Veto players and the structure of budgets in advanced industrialized countries

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Abstract. This article tests expectations generated by the veto players theory with respect to the over time composition of budgets in a multidimensional policy space. The theory predicts that countries with many veto players (i.e., coalition governments, bicameral political systems, presidents with veto) will have difficulty altering the budget structures. In addition, countries that tend to make significant shifts in government composition will have commensurate modifications of the budget. Data collected from 19 advanced industrialized countries from 1973 to 1995 confirm these expectations, even when one introduces socioeconomic controls for budget adjustments like unemployment variations, size of retired population and types of government (minimum winning coalitions, minority or oversized governments). The methodological innovation of the article is the use of empirical indicators to operationalize the multidimensional policy spaces underlying the structure of budgets. The results are consistent with other analyses of macroeconomic outcomes like inflation, budget deficits and taxation that are changed at a slower pace by multiparty governments.

The purpose of this article is to test empirically the expectations of the veto players theory in a multidimensional setting. The theory defines 'veto players' as individuals or institutions whose agreement is required for a change of the status quo. The basic prediction of the theory is that when the number of veto players and their ideological distances increase, policy stability also increases (only small departures from the status quo are possible) (Tsebelis 1995, 1999, 2000, 2002). The theory was designed for the study of unidimensional and multidimensional policy spaces. While no policy domain is strictly unidimensional, existing empirical tests have only focused on analyzing political economy issues in a single dimension. These studies have confirmed the veto players theory's expectations (see Bawn (1999) on budgets; Hallerberg & Basinger (1998) on taxes; Tsebelis (1999) on labor legislation; Treisman (2000) on inflation; Franzese (1999) on budget deficits).

This article is the first attempt to test whether the predictions of the veto players theory hold in multidimensional policy spaces. We will study a phenomenon that cannot be considered unidimensional: the 'structure' of budgets – that is, their percentage composition, and the change in this composition over

Published by Blackwell Publishing Ltd., 9600 Garsington Road, Oxford, OX4 2DQ, UK and 350 Main Street, Malden, MA 02148, USA

time. In addition, unlike Tsebelis (1999) who studies government policies (i.e., the specific legislation produced by governments) and like most of the political economy literature (Bawn 1999; Treisman 2000; Franzese 1999), we focus here on outcomes (i.e., the amounts of money spent for different items) (for a detailed discussion of the distinction between policies and outcomes, see Tsebelis 2002). We base our study on 19 OECD countries in the period from 1973 to 1995 (including Austria, Australia, Belgium, Canada, Denmark, Finland, France, Germany, Iceland, Ireland, Italy, Luxembourg, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden and United Kingdom). Changes in budget allocation do not occur in a single dimension. A budget is composed of several items, and money not only shifts from one item to another, but items also increase or decrease independently. As a result of this shift in composition, multidimensionality is a serious concern and we have constructed indicators that address the problem of change in multiple dimensions.

We argue that budgets are altered in two different ways. First, deliberately, in the sense that the current government wants to increase or decrease budgetary spending and spend a higher or lower percentage of it in some area: one example might be increasing the education budget by shifting expenses from defense to education. Second, budgets are adjusted automatically in the sense that existing legislation (whether introduced by the current or previous government) has economic consequences: increasing unemployment affects the composition of the budget because provisions in social security legislation specify who is entitled to unemployment benefits, for how long, and so on. Of course, the size of the budget change depends on the specific provisions of legislation in each country.

We differentiate between deliberate and automatic structural change of budgets. Our basic finding is that deliberate change in the structure of budgets in advanced industrialized countries depends on the composition of governments (what we call the 'ideological distance' of government coalitions), and the difference in ideological position between the previous and the current government (what we call 'alternation'). Specifically, the more ideologically diverse the government coalition (the larger the ideological distances among parties), the less change occurs in the structure of budgets. In addition, the larger the alternation, the more significant the change in structure. These findings are consistent with the expectations of the veto players theory and congruent with other empirical studies that analyze problems in a single dimension.

Moving the analysis from a single dimension to multiple dimensions is not a trivial matter – theoretically or empirically. It is well known at the theoretical level that results obtained when the underlying space is unidimensional (like the median voter theorem) simply disappear in multiple dimensions.

Indicators for the study of multidimensional macroeconomic phenomena have not yet been developed at the empirical level. While numerous scholarly efforts seek to relate the *level* of budget deficits or government spending to the ideological attributes of government structure (Cameron 1978; Roubini & Sachs 1989a, 1989b; Cosetti & Roubini 1993; Blais et al. 1993, 1996; Spolaore 1993; Alesina & Perotti 1995; Von Hagen & Harden 1995; Hallerberg & Von Hagen 1999), we hope to enrich our understanding of public finance by focusing on the *composition* or structure of the budget.

This article is divided into four parts. First we review different literatures on budgets that deal primarily with the size of government budget deficits. Second, we present a multidimensional theory of policy stability and change. Third, we discuss our dataset, which includes the budget structures of 19 OECD countries and the political composition of their governments. The final part presents our findings.

Empirical literature on budget deficits and budgets

While government budgets have been analyzed in the literature, only two articles (Bawn 1999 and König & Tröger 2001) deal with policy stability of several budget items. The literature on budget deficits is much more prolific, and the discussion focuses on the causes of such deficits. We will discuss the literature underlying the discussions that focus on the stability of some feature of the budget. There is a vast political economy literature on budget deficits. Necessarily, we will be terse in the presentation of their arguments. In a broad sense, this literature can be divided into two different categories. The first explanation of budget deficits is that large government coalitions cause large deficits because of a collective action problem. The second – known as 'delayed stabilization' of budget deficits – expects large coalition governments to be slow in their adjustment, so it has expectations similar to the veto players theory. The time periods covered by empirical studies (post-oil shock stabilization) make it difficult to distinguish between these two theories.

A key notion in the *collective action* literature is the 'common pool problem'. The essence of this is that, in a decentralized policymaking government where each spending ministry only has authority over its own portfolio, the cost of overspending is shared with other ministries. Therefore, each ministry is motivated to overspend to please its constituency at the expense of other ministries. In other words, since each ministry internalizes only part of the cost of rising spending on their own goods, all of the groups have an incentive to spend more than the optimum or to appropriate resources for their own benefit. Thus, individual rationality leads to collective irrationality, where

the resultant budget deficit is radically different from the cooperative solution. In sum, the collective action literature argues that the more dispersed the decision-making authority, the higher the budget deficit. The proposed solution, accordingly, is to completely centralize decision-making authority by delegating the decision-making power to an independent agent, such as a strong treasury minister. The collective action approach has received some empirical support (Kontopoulos & Perroti 2000; Roubini & Sachs 1989b). Scholars have demonstrated that countries with multiparty governments have higher deficits. However, the empirical evidence in Von Hagen and Harden (1995) and Hallerberg and Von Hagen (1999) suggest that countries where budgetary decision-making authorities are centralized are less likely to suffer from budget deficits.

The *delayed stabilization* approach emphasizes the possibility a consensus to change an unsustainable status quo may not exist when there are too many parties in government. Alesina and Drazen (1991) first developed a 'war-of-attrition' model of delayed stabilization and demonstrate the difficulty in reaching a collective decision to implement fiscal adjustments due to the disagreement among different social groups about how to distribute the fiscal burden. Spolaore (1993) extends the war-of-attrition model to coalition governments and shows that a coalition government is more likely to delay fiscal adjustment than a single-party government. The rationale is that, unlike the ruling party in a single-party government that can easily shift costs to outside members, governing parties in a coalition government are more likely to disagree or veto any fiscal policy that contradicts their constituencies' interests. Accordingly, the policy inertia approach argues that delays in the adjustment or the elimination of existing deficits might result from struggles between coalition partners (or the social groups they represent) about who will bear the necessary costs/cuts in budget spending, even if these players agree that the current debt requires adjustment.

The empirical evidence for the policy inertia approach is strong but not unanimous. Roubini and Sachs (1989a) find that large deficits are positively associated with weak governments in OECD countries. Cosetti and Roubini (1993) and Alesina and Perotti (1995) expand on Roubini and Sachs' structural model and confirm their findings. Regarding American politics, Poterba (1994), Alt and Lowry (1994) and Krause (2000) present evidence in favor of the 'delayed stabilization' thesis, but there is also some empirical evidence that disputes Roubini and Sachs' findings (Edin & Ohlsson 1991; De Haan & Sturm 1997).

All empirical analyses of delayed stabilization test cases where governments aimed at reducing budget deficits, so while the argument at the theoretical level is different from the collective action literature, the empirical

finding of slow deficit reduction under government coalitions is consistent with both theories. There is, however, a particular analysis that goes one step further and identifies a 'crucial experiment' to discriminate between the two. Franzese (1999) conducts a study that covers 21 countries for the post-Second World War period. When testing for the size of a deficit as a function of the size of debt (this in fact is nothing but accumulated deficits), he concludes that: 'a standard deviation increase in fractionalization (+1.2 parties) centered on the mean (i.e., from 1.5 to 2.7 parties) increases deficits 0.18 per cent of GDP, but the same number of parties increase produces a 0.19 per cent GDP deficit reduction at low debt'. In other words, Franzese finds that multiple parties in government preserve the status quo more effectively regardless of whether deficits continue to be high (in countries like Italy) or low (in countries like Switzerland).

In terms of *budget composition*, although several studies have focused on institutional and partisan effects on a specific budget characteristic like welfare spending (Swank 1988; Hicks & Swank 1992), only two have focused on the effects of government composition on a variety of budgets items (Bawn 1999; König & Tröger 2001). Bawn studies specific items in the budget of the Federal Republic of Germany from 1961 to 1989. She describes items in the budget based on a system of two-digit categories, and from these categories she identifies items favored by the Socialists (SPD) and the Christian Democrats (CDU-CSU). In the first category she includes spending on educational grants and loans, professional education, art and cultural education, labour market policy, sports, the environment, municipal community service, urban renewal, mining and manufacturing, and aid to East Germany. In the second, she includes defense, non-university research and development, housing, improvements in agricultural structure, infrastructure investments, roads, rivers and harbors, aviation and shipping. (All of the latter items were included on the grounds that they are infrastructure/business pork barrel items.) She also identifies a series of ambiguous items, but these did not affect her analysis.

Throughout this analysis, the Liberal Party was assumed to want to minimize spending. As a result, the preferences on SPD items range from low (the Liberals) to middle (the Christian Democrats), to high (the Socialist), while on CDU-CSU items the preferences range from low (Liberals) to middle (Socialists) to high (Christian Democrats). Bawn's analysis identifies the range for each of the coalition governments, and the items that an increase or decrease in budget one would expect with a change in government. For example, when the SPD entered into government in 1966, replacing the Liberals, budget items in the SPD list were expected to increase because the country moved from a coalition desiring low spending on these items to a coalition requiring high spending. On the contrary, when the SPD/CDU-CSU

coalition was replaced in 1969 by the SPD/Liberal coalition, no change in the SPD budget items was expected (despite the fact that the SPD gained control of the chancellorship). Bawn forms a series of expectations on the basis of veto players analysis in a single dimension, and all of these expectations are corroborated.

König and Tröger (2001) essentially replicate Bawn's findings over a longer period of time, and use the estimated preferences of the different parties. Their approach is an improvement on that of Bawn because, instead of assuming that the Liberals want to minimize spending, the authors take them at their word and estimate that they are willing to spend on some budget items. Both Bawn's and König and Tröger's analyses, however, cover only one country and they reduce the policy space to a single dimension. Such a choice is impossible when one considers all the budget items.

All of the studies mentioned above, whether dealing with budget deficits or with specific budget items are supported by unidimensional models: whatever the dependent variable, parties are classified on the basis of how high or low they prefer the level of the corresponding variable to be, and coalitions are more or less diversified, with more or less members, etc. No other dimensions of variation among these variables are considered. In what follows, we will try to address the findings of these different lines of inquiry by connecting the issues of policy stability and budget composition (as opposed to the size of the deficit). Positive findings in our empirical analysis will support other policy stability results reported in the literature (Tsebelis (1999) on labor policies, Hallerberg & Basinger (1998) on taxes, Treisman (2000) on inflation, Henisz (2000) on growth). They will also be congruent with the inertia approach to explaining budget deficits, but will not contradict the collective action approaches described above.

Theory and data

Our dependent variable is the change in the structure of budgets in advanced industrialized countries. The budget of each country allocates resources in a series of areas. Consequently, we conceptualize it as a vector in an n-dimensional Euclidean issue space. It consists of a sequence of percentages (in order to control for size) allocated to different jurisdictions: $(a_1, a_2, \ldots a_n)$. Each year there is a different budget allocation, so the above sequence should be indexed by the time it was selected. As a consequence, the difference between two budgets can be represented by the distance between the points that represent them in the n-dimensional Euclidean space. In algebraic terms, the dependent variable, budget distance (BD) will be:

$$BD = \sqrt{(a_{1,t} - a_{1,t-1})^2 + (a_{2,t} - a_{2,t-1})^2 + \ldots + (a_{n,t} - a_{n,t-1})^2} \quad (1)$$

Figure 1 is a graphical representation of the difference in the structure of three-item budgets in two different countries. In each country the composition of the budget is presented by two variables (the third item is by definition allocated the remainder of the budget). In country A, the differences from time t to time t + 1 are small, while in country B the structure of the budget is significantly altered.

While the dependent variable in our study is a scalar, it is clearly generated in a multidimensional space and depends on the differences between parties not only in one particular dimension (be it left-right), but also in all the dimensions that compose the budget. Therefore, what we need is a multidimensional theory of policy change. The theory we will test here is presented in Tsebelis (2002: Chapter 1). Tsebelis has argued that when the unanimity core of a political system contains the unanimity core of another political system, changes in the status quo will be more difficult in the first case than in the second (the definition of 'unanimity core' or 'Pareto set' is the set of positions of the status quo that cannot be altered by unanimity of decision makers).

Figure 2 provides a visual representation of the argument. The system represented by the three veto players A1, A2 and A3 is more stable than the

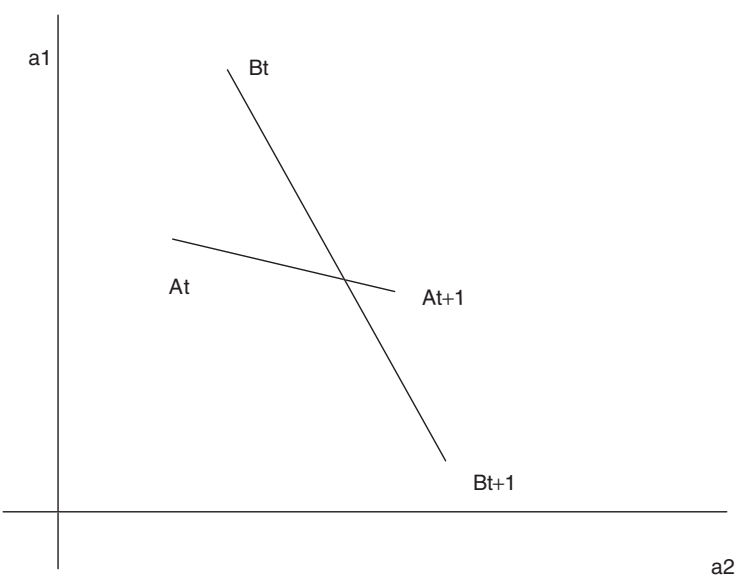

Figure 1. Budget in country A changes less over time than budget in country B.

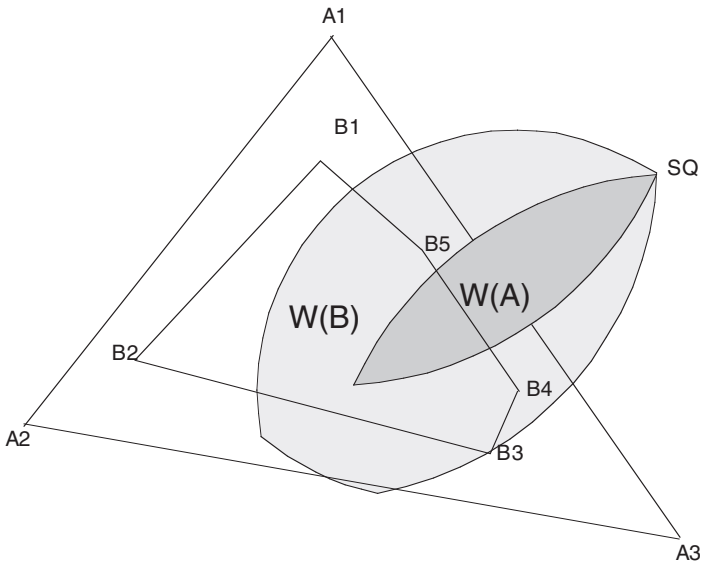

Figure 2. Veto players A1–A3 produce more policy stability than B1–B5 (no matter where the status quo is).

system B1, B2, B3, B4 and B5 – that is, significant changes of the status quo are more difficult in the first system than in the second. Indeed, if the status quo is located inside the system of Bs, no change is possible in either system; if it is located inside the system of As, but outside the Bs, a change is possible in the system of Bs but not in the system of As; if it is located outside the system of As, then change is possible but the points considered acceptable by the system of As are included within the points deemed acceptable by the system of Bs as the figure indicates.[1]

The proposition that we will test is that parties located further from each other in a multidimensional space cannot modify the status quo as significantly as a coalition with less diversified parties. This is a different analysis from many articles in the literature that argue policy stability depends on the sheer *number* of veto players (or veto points) (Birchfield & Crepaz 1998; Henisz 2000; Keefer & Stasavage 2000). We expect higher policy stability to be associated with system A rather than system B, despite the fact that it has fewer veto players.

Another proposition that we will test is the possibility that change is a function of the position of the status quo: the further away the status quo is from the preferences of the veto players, the greater the possible departure from the status quo. Given that, with respect to the budget, the status quo is created

by the previous government, we will consider the average position of the veto players in the previous government as a proxy for the position of the status quo. According to the theory tested, *BD* is a decreasing function of the ideological distances of the existing veto players (ID), and an increasing function of the ideological distance between successive groups of veto players, that we will call 'alternation' (A). In other words, $BD = f\ (ID^{-}, A^{+}, Z)$ where Z stands for the set of control variables.

At this point we need to address an important question raised by our conceptualization: it implies that we consider the current government responsible for the budget, while budgets are voted the previous year. We consider the current government responsible for the realization of the budget because, according to the literature, the current government has the means to alter the existing budget. In particular, a comprehensive study of budget rules in European Union (EU) countries by Hallerberg et al. (2001) identifies a series of ways a current government can amend the budgetary structure. First, finance ministers in most EU countries (the exceptions being Finland, the Netherlands, Spain and Sweden) can either block expenditure or impose cash limits. They also have the power to allow funds to be transferred between chapters, and the disbursement of the budget in the implementation stage is subject to finance minister approval. Second, there is a set of formal rules that enables governments to deal with unexpected expenditure and revenue shocks. In particular, 11 of 15 EU states grant governments the power to take necessary actions if they encounter unexpected fiscal shocks. For example, Denmark requires governmental action to correct the structure of the budget either if expenditure is higher than expectation or revenues are lower. Finally, most EU countries (with the exception of Finland, Greece and Luxembourg) allow a carryover of funds into the next budgetary year. Hallerberg et al. (2001) also note that the degree of governmental discretion over the current budget can be substantial: in theory, the United Kingdom allows 100 per cent of unspent funds to be moved forward into the following year.

Hallerberg et al.'s findings are replicated in presidential systems. In the study of state governments' budgetary policy in the United States, Alt and Lowry (1994) and Porteba (1994) suggest that the current government determines the final formation of budget allocation. Their argument is that, after a budget is passed, revenues and expenditures may diverge from expectations and lead to unexpected deficits. Under such a scenario, the current government can alter the budget decision so that the unexpected deficits can be avoided. Specifically, Porteba suggests that many state constitutions prevent state governments from running deficits, and they also vary in the policies that are available to eliminate a deficit and satisfy balanced-budget rules. For example, some states are allowed to borrow and close the current budgetary

gap. Some can also draw down their general fund balances to cover budget deficits. Similarly, Alt and Lowry argue that the states have a variety of balanced-budget laws that might influence fiscal policy, and some of the laws explicitly nullify unfunded expenditures. In sum, we posit that it is the configuration of current government rather than the previous one that accounts for the budget structures. As we shall elaborate later, we test for both current and previous government affecting budget structures, and our empirical findings corroborate our expectation.

In order to test whether there is a systematic impact on the structure of budget of the ideological distance within government and the ideological alternation between governments, we created a cross-sectional time-series dataset by merging data on the composition of budgets provided by the International Monetary Fund (IMF) with data on the composition of governments. Our dataset covers 19 OECD countries in the period 1973 to 1995, where data availability allows the most reliable analysis.

For s*tructure of budgets*, we used the government budget expenditure data from the IMF's *Government Finance Statistics Yearbook* to construct our measurement of budget structures in a cross-nationally and cross-temporally comparable way. In this dataset, all budgetary expenditures for each individual country are itemized into detailed categories, as shown in Table 1. Specifically, Table 1 demonstrates that a country's budget composition can be described by nine main categories: general public service, defense, education, health, social security and welfare, housing and community amenities, other community and social services, economic services and others. Obviously this categorization is crude and there are significant changes in the budget that are not reflected in it. For example, a shift of funds from primary to secondary education would not appear in our variable, although it is an important political change. Such a change would be captured by the sub-categories in Table 1, but one can imagine changes that would require an even finer partition of budgets to be captured. Unfortunately, our dataset does not permit any finer partition of budgets without losing countries or years from each one of them.

Following our conceptualization of budget structure in Equation 1, we constructed our dependent variable, *BD*, for each government per year. In fact, since there were missing values for some budget items, we calculated the *average* budget distance from year to year (i.e., we divided by the number of available items). This way we have a more reliable estimate of budget distances.

Figure 3 plots the change of the budget structure for 19 OECD countries from 1973 to 1995. The units we use are the number of percentage points of budget change divided by the number of relevant dimensions (usually 9). First, we note that there is no discernible trend in *BD* over time for each country.

Table 1. Classification of budgetary expenditures listed in IMF database

Main category	Sub-category
General public services	
Defense	
Education	Schools Universities and colleges
Health	Hospital and clinics Individual health services
Social security and welfare	Social security and assistance Welfare spending
Housing and community amenities	Housing Community development Sanitary service
Other community and social services	
Economic services	Electricity, gas and water Agriculture, forestry, fishing and hunting Mining, manufacturing and construction Roads Transportation and communication
Others	

Second, we can see that the change in the budgetary structure *per category* is mostly less than one percentage point. However, the intra-country changes (which are the focus of this study) are significant. In particular, Iceland, Italy, New Zealand, Portugal and Spain appear to have more changes in their budgetary structure (in other words, they have the highest mean of budget distance), and Belgium, France, Norway, Portugal and Spain tend to have the most volatile budgetary structure (i.e., the highest variability in budget distance).

For *governments*, we used the veto players dataset available at www.polisci.ucla.edu/tsebelis/ to construct a dataset measuring the veto players structure of 19 OECD countries. The unit of analysis is country-year, where a government that is responsible for the budget structure in a given year is the weighted average of actual governments in power that year.

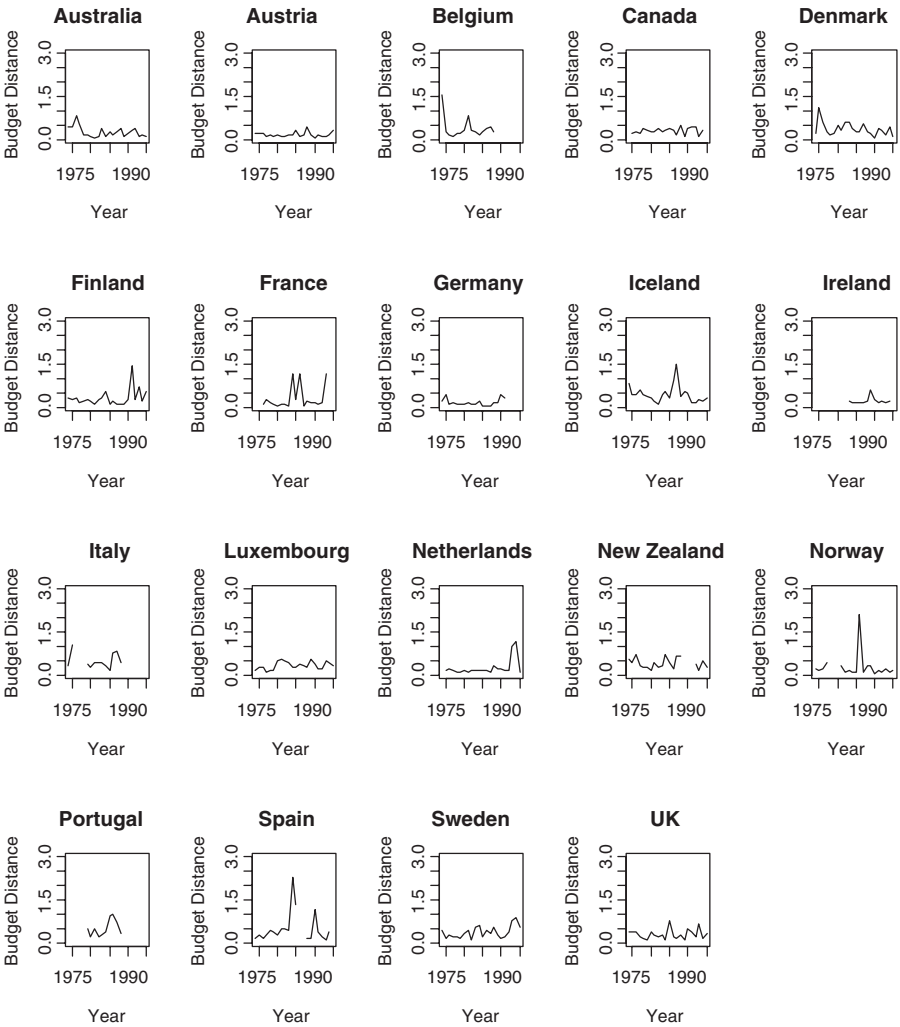

Figure 3. Change of budget structure in 19 OECD countries, 1973–1995.

Before moving to our empirical analysis, our first task is to identify all the veto players in our dataset. Tsebelis (1995) summarizes the counting rule of veto players as: 'a veto player is any player, institutional or partisan, who can block the adoption of a policy'. Following this definition, we take into account two types of veto players: *partisan veto players* (the political parties that are in the ruling governmental coalition) and *institutional veto players* (the political actors whose formal veto powers are legally specified by the constitution).

Specifically, institutional veto players include the upper chambers in Germany, Australia and Canada, and the President of Portugal (but not France), who has veto power over legislation. Partisan veto players are the parties required for a majority vote inside each institutional player. As a result, when the upper and lower chamber in a country, or the president and the legislature, have different majorities the number of veto players increases.

On the basis of four indices, we located each veto players' ideological position[2] in a two-dimensional policy space (the maximum for which data was available for all our countries). The first dimension expresses the positions of different parties on the left-right dimension on the basis of three different sources (the calculations generating the variable replicate – Tsebelis 1999). The second dimension is based on the position of different parties versus the Soviet Union. Information on this issue exists only in one source. Different countries had different issues as their third dimension, so we did not consider it appropriate to include additional (different per country) dimensions in our dataset.

The placement of governments in the first dimension (left-right) is calculated on the basis of three different measurements. The first is from Warwick (1994). This index was generated from 40 different measurements developed from experts, party manifestos and survey sources. Of the governments included in this dataset, the index ranged from a low of –6 (left) to a high of +5 (right). The second measurement for the left-right dimension is provided by Castles and Mair (1984). These ideological scores were generated from a questionnaire survey of more than 115 political scientists from Western Europe and the United States (Castles & Mair 1984: 75). The questionnaire asked each respondent to place all of the parties holding seats in his or her national legislature on the left-right political spectrum ranging from zero (ultra-left) to 10 (ultra-right), with 2.5 representing the moderate left, 5 the center, and 7.5 the moderate right. Castles and Mair present the results from those countries that had at least three respondents. The ideological score reported for each party was the average of available responses. Given the 10-point scale, the potential range of responses was zero to 10. Of the parties analyzed here, however, the low score was 1.4, received by the Communist Party of France, and the high was 8.2, received by the Gaullist Party.

The third left-right measurement is drawn from Laver and Hunt's (1992) first dimension variable 'increase services versus cut taxes'. Each respondent was asked to locate the policy positions of both the party leaders and voters for each party in his or her country on the left-right spectrum. In addition to the parties that won seats in the most recent election, respondents were asked to evaluate every party that won at least 1 per cent of the national

vote, as well as any significant regional parties. Laver and Hunt adopted a 20-point scale (to accommodate countries included in their study that contained up to 14 parties). Respondents assigned each party a score for the first dimension (taxes versus public services) ranging from 1 ('promote raising taxes to increase public services') to 20 ('promote cutting public services to cut taxes').

The only available measurement for a second dimension is in Laver and Hunt. They scored parties on the basis of their 'pro-friendly relations to USSR versus anti'. This dimension is not so clearly relevant to budget matters (with the possible exception of defense), and probably lacks political significance after the collapse of the Soviet Union. We will see that ultimately this second dimension does not add explanatory power to the model. However, in order to make this assessment, the estimation of a multidimensional model is necessary. Only Laver and Hunt included all 19 countries in their study. Warwick did not code the parties of Australia, Canada, France and New Zealand. In addition, some government parties in Ireland, Italy, Spain and Sweden were not scored. Castles and Mair did not include Iceland, Luxembourg and Portugal.

On the basis of each of these measures of ideology, we were able to construct new variables representing the 'ideological distance' of the parties in each government in two different dimensions, as well as the 'alternation' from one government to the next in two dimensions. Ideological distances in each dimension were created by taking the absolute value of the distance between the most extreme parties of a coalition on each of the two axes. These two parties were usually (but not always) the same for different indices in the first dimension. The second dimension did not present any problems because only one indicator was available. The alternation variable in each dimension was calculated by finding the mid-range position of each government and taking the difference between two successive governments.[3]

We used the above-mentioned four ideological distance and alternation variables to create new indices that measure *ideological distance* and *alternation* in the two-dimensional setting. In order to preserve the size of our dataset, as well as use all the available information, we constructed measures of ideological distance and alternation in the left-right dimension based on the values of *all* available indices, and then combined them with measurements in the second dimension to create multidimensional indices. In order to create the ideological distance and alternation measures in the left-right dimension using all the available information, we normalized each of the indices and then took the average of the normalized scores that were available for each government case. For normalization, we used only the values of the variables for the countries covered by all three indices. This normalization and averaging

procedure was run separately on all three ideological distance and alternation variables.

The average ideological distance and alternation variables for the left-right dimension used all the available information in the following way. Where all three indices existed, the average was calculated on the basis of all of them; for countries with only two or three indices, the average was calculated on the basis of the two or three standardized indices; in the cases covered by a single analyst, we used that one standardized score. For the second dimension, we used the only available index of ideological distance and alternation. The construction of indices in the first dimension is identical to Tsebelis (1999) with the exception that his unit of analysis is governments, while ours is years (the weighted average of indicators of all governments in place for a year).

The multidimensional indices were created in the following way. On the basis of our model, we needed to know whether the unanimity core of one government is included in the unanimity core of another. In a single dimension this is an easy task: one compares the length of the core (i.e. the 'range') of two coalitions. The use of the range of a coalition, therefore, is a straightforward application of our model. In two or more dimensions, however, such a straightforward measure does not exist. For example, it is *not* true that if the unanimity core of coalition A covers a larger area than the unanimity core of coalition B, then A necessarily includes B (which is the relevant criterion, according to Figure 2). For example, if a coalition has two distant members (which by definition means that its unanimity core is a straight line and therefore covers an area of zero), it does not mean that it can make decisions more easily than a coalition with three members located close to each other (which covers a small but positive area). As a result, we approximate the ideological distances of different coalitions in two dimensions by using the range of these coalitions in each dimension, and then by calculating their average. Obviously, this approach is an approximation and as such it is not immune to criticism. We use it because there is no other variable that can accurately capture the argument made in Figure 2. Also, our variable has two advantages: it is simple and easily generalizable in any number of dimensions.

For alternation, the selection of the indicator was easier because we know the position of the middle point of the range in each dimension, so the distance between two governments can be calculated by the Pythagorean Theorem where A_1 is the alternation in the first dimension, and A_2 is the alternation in the second dimension: $A_{12} = \sqrt{A_1^2 + A_2^2}$. Note that this formula produces positive distances regardless of whether the successor government is to the left or right of the predecessor in the first dimension, or their relative positions in the second dimension.

Testing the implications of the veto players theory

Model specification

We now present our model specification which tests the hypothesis outlined in the previous section. In our empirical model we attempt to differentiate between deliberate changes of the budgetary structure affected by ideological distance and alternation of governments from possible driving forces that lead to the automatic change of budgetary structure. Toward this end, we first took the absolute value of the change in unemployment rate (ΔUE) and the absolute value of the change in the proportion of elderly people above 65 years of age ($\Delta POP65$) into consideration.[4] Many scholarly works on redistributive politics in advanced economies have identified that societal demand for the reallocation of public spending is driven largely by these two variables (Tabellini 1990, 1991).

Moreover, it has been argued that different types of governments will have different effects on policy stability. In particular, Strom (2000) expects minority governments to have higher policy stability than minimum winning coalitions since they require legislative support from parties not in government. Similarly, oversized coalitions do not require a positive vote from all government members, which leads to the expectation that policy stability under oversized coalitions will be lower than under minimum winning coalitions. Operationally, we construct two dummy variables, *MINORITY* and *OVERSIZE* (with the baseline of comparison being the minimum-winning coalition government), to capture the potential effects of these two government characteristics. Following the standard convention, *MINORITY* takes a value of 1 if a government fails to obtain a majority in the legislature. Similarly, we set the value of *OVERSIZE* to 1 if a government still enjoys a majority in the legislature after dropping its smallest coalition partner from the government. Recall that the unit of analysis in this study is country-year. Therefore, if a country experiences any government change in some particular year, the values of these two variables will be the average from each government weighed by their duration within that year. Furthermore, following our previous discussion, we include the configuration of veto players (ideological distance and alternation) of both current and past governments to examine which has the dominating effect on budget structure. Accordingly, both *ID* and *A* are lagged and included in the model specification.

Finally, we also use a set of dummy variables, one for each country, to pick up any country-specific intercept that is not accounted for in our model and to prevent any unobserved heterogeneity. Specifically, we include the dummies because different countries have significantly different provisions about

policies that are reflected in the budget. For example, the time period that workers qualify for unemployment benefits may have a significant impact on the budget. Given that we cannot control for all possible variables that would affect the budget, the inclusion of dummies was necessary. At the empirical level, an F-test (with no difference of intercepts as the null hypothesis) resulted in highly significant result, which suggested that our choice was empirically justified. Similarly, we also take into account for the year fixed effect by including a set of year dummy variables. Note that by including the year dummies, we are able to eliminate any bias caused by unaccounted trends and external shocks to which all these OECD countries might be commonly exposed. By including the country fixed effect and the year fixed effect in a unifying model, we are able to guard against the potential threat of all the country-specific or time-specific unobserved exogenous shocks and estimate the pure effect of veto players configuration on the budget structure.

Before moving to estimate our model, we first investigate the dynamics of our dependent variable. Our first task is to examine whether our dependent variable is stationary. When it is not stationary, the underlying data generating process will not remain constant over time. As a result, the usual t statistics will follow nonstandard distributions and lead to misleading inference. Our preliminary visual inspection of the dependent variable, as plotted in Figure 3, finds no discernible trend and we speculate that the dependent variable seems to be stationary. To test whether unit roots are present in a more systematic way, we implement the Levin-Lin test (Levin et al. 2002) for unit roots in our cross-sectional time-series data (see Baltagi 2001 and Fabian 2002 for details). The results, as shown in the upper panel of Table 2, indicate no evidence of non-stationarity. Since it is often argued that the Levin-Lin test (and other tests for unit roots in general) only enjoy limited explanatory power, we double-check the stationarity by using another test (Im-Pesaran-Shin test). Again, as we can see the results from the lower panel of Table 2, our dependent variable appears to be stationary. Therefore, we feel very confident in proceeding with our analysis by specifying the dependent variable in levels rather than in first differences.

In short, our model can be represented by the following equation:

$$\begin{aligned} BD_{i,t} = {} & \beta_1 D_{i,t-1} + \beta_2 ID^-_{i,t} + \beta_3 A^+_{i,t} + \beta_4 ID^-_{i,t-1} + \beta_5 A^+_{i,t-1} + \beta_6 \Delta UE^+_{i,t} \\ & + \beta_7 \Delta POP65^+_{i,t} + \beta_8 MINORITY^-_{i,t} + \beta_9 OVERSIZE^+_{i,t} \\ & + \sum_{i=1}^{19} \gamma_i Country_i + \sum_{t=1973}^{1995} \varphi_t Year_t + \varepsilon_{i,t} \end{aligned} \tag{2}$$

where i stands for countries, t stands for year and $\varepsilon_{i,t}$ is the error term. Note that it is very reasonable to expect budgetary inertia in the budget structure.

Table 2. Panel unit roots tests

	Budget distance: levels			
Levin & Lin tests	coefficient	t-value	t*	p
constant	–0.9189	–14.463	–7.689	0.000***
constant, trend	–1.0227	–15.648	–5.969	0.000***
Im, Pesaran & Shin tests	t-tar	CV1%	Ψ	p
demeaned, no trend	–3.427	–1.99	–8.498	0.000***
demeaned, trend	–3.587	–2.62	–6.653	0.000***
not demeaned, no trend	–3.444	–1.99	–8.576	0.000***
not demeaned, trend	–3.619	–2.62	–6.804	0.000***

Notes: Levin & Lin tests augmented by 1 lag. H_0: nonstationarity; coefficient: Coefficient on lagged levels; t-value: t-value of coefficient; t*: transformed t-value. t* ~ N(0, 1); p: p-value of t*; *** $p < 0.001$. Im, Pesaran & Shin Tests augmented by 1 lag. H_0: nonstationarity; t-bar: mean of country-specific Dickey-Fuller tests; CV1%: 1 per cent critical value of the Im, Pesaran & Shin Test; X: transformed t-bar statistics. Ψ ~ N(0, 1); p: p-value of Ψ; *** $p < 0.001$.

Therefore, we correct the first-order autocorrelation with the lagged dependent variable. We also report panel-corrected standard errors to guard against potential problems of panel heteroskedasticity across countries and contemporaneous correlation of error[5] (Beck & Katz 1995, 1996).

Estimation results

The results of estimation are summarized in Model 1 in Table 3. From Table 3 we can see that our model performs well. First notice that the result of the Wald test allows us to reject the null hypothesis that there is no relationship between the explanatory variables and the dependent variable at less than 0.0001 level. The adjusted R^2 of more than 60 per cent also indicates that our model explains more than half of the variation in the dependent variable. More importantly, the results are consistent with the predictions of the veto players theory: both coefficients of ideological distance and alternation in this model are significant and signed according to expectation. This result clearly confirms our hypothesis that the deliberate change in the structure of budgets depends on the composition of governments and the ideological distance between two successive governments. Moreover, in comparision to other coefficients, the size of standardized coefficients of ideological distance (–0.2630) and alternation (0.1705) suggests that the effect of the veto players structure

Table 3. Estimated results on budget structure in 19 OECD countries, 1973–1995 (estimated by fixed-effect cross-sectional time-series model with panel-corrected standard errors)

	Model 1		Model 2
Variable	Estimated coefficient	Test of joint hypothesis	Estimated coefficient
Lagged BD	0.1183		0.0580**
	(0.1170|0.1395)		(0.0878|0.1279)
ID	–0.0851***		–0.0420**
	(–0.2630|0.0344)		(–0.1279|0.0195)
Alternation	0.0553***		0.0438**
	(0.1705|0.0198)		(0.1283|0.0203)
Lagged ID	0.0346		
	(0.1068|0.0324)		
		0.5572	
Lagged alternation	–0.0018		
	(–0.0058|0.0203)		
Minority	0.0027		
	(0.0038|0.1047)		
		0.3488	
Oversize	0.0964		
	(0.1142|0.0648)		
ΔUnemployment	0.0555*		0.0490
	(0.1578|0.0346)		(0.1371|0.0364)
ΔAge > 65	–0.0507		
	(–0.0233|0.1835)		
N	329		336
Adjusted R^2	0.6400		0.6294
F (Country)	861.65***		1870.08***
F (Year)	257.47***		1185.29***

Note: *p < 0.1; **p < 0.5; ***p < 0.01, all tests are two-tailed. (Standardized coefficient|panel-corrected standard errors) in parentheses. F (Country) = F-test for the inclusion of country dummies; F (Year) = F-test for the inclusion of year dummies; fixed effects are suppressed to facilitate the presentation.

is not only statistically significant, but also substantively important. Moreover, from Table 3, we did not find any supporting evidence for the government that voted the budget: both lagged *ID* and lagged *A* are not significant, and they are not jointly significant either. We believe that the reason for this null relationship between the composition of the budget and the government that voted for it is due to two factors: (1) our variable is a consistent, but not very

sensitive, indicator of changes in the budget: moving expenses from one item to another will not be reflected in our variable unless they cross the categories of the IMF database (Table 1); and (2) these modifications are very slow (at the margin) and current governments clearly have the means to impose such modifications. Finally, the type of government (minority or oversized coalition) does not have any significant effect on policy stability since these two coefficients in Model 1 are insignificant, and we cannot reject the null hypothesis that these two variables are jointly different from zero.

Robustness check. To assure the validity of our empirical results, we implement a series of robustness checks. First, we drop the variables lagged *ID*, lagged *A*, $\Delta POP65$, *MINORITY* and *OVERSIZE* and rerun the regression to check the robustness of our results. As shown in Model 2 in Table 3, all of our main variables remain stable. Second, one might object to the use of country dummy variables and year dummy variables. It is suggested that the cross-sectional time-series analysis is not very reliable since it can be very sensitive to the inclusion of country dummies and year dummies (Kittel & Winner 2002). Moreover, the country dummies might remove all cross-national variation that otherwise would have been captured by the main independent variables. Accordingly, we rerun our estimation with all the possible model specification with fixed effects and report our results in Table 4. From Table 4, we can see clearly that the coefficients of ideological distance and alternation remain significant across all the model specifications with the expected sign, and the magnitude of the coefficients remain more or less unchanged. These results provide clear evidence that our empirical model is well-specified and our empirical results are not contingent upon the arbitrary choice of fixed effects.

Finally, we pay special attention to autocorrelation that might be inherent in our dataset. As Achen (2000) notes, OLS estimates in the model with a lagged dependent variable will not be unbiased and consistent if autocorrelation still exists in the residuals. More importantly, combining the lagged dependent variable with the country fixed effects might increase the bias because correlation exists between the lagged dependent variable and the country dummies (Kittel & Winner 2002). In fact, as Beck and Katz (1995) emphasize, researchers should get rid of any autocorrelation before applying their panel corrected standard errors. Accordingly, to assure that we have adequately handled autocorrelation, we calculate Durbin's h test (Durbin 1970). As we can see from the results reported in the last rows of Table 4, none of the Durbin h-statistics is significant, indicating that our empirical results are free of autocorrelation. This result is further confirmed by regression of the residual on the lagged residuals where we find the coefficients of the lagged residuals are not significant across all models.

Table 4. Robustness checks

	Model 4 FE(CY)	Model 5 FE(C)	Model 6 FE(Y)	Model 7 Pool
Lagged BD	0.1183	0.1379	0.2305	0.2431**
	(0.1395)	(0.1015)	(0.1483)	(0.0992)
ID	–0.0851***	–0.0883***	–0.0679*	–0.0667**
	(0.2630)	(0.0341)	(0.0374)	(0.0329)
Alternation	0.0553***	0.0498***	0.0475***	0.0429**
	(0.0198)	(0.0185)	(0.0181)	(0.0187)
Lagged ID	0.0346	0.0366	0.0459	0.0479
	(0.0324)	(0.0328)	(0.0328)	(0.0317)
Lagged alternation	–0.0018	0.0030	–0.0148	–0.0081
	(0.0203)	(0.0171)	(0.0190)	(0.0175)
Minority	0.0027	–0.0205	0.0458	0.0371
	(0.1047)	(0.0913)	(0.0518)	(0.0474)
Oversize	0.0964	0.0201	0.0759	0.0504
	(0.0648)	(0.0714)	(0.0469)	(0.0481)
ΔUnemployment	0.0555*	0.0399*	0.0528	0.0444**
	(0.0346)	(0.0209)	(0.0369)	(0.0220)
ΔAge > 65	–0.0507	–0.0231	0.1330	0.1351
	(0.1835)	(0.1342)	(0.2307)	(0.1488)
N	329	329	329	329
Adjusted R^2	0.6400	0.6273	0.6188	0.0877
F (Country)	861.65***	82.35***		
F (Year)	257.47***		258.74***	
P (Lagged residuals)	0.633	0.829	0.682	0.887
P (Durbin's h)	0.4534	0.7848	0.674	0.3189

Note: FE(CY) = fixed country and time effects; FE(C) = fixed country effects; FE(Y) = Fixed year effects; Pool = simple OLS on pooled specification. F (Country) = F-test for the inclusion of country dummies; F (Year) = F-test for the inclusion of year dummies; P (Lagged residual) = P-value of lagged residuals resulting from regressing the residuals on the lagged residuals; P (Durbin's h) = p-value of the Durbin's h-statistics. *p < 0.1; **p < 0.5; ***p < 0.01, all tests are two-tailed. Constant and fixed effects not shown. Panel-corrected standard errors in parentheses.

To sum up, all the empirical evidence presented here validates our hypothesis that, despite the factors that account for the automatic change of budgetary structure, the deliberate change of budgetary structure can be explained by governmental ideological distance and ideological differences between governments. Specifically, a government coalition will be associated with more significant change in the budget if the members of this government are less

ideologically diverse or if its ideological position is more divergent from the previous government. In other words, the budgetary structure tends to lock itself into the existing pattern in political systems with ideologically distant veto players; in contrast, the budgetary structure tends to be more flexible in political systems with ideologically similar veto players.

It is interesting to investigate how ideological distance and alternation affect the budget structure on a disaggregated level. Therefore, we perform a series of regressions on each individual budget category by using the same model structure as in Model 2. The results, which are summarized in Table 5, suggest that ideological distance and alternation also explain change in each of the budget categories very well. In fact, there is only one out of the nine budget items (housing and community amenities) with the wrong sign (though it is not statistically significant) and the rest have the expected signs. In all cases other than defense, at least one coefficient is significant. Ideological range significantly affects six out of nine budget categories, and alternation also has a significant effect on six out of nine budget categories. We use a one-

Table 5. Estimated results for each budget category

Budget category	Ideological distance	Alternation
General public services	–0.0665*	0.0522*
	(0.0466)	(0.0359)
Defense	–0.0116	0.0208
	(0.0232)	(0.0176)
Education	–0.1117**	0.0874**
	(0.0663)	(0.0419)
Health	–0.2392***	0.2696****
	(0.0914)	(0.0684)
Social security and welfare	–0.2774***	0.0918*
	(0.0974)	(0.0716)
Housing and community amenities	0.1068	–0.0001
	(0.1191)	(0.0361)
Other community and social services	–0.0106	0.0109**
	(0.0144)	(0.0051)
Economic services	–0.1801*	0.1674***
	(0.1300)	(0.0600)
Others	–0.2007*	0.1476
	(0.1585)	(0.1191)

Note: Estimated coefficients for country dummies, change in unemployment rate and lagged dependent variable are surpassed to facilitate the presentation. Panel-correction standard errors are in parentheses. * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$; **** $p < 0.001$, all tests are one-tailed.

tailed test since our theory explicitly predicts the direction of the effect. We acknowledge the fact that our empirical findings appear relatively weak on the basis of significance at the 0.1 level in a one-tailed test. If we use the conventional two-tailed test, our ideological distance variables still works on four out of nine categories, while the alternation variable only works on one category. In particular, we find that among these nine budget categories, changes to education, health, social security and welfare, and economic service are especially sensitive to both the effects of ideological distance and alternation.

Finally, we compare one- and two-dimensional analyses of budget structure. Table 6 summarizes our results and shows that the two-dimensional index performs approximately at the same level as the first dimension index, but significantly better than the second dimension one. At the item-by-item level, the two-dimensional model performs better than the second dimension (particularly on the 'ideological distance' variable). We remind the reader that most theoretical and empirical analyses use only one dimension for the analysis. The only second dimension existing in the empirical literature is the attitude towards the Soviet Union in Laver and Hunt (1992). It is unfortunate that a second dimension more reflective of budget differences does not exist in the literature.

Conclusions

This article studies and defines the budgetary structure in advanced industrialized countries. Instead of focusing on budget deficit or government spending as is done quite often in the political-economy literature, we examine budgetary structure from another theoretical angle and the question of interest is what accounts for the change of budgetary structure over time. Budget composition is almost by definition a multidimensional problem, so unlike the existing political economy literature that uses only a left-right scale, we rely on multidimensional analytical models, and provide measurements of multidimensional variables.

We constructed two such multidimensional indicators: one representing the ideological distances among veto players and the other representing the ideological distances between two successive governments. The changes in budgetary structure were identified in a cross-sectional time-series model. In that model, we distinguished between automatic changes of the budget generated by variables like unemployment or percentage of older people in the population, as well as legislative differences among countries (by controlling for dummy variables). The essential finding is that despite the factors that account for the automatic change of budgetary structure, the deliberate change of

Table 6. Comparison of explanatory power of unidimensional and two-dimensional analyses

Budget	ID in 1st dimen.	Alt in 1st dimen.	ID in 2nd dimen.	Alt in 2nd dimen.	ID in two-dimen.	Alt in two-dimen.	Improvement			
							ID 1st → 2	Alt 1st → 2	ID 2nd → 2	Alt 2nd → 2
Total	***	***	*	***	**	***	↓		↑	
General public services	**	*	C	*	*	*	↓		↑	
Defense	C	**	C	**	C	C		↓↓		↓↓
Education	**	**	C	***	**	**			↑↑	↓
Health	***	****	**	****	***	****			↑	
Social security and welfare	***	*	***	C	***	*				↑
Housing	W	C	W	*	W	C				↓
Other community and social services	*	C	C	C	C	**	↓	↑↑		↑↑
Economic services	**	*	C	**	*	***	↓	↑↑	↑	↑
Others	C	**	**	*	*	C	↑	↓↓	↓	↓

Note: W denotes 'wrong sign'; C, 'correct sign'; * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$. ↑ denotes results improving by one step (from W to C, * to **, etc.); ↑↑ denotes results improving by two steps (W to *, to ***, etc.); ↑↑↑ denotes results improving by three steps (C to ***, etc.); ↓ denote worsening results.

budgetary structure can be well explained by the ideological distances among veto players and the alternation between governments. In particular, the composition of the *current* government accounts for the realization of the budget. The relationship has the following features: the budgetary structure tends to lock itself into the existing pattern in political systems with ideologically distant veto players; by contrast, the budgetary structure tends to be more flexible in political systems with ideologically similar veto players. In addition, significant changes in government composition from one year to the next lead to significant changes in the composition of the budget.

Acknowledgments

The authors wish to thank Geoffrey Garrett, Mark Hallerberg and William Keech for their comments on earlier drafts of this article. George Tsebelis would like to acknowledge financial support from the Russell Sage Foundation, and thank Liz McDaniel for research assistance.

Notes

1. The reason for the inclusion of the winset of As in the winset of Bs (regardless of the position of the status quo) is demonstrated by Tsebelis (2002: Chapter 1) who proved that any point accepted by all three of the As is also accepted by any one of the Bs. Indeed, if a point X belonged in the winset of the status quo for all the As and did not belong in the winset of the status quo of one of the Bs, one could draw a hyperplane H through the middle of SQX and perpendicular to it. All points A would be on one side of this plane, while one of the Bs would be on the other side of it. This contradicts the assumption that the Bs belong to the unanimity core of the As.
2. How to estimate party ideological positions appropriately has been subject to numerous studies and was hotly debated recently. For instance, Gabel and Huber (2000) suggest a plain vanilla method of using principal component analysis on the manifestos data to estimate party left-right position. On the other hand, Laver and Garry (2000) propose an alternative estimating procedure that uses computer-coding techniques for estimating the party policy positions. As Laver and Garry (2000) argue, however, to date there is no universally accepted method, and each method used has its advantages and disadvantages. The indices adopted in this article are mostly created by using an expert survey approach, which has the advantage of yielding decidedly positional estimates for parties. Another advantage of our measurement technique is that it is comparable to Tsebelis (1999) since we have the same data and analyze it in a similar way (with the addition of the second dimension). Meanwhile, we recognize the fact that the positional indices we use are time invariant, and seek to provide a general representative position for each party during the 1973–1995 period by taking the average of three different indices created at different times. It turns out that the use of manifesto data corroborates our results: 'I find that even though the size of the budget may be stable, changes in government policy positions are

strongly related to rearrangements in the government expenditure. This is in line with the recent findings of the authors of this article, but using different, manifesto-based measures of the ideological conflict between parties' (Braeuninger 2002).

3. The formula we used was: 1/2((maxgovt1 + mingovt1) – (maxgovt2 + mingovt2)), where max-govt1 and min-govt1 are the ideological scores of the preceding government, and max-govt1 and min-govt2 are the ideological scores of the 'current' government. Therefore, if a government succeeded (or 'replaced') a government with the same party structure, then all of the alternation variables would equal zero. The following example is useful. Consider a hypothetic system experiencing a government change. The previous government consisted of two parties that scored 2 and 8, respectively. The current government consists of two parties with ideological positions of 4 and 6. According to our operationalization, the alternation is $1/2((2 + 8) - (4 + 6)) = 0$. However, since the ideological distance of the government is smaller in the current government $|6 - 4| = 2 < |8 - 2| = 6$, our theory predicts more potential for change in the budget structure in the current government due to the effect of ideological distance.
4. It should be noted that we take the absolute value of the difference of these control variables, since we are interested in the effect of a change in these variables regardless of direction. We have also controlled for the potential effects of the change in economic growth (ΔG*ROWTH*) and the change in inflation (Δ*INF*) on the budgetary structure, since it has been suggested that growth might induce governments to spend more on the social dimensions of the budget, while inflation might lead them to reduce spending. These two variables are excluded here because the theoretical justification for them is weak, their coefficients are insignificant and they do not affect the key variable of interest. In other words, ID and A are insenstive to the inclusion or exclusion of ΔG*ROWTH* and Δ*INF*.
5. Using Monte Carlo simulation, Beck and Katz (1995) show that OLS with PCSE is a generally superior method for analyzing a cross-sectional time-series model than the traditional Parks-Kmenta method, which tends to exhibit egregious bias towards finding statistical significance.

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