistence of preferences for extreme parties in the opposition are (much) lower than the corresponding estimates for moderate parties, that is, $\pi_e^o < \pi_m^o$ across all models. No such pattern is apparent for the corresponding inequality ($\pi_e^g < \pi_m^g$) for parties in government, with the possible exception of Australia in the models with probabilistic shocks.

Despite the evidence from the point estimates reported in Table 2, a proper test of the hypothesis that the persistence of preference probabilities satisfy the inequality $\pi_e^o < \pi_m^o$ (or $\pi_e^g < \pi_m^g$) must take account of the correlation among these variables in the posterior distribution. We properly evaluate the evidence in support of these comparisons using Bayes factors, and report the relevant figures in Table 3. Note that the Bayes' factors, being the ratio of posterior to prior odds for the hypothesis, reflect additional evidence in favor/against the hypothesis obtained from the data (compared to the prior evidence). For the models without electoral shocks ($s = 0$) there is substantial evidence against the hypothesis that $\pi_e^o > \pi_m^o$ (with the exception of the UK) implying that, indeed, the claim that moderation is stickier than extremism is supported for oppositions. The evidence is strong(er) in the case of the pooled models for all levels of electoral shocks, but becomes weak in the country specific models with electoral shocks. By contrast, the claim about governments being more likely to remain extreme than moderate is merely supported, or there is 'minimal evidence' against it, for all the models with two exceptions: there is substantial evidence against this hypothesis (or evidence in favor of the hypothesis that extreme government parties switch preference more often than their moderate counterparts) in the case of the pooled model without electoral shocks (Bayes factor of 0.22 when $s = 0$), and for Australia with electoral shocks (Bayes factor of 0.29 when $s = 0.2$).

Over all, the estimates of the probabilities of party preference persistence in Table 2 portray a mixed picture of internal party politics. On the one hand, these estimates indicate that parti-

		Australia	Greece	Malta	New Zealand	UK	Pooled	Pooled(4)
$s = 0$	$\pi_e^o > \pi_m^o$	0.17	0.23	0.19	0.17	0.37	0.00	0.01
	$\pi_e^g > \pi_m^g$	0.60	1.42	0.91	0.98	1.66	0.22	0.40
$s = 0.1$	$\pi_e^o > \pi_m^o$	0.58	0.48	0.50	0.46	0.50	0.05	0.09
	$\pi_e^g > \pi_m^g$	0.43	1.87	1.11	1.49	2.06	0.45	0.69
$s = 0.2$	$\pi_e^o > \pi_m^o$	0.68	0.69	0.70	0.64	0.70	0.20	0.27
	$\pi_e^g > \pi_m^g$	0.29	1.69	0.93	1.19	1.79	0.64	1.20

Table 3: Bayes factors for comparison of preference persistence in moderate and extreme parties

The inequalities as expressed in the table can be treated as 'model 1' while the alternative, with the inequality reversed, is 'model 2'. Following a classification of Jeffreys (1961), a value B (or B^{-1} for model 2) can be interpreted as follows:
$1 > B \geq 10^{-\frac{1}{2}}$: Minimal evidence against model 1;
$10^{-\frac{1}{2}} > B \geq 10^{-1}$: Substantial evidence against model 1;
$10^{-1} > B \geq 10^{-2}$: Strong evidence against model 1;
$10^{-2} > B$: Decisive evidence against model 1.

san preferences are fairly persistent. On the other hand, parties with extreme preferences in the opposition tend to switch preferences with much higher probability, which sensibly suggests that opposition parties become more competitive after a spell out of power. Perhaps contrary to the first finding, the second finding suggests that there are forces within political parties that trigger or permit a possibly strategic shift that enhances the party's electoral prospects. Note that such (potentially) strategic moves only occur when the party is in the opposition and, even if political parties consciously attempt a shift in preferences from extremism to moderation, we find no support for the hypothesis that they can effect that shift with probability one: all estimates of the relevant parameter π_e^o are larger than zero.[12] Thus, it appears that the ability of the parties to strategically move towards moderation is tempered either by certain collective action failures within the party, or simply because parties do not possess the perfect institutional instruments that can bring about reform and a credible shift in the underlying party preferences.

Turning to the last row of Table 2 for each level of electoral shock, consider the estimates of the value of government (G). Recall that a value of G equal to one implies that party extremists would be willing for *their* party to enact the policies of the *opposition's* extremists so long as they could remain in government. That is, as G grows large, even hardcore party ideologues are willing to 'sell out' to hold power. According to our estimates, Greek politicians are keenest on office

[12]Specifically, the left boundary of the 95% highest posterior density or credible interval is well above zero in all cases with a minimum value of 0.11 for the country-specific estimates and 0.05 for the pooled models.

Model	$s = 0$	$s = 0.1$	$s = 0.2$
Country-specific	157.06	193.05	191.14
Pooled	183.32	247.58	277 13
Country-specific(4)	116.21	150.47	156.02
Pooled(4)	132.36	177.67	199.97

Table 4: Country-specific vs pooled estimation.

Deviance Information Criterion (DIC) for pooled and country-specific models. Lower DIC implies better fit. The country-specific models outperform the pooled alternative.

rents ($G = 1.19$ when $s = 0$, 1.54 for $s = 0.1$, and 2.04 when $s = 0.2$) with Malta, New Zealand, and the UK following in that order. Australia once more exhibits an unusual pattern. Australian politicians seem relatively uninterested in the trappings of high office ($G = 0.61$) in the model without electoral shocks. It is tempting to relate this low valuing of office to the fact Australia is the only country with a significant federal structure, which makes holding central office somewhat less valuable *ceteris paribus* than in a unitary state like Greece, Malta, or Britain. Nevertheless, the estimates of the value of office for Australia for the models with electoral shocks are more in par with those for the remaining countries. Finally, note that there is a notable increase in the estimated value of office across models as the electoral surprise parameter s increases. As discussed at the end of the previous section, this is largely a consequence of the parameter restriction (2), which makes the estimated values of G across models with different values of s not comparable.

4.2 Country-specific differences and electoral shocks

The differences in point estimates across countries, notably between Australia and the remaining countries, raises the question as to whether political competition across these countries is characterized by common structural parameters. To investigate this possibility we fit *pooled* models, wherein the structural parameters $\pi_e^o, \pi_m^o, \pi_e^g, \pi_m^g, G$ are assumed identical across a set of countries. While the model labeled 'Pooled' includes all five countries, model 'Pooled(4)' excludes Australia which, as we already discussed, appears to stand out from the remaining countries in our sample. As is to be expected, the pooled models that combine more data observations lead to sharper inferences as is evident by the smaller standard errors in Table 2, or the Bayes factors in the last two columns of Table 3.

Whether we should prefer this pooled arrangement requires a comparison of models with different numbers of parameters. One possible way to perform this comparison in the Bayesian framework of our study is the Deviance Information Criterion (DIC) (Spiegelhalter, Best, Carlin, and van der Linde, 2002).[13] This statistic is reported in Table 4. The country-specific estimation with five countries certainly does better than the pooled model (DIC difference is greater than 20) for all levels of the electoral shock parameter. The four-country pooled model (excluding Australia) is also outperformed by the corresponding country-specific model, although the contest is closer in this case in the models without electoral shocks ($s = 0$). Hence, we conclude that, though the pooled model is certainly theoretically plausible, there appears to be substantial evidence to the effect that national political idiosyncracies apparently account for the variation in the observed data, and that a significant component of this country-specific variation is associated with Australia.

Note that the potential explanations for the Australian exceptionalism we have identified differ depending on whether we assume no electoral shocks ($s = 0$), in which case the primary difference with Australia is lower estimated value of the prize that comes from winning power, or whether we assume positive electoral shocks, in which case an alternative explanation is the superior ability of Australian governing parties to shed a reputation for being extreme, or maintain a reputation for being moderate.[14] A comparison of these two models (with and without electoral shocks) is also possible using the DIC scores reported in Table 4. Recalling that a lower DIC implies a better fit, we see that the model without electoral shocks consistently outperforms models with positive electoral shocks by wide margins. In fact, although these numbers are not reported directly in Table 4, there is only one country-specific model for which the computed DIC is comparable or smaller for the models with electoral shocks ($s = 0.1, s = 0.2$) than the model without such shocks: Australia alone. Over all, with the possible exception of the Australian case, the performance of the models with electoral shocks appears inferior (in terms of the sharpness of inferences and the reported DIC) to the model without electoral shocks.

[13] For a model with data from any set of countries $\mathcal{C} \subseteq \{1, \ldots, C\}$, denote the deviance as $D(\theta, \{b_c^1, \mathbf{x}_c\}_{c\in\mathcal{C}}) = -2 \log \mathcal{L}(\theta, \{b_c^1, \mathbf{x}_c\}_{c\in\mathcal{C}} \mid \{\mathbf{P}_c\}_{c\in\mathcal{C}})$. DIC is the expected deviance, computed as the average of the posterior sample of size M, $\bar{D} = \frac{1}{M}\sum_{i=1}^{M} D(\theta_i, \{b_{c,i}^1, \mathbf{x}_{c,i}\}_{c\in\mathcal{C}})$, plus a penalty for the effective number of parameters, p_D. Thus $DIC = \bar{D} + p_D$, with the implication that models with lower DIC are preferred. We approximate p_D by $p_V = \frac{\mathrm{var}(D(\theta,\{b^1,\mathbf{x}_c\}_{c\in\mathcal{C}}))}{2}$ (see Gelman, Carlin, Stern, and Rubin (2004)), which is very straightforward to calculate from the posterior sample.

[14] The reader might speculate that either of these differences is due to the larger number of observations we have available for Australia (26 compared to 17 for the immediately smaller sample size). To dispel this possibility, we re-estimated all models using only the first 17 or only the last 17 Australian observations. These estimates are reported in Supplementary Appendix E and are almost identical to those reported for the complete 26 period data; otherwise put, more data is not the explanation for Australia's distinctiveness.

5 Model Fit

In this section we scrutinize the model's fit by comparing observables with model predictions. In particular, we use the posterior predictive probability distribution over hypothetical replications of the data assuming they are generated from the estimated model. We draw a large number of replicated sequences of the data, $\mathbf{P}^{rep}$, i.e., the sequence of election winners in the sample period, using this posterior predictive distribution, $P(\mathbf{P}^{rep} \mid \mathbf{P})$, and then we compute p-values for various statistics based on the value of this quantity observed in the data.[15] The two statistics that we focus on are (1) the number of alternations in office, and (2) the number of electoral victories of the Left. These two statistics provide a good summary of the data and the level of party competition.

	Australia	Greece	Malta	New Zealand	UK	Pooled	Pooled(4)
Observed	6	5	4	7	7	29	23
$s = 0$	0.7514	0.6983	0.7608	0.7108	0.7159	0.6125	0.6316
$s = 0.1$	0.2784	0.6030	0.5503	0.6352	0.6984	0.5788	0.5864
$s = 0.2$	0.4570	0.4896	0.4468	0.5003	0.6407	0.6044	0.5844

Table 5: p-values from posterior predictive distribution of number of alternations in office

	Australia	Greece	Malta	New Zealand	UK	Pooled	Pooled(4)
Observed	10	6	5	5	8	34	24
$s = 0$	0.8484	0.2429	0.5609	0.9794	0.8327	0.9184	0.8635
$s = 0.1$	0.7162	0.3276	0.6179	0.9473	0.7953	0.9055	0.8574
$s = 0.2$	0.7637	0.3624	0.6027	0.9123	0.7389	0.8941	0.8415

Table 6: p-values from posterior predictive distribution of number of victories by Left party

We report these calculations in Tables 5 and 6. As usual, p-values that are too close to zero or one suggest that the data are unlikely to have arisen from the model and would be a cause for concern. Looking at the p-values corresponding to the predictive distribution of the number of alternations in government, we see that the model does very well, with none of the models coming close to a p-value that would raise alarm given conventional levels of significance. Turning to the number of victories for the Left party, we see a similar picture with the only possible exception being New Zealand. In that case the model tends to overpredict the number of victories of the

[15] Specifically, for each one of the 10,000 parameter configurations in the posterior sample, we replicate one data sequence $\mathbf{P}^{rep}$ of election winners for the T periods in the sample $\mathbf{P}$. Let $f(\mathbf{P})$ be some statistic of the data. We compute p-values by evaluating the probability $\mathsf{Pr}(f(\mathbf{P}^{rep}) \geq f(\mathbf{P}))$. This methodology is standard (e.g., see Gelman, Carlin, Stern, and Rubin (2004), chapter 6).

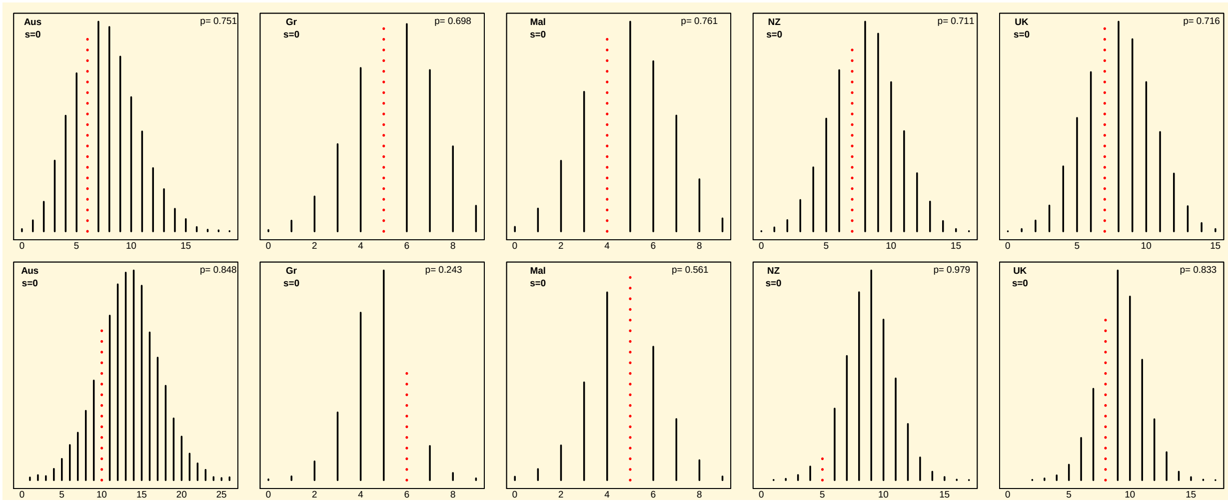

Figure 1: Posterior predictive distribution of number of alternations in office and number of victories by the Left and corresponding p-values ($s = 0$ model)

First row corresponds to number of alternations, second row corresponds to number of Left party victories. The observed number of alternations is indicated with red dotted lines. p-values appear in the top right corner of each graph. Posterior predictive distributions are computed using $10,000$ replicated data sequences using parameters from the posterior sample of the country specific models.

Left compared to the observed number of such victories in the sample, notably, in the case of the highest p-value of 0.9794 for the New Zealand model with $s = 0$.[16] Note that it does not follow that the model cannot accommodate protracted periods of asymmetric competition between the two parties (even though the two parties are treated symmetrically in the model), as the p-values for the Australian data indicate reasonable fit. In Figure 1 we provide a graphical depiction of the entire posterior predictive distribution for the country-specific models without electoral surprises, in order to allow the reader to obtain a clearer idea of the range of predictions allowed by the posterior sample. Note that in many cases the observed value is at or near the mode of the posterior predictive distribution.

Moving beyond these two aggregate statistics summarizing the observed data, we now turn to an evaluation of the model's fit with regard to individual elections. We, once more, use the posterior predictive distribution and record goodness-of-fit with standard criteria such as the 'proportion of elections correctly predicted' and the metrics designed by Efron (1978) (pseudo-R^2) and Cramer

[16] If we compute the p-value as the posterior predictive probability of *lower* or equal number of Left victories ($\mathsf{Pr}(f(\mathbf{P}^{rep}) \leq f(\mathbf{P}))$ instead of $\mathsf{Pr}(f(\mathbf{P}^{rep}) \geq f(\mathbf{P}))$), then the p-value is 0.0514.

		$s=0$	$s=0.1$	$s=0.2$	Static	$s=0$ (adj.)	$s=0.1$ (adj.)	$s=0.2$ (adj.)	AR1a	AR1b
Australia	CorP	0.81	0.81	0.77	0.62	0.80	0.80	0.76	0.76	0.76
	R^2	0.55	0.35	0.28	0.00	0.52	0.32	0.26	0.21	0.23
	λ	0.40	0.37	0.29	0.00	0.37	0.37	0.29	0.25	0.23
Greece	CorP	0.90	0.90	1.00	0.60	0.89	0.89	1.00	0.56	0.67
	R^2	0.45	0.30	0.14	0.00	0.34	0.21	0.05	-0.11	0.03
	λ	0.35	0.20	0.09	0.00	0.25	0.17	0.08	0.02	0.03
Malta	CorP	1.00	0.80	0.70	0.50	1.00	0.78	0.67	0.56	0.56
	R^2	0.37	0.29	0.15	0.00	0.29	0.27	0.14	0.00	0.01
	λ	0.26	0.17	0.10	0.00	0.18	0.16	0.10	0.01	0.01
NZ	CorP	1.00	0.88	0.65	0.71	1.00	0.88	0.63	0.56	0.75
	R^2	0.26	0.14	-0.01	0.00	0.13	0.03	-0.12	-0.31	0.01
	λ	0.26	0.14	0.06	0.00	0.17	0.12	0.05	-0.01	0.01
UK	CorP	1.00	1.00	0.88	0.53	1.00	1.00	0.88	0.56	0.56
	R^2	0.45	0.32	0.18	0.00	0.41	0.27	0.14	0.00	0.02
	λ	0.30	0.19	0.11	0.00	0.26	0.16	0.09	0.02	0.02
Pooled	CorP	0.87	0.91	0.91	0.59	0.86	0.90	0.90	0.56	0.64
	R^2	0.32	0.25	0.16	0.05	0.30	0.24	0.16	0.00	0.11
	λ	0.20	0.15	0.09	0.05	0.18	0.14	0.09	0.01	0.11
Pooled(4)	CorP	0.71	0.70	0.69	0.60	0.69	0.68	0.67	0.63	0.68
	R^2	0.31	0.26	0.18	0.04	0.29	0.25	0.17	0.07	0.15
	λ	0.21	0.17	0.12	0.04	0.19	0.16	0.12	0.09	0.15

Table 7: Goodness-of-fit of formal model and naive alternatives

CorP: Fraction correctly predicted. R^2: Efron's (1978) pseudo-R^2. λ: Cramer's (1999) λ. The Static model predicts according to observed frequency, AR1a and AR1b according to observed frequency conditional on previous period's government. AR1a assumes probability of alternation is identical for Left and Right parties, while AR1b is based on different probabilities of alternation for the two parties. For comparability with AR1a and AR1b, the first observation of each data series is not included in the computation of these statistics in the (adj.) columns corresponding to the formal model. 'Pooled(4)' refers to estimates with observations pooled across four countries, excluding Australia.

(1999) (λ). Unlike the p-values reported above with which we aimed to flag out any discrepancies between the model and observations, these statistics constitute cardinal measures of the model's fit. In order to provide benchmarks for comparison, we also compute these measures for three heuristic non-strategic or naive models. We label the first of the three models 'static' as it predicts the same probability of victory of the Left across periods based on the frequency of such victories in the data. In the second naive model, 'AR1a', the probability of victory of the Left (Right) is computed on the basis of empirical frequency conditional on the identity of previous period's winner (i.e., $\Pr[P^t = L|P^{t-1} = L] \neq \Pr[P^t = L|P^{t-1} = R]$) but maintains the same probability of alternation in government for the two parties ($\Pr[P^t = L|P^{t-1} = L] = \Pr[P^t = R|P^{t-1} = R]$). Finally, in the third version, 'AR1b', we maintain the dependence of victory probabilities on the identity of the winner in the previous election but compute separate probabilities of alternation for the two parties

($\Pr[P^t = L | P^{t-1} = L] \neq \Pr[P^t = R | P^{t-1} = R]$). Supplementary Appendix E provides details of these measurement strategies and the competing alternatives to the model.

Table 7 reports the relevant figures; the fit for each of the three models ($s = 0, 0.1, 0.2$) is given in columns 1 to 3 and 5 to 7, and the appropriate contrast is to columns 4, and 8 to 9, respectively. A larger number implies a better fit and, as is obvious from the table, the models with levels of electoral shocks $s = 0$ and $s = 0.1$ outperform the naive alternatives in *every one* of the twenty-one pairwise comparisons. When the probability of electoral shocks is $s = 0.2$ there are two exceptions (New Zealand and pooled(4) models), but even in these rare cases where the naive models do better, the absolute difference is very small. Of note, the pooled formal models outperform the dynamic naive model, even though we use country-specific predictive probabilities for the naive model (i.e., we do not impose the restriction that conditional predictive probabilities for the dynamic naive model must coincide across the countries. See Supplementary Appendix E). The reader may wonder what accounts for the superior performance of the formal model compared to the AR1 models. The main difference stems from the fact that the AR1 model's predictive probability is constant during a party's spell in office once that party is in power. On the contrary, in the formal model the likelihood of electoral success of the incumbent party changes systematically as that party's spell in office becomes longer. This dynamic adjustment becomes evident by our discussion of the evolution of government policies and reputations in the next section.

6 Predictions on unobservables

The model may also be judged with respect to quantities that are not part of the data used for estimation. In this section we consider first the policies (moderate or extreme) implemented by the governments in our sample; second, the reputations of the two political parties, that is, the beliefs of the electorate about the true preferences prevailing within parties in each period. Precisely because we lack consistent direct measures of these quantities, this operation is more difficult and our analysis more discursive here. Nevertheless, we are able to bring more 'hard data' to bear on our evaluation of the model when it comes to party reputations which we contrast with survey data from the UK and historical vote margins across the five countries in our study.

We report posterior estimates of party reputations and posterior probabilities of extreme government policies for the UK models in Table 8. Space considerations do not allow us to report

	$s = 0$			$s = 0.1$			$s = 0.2$			Left-Right		Extremism	
t	b_L^t	b_R^t	x^t	b_L^t	b_R^t	x^t	b_L^t	b_R^t	x^t	Lab	Con	Lab	Con
1945	0.39	0.69	0.00	0.40	0.65	0.08	0.41	0.61	0.14	–	–	–	–
	(0.22)	(0.22)		(0.24)	(0.25)		(0.26)	(0.27)					
1950	0.42	0.45	0.38	0.45	0.47	0.32	0.46	0.47	0.28	–	–	–	–
	(0.17)	(0.17)		(0.21)	(0.18)		(0.23)	(0.19)					
1951	0.57	0.41	0.00	0.56	0.44	0.02	0.54	0.45	0.05	–	–	–	–
	(0.28)	(0.18)		(0.27)	(0.19)		(0.27)	(0.20)					
1955	0.44	0.41	0.00	0.46	0.45	0.07	0.46	0.46	0.12	–	–	–	–
	(0.19)	(0.19)		(0.20)	(0.20)		(0.20)	(0.22)					
1959	0.40	0.40	0.34	0.43	0.46	0.33	0.45	0.48	0.28	–	–	–	–
	(0.18)	(0.18)		(0.20)	(0.22)		(0.21)	(0.25)					
1964	0.39	0.53	0.00	0.43	0.55	0.02	0.44	0.53	0.05	–	–	–	–
	(0.18)	(0.29)		(0.20)	(0.28)		(0.21)	(0.28)					
1966	0.41	0.42	0.13	0.45	0.45	0.19	0.46	0.46	0.20	–	–	–	–
	(0.19)	(0.20)		(0.21)	(0.21)		(0.23)	(0.21)					
1970	0.44	0.39	0.08	0.50	0.43	0.10	0.50	0.45	0.11	–	–	–	–
	(0.24)	(0.19)		(0.26)	(0.20)		(0.27)	(0.21)					
1974	0.39	0.42	0.37	0.44	0.47	0.32	0.45	0.48	0.25	–	–	0.46†	0.41†
	(0.19)	(0.22)		(0.21)	(0.23)		(0.22)	(0.24)				(0.12)	(0.12)
1979	0.55	0.39	0.00	0.55	0.43	0.01	0.53	0.45	0.03	–	–	–	–
	(0.29)	(0.19)		(0.28)	(0.20)		(0.28)	(0.21)					
1983	0.43	0.40	0.00	0.46	0.44	0.02	0.46	0.45	0.07	3.24*	5.78*	0.56	0.55
	(0.20)	(0.19)		(0.21)	(0.21)		(0.22)	(0.22)		(1.97)	(1.77)	(0.23)	(0.23)
1987	0.40	0.40	0.00	0.43	0.44	0.08	0.45	0.46	0.13	–	–	0.56	0.54
	(0.19)	(0.19)		(0.20)	(0.21)		(0.21)	(0.24)				(0.23)	(0.23)
1992	0.39	0.39	0.38	0.43	0.46	0.36	0.44	0.48	0.31	–	–	0.33	0.35
	(0.19)	(0.19)		(0.20)	(0.23)		(0.21)	(0.26)				(0.21)	(0.22)
1997	0.39	0.55	0.00	0.42	0.57	0.01	0.44	0.54	0.03	3.97	7.16	0.21	0.45
	(0.19)	(0.30)		(0.20)	(0.29)		(0.21)	(0.29)		(2.39)	(2.64)	(0.14)	(0.22)
2001	0.40	0.43	0.00	0.44	0.46	0.05	0.46	0.47	0.09	5.00	6.61	0.20	0.42
	(0.19)	(0.20)		(0.21)	(0.21)		(0.22)	(0.22)		(1.91)	(2.20)	(0.14)	(0.22)
2005	0.40	0.40	0.25	0.45	0.43	0.26	0.47	0.45	0.24	4.85	6.80	–	–
	(0.19)	(0.19)		(0.22)	(0.20)		(0.24)	(0.21)		(2.16)	(2.06)		
2010	0.49	0.39	0.10	0.52	0.43	0.10	0.52	0.44	0.09	–	–	–	–
	(0.27)	(0.19)		(0.28)	(0.20)		(0.28)	(0.21)					

Table 8: Posterior estimates of party reputations and policy choices in the UK, and BES data on party reputations.

Columns for party reputations b_L^t, b_R^t report posterior means with standard errors in parenthesis. Column on policies x^t reports fraction of extreme policy choices in posterior sample of 10,000. Left-Right columns report means (with standard errors in parenthesis) of BES survey respondents' placement of parties on a $0-10$ left-right scale (* 1983 data converted to a 0-10 scale from the original 21 point scale). Extremism columns report means (with standard errors in parenthesis) of BES survey responses on whether each party is extreme (1), moderate (0), or neither/both (0.5) († October 1974 data computed using a five option version of this question).

these posterior estimates for all countries, but we make them available in Supplementary Appendix H. With regard to model predictions on government policies, two general patterns emerge from Table 8 that also generalize for the remaining countries. First, posterior probabilities of extreme policy in any given period are generally small (specifically, ≤ 0.38 for the UK). Second, and perhaps more important, is the clear monotonic pattern of increasing probability of extreme policy choices during a party's tenure in office. In the models with electoral surprises, there is positive but small probability of extreme policy choice in most all periods, and this probability increases steadily and becomes significant only toward the end of a party's spell in office. In the model without electoral shocks extreme policies are estimated with positive probability *only* in the last period of a party's spell in office.

This pattern is closely connected with the evolution of parties' reputations implied by the equilibrium as we explain using Labour's last spell in office in the UK from 1997 to 2010 as an illustration. According to Table 8, Labour enjoyed an advantage over the Conservative party in 1997 when it first came to power (e.g., posterior reputation of 0.39 versus 0.55 for the Tories in the model without electoral shocks). This is typical as parties that first win elections enjoy (on average) a better reputation than opposition parties. Given this advantage, it does not pay even for party extremists to sacrifice almost certain reelection in order to pursue an extreme policy, so that early policy choices are likely to be moderate. But this initial reputation advantage of the governing party tends to gradually vanish (again looking at the case $s = 0$ from Table 8, both parties' reputations equal 0.39 by the 2005 election).[17] Without a reputation advantage, extremists, if they control the party, sacrifice less by choosing their preferred (extreme) policy (since they face a close election they are no longer guaranteed to win). Thus, according to the posterior estimates, the last Labour government of Gordon Brown was much more likely to pursue a left-wing policy in the 2005-2010 interelection period (with a probability of about 0.25) compared to the early Labour governments of Tony Blair. Skeptical readers may think such a pattern is at best a stretch of reality. This is *a fortiori* true with regard to the previous spell in office by the Conservative party, for which the model predicts a low probability of extremism in the initial terms in office for Thatcher's governments and a higher probability for the (last) John Major government. Of course, we cannot evaluate a model by cherry-picking the cases it does or does not predict well.

[17] This is typical in the estimated models, but see Proposition 1 and the next section's test of inequality (4).

Regrettably, we are limited in our ability to perform a more rigorous evaluation of the overall performance of the model without more objective measures of these government policies' appeal to the median electorate across time.

	Left-Right		Extremism	
	Lab	Con	Lab	Con
$s = 0$	0.56†	0.68	0.56	-0.03
$s = 0.1$	0.13†	0.66	0.50	-0.18
$s = 0.2$	-0.10†	0.64	0.48	-0.28
Manifesto	0.75	-0.40	0.74†	0.48

Table 9: Correlation between mean survey responses and mean posterior party reputations and left-right manifesto data for the UK.

Column variables (Left-Right and Extremism, Lab and Con) corrspond to survey data reported in the last four columns of Table 8. Manifesto refers to left-right party placements reported from the manifesto data (Budge, Klingemann, Volkens, Bara, and Tanenbaum, 2001).
† These correlations are multiplied by -1 so that positive correlations indicate good performance for the model or the manifesto data.

	Australia			Greece			Malta			New Zealand			United Kingdom		
s	0	0.1	0.2	0	0.1	0.2	0	0.1	0.2	0	0.1	0.2	0	0.1	0.2
All t	0.56	-0.06	-0.19	0.90	0.80	0.73	0.74	0.71	0.58	0.71	0.76	0.72	0.60	0.56	0.51
$P^t = 1$	0.62	-0.57	-0.58	0.85	0.93	0.96	0.48	0.49	0.23	0.76	0.92	0.87	0.54	0.52	0.52
$P^t = 0$	0.30	-0.48	-0.46	0.70	0.03	-0.18	0.54	0.37	-0.42	0.45	0.51	0.48	-0.44	-0.42	-0.41

Table 10: Correlation between observed vote margins and mean posterior difference in party reputations.
Correlation of difference in vote shares between the party of the left and the party of the right, with the posterior mean difference in the reputation of the right party and the left ($b_R^t - b_L^t$). First row correlations are computed using all periods, second row using only elections in which the Left won ($P^t = 1$), third row using using only elections in which the Right won ($P^t = 0$).

We can bring additional evidence to evaluate the model more systematically in the case of party reputations. First, we look at the average responses of voters to two survey questions included in the British National Election Studies (BES) from 1974 to 2005. The first question asks voters to place the two political parties on a left-right scale; the second asks whether they believe either party to be extreme. Both quantities (directly or indirectly) capture voter's perception of the two parties' extremism. Though we have relatively little data here, we do note that a positive correlation exists between these mean opinions and the average posterior reputations estimated from the model (see Table 9), more so for the model without electoral shocks, with the notable exception of the low and negative correlation between the posterior reputation of the Conservative party and the survey responses on the extremism of this party in the BES data (values of -0.03 for the model with $s = 0$ up to -0.28 for the model with $s = 0.2$). To offer some perspective on these figures, we also report in Table 9 the correlation of the left-right Manifesto scales (Budge, Klingemann, Volkens, Bara, and Tanenbaum, 2001) with these survey data. Perhaps ironically, the manifesto data correlate quite well with the survey question on parties' extremism, but exhibit a negative correlation of -0.40 in the case of the voters' placement of the Conservative party on a left-right scale.

As a final check on the model's performance in capturing party reputations, we studied the correlation between the vote margin (the difference in received vote shares between the Left and the Right parties) in the elections in our sample and the posterior difference in party reputations. Our reasoning in looking at these correlations is straightforward. If the posterior party reputations from the model recover (in part) the beliefs of the electorate about parties' true preferences, then bigger differences in the reputations of these parties should translate to (on average) wider margins of victory for the party with the better reputation. These correlations are reported in Table 10. In most all cases,[18] these correlations are positive and large indicating that differences in posterior party reputations indeed capture cardinal differences in the sentiment of the electorate for the two parties. In fact, the correlations remain positive and of similar size when we compute them separately on the subset of the data for which the Left or the Right won (second and third rows of Table 10). This is notable as the data we used to estimate the model cannot possibly contain any information on the margin of the electoral victories of the Left (Right), conditional on the Left (Right) winning the election.

7 Long-run dynamics

We now turn our attention to the implications of the reported estimates for long-term policy and electoral dynamics. Recall that from parts 3 and 4 of Proposition 1, we expect distinct patterns of party alternation in government and extreme policy choices depending on whether inequality (4) holds or not. We test whether this inequality holds using Bayes factors which are reported in Table 11. The data substantially (if not overhwelmingly) support the hypothesis that inequality (4)

[18] Important exceptions are the Australian models with electoral shocks, and the UK models in elections won by the Conservatives. The split in the Labour Party, and the rise in electoral popularity of the breakaway Social Democratic Party, lead to an unusually low vote share for Labour in the 1983 and 1987 UK elections. This largely accounts for the negative correlation in the latter period. This aberrant performance of the model is also consistent with the negative correlations in the last column of Table 9.

holds across both pooled and country specific models, with one exception: there is strong evidence against inequality (4) in Australia when the level of the electoral shocks is $s = 0.2$ (and weak evidence to that effect in the model with $s = 0.1$). Thus, according to parts 3 and 4 of Proposition 1, in the countries of our sample, with the possible exception of Australia, we find evidence that both extreme policies and alternation of parties in government are a regular equilibrium phenomenon (and not the consequence of electoral surprises) in the long-term. On the other hand, Australian politics may exhibit a strong incumbency advantage and low incidence of policy extremism in the long run.

	Australia	Greece	Malta	New Zealand	UK	Pooled	Pooled(4)
$s = 0$	0.00	0.00	0.03	0.00	0.00	0.00	0.00
$s = 0.1$	2.06	0.02	0.21	0.04	0.00	0.00	0.00
$s = 0.2$	16.90	0.15	0.74	0.39	0.15	0.03	0.00

Table 11: Bayes factors for long-term dynamics (inequality (4))

Low values indicate support for inequality (4). See Table 3 for interpretation.

With that evidence in mind, we use the model and the posterior sample for each specification in order to estimate the *long run* probability of an extreme policy. This is done by simulating long sequences of equilibrium play (2,000 periods) for the model corresponding to each point in the posterior sample separately, and then averaging the frequency of extreme policies across periods (discarding the first 1,000 periods as burn-in). The relevant predictions are displayed in Table 12. These calculations suggest none of the nations under study are exceedingly likely to undertake radical policies in the long-run, with the frequency of such policies ranging roughly between one in every two hundred to one in every seven elections. Notice that the model predicts Australia has the lowest long-run probability of witnessing such extreme policies— 0.57 every ten elections in the model without electoral surprises and as low as 0.05 every ten elections in the model with $s = 0.2$. The highest of these values is considerably lower—almost half—than that of the nearest country estimate, which is Malta. The highest probabilities occur in Greece and the UK, with predictions that range between one in eight and one every seven elections across levels of electoral shocks.

Using the same procedure, we generate long term predictions for the expected number of party alternations in government for the various countries. Table 13 contains the relevant figures. In keeping with the pattern we encountered in Table 12, notice that Australia stands out from

	Australia	Greece	Malta	New Zealand	UK	Pooled	Pooled(4)
s=0	0.57	1.21	0.95	0.93	1.25	0.51	0.71
	(0.45)	(0.99)	(0.88)	(0.72)	(0.85)	(0.42)	(0.56)
s=0.1	0.18	1.51	0.95	1.06	1.36	0.63	0.90
	(0.39)	(1.20)	(1.01)	(0.85)	(0.95)	(0.50)	(0.68)
s=0.2	0.05	1.56	0.80	0.92	1.36	0.71	1.10
	(0.21)	(1.45)	(1.13)	(1.03)	(1.21)	(0.57)	(0.81)

Table 12: Extreme policies in the long-run

Posterior predicted average number of extreme policies per ten elections in the long-run. Standard errors in parenthesis.

	Australia	Greece	Malta	New Zealand	UK	Pooled	Pooled(4)
s=0	2.89	5.86	5.06	4.84	5.12	3.90	4.73
	(0.84)	(1.31)	(1.61)	(1.10)	(1.09)	(0.57)	(0.70)
s=0.1	1.73	5.64	4.50	4.77	5.07	3.95	4.75
	(1.15)	(1.31)	(1.89)	(1.21)	(1.05)	(0.56)	(0.69)
s=0.2	2.15	5.02	3.84	4.14	4.72	4.03	4.78
	(0.52)	(1.46)	(1.66)	(1.45)	(1.30)	(0.61)	(0.65)

Table 13: Long-run predicted number of alternations in office

Posterior predicted average number of alternations of party in government per ten elections in the long-run. Standard errors in parenthesis.

the remaining countries with its governments enjoying much longer life in power, with roughly two alternations in office every ten elections or an average duration in office of five periods compared to, say, over five government alternations in Greece for every ten elections amounting to average spells in office that last short of two full periods. Most other countries exhibit a number of alternations that is intermediate between those encountered in Greece and Australia, with most all cases being closer to the former rather than the latter.

8 Conclusion

The ability of parties to rebuild or maintain a good reputation regulates political competition in two-party parliamentary systems, and this paper sought to explore these reputation dynamics by matching data with an equilibrium model. We found that latent party preferences are sticky, suggesting the presence of significant inertia within party organizations, but at the same time found that extreme parties that lose elections are able to switch to moderate preferences relatively swiftly (but not with probability one), a finding consistent with the presence of strategizing forces within political parties. Despite these common general patterns, we found significant country-specific differences, with an especially distinct pattern of political competition emerging for Australia.

The theoretical model fits the observed data well as documented by various statistics based on the posterior predictive distributions. Looking at posterior probability estimates of extreme policies by governments in the sample, we found that parties tend to suppress any latent policy extremism at the beginning of a government's spell in office, while governing parties are increasingly likely to show their true colors as they reach the end of their spell in office. The posterior estimates of party reputations (the unobserved beliefs of the electorate about extremism within parties) correlate well with data (not used for estimation) on the electoral performance of the parties.

Overall, our estimates suggested a healthy level of long-run equilibrium competition between the two political parties in the countries of study, with reasonably short spells of control of the government by each party (with the exception of Australia), and predominantly moderate policies in the long-run. We view these findings as a first step in a fruitful research direction. Future research can benefit significantly from the enrichment of both the theoretical model, and the data that is used for estimation. When it comes to the theoretical model, much richer patterns of policy-making are possible if we incorporate finer distinctions in the policy preferences of various party groups. In terms of data, we would seek to add more information about the policy choices of the governments.

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APPENDIX

A. Equilibrium

In this appendix, we introduce necessary notation and state the exact form of the equilibrium of the model described in section 2. A strategy for the extreme type, e, of party $P \in \{L, R\}$ is a function $\sigma_P : [0,1]^2 \to [0,1]$ that maps the pair of party reputations to the probability of an extreme policy choice. A strategy for the voter is a function $\sigma_M : [0,1]^2 \times \{L, R\} \to [0,1]$ that maps the pair of party reputations and the identity of the incumbent party in government to a re-election probability. An equilibrium is a trio of strategies $\sigma = (\sigma_L, \sigma_R, \sigma_M)$ that satisfy sequential rationality, and a belief updating rule

$$b' : [0,1]^2 \times \{L, R\} \times X \to [0,1]^2, \tag{6}$$

that maps reputations b^t and observables P^t, x^t in period t to updated reputations $b^{t+1} = b'(b^t, P^t, x^t)$ in period $t+1$ and is obtained via Bayes rule whenever possible.[19] In addition to the above standard equilibrium conditions, we focus on equilibria in which the strategy of the voter takes the following intuitive form:

$$\sigma_M(b, P) = \begin{cases} 1 & \text{if} \quad b_P < b_{-P} \\ 0 & \text{if} \quad b_P > b_{-P}. \end{cases} \tag{7}$$

Kalandrakis (2009) refers to such equilibria as *responsive*. The following is a straightforward extension of Propositions 3 and 4 in Kalandrakis (2009):

Proposition 2 *Every* responsive *equilibrium is such that:*

1. *The strategy of extreme type of party P is given by*

$$\sigma_P(b) = \begin{cases} \dfrac{\mathcal{T}_g(b_P) - \mathcal{T}_o(b_{-P})}{b_P(\pi_e^g - \mathcal{T}_o(b_{-P}))} & \textit{if} \quad \mathcal{T}_g(b_P) > \mathcal{T}_o(b_{-P}) \\ 0 & \textit{otherwise.} \end{cases} \tag{8}$$

2. *The strategy of voter M satisfies (7) and*[20]

$$\sigma_M(b, P) = \frac{1 + s\pi_e^g\big(\sigma_R(b_R, \pi_e^g) - \sigma_R(b)\big) + (1-s)b_R\big(\sigma_R(b) - \sigma_R(\pi_e^g, b_R)\big)}{(1-2s)\big(2G + \sigma_R(b)(\pi_e^g + b_R)\big)}, \textit{ if } b_P = b_{-P} \in [1 - \pi_m^o, \pi_e^o]. \tag{9}$$

3. *The updated reputation of party P following a period with government party P' and implemented policy x satisfies*

$$b'_P(b, P', x) = \begin{cases} \mathcal{T}_o(b_P) & \textit{if} \quad P \neq P' \\ \mathcal{T}_g\left(\frac{(1-\sigma_P(b))b_P}{1-\sigma_P(b)b_P}\right) & \textit{if} \quad P = P',\ x = x_M, \textit{ and } b_P < 1 \\ \mathcal{T}_g(1) & \textit{if} \quad P = P' \textit{ and } x = x_P. \end{cases} \tag{10}$$

The functions $\mathcal{T}_o, \mathcal{T}_g$ in Proposition 2 are defined as $\mathcal{T}_o(b) = \pi_e^o b + (1 - \pi_m^o)(1 - b)$ and $\mathcal{T}_g(b) = \pi_e^g b + (1 - \pi_m^g)(1 - b)$, respectively.

[19] There is minimal need for additional restrictions on out-of-equilibrium beliefs, namely players infer that the type of the government is extreme whenever they observe an extreme policy.

[20] In deriving this expression, use is made of the fact that $\sigma_L(b', b'') = \sigma_R(b'', b')$.

B. Likelihood

In this appendix, we explicitly derive the likelihood function used in our estimator. Note that due to (7) and the fact that $b^1_L \neq b^1_R$ the probability of observing a government by party P^1 in period 1 is given by

$$\mathsf{Pr}(P^1|s,\theta,b^1) = \begin{cases} 1-s & \text{if} \quad b^1_{P^1} < b^1_{-P^1} \\ s & \text{if} \quad b^1_{P^1} > b^1_{-P^1}. \end{cases} \tag{11}$$

Since it is possible that $b^t_L = b^t_R$ in periods $t = 2,\dots,T$, the probability of a government by party P^t in period $t > 1$ is

$$\mathsf{Pr}(P^t \mid \mathbf{P}^{t-1}, \mathbf{x}^{t-1}, s, \theta, b^1) = \begin{cases} (1-s)\sigma_M(b^t, P^{t-1}) + s(1-\sigma_M(b^t, P^{t-1})) \text{ if } P^{t-1} = P^t \\ s\sigma_M(b^t, P^{t-1}) + (1-s)(1-\sigma_M(b^t, P^{t-1})) \text{ if } P^{t-1} \neq P^t. \end{cases} \tag{12}$$

The function σ_M that appears in (12) is defined in (9), while the party reputations, b^t, that enter as its arguments are obtained inductively using the initial exogenously given reputations b^1 by repeated application of (10).[21] We can similarly obtain the probability that the governing party, P^t, implements policy x^t in period $t = 1,\dots,T$, as

$$\mathsf{Pr}(x^t \mid \mathbf{P}^t, \mathbf{x}^{t-1}, s, \theta, b^1) = \begin{cases} b^t_{P^t}\sigma_{P^t}(b^t) & \text{if} \quad x^t = x_{P^t} \\ 1 - b^t_{P^t}\sigma_{P^t}(b^t) & \text{if} \quad x^t = x_M, \end{cases} \tag{13}$$

and we use the convention of rewriting the probability of the policy choice in the first period by $\mathsf{Pr}(x^1|P^1,s,\theta,b^1) = \mathsf{Pr}(x^t|\mathbf{P}^1,\mathbf{x}^0,s,\theta,b^1)$. The function $\sigma_{P^t}(b^t)$ in (13) is defined in (8). We can now combine the probabilities in (11), (12), and (13) in order to write the likelihood as expressed in (5). The likelihood in (5) can be used for estimation purposes with data from a single country. If data from a total of C countries is available, we index countries by $c = 1,\dots,C$ and denote data from the c-th country as a vector $\mathbf{P}_c$ and $\mathbf{x}_c$. If we assume that all countries in a subset $\mathcal{C} \subseteq \{1,\dots,C\}$ share the same structural parameters θ, but not necessarily the initial reputations— which we denote by b^1_c for the c-th country—then we can pool data for estimation purposes using (5) to obtain the combined likelihood

$$\mathcal{L}(s,\theta,\{b^1_c\}_{c\in\mathcal{C}} \mid \{\mathbf{P}_c,\mathbf{x}_c\}_{c\in\mathcal{C}}) \quad = \quad \prod_{c\in\mathcal{C}} \mathcal{L}(s,\theta,b^1_c \mid \mathbf{P}_c,\mathbf{x}_c). \tag{14}$$

[21] Note that (12) specifies $\sigma_M(b,b)$ only for values of $b \in [1-\pi^o_m, \pi^o_e]$ which is sufficient since the reputation of the opposition party is confined in $[1-\pi^o_m, \pi^o_e]$ in all periods $t > 1$ by (10).

SUPPLEMENTAL APPENDIX

C. MCMC implementation details

We constructed priors starting with independent uniform distribution in the open unit interval for the probabilities π^g_τ, π^o_τ, and b^1 . For the value of office parameter G we assume a vague independent Gamma prior distribution with unit parameters, $\Gamma(1,1)$. These distributions are then truncated according to the inequality restrictions (1) and (2).[22] For any model comprising data from a subset of countries $\mathcal{C} \subseteq \{1,\ldots,C\}$, we use a Gibbs sampling scheme to obtain a sample from the posterior distribution, i.e., we specify initial values for the parameters $\theta^0, \{b^{1,0}\}_{c\in\mathcal{C}}$, and the unobserved policies $\mathbf{x}^0$, and then at the m-th iteration, we first sample $\theta^m, \{b^{1,m}\}_{c\in\mathcal{C}}$ from the posterior distribution of the parameters conditional on policies $\mathbf{x}^{m-1}_c, c \in \mathcal{C}$. We then sample from the distribution of policies $\mathbf{x}$ by successively sampling individual policies $x^{t,m}_c, t = 1,\ldots,T_c, c \in \mathcal{C}$, conditional on the parameter values $\theta^m, \{b^{1,m}\}_{c\in\mathcal{C}}$ and the remaining policies. We use the Metropolis algorithm in order to sample from the conditional distribution of the parameters $\theta, \{b^1\}_{c\in\mathcal{C}}$ running 100 inner iterations of the Metropolis algorithm in order to get a sample from the distribution of $\theta, \{b^1\}_{c\in\mathcal{C}}$ conditional on policies in the m-th outer iteration. We adjusted suitable (normal) driver densities for these Metropolis steps using preliminary runs. Then we ran the chains for $1,100,000$ outer iterations each. We removed the first $100,000$ as 'burn-in,' and we thinned the remaining $1,000,000$ outer iterations, taking one in every 100 posterior draws in order to reduce dependence, thus obtaining a sample of $10,000$ from which the relevant statistics were computed. These choices were rather conservative compared to the recommended 'burn-in' and 'thinning' choices suggested by standard convergence diagnostics such as Raftery and Lewis (1992).

D. Parameter Estimates: Medians

Table 14 below is identical to Table 2 in the main text, but report posterior medians instead of posterior means.

[22] Due to this truncation, the parameters are not independently distributed *a priori*.

		Australia	Greece	Malta	New Zealand	UK	Pooled	Pooled(4)
$s=0$	π_e^g	0.84	0.88	0.85	0.86	0.87	0.81	0.83
		(0.15)	(0.14)	(0.15)	(0.14)	(0.14)	(0.17)	(0.16)
	π_m^g	0.89	0.84	0.87	0.87	0.82	0.94	0.91
		(0.09)	(0.12)	(0.11)	(0.10)	(0.12)	(0.06)	(0.08)
	π_e^o	0.44	0.46	0.43	0.45	0.54	0.25	0.30
		(0.18)	(0.18)	(0.18)	(0.18)	(0.18)	(0.13)	(0.14)
	π_m^o	0.77	0.75	0.77	0.77	0.73	0.89	0.86
		(0.14)	(0.15)	(0.15)	(0.14)	(0.15)	(0.08)	(0.10)
	G	0.59	1.02	0.86	0.80	0.84	0.66	0.77
		(0.09)	(0.62)	(0.45)	(0.31)	(0.39)	(0.11)	(0.19)
$s=0.1$	π_e^g	0.75	0.88	0.87	0.88	0.88	0.83	0.86
		(0.18)	(0.13)	(0.15)	(0.14)	(0.13)	(0.16)	(0.15)
	π_m^g	0.90	0.81	0.85	0.84	0.81	0.92	0.89
		(0.11)	(0.14)	(0.13)	(0.12)	(0.13)	(0.08)	(0.09)
	π_e^o	0.52	0.52	0.54	0.53	0.56	0.34	0.38
		(0.21)	(0.19)	(0.21)	(0.20)	(0.19)	(0.17)	(0.18)
	π_m^o	0.68	0.70	0.72	0.72	0.70	0.85	0.82
		(0.19)	(0.17)	(0.18)	(0.17)	(0.17)	(0.11)	(0.12)
	G	1.08	1.35	1.22	1.10	1.11	0.89	1.11
		(0.89)	(0.81)	(0.79)	(0.51)	(0.53)	(0.15)	(0.29)
$s=0.2$	π_e^g	0.71	0.87	0.84	0.86	0.88	0.85	0.86
		(0.18)	(0.14)	(0.16)	(0.15)	(0.14)	(0.15)	(0.14)
	π_m^g	0.90	0.80	0.87	0.85	0.82	0.90	0.86
		(0.12)	(0.15)	(0.14)	(0.13)	(0.14)	(0.09)	(0.11)
	π_e^o	0.53	0.55	0.57	0.57	0.57	0.42	0.45
		(0.20)	(0.20)	(0.21)	(0.21)	(0.19)	(0.20)	(0.19)
	π_m^o	0.64	0.67	0.68	0.69	0.68	0.81	0.78
		(0.20)	(0.18)	(0.19)	(0.18)	(0.18)	(0.14)	(0.14)
	G	1.67	1.78	1.68	1.58	1.58	1.25	1.48
		(1.00)	(0.94)	(0.90)	(0.80)	(0.74)	(0.30)	(0.47)

Table 14: Parameter estimates (models with $s = 0$, $s = 0.1$, and $s = 0.2$)

Point estimates are posterior medians, with posterior standard deviations in parenthesis. The first five columns correspond to country-specific estimates. The 'Pooled' model pools observations across the five countries, and 'Pooled(4)' pools observations across the four countries, excluding Australia.

E. Goodness-of-Fit Measures

We assess model fit by how well the model predicts the winner of each election. We can use the simulated replicated data sequences used to compute p-values in Tables 5 and 6, but in the case of individual elections we have a more accurate (but equivalent) alternative. To compute individual election predictive probabilities for a model with data from any set of countries $\mathcal{C} \subseteq \{1, \ldots, C\}$, for each of the parameter vectors $\theta_i, \{b_i^1, \mathbf{x}_{c,i}\}_{c\in\mathcal{C}}, i = 1, \ldots, M$, in the posterior sample of size M, we calculate probabilities of victory for the left, $\hat{P}_{c,i}^t$, for each period t and country c applying equations (11) and (12), and then obtain the predictive probability $\hat{P}_c^t$ for period t and country c

by computing the mean

$$\hat{P}_c^t = \sum_{i=1}^{M} \frac{\hat{P}_{c,i}^t}{M}.$$

We compare the predictive ability of the formal model against three null models. First, we consider a naive *static* model in which the predicted probability of a victory for the left is simply the number of victories that the left achieved during the time series, divided by the number of elections. A more sophisticated 'dynamic' model posits that the election winner follows a two-state Markov chain summarized by two probabilities, $\Pr[P^t = L|P^{t-1} = L]$ and $\Pr[P^t = L|P^{t-1} = R]$, that condition on the identity of the winner in the previous period. In one version of this dynamic model (AR1a) we require that $\Pr[P^t = L|P^{t-1} = L] = \Pr[P^t = R|P^{t-1} = R]$, whereas for the other version (AR1b) we allow $\Pr[P^t = L|P^{t-1} = L] \neq \Pr[P^t = R|P^{t-1} = R]$. We compute these probabilities on the basis of the empirical frequency of these events for each country $c = 1, \ldots, C$ in the sample. Since these model involve prediction on the basis of the winner in the previous election, they do not yield a prediction for the first election in each country, and we calculate goodness-of-fit statistics with the remaining periods. In sum, each of the three models, the formal model along with the static and dynamic naive models, yield an estimated probability, $\hat{P}_c^t$, of a left government for country c and period t.

We then evaluate model fit by using three different measures of goodness-of-fit for binary outcomes. The first measure of fit is the 'proportion of elections correctly predicted' which is simply the count of those periods that the model predicts correctly as leftist, plus those it correctly predicts as rightist, divided out by the total number of periods. In keeping with standard practice, a left government is predicted if $\hat{P}_c^t \geq \frac{1}{2}$, and a right government is predicted otherwise. Such metrics are often over-interpreted in favor of the fitted model for binary data (see Greene, 2002, 685), so, in addition, we compute a pseudo-R^2 suggested by Efron (1978), and a more nuanced measure that accounts for the model's ability to predict both types of winners proposed by Cramer (1999).

We calculate all of the above goodness measures for the country specific and pooled models separately. In the case of the pooled models, a difficulty arises in setting up suitable null models for comparison: that is, in generating appropriate predicted probabilities for the 'pooled' static and dynamic naive models. In order to subject the formal model to the least favorable comparison, we employ the same country-specific probabilities in order to predict the pooled data for the static and dynamic null models as those that are used for the country-specific comparisons. As a consequence, the pooled formal model generates predictive probabilities by assuming common structural parameters θ across countries, which must then compete against null models that allow for variation in the predictive probabilities as we move across countries.

F. Australian distinctiveness is not due to data length

We have more observations available for Australia than for any other country in our data set (26 elections). Readers may be concerned that Australia's marked estimated 'different-ness' comes in large part from the fact that the time series is longer for this country. This is false. To support this claim, we re-estimated all models (baseline, and $s > 0$) with subsets of the Australian data. In particular, we took the first and last 17 periods of the entire Australian sequence as two new pseudo-data-sets which are identical in length to the data used for New Zealand and the UK. The parameter estimates, and their variances, are almost identical to those reported for the original 26 period case, as can be readily seen from Table 15.

	model	π_e^g	π_m^g	π_e^o	π_m^o	G
full 26		0.80	0.87	0.45	0.75	0.61
		(0.15)	(0.09)	(0.18)	(0.14)	(0.09)
periods 1–17	$s = 0$	0.81	0.87	0.43	0.75	0.64
		(0.15)	(0.09)	(0.18)	(0.14)	(0.12)
periods 10 – 26		0.83	0.87	0.48	0.77	0.70
		(0.14)	(0.09)	(0.18)	(0.13)	(0.18)
full 26		0.73	0.87	0.52	0.65	1.42
		(0.18)	(0.11)	(0.21)	(0.19)	(0.89)
periods 1–17	$s = 0.1$	0.73	0.87	0.53	0.65	1.49
		(0.18)	(0.12)	(0.22)	(0.19)	(0.90)
periods 10 – 26		0.78	0.86	0.50	0.71	1.27
		(0.17)	(0.11)	(0.20)	(0.17)	(0.77)
full 26		0.69	0.87	0.53	0.62	2.00
		(0.18)	(0.12)	(0.20)	(0.20)	(1.00)
periods 1–17	$s = 0.2$	0.71	0.87	0.54	0.63	1.95
		(0.18)	(0.12)	(0.20)	(0.19)	(0.97)
periods 10 – 26		0.73	0.86	0.52	0.65	1.88
		(0.18)	(0.12)	(0.20)	(0.19)	(0.93)

Table 15: Parameter estimates: means (standard deviations) for Australia, data split and models re-estimated. Note separation into first 17 and last 17 periods (of 26). Both estimates and their variances are very similar to those resulting from estimation using full data.

G. Identification Issues

In this supplemental appendix we discuss issues related to the identification of the model parameters given observed data, i.e., the sequence of election winners. We focus on the case without probabilistic shocks ($s = 0$), although most of these arguments extend to the case with probabilistic elections without any modification. We show that the model satisfies a necessary condition for identification under the restriction that inequality (4) holds, that is, under the assumption that equilibrium parameters induce the equilibrium dynamics of part 4 of Proposition 1. In a nutshell, observed data allow us to estimate the transition probabilities of a certain stochastic process on the length of spells in government for a party. The number of entries in the transition matrix of this stochastic process (far) exceeds that of the number of parameters to be estimated. Under a (reasonable) extra assumption on the manner these probabilities vary with model parameters, these parameters are identified from the observed data. Monte Carlo simulations support this notion, exhibiting the correct asymptotic reduction in individual parameter estimation bias as the number of election periods increases.

Recall that we represent observed data of parties in government over T periods by a vector $\mathbf{P} = (P^1, P^2, \ldots, P^T)$. From these data we can derive a vector

$$\mathbf{S} = (S^1, S^2, \ldots, S^k), k \leq T,$$

where S^1 represents the duration of the spell in office for the first party in government until the first alternation occurred, S^2 is the duration of the second spell in office, i.e., the number of periods between the first alternation and the second, etc. With the vector $\mathbf{S}$ recorded that way, we have $\sum_{j=1}^{k} S^j = T$. We can then interpret the data $\mathbf{S}$ as arising from a certain stochastic process and we can compute, for any integer $m \geq 4$, an $m \times m$ empirical transition matrix $Q(\mathbf{S}) = [q_{ss'}(\mathbf{S})]$ with entries

$$q_{ss'}(\mathbf{S}) = \begin{cases} \frac{\sum_{j=2}^{k} I_{\{s'\}}(S^j) I_{\{s\}}(S^{j-1})}{\sum_{j=1}^{k-1} I_{\{s\}}(S^j)} & \text{if} \quad 1 \leq s', s < m, \sum_{j=1}^{k-1} I_{\{s\}}(S^j) > 0 \\ \frac{\sum_{j=2}^{k} I_{\{y \geq s'\}}(S^j) I_{\{s\}}(S^{j-1})}{\sum_{j=1}^{k-1} I_{\{s\}}(S^j)} & \text{if} \quad 1 \leq s < m, s' = m, \sum_{j=1}^{k-1} I_{\{s\}}(S^j) > 0 \\ \frac{\sum_{j=2}^{k} I_{\{s'\}}(S^j) I_{\{y \geq s\}}(S^{j-1})}{\sum_{j=1}^{k-1} I_{\{y \geq s\}}(S^j)} & \text{if} \quad 1 \leq s' < m, s = m, \sum_{j=1}^{k-1} I_{\{y \geq s\}}(S^j) > 0 \\ \frac{\sum_{j=2}^{k} I_{\{y \geq s'\}}(S^j) I_{\{y \geq s\}}(S^{j-1})}{\sum_{j=1}^{k-1} I_{\{y \geq s\}}(S^j)} & \text{if} \quad s' = s = m, \sum_{j=1}^{k-1} I_{\{y \geq s\}}(S^j) > 0 \\ 0 & \text{otherwise.} \end{cases}$$

Our identification arguments are based on the following claim:

Claim 1 *Assume data*

$$\mathbf{S} = (S^1, S^2, \ldots, S^k), k \leq T,$$

generated from the model with parameters θ, b^1 and $s = 0$ over T periods.

1. *For all $m \geq 4$, if $\dfrac{1-\pi_e^g}{1-\pi_m^g} < \dfrac{1-\pi_e^o}{1-\pi_m^o}$, then $\lim_{T \to +\infty} k = +\infty$ and $\lim_{T \to +\infty} Q(\mathbf{S}) = \hat{Q}$ and $\hat{q}_{ss'} > 0$ for all $s, s' = 1, \ldots, m$, and*

$$\lim_{T \to +\infty} \textsf{Bayes Factor}\left(\tfrac{1-\pi_e^g}{1-\pi_m^g} < \tfrac{1-\pi_e^o}{1-\pi_m^o}\right) = 0.$$

2. *If $\dfrac{1-\pi_e^g}{1-\pi_m^g} > \dfrac{1-\pi_e^o}{1-\pi_m^o}$, then $\lim_{T \to +\infty} k < +\infty$ and*

$$\lim_{T \to +\infty} \textsf{Bayes Factor}\left(\tfrac{1-\pi_e^g}{1-\pi_m^g} < \tfrac{1-\pi_e^o}{1-\pi_m^o}\right) = +\infty.$$

The argument for the first part of the claim is based on the observation that the model induces a Markov chain over observed and unobserved quantities (P^t, x^t, b^t), which is irreducible and aperiodic with support on a countable infinity of electoral winners, policies, and belief combinations. This ensures that the limit of the frequency estimator represented by the matrix Q is indeed attained. To see that all entries of that matrix are positive, note that from all states in that state space, there is positive probability of reaching certain combinations with $P^t \in \{L, R\}$, $x^t = 0$, and $b_L^t = b_R^t \in (b^g, b^o)$. At such beliefs, the voter mixes between replacing the incumbent or not, and the government mixes between implementing an extreme policy or not. As a result there is positive probability of moving to $P^{t+1} = P^t$, $x^t = 0$, and $b_L^{t+1} = b_R^{t+1} \in (b^o, b^g)$ but also positive probability of $P^{t+1} \neq P^t$, $x^t = 0$, and $b_L^{t+1} = b_R^{t+1} \in (b^o, b^g)$. Thus, any possible duration of spells in office may be succeeded by a spell of any length with positive probability. The entries of the quasi-transition matrix $\hat{Q}$ are (complicated) functions of the model parameters θ (these reflect integrations over alternative paths of play, assuming the chain over (P^t, x^t, b^t) has converged to its invariant distribution). With $m \geq 4$, these expressions define more equations than unknown parameters, thus satisfying a necessary condition for identification of θ. If the entries of the matrix $\hat{Q}$ vary sufficiently richly with θ, then these parameters are identified.[23]

It is easy to see that these necessary conditions for identification fail when inequality (4) is not satisfied since equilibrium dynamics are characterized by part 5 of Propositition 1. This suggests that estimation of the model parameters is justified under the restriction that inequality (4) holds. Nevertheless, it is easy to show that the observations from a single sequence of election winners allow us to test whether inequality (4) is satisfied or not. In particular, as stated in the Claim, the Bayes factor testing whether inequality (4) holds takes the correct value as the number of election periods increases. We have thus chosen to report posterior point estimates of the model parameters without imposing the restriction that inequality (4) is satisfied. From a Bayesian viewpoint, these posteriors are well-defined given that we assume proper prior distributions. According to the Claim above, either inequality (4) is satisfied, in which case these posteriors recover consistent point estimates as the number of periods in the data increases, or inequality (4) is not satisfied in which case these posteriors will correctly reflect an area of the parameter space in which the model parameters lie.

H. Party Reputations and Policies for non-UK cases

The tables that follow in this Appendix are as Table 8 except that we have no suitable National Election Study data with which to compare the estimates.

[23]This type of assumption is common in the literature on the identification Markovian models.

	$s = 0$			$s = 0.1$			$s = 0.2$		
t	b_L^t	b_R^t	x^t	b_L^t	b_R^t	x^t	b_L^t	b_R^t	x^t
1946	0.61	0.77	0.61	0.56	0.6	0.39	0.53	0.56	0.31
	(0.20)	(0.18)		(0.28)	(0.29)		(0.29)	(0.29)	
1949	0.64	0.4	0.00	0.5	0.44	0.00	0.45	0.46	0.00
	(0.26)	(0.18)		(0.26)	(0.19)		(0.24)	(0.2)	
1951	0.38	0.37	0.00	0.44	0.39	0.00	0.45	0.39	0.00
	(0.18)	(0.18)		(0.21)	(0.2)		(0.2)	(0.20)	
1954	0.34	0.34	0.00	0.43	0.36	0.00	0.45	0.36	0.00
	(0.18)	(0.18)		(0.22)	(0.21)		(0.21)	(0.21)	
1955	0.32	0.32	0.00	0.43	0.35	0.00	0.44	0.34	0.00
	(0.18)	(0.18)		(0.22)	(0.21)		(0.22)	(0.21)	
1958	0.32	0.32	0.00	0.43	0.34	0.00	0.44	0.33	0.00
	(0.18)	(0.18)		(0.22)	(0.22)		(0.22)	(0.22)	
1961	0.32	0.32	0.00	0.42	0.34	0.00	0.44	0.32	0.00
	(0.18)	(0.18)		(0.22)	(0.22)		(0.22)	(0.22)	
1963	0.32	0.32	0.00	0.42	0.34	0.00	0.44	0.32	0.00
	(0.18)	(0.18)		(0.22)	(0.22)		(0.22)	(0.22)	
1966	0.32	0.32	0.00	0.42	0.33	0.01	0.44	0.32	0.00
	(0.18)	(0.18)		(0.22)	(0.23)		(0.22)	(0.22)	
1969	0.32	0.32	0.20	0.42	0.33	0.05	0.44	0.31	0.01
	(0.18)	(0.18)		(0.22)	(0.23)		(0.22)	(0.22)	
1972	0.32	0.4	0.00	0.42	0.35	0.01	0.44	0.31	0.01
	(0.18)	(0.26)		(0.22)	(0.25)		(0.22)	(0.23)	
1974	0.32	0.33	0.24	0.39	0.42	0.06	0.39	0.43	0.01
	(0.18)	(0.18)		(0.22)	(0.22)		(0.22)	(0.22)	
1975	0.42	0.32	0.00	0.39	0.42	0.00	0.37	0.44	0.00
	(0.28)	(0.18)		(0.25)	(0.22)		(0.23)	(0.22)	
1977	0.34	0.32	0.00	0.42	0.39	0.01	0.43	0.39	0.00
	(0.19)	(0.18)		(0.22)	(0.22)		(0.22)	(0.22)	
1980	0.32	0.32	0.26	0.42	0.37	0.07	0.44	0.36	0.01
	(0.18)	(0.18)		(0.22)	(0.23)		(0.22)	(0.22)	
1983	0.32	0.43	0.00	0.42	0.38	0.00	0.44	0.35	0.00
	(0.18)	(0.28)		(0.22)	(0.25)		(0.22)	(0.23)	
1984	0.32	0.34	0.00	0.39	0.42	0.00	0.39	0.43	0.00
	(0.18)	(0.19)		(0.22)	(0.22)		(0.22)	(0.22)	
1987	0.32	0.32	0.00	0.37	0.43	0.00	0.36	0.44	0.00
	(0.18)	(0.18)		(0.22)	(0.22)		(0.22)	(0.22)	
1990	0.32	0.32	0.00	0.36	0.43	0.01	0.35	0.44	0.00
	(0.18)	(0.18)		(0.23)	(0.22)		(0.22)	(0.22)	
1993	0.32	0.32	0.25	0.35	0.42	0.07	0.33	0.44	0.01
	(0.18)	(0.18)		(0.23)	(0.22)		(0.22)	(0.22)	
1996	0.43	0.32	0.00	0.37	0.42	0.00	0.33	0.44	0.00
	(0.28)	(0.18)		(0.26)	(0.22)		(0.23)	(0.22)	
1998	0.34	0.32	0.00	0.42	0.39	0.00	0.43	0.39	0.00
	(0.19)	(0.18)		(0.22)	(0.22)		(0.22)	(0.22)	
2001	0.32	0.32	0.25	0.42	0.37	0.07	0.44	0.36	0.01
	(0.18)	(0.18)		(0.22)	(0.23)		(0.22)	(0.22)	
2004	0.32	0.32	0.00	0.42	0.36	0.00	0.44	0.35	0.00
	(0.18)	(0.18)		(0.22)	(0.23)		(0.22)	(0.22)	
2007	0.32	0.42	0.06	0.42	0.37	0.02	0.44	0.34	0.00
	(0.18)	(0.28)		(0.22)	(0.25)		(0.22)	(0.23)	
2010	0.32	0.34	0.00	0.39	0.42	0.00	0.39	0.43	0.00
	(0.18)	(0.19)		(0.22)	(0.22)		(0.22)	(0.22)	

Table 16: Posterior estimates of party reputations, Australia. Columns for party reputations b_L^t, b_R^t report posterior means with standard errors in parenthesis. Column on policies x^t reports fraction of extreme policy choices in posterior sample of $10,000$.

t	$s=0$			$s=0.1$			$s=0.2$		
	b_L^t	b_R^t	x^t	b_L^t	b_R^t	x^t	b_L^t	b_R^t	x^t
1977	0.74	0.57	.54	0.62	0.57	0.49	0.57	0.55	0.44
	(0.19)	(0.21)		(0.27)	(0.25)		(0.29)	(0.28)	
1981	0.42	0.64	0.00	0.44	0.63	0.04	0.46	0.60	0.08
	(0.17)	(0.26)		(0.18)	(0.27)		(0.19)	(0.26)	
1985	0.4	0.41	0.12	0.45	0.45	0.17	0.47	0.47	0.19
	(0.18)	(0.18)		(0.20)	(0.19)		(0.22)	(0.20)	
1990	0.41	0.36	0.35	0.47	0.41	0.31	0.5	0.45	0.24
	(0.23)	(0.18)		(0.25)	(0.2)		(0.26)	(0.21)	
1993	0.35	0.52	0.00	0.41	0.54	0.02	0.45	0.53	0.06
	(0.18)	(0.3)		(0.2)	(0.29)		(0.21)	(0.27)	
1996	0.36	0.39	0.00	0.42	0.43	0.08	0.46	0.46	0.13
	(0.19)	(0.2)		(0.21)	(0.21)		(0.23)	(0.22)	
2000	0.36	0.36	0.29	0.44	0.41	0.33	0.48	0.45	0.29
	(0.18)	(0.18)		(0.24)	(0.2)		(0.26)	(0.21)	
2004	0.48	0.35	0.00	0.53	0.4	0.04	0.53	0.44	0.07
	(0.29)	(0.18)		(0.3)	(0.2)		(0.29)	(0.21)	
2007	0.37	0.36	0.22	0.43	0.43	0.26	0.46	0.47	0.24
	(0.19)	(0.19)		(0.21)	(0.23)		(0.22)	(0.24)	
2010	0.35	0.44	0.11	0.41	0.5	0.12	0.44	0.52	0.12
	(0.18)	(0.28)		(0.2)	(0.28)		(0.22)	(0.28)	

Table 17: Posterior estimates of party reputations, Greece. Columns for party reputations b_L^t, b_R^t report posterior means with standard errors in parenthesis. Column on policies x^t reports fraction of extreme policy choices in posterior sample of $10,000$

t	$s=0$			$s=0.1$			$s=0.2$		
	b_L^t	b_R^t	x^t	b_L^t	b_R^t	x^t	b_L^t	b_R^t	x^t
1966	0.74	0.57	0.55	0.57	0.57	0.49	0.53	0.54	0.40
	(0.19)	(0.21)		(0.29)	(0.27)		(0.29)	(0.29)	
1971	0.39	0.62	0.00	0.43	0.6	0.00	0.45	0.54	0.00
	(0.17)	(0.27)		(0.19)	(0.27)		(0.19)	(0.28)	
1976	0.37	0.38	0.00	0.42	0.46	0.02	0.43	0.47	0.03
	(0.17)	(0.18)		(0.2)	(0.21)		(0.21)	(0.21)	
1981	0.33	0.33	0.00	0.41	0.43	0.03	0.42	0.46	0.05
	(0.17)	(0.17)		(0.21)	(0.21)		(0.22)	(0.22)	
1987	0.32	0.32	0.10	0.41	0.42	0.12	0.42	0.45	0.12
	(0.17)	(0.17)		(0.22)	(0.21)		(0.24)	(0.22)	
1992	0.36	0.32	0.09	0.44	0.41	0.08	0.44	0.45	0.07
	(0.23)	(0.17)		(0.26)	(0.22)		(0.27)	(0.22)	
1996	0.32	0.35	0.25	0.42	0.44	0.23	0.44	0.45	0.14
	(0.18)	(0.22)		(0.22)	(0.24)		(0.22)	(0.24)	
1998	0.43	0.31	0.00	0.49	0.41	0.01	0.47	0.45	0.02
	(0.29)	(0.17)		(0.28)	(0.22)		(0.27)	(0.22)	
2003	0.34	0.32	0.00	0.43	0.41	0.03	0.45	0.43	0.04
	(0.19)	(0.18)		(0.22)	(0.22)		(0.22)	(0.23)	
2008	0.32	0.32	0.10	0.42	0.41	0.11	0.45	0.43	0.10
	(0.18)	(0.18)		(0.22)	(0.23)		(0.22)	(0.24)	

Table 18: Posterior estimates of party reputations, Malta. Columns for party reputations b_L^t, b_R^t report posterior means with standard errors in parenthesis. Column on policies x^t reports fraction of extreme policy choices in posterior sample of $10,000$

t	$s=0$ b_L^t	b_R^t	x^t	$s=0.1$ b_L^t	b_R^t	x^t	$s=0.2$ b_L^t	b_R^t	x^t
1946	0.58	0.74	0.54	0.57	0.59	0.51	0.55	0.53	0.42
	(0.20)	(0.19)		(0.26)	(0.28)		(0.28)	(0.29)	
1949	0.63	0.4	0.00	0.62	0.43	0.00	0.57	0.45	0.03
	(0.27)	(0.17)		(0.27)	(0.18)		(0.28)	(0.19)	
1951	0.39	0.38	0.00	0.45	0.42	0.02	0.47	0.44	0.05
	(0.18)	(0.18)		(0.2)	(0.19)		(0.21)	(0.20)	
1954	0.34	0.34	0.07	0.42	0.41	0.10	0.45	0.43	0.11
	(0.17)	(0.17)		(0.21)	(0.21)		(0.22)	(0.22)	
1957	0.33	0.36	0.30	0.41	0.44	0.26	0.45	0.45	0.17
	(0.17)	(0.21)		(0.21)	(0.24)		(0.22)	(0.25)	
1960	0.47	0.32	0.00	0.51	0.40	0.01	0.5	0.44	0.02
	(0.30)	(0.17)		(0.28)	(0.21)		(0.27)	(0.22)	
1963	0.36	0.33	0.00	0.43	0.41	0.01	0.46	0.44	0.03
	(0.19)	(0.18)		(0.22)	(0.21)		(0.23)	(0.23)	
1966	0.33	0.33	0.00	0.41	0.41	0.02	0.45	0.43	0.05
	(0.18)	(0.18)		(0.21)	(0.22)		(0.23)	(0.23)	
1969	0.32	0.32	0.08	0.4	0.41	0.10	0.45	0.43	0.11
	(0.17)	(0.17)		(0.21)	(0.22)		(0.23)	(0.24)	
1972	0.32	0.35	0.28	0.4	0.43	0.26	0.45	0.45	0.17
	(0.17)	(0.21)		(0.21)	(0.24)		(0.23)	(0.26)	
1975	0.46	0.32	0.00	0.5	0.4	0.01	0.49	0.44	0.03
	(0.29)	(0.17)		(0.28)	(0.21)		(0.27)	(0.22)	
1978	0.35	0.33	0.00	0.43	0.41	0.05	0.46	0.44	0.07
	(0.19)	(0.18)		(0.22)	(0.22)		(0.23)	(0.23)	
1981	0.33	0.33	0.25	0.41	0.42	0.25	0.45	0.45	0.18
	(0.18)	(0.18)		(0.21)	(0.23)		(0.23)	(0.25)	
1984	0.32	0.44	0.00	0.4	0.50	0.03	0.45	0.48	0.05
	(0.17)	(0.29)		(0.21)	(0.29)		(0.23)	(0.28)	
1987	0.33	0.34	0.28	0.42	0.42	0.27	0.46	0.45	0.20
	(0.18)	(0.19)		(0.23)	(0.22)		(0.24)	(0.23)	
1990	0.45	0.33	0.00	0.51	0.40	0.02	0.50	0.44	0.03
	(0.29)	(0.18)		(0.29)	(0.21)		(0.28)	(0.22)	
1993	0.35	0.33	0.10	0.42	0.42	0.12	0.46	0.45	0.10
	(0.19)	(0.18)		(0.22)	(0.22)		(0.23)	(0.24)	

Table 19: Posterior estimates of party reputations, New Zealand. Columns for party reputations b_L^t, b_R^t report posterior means with standard errors in parenthesis. Column on policies x^t reports fraction of extreme policy choices in posterior sample of $10,000$.