Political fragmentation, fiscal deficits and political institutionalisation

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Abstract

One line of research finds the size of the deficit to be positively correlated with the number of political actors. This 'political fragmentation' hypothesis has been tested on OECD countries. We successfully replicate Volkerink and de Haan's (2001) model on an OECD sample. However, when we add ten non-OECD countries, the effect of political fragmentation disappears. We argue that the importance of political fragmentation varies according to the institutionalization of political systems. When we interact the age of a democracy with political fragmentation, we find that legislative fractionalisation increases the budget deficit as a democracy becomes more institutionalised.

Keywords: common-pool resource problem; fiscal deficit; political fragmentation; institutionalisation

1. Introduction

There is now a considerable body of work that views fiscal performance as an example of the common-pool resource problem. One line of research focuses on the political fragmentation hypothesis. Given finite resources and in the absence of a central co-ordinator, opportunistic actors will seek to maximise their individual benefit and externalise the overall cost to the group as a whole. From this perspective, the size of the budget deficit is positively correlated with the number of politically relevant actors. So, Roubini and Sachs (1989) demonstrated that, all else equal, coalition governments are associated with greater deficit spending than single-party governments. Kontopolous and Perotti (1999) showed that the number of spending ministers affects the level of the budget deficit. De Haan, Sturm and Beekhuis (1999) stressed the number of parties in government: the greater the number of parties, the higher the central government debt-to-GDP ratio. Volkerink and de Haan (2001) confirmed that the number of ministers was significant and showed that the government's level of parliamentary support was also a determinant. These studies share the intuition that fiscal performance can be explained using a common-pool resource approach. They test this intution on a panel of OECD countries and they all find some evidence to support the political fragmentation hypothesis. However, these studies vary in terms of the political variables they stress. Moreover, the models are always highly sensitive to the sample of countries chosen and the time period under consideration.

In this paper, we aim to determine whether there is evidence to support the political fragmentation hypothesis when it is applied to both OECD and non-OECD countries. Given that the approach is based on a general theory, then there should be evidence to support it. To this end, we apply Volkerink and de Haan's (2001) model of fragmented government and fiscal performance to a panel of OECD and non-OECD democracies. We begin by replicating their study on a panel of OECD countries. Consistent with their study, we confirm the finding that the larger the number of spending ministers the lower the budget surplus. We then add twelve new countries to our sample, ten of them long-standing democracies from outside the OECD. When we do so, none of the measures of fragmented government approaches statistical significance. Thus, we have a puzzle. Why does a model based on a general theory apply to one set of countries but not another? Since we have included a battery of socio-economic controls we argue that the essential difference between the two samples is political. We think that the level of overall institutionalization explains the differential importance of measures of political fragmentation in the two samples. We test this hypothesis by interacting the political fragmentation variables with the age of the democracy, which is our proxy for institutionalization. In so doing, we find that legislative fractionalization reduces the budget surplus, in line with the theory.

Overall, we contribute to the existing literature by demonstrating that the political fragmentation hypothesis applies beyond just a group of rich long-standing democracies. Like previous studies, our findings are sensitive to sample composition. However, unlike previous studies, we provide an explanation as to why our results are sensitive to the sample of countries. Thus, we demonstrate how to take account of political factors when extending a general theory from one context to another quite different context.

2. The debate

In the standard neo-classical model, public debt varies as a function of temporary increases in government spending, for example during wartime, and as counter-cyclical responses to changing government income (Barro 1979). While there was some empirical support for this model, the debt crisis in OECD countries in the late 1970s and 1980s suggested that the explanatory power of the standard model was limited (Roubini and Sachs 1989). In response, alternative models of public debt emerged. Common to these models is that fiscal performance can be seen as an example of the tragedy of the commons, or a common-pool resource problem. Velasco (2000, 122) has formalised this model and described it as follows:

> If we move beyond the view of government as a monolithic entity that behaves like a single individual, economics must provide an account of how economic decisions are made among government groups, and how politics both frames and determines those decisions … [G]overnment net income is a 'commons' from which interest groups can extract resources. This setup has striking macroeconomic implications. Transfers are higher than a benevolent planner would choose them to be; fiscal deficits emerge even when there are no reasons for intertemporal smoothing, and in the long run government debt tends to be excessively high …

This approach to the study of public debt has become popular. In a recent literature review, Alt (2002, 160) sees the common-pool resource problem as one of two approaches that have dominated theoretical research on the fiscal effects of institutions. For Alt, the core intuition of this approach is that "politicians spend more on their constituencies to the extent that they do not internalize the full costs of their spending and taxing decisions. Multiplicity … matters in this model. Competition between claimants on the budget generates a spending bias because each of *n* claimants internalizes on *1/n* of the cost of financing an additional unit of spending" (ibid). In short, much of the contemporary work on budget deficits relies explicitly or implicitly on assumptions consistent with a common-pool resource approach.

There is now a considerable amount of empirical support for the common-pool resource approach to the study of public debt. These studies are underpinned by a common theoretical intuition, but there is disagreement as to which political institutional variables best capture cross-national and intertemporal variations in public debt. In the earlier work, Roubini and Sachs (1989) focused on the impact of coalition versus single-party government. They argued that tough decisions to reduce the level of public debt were likely to be more difficult to achieve under coalition governments than single-party governments. This became known as the 'weak government' hypothesis. While this argument was not framed in the context of a common-pool problem, it is entirely consistent with it. Using a power dispersion index that captured whether or not there was a coalition and, if so, the number of parties comprising the coalition, they tested their hypothesis on 14 OECD countries from 1960 to 1985 and found supporting evidence.

Subsequent work on the weak government hypothesis was less supportive and different political variables were stressed. For example, using a corrected version of the power dispersion index and using data on a sample of 21 OECD countries from 1982 to 1992, de Haan and Sturm (1997) failed to reject a null hypothesis of no difference between coalition and single-party governments. In subsequent work, de Haan, Sturm and Beekhuis (1999) further confirmed that there was little evidence to support the weak government hypothesis. At the same time, on the basis of a sample of 21 OECD countries from 1979 to 1995, they found support for what they termed a 'size fragmentation' hypothesis. Again, while de Haan, Sturm and Beekhuis did not frame their hypothesis explicitly in the form of a common-pool problem, the theoretical exposition of their model is entirely consistent with it (ibid, 165). They found that there was a positive correlation between the number of parties in government and the size of the public sector deficit measured in terms of central government's debt-to-GDP ratio.

For their part, Crain and Muris (1995, 311) explicitly referred to the common-pool problem. Based on a study of US state legislatures, they found (ibid, 326) that revenues were higher in states that merged spending and taxation authority in a single committee compared to those that dispersed authority across a number of committees. They also found (ibid, 328) that expenditure was higher in legislatures where spending authority was spread across multiple committees rather than where it was centralised in one committee.

A further size fragmentation hypothesis was developed by Perotti and Kontopolous (2002). Interestingly, they explicitly placed their contribution in the context of the common-pool resource problem (ibid, 195). They used a panel of 19 OECD countries in the period 1970-95 and adopted a slightly different definition of the budget deficit than de Haan, Sturm and Beekhuis (1999). On this basis, Perotti and Kontopolous found that there was a strong and robust correlation between cabinet size – the number of spending ministers in the government – and the budget deficit. Perotti and Kontopolous also found some support for the relationship between the number of parties in the government coalition and the budget deficit, but only in certain models and to a lesser degree of significance than the finding for the number of spending ministers.

In subsequent work, the significance of the number of spending ministers has been confirmed. In particular, Volkerink and de Haan (2001) used a panel of 22 OECD countries for the period 1971 to 1996 and confirmed that size fragmentation matters for the budget deficit measured, as per their previous study, and in contrast to the measure used by Perotti and Kontopolous, by the budget deficit-to-GDP ratio of central government. Thus, the number of spending ministers appears to be an important determinant of government debt. In addition, they found support for another variable consistent with a common-pool resource problem. Specifically, they hypothesised that there would be a correlation between the government's parliamentary majority and the budget deficit: the greater the government's majority, the lower the deficit/higher the surplus. The intuition here is that as the number of excess majority seats rises, the bargaining power of coalition parties decreases. Again, there was some, albeit limited, empirical support for this hypothesis.

This review has demonstrated that the logic of the common-pool resource problem has underpinned either explicitly or implicitly much of the recent work on the political sources of fiscal performance. It has also shown that there is considerable empirical support for the political fragmentation hypothesis. However, the review has demonstrated that there is disagreement as to which political variables are relevant to fiscal performance. It has also demonstrated that time and time again the results are sensitive to the measurement of the variables as well as period and country selections. Finally, the review has indicated that work on the political fragmentation hypothesis has relied almost entirely on evidence from OECD countries. To date, the only studies based on evidence outside the OECD remain Jones et al.'s (2000) study of Argentina, Stein et al.'s (1999) study of Latin America and Woo's global study (2003).

In this article, we wish to determine whether or not the the political fragmentation hypothesis applies to both OECD countries and long-standing non-OECD democracies. In theory, it should; after all, the common-pool resource problem is a general problem rather than a region-specific problem. If it does, then this would reinforce the policy recommendation that usually flows from the common-pool resource approach, namely that budgetary processes should be reformed so as to establish a hierarchical system in which the Finance Minister can set the budgetary agenda and enforce collective decisions (Hallerberg and von Hagen 1999, 214-216). If budgetary decisions can be delegated to a single, central figure, such as the Finance Minister, then the costs of spending can be more effectively internalised among the set of spending ministers and, assuming adherence to collective decisions can be enforced, the basic prisoner's dilemma at work in the budgetary negotiation game can be resolved (Alesina and Poterba 1999). In the next section, we specify the model we wish to test.

3. Data

Our population is defined by a single criterion: democracy. The indicators of political fragmentation are essentially measures of executive and legislative institutions. In democratic states, it seems likely that these institutions are responsible for economic decision-making. In undemocratic states, such institutions sometimes do not exist and, when they do, they are often supplanted by other, especially informal, decision-making systems (Helmke and Levitsky 2004). We use Freedom House's "Free" classification as our proxy for democracy (Freedom House. 2005). This is a widely used indicator of democracy in both economics and political science (Acemoglu and Robinson 2006; Persson and Tabellini 2003). It is based on an extensive list of political and civil rights. Crucially, it assesses the presence of actual rights, not just legal rights or institutions.[1] We include states in our population if they were continuously classified as “Free” from 1994 or earlier until 2005. This period of ten years entails at least two general elections in most states. This criterion gives us 27 OECD states and 40 non-OECD states, Mexico and Slovakia being OECD members but not having a continuous “Free” rating.

Our sample includes all democracies for which we could find ten observations without missing data. This criterion eliminates three quarters of our non-OECD states. Many of the non-OECD states for which we could not get sufficient data are microstates. In addition, Hungary, Poland, and South Korea are eliminated from the OECD sample. Since many of our measures are derived from the 2004 version of the World Bank’s Database of Political Institutions we begin our observations in 1975 (Beck, et al 2001). Thus, our dataset consists of an unbalanced panel of 34 states between 1975 and 2004. This is a much larger and more diverse dataset than the narrow OECD sample on which the vast majority of the existing empirical literature is based. The ten-year democracy criterion ensures that it is also an appropriate sample.

We replicate the literature on the OECD using Volkerink and de Haan’s (2001) article as a benchmark. They identify three groups of variables.

The first group of variables comprises measures of the size fragmentation of government. This can be thought of in terms of parties and ministers. Volkerink and de Haan use the effective number of parties in government, while

[1] One criterion it does not include is the necessity of a democratic turnover of government (For example, see Przeworski et al. 2000, 23-28). Two of our sample states, Botswana and Namibia, do not meet this criterion. However, the exclusion of either of these countries does not change our results.

we employ Rae's (1971) fractionalistion index to the number and size of parties in government.

$$Fractionalisation = 1 - \sum_{i=1}^{N} p_i^2$$

where N is the number of parties in the government, and p is the proportion of seats held by party i. We obtained this measure from the Database of Political Institutions (DPI), with some supplementary research to fill in a small number of missing observations.

Another measure of size fragmentation is the number of spending ministers. Like Volkerink and de Haan, we count all ministers excluding the finance minister and prime minister. While their source is Woldendorp, Keman and Budge (Woldendorp, et al. 1993; 1998), we have counted the ministers ourselves using Keesing's online (http://www.keesings.com/).[2]

The second group of variables comprises measures of legislative fragmentation. One such measure is the division between government legislators and others. Like Volkerink and de Haan, we use a scaled measure of the government's excess seats in parliament. Volkerink and de Haan also include the effective number of parliamentary parties. Our equivalent is legislative fractionalisation, again as defined by Rae and found in the DPI. Volkerink and de Haan provide a third measure of legislative fragmentation for which we have no equivalent. This is the political fragmentation of parliament, which takes into account the distance between parties on a ten-point left-right scale. No such scale is available for our non-OECD states, largely for the very good reason that in

[2] The Canadian measures are derived from a dataset kindly provided by Matthew Kerby of Memorial University. The UK and Spanish measures come from *Keesing's Contemporary Archive* (http://www.keesings.com/) and *Keesing's Record of World Events* (http://www.keesings.com/keesings_record_of_world_events).

most of them it cannot be assumed that the Western left-right scale is a dominant, or even important, dimension of party competition.

The third group of variables relates to the political fragmentation of government. This aims not just to capture the number and relative size of decision-makers but also the level of dissensus between them. For Volkerink and de Haan this is the equivalent of the previous measure (political fragmentation of parliament) and is again based upon the left-right scale, which is unavailable and inapplicable for us. Our equivalent uses a nominal, as opposed to a scale, measure of political differences. The DPI codes the party types of the three largest governing parties. Multiple left, centre, right, and rural parties were counted as one type. Multiple regional and nationalist parties were counted separately, on the assumption that more often than not multiple parties in these categories will represent different nations or regions. The different religions were counted as separate types, with the category of "not specified" also counted as a separate religion. Once we had identified the number of seats held by the different types of governing parties, we once again applied Rae's fractionalisation index. This measure takes into account the various and multiple dimensions of party competition across the globe. Volkerink and de Haan's second measure of political fragmentation is the maximum ideological distance between governing parties. This measure is again unavailable and inappropriate. Therefore, we provide no equivalent. While these measures do target interrelated phenomena their correlations are well short of the norm for potential severe mulicollinearity.

Volkerink and de Haan control for GDP growth and the change in actual debt-servicing costs. Since the latter measure is unavailable for our sample, we substitute inflation (XZF-IFS from the IMF). We use the same dependent variable: the budget surplus as a proportion of GDP (ZF-IFS from the IMF).

We add several controls appropriate to more diverse sets of states than those contained in Volkerink and de Haan's narrow OECD sample (Persson and

Tabellini 2004; Woo 2003). The new variables are the log of GDP per capita, openness (exports and imports over GDP), log of population, the proportion of population between 15 and 64 and the proportion of population at 65 or over. All are drawn from the World Bank's World Development Indicators except GDP, which is taken from the Penn World Tables. GDP is used to control for the potential effects of economic underdevelopment on the budget. Poor countries may have relatively inefficient tax and spending systems and therefore be more susceptible to budget deficits. Since Cameron (1978) openness has often been shown to be associated with larger government (Cameron 1978) but there is disagreement as to whether it increases or decreases the deficit. Rodrik (1998) argues that the economic vulnerability associated with openness increases the demand for social insurance, thereby increasing the deficit. In contrast, Goode (1984) suggests that openness offers an opportunity for revenue generation through the taxing of trade, thereby decreasing the deficit. Alesina and Wacziarg have argued that government spending is influenced by country size, which determines the scope of economies of scale and the heterogeneity of voters' preferences (1998). Lower deficits should be associated with a larger working age population due to a potentially larger taxable income. Larger deficits are hypothesized to be associated with older populations because of health and social welfare spending.

4. Testing the OECD model

Volkerink and de Haan's dataset contains observations from all "old" OECD countries (less Luxembourg) between 1971 and 1996. We proceed to compare our dataset to that of Volkerink and de Haan. We use an overlap sample, thereby excluding the pre-1975 observations of Volkerink and de Haan and the post-1996, new OECD, non-OECD and Luxembougian observations of our own sample. We begin with an equivalent model, which employs all of Volkerink and de Haan's variables for which we were able to measure equivalents for our model.[3] Given the length of the panels in this sample, we can employ straightforward OLS with a lagged dependent variable. We include dummies for country effects and two periods of 11 years each.

Firstly, we replicated Volkerink and de Haan's model on their dataset, while restricting the sample to those countries and years that were also included in our dataset. As Table 1 shows, the overlap sample and equivalent model with lagged dependent variable reproduces the chief result of Volkerink and de Haan (Volkerink and de Haan 2001): the greater the number of spending ministers, the lower the budget surplus; the greater the level of excess seats, the higher the surplus. Consistent with their results, the other measures of fragmented government are insignificant when included in the full model. Next, we move on to testing the model on our data. Again, the observations are those shared by both datasets. In Table 2, we present results of the Volkerink and de Haan model run on our own data and measures. None of the political fragmentation measures is significant. Also in Table 2, we run our own model, with its expanded set of socio-economic controls, on the overlap sample in our dataset. The number of spending ministers is significant and in the right direction.

[3] We also tested all the fragmentation variables one-by-one for all of the models shown below. For the model in Table 5, we included each fragmentation variable and its interaction with political institutionalization. Also, since they are emphasized by Volkerink and de Haan, we tested Spending Ministers and Excess Seats together without the other three measures. These procedures produced only one variation in our results. In the OLS column of Table 5, Legislative fractionalization was insignificant and its interaction was only significant at the ten per cent level. However, both these variables remained highly significant in the more rigorous GMM version of this model.

Overall, these are very similar results to those of Volkerink and de Haan and reassure us that our measures represent the literature accurately.

[Tables 1 and 2 about here]

Substantively, these results provide some evidence for Volkerink and de Haan's positive but limited finding of the relevance of the fragmented government effects in the OECD. The relevance of spending ministers is confirmed. Methodologically, these results suggest that our global measures are similar enough to those used in the literature restricted to the OECD to reproduce the same results when used in equivalent models and samples.

In Table 3, we present the results for the whole sample, adding on the Czech Republic, Luxembourg and the ten non-OECD countries. We have removed 19 observations where the budget surplus or deficit exceeded 10 per cent.[4] Since several of the non-OECD panels are relatively short, we test the robustness of our OLS results by presenting GMM estimates according to the Arellano-Bond (1991) procedure. For this global sample, none of the measures of fragmented government approaches statistical significance in either OLS or GMM equations. In other studies, the economic and demographic controls we have included facilitated the isolation of significant political effects. In the OECD equations reported above, including the Volkerink and de Haan replication, our political variables were able to reproduce the literature's finding that a larger number of spending ministers increases the budget deficit. Our sample adds twelve countries and 252 observations to Volkerink and de Haan's authoritative work on the OECD. It does so while defining the population in such a way that we can be confident of the applicability of the institutional measures of government fragmentation. Thus, we think our failure to reject the null hypothesis for the global sample is a robust finding.

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[4] All twelve extreme deficits are from Israel and all seven extreme surpluses are from Botswana.

[Table 3 about here]

5. Institutionalisation

As a whole, the literature suggests that findings on political fragmentation are sensitive to the measurement of variables and the inclusion of countries in the sample. We have shown that our variables can produce similar results to Volkerink and de Haan's in an equivalent sample. However, when extended to a new sample, even with a wide range of socio-economic controls, the significance of political fragmentation disappears. Thus, the difference between the two samples may be political rather than socio-economic. Our measures of political fragmentation are institutional. Therefore, we investigate whether political fragmentation matters differently according to the level of institutionalisation of a democracy. The greater the level of institutionalisation, the greater the extent to which formal institutions structure politics. The more institutionalised a democracy the more important we expect political fragmentation effects to be.

Our proxy for institutionalisation is democratic age, specifically the natural logarithm of the number of years since (re-)democratisation (Clague et al 1996, 253; Keefer 2005, 14).[5] Our understanding of young democracies has a lot in common with that of Keefer. He emphasises that politicians in young democracies are less able to make broadly credible promises to voters. Instead, they are only credible when they make very narrow clientelist promises. This

[5] For those countries that were not classified as democracies in 1975 by Freedom House, we begin counting at the year they achieve this classification. For all others, we assume that democratization took place in 1950, except for Barbados, which we take to have been democratic since independence in 1966. Since we have taken the log of democratic age, the fact that we have truncated the age of several of our democracies makes little difference.

means that legislative parties, parties in government and ministries in government are not the fundamental measures of fragmentation in such a political system and do not define the number of actors with access to the common pool. However, as the democracy ages, it institutionalises. Increasingly, institutions become valid measures of fragmentation and define the number of actors that must co-ordinate to manage the common pool.

In Table 4, we proceed to test the power of this institutionalisation argument by incorporating a measure of the years of democracy in our equation, and interacting it with each of our five measures of political fragmentation. Three political variables are significant in both the OLS and GMM equations: the age of the democracy, legislative fractionalisation and the interaction of age and fractionalisation. As the democracy ages, the budget surplus tends to increase. Legislative fractionalisation increases the surplus, while legislative fractionalisation interacted with the age of the democracy reduces the surplus. The second effect is much larger (by 66% if we calculate beta coefficients) and is consistent with our expectations. This aspect of political fragmentation matters more as the democracy becomes more institutionalised. In the early years of a democracy, institutional fragmentation does not deplete the common pool. As a democracy ages, and institutions become embedded, a higher number of instititutionally-defined actors does deplete the common pool of the central government budget.

In line with previous studies, we have found that the effect of political fragmentation on the budget deficit is sensitive to the sample being used. However, in contrast to previous studies, we have provided a theoretical basis as to why the relevance of political fragmentation varies according to context. We have shown that the level of institutionalisation explains the difference between its performance in a sample of rich established countries and a sample of poorer newer democracies.

[Table 4 about here.]

6. Conclusion

The logic of the common-pool resource problem has generated the so-called 'political fragmentation hypothesis', whereby the number of political relevant actors is positively correlated with the level of the budget deficit: the greater the number of actors, the greater the deficit. Various models have found empirical support for the fragmentation hypothesis. However, these models are highly sensitive to political variables, sample size, time periods, and measures of the budget deficit itself. Moreover, to date, work on the political fragmentation hypothesis has been based almost entirely on evidence from OECD countries.

In this article, we replicated an existing fragmentation model on a sample of OECD countries and confirmed the positive results: the size of the budget deficit was correlated with the number of spending ministers and the size of the government's majority in the legislature. However, when we extended this model to include Luxembourg, the Czech Republic and non-OECD democracies, we found that neither measure of size fragmentation was statistically significant. This demonstrated, once again, that the empirical success of the political fragmentation is very sensitive to sample composition. In contrast to previous studies, though, we go further by offering an explanation of the difference in the findings between the two samples. Our explanatory variable is political institutionalisation. Variations in institutionalisation capture the extent to which institutions matter. We find that in the older (more institutionalised) democracies, legislative fractionalisation reduces the budget surplus. Therefore, when controlling for political context, we find that the general theory of the common-pool resource problem can explain variations in fiscal performance outside the usual set of OECD countries.

If correct, this conclusion has important policy implications relating to the organization of the budget process. Based on the logic of the common-pool resource model and supported by empirical studies of OECD countries, there is evidence that a hierarchical and centrally coordinated budgetary process can offset some of the problems associated with fragmentation effects. Our study suggests that such a policy recommendation may be more appropriate in some contexts than others. Specifically, this recommendation may be more appropriate in more institutionalized countries and less so in less institutionalized non-OECD countries.

***Table 1*. Volkerink and DeHaan dataset overlap sample equivalent model (OLS)**

Lagged surplus	0.7387812	(0.0351833)**
GDP growth	0.3002017	(0.0539653)**
Inflation	0.0299514	(0.0150971)
Government fragmentation	-0.0021409	(0.0026597)
Ideological frag. of govt.	0.0003493	(0.0013044)
Parliamentary fragmentation	0.0000653	(0.0013491)
Spending ministers	-0.0008763	(0.0003152)*
Excess seats	0.0265879	(0.0136476)
Constant	-0.0138255	(0.0108602)
No. Obs.	446	
R^2	0.7819	

Notes: Period: 1975-1996 (inclusive). Model is OLS. This regression includes dummies for N-1 states and a dummy the panel into dividing into two eleven-year periods. Standard errors (in parentheses) are clustered by country.

Table **2. 'Our dataset' overlap sample (OLS)**

	Volkerink and de Haan model	'Our model'
Surplus (lag)	0.7288117	0.648946
	(0.0387533)**	(0.0380914)**
GDP growth	0.0020279	0.0020362
	(0.000358)**	(0.0004113)**
Inflation	-0.0001212	0.0002636
	(0.0001933)	(0.0001289)
GDP pc (log)	-	0.0618672
		(0.0147118)**
Openness	-	-0.0003348
		(0.0001462)*
Population (log)	-	-0.0053282
		(0.0105144)
Working age population	-	0.0003259
		(0.0009251)
Population over working age	-	-0.0017679
		(0.0008934)
Government fractionalisation	-0.0026792	0.0042179
	(0.0086534)	(0.0112223)
Political fractionalisation	-0.0022629	-0.0097439
	(0.0094119)	(0.009918)
Legislative fractionalisation	-0.00263	-0.0244969
	(0.0243266)	(0.0338449)
Spending ministers	-0.0001363	-0.0011416
	(0.000377)	(0.0005257)*
Excess seats	0.000045	0.0000289
	(0.0000795)	(0.0000733)
Constant	-0.0098602	-0.5272494
	(0.0173675)	(0.1600267)**
No. obs.	443	443
R^2	0.738	0.757

Notes: Period: 1975-1996 (inclusive). Dummies for N-1 states and a dummy dividing the panels into two eleven-year periods are included in all regressions. Robust standard errors (in parentheses) are clustered by country. * significant at 5%; ** significant at 1%.

Table 3. **Global Sample**

	OLS	GMM
Lagged surplus	0.321326 (0.070568)**	0.003851 (0.164999)
GDP growth	0.045812 (0.024369)	0.036787 (0.017919)*
Inflation	0.007739 (0.001183)**	0.015781 (0.015611)
GDP per capita (log)	0.861441 (0.514574)	3.396473 (1.672466)*
Openness	0.002909 (0.006298)	-0.006424 (0.010308)
Population (log)	-0.239455 (0.108539)*	1.588479 (2.227929)
Population (15-64)	0.002693 (0.054534)	0.087336 (0.102567)
Population (65 plus)	-0.037852 (0.046280)	-0.098375 (0.080972)
Government fractionalisation	-0.063633 (0.350545)	-0.376454 (0.403475)
Political fractionalisation	-0.155009 (0.257410)	-0.216160 (0.242959)
Legislative fractionalisation	-0.008314 (1.291967)	3.487557 (1.827505)
Spending ministers	-0.003178 (0.015908)	-0.016117 (0.025137)
Excess seats	0.000769 (0.001742)	0.000695 (0.001698)
Constant	-6.799940 (4.810742)	-0.061642 (0.029795)*
No. obs.	721	649
R^2	0.68	-
2^{nd} Order test	-	0.2873

Notes: Notes: Period: 1975-2004 (inclusive). Dummies for N-1 countries and two dummies dividing panels into three ten-year periods are included in all regressions.

Robust standard errors in parentheses are clustered by country in OLS. * significant at 5%; ** significant at 1%. 2nd order test is a P-value for rejecting the null hypothesis that there is no second order correlation in the first-difference residuals.

Table 4. **Global Sample with Institutionalisation**

	OLS	GMM
Lagged surplus	0.286719 (0.067105)**	-0.007057 (0.137685)
GDP growth	0.049531 (0.025046)	0.037363 (0.019835)
Inflation	0.006773 (0.001142)**	0.012536 (0.015413)
GDP per capita (log)	0.476686 (0.485656)	2.149774 (1.342903)
Openness	0.009005 (0.007258)	-0.000422 (0.009945)
Population (log)	-0.305878 (0.106751)**	2.731416 (1.849750)
Population (15-64)	-0.012183 (0.053249)	0.046405 (0.067036)
Population (65 plus)	-0.014052 (0.037458)	0.090151 (0.065073)
Government fractionalisation	1.005341 (1.652241)	-0.377070 (2.571864)
Political fractionalisation	-5.135380 (2.809330)	-4.575895 (3.441017)
Legislative fractionalisation	11.232288 (5.061100)*	28.732208 (9.626791)**
Spending ministers	0.057424 (0.112889)	0.093608 (0.149756)
Excess seats	0.017074 (0.016217)	0.016960 (0.010511)
Years of Democracy (log)	2.664271 (1.074375)*	4.784122 (1.898319)*
Yrs. of Demo. * Gov. Frac.	-0.291134 (0.440554)	0.007135 (0.761048)
Yrs. of Demo. * Pol. Frac.	1.518393 (0.813888)	1.272336 (1.000629)
Yrs. of Demo. * Leg. Frac.	-3.630929 (1.432647)*	-7.739575 (2.637629)**

Yrs. of Demo. * No. of Mins.	-0.019479 (0.033321)	0.020630 (0.043839)
Yrs. of Demo. * Excess Seats	-0.005132 (0.004745)	-0.004550 (0.002990)
Constant	-10.633789 (5.475545)	0.045949 (0.023622)
No. obs.	721	649
R^2	0.69	-
2nd Order test	-	0.3515

Notes: Notes: Period: 1975-2004 (inclusive). Dummies for N-1 countries and two dummies dividing panels into three ten-year periods are included in all regressions. Robust standard errors in parentheses are clustered by country in OLS. * significant at 5%; ** significant at 1%. 2nd order test is a P-value for rejecting the null hypothesis that there is no second order correlation in the first-difference residuals.

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