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Fabrizio Carmignani

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POLITICAL BIAS IN FISCAL POLICY FORMATION: AN ECONOMETRIC ANALYSIS OF COALITION SYSTEMS

Fabrizio Carmignani*

Dipartimento di Economia Politica, Università Statale di Milano Bicocca

Department of Economics, Glasgow University

Abstract

This paper provides a comprehensive econometric analysis of some debated issues concerning the political and institutional determinants of fiscal policy outcomes. Several innovative results are obtained. It turns out that a significant effect on fiscal policy formation can be traced back to the ideological orientation of the policymaker, to the degree of cabinet instability, to cross-country differences in electoral and budgetary institutions and to the dispersion of political power within the ruling coalition. Instead, the preferences of the median voter appear to have little importance. The evidence also rejects the theory of fiscal illusion in decision-making.

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* Corresponding address: Dipartimento di Economia Politica, Edificio U6 III Piano, Università Statale di Milano Bicocca, Piazza dell'Ateneo Nuovo 1, 20126 Milano, e-mail: fabrizio.carmignani@unimib.it I wish to thank Anton Muscatelli, Luca Stanca, Patrizio Tirelli, Ulrich Woitek and participants at seminars at Università Statale di Milano Bicocca and Università Cattolica del Sacro Cuore for helpful comments. All remaining errors are solely my own responsibility.

Introduction

The objective of this paper is to provide a comprhensive econometric analysis of some open issues concerning the political and institutional aspects of fiscal policy formation in western European coalition systems. To this purpose, a newly constructed data-set of political indicators is used. Several interesting and innovative results are obtained. First, there is no evidence that fiscal policy is systematically manipulated when elections are approaching. Second, the ideological orientation of the policymaker does matter: leftist governments appear to spend and tax more than rightist-ones. However, in aggregate, leftist governments appear to sustain smaller fiscal deficits. Third, dispersion and fragmentation of political power within the decision making process affect the ability of the incumbent to run a tight fiscal policy. Fourth, there is a significant positive correlation between the degree of government volatility and the size of fiscal deficit, especially when government changes involve high alternation of parties in office and volatility of portfolios. Fifth, electoral rules and budgetary processes do play some role. Finally, the impact of the preferences of the median voter, when contrasting with those of the government in office, appears to be negligible.

The rest of the paper is organised as follows. Section 1 introduces the issues at stake. Section 2 defines the econometric framework and provides a necessarily brief description of the political data-set. For further details on this data-set, the interested reader is invited to refer to Carmignani (1999). Section 3 presents the econometric results. Section 4 concludes. The tables with the full set of econometric results can be found in the Appendix.

Section 1. Sources of political bias in fiscal policy formation.

A widely accepted theory of optimal budget policy based on strictly economic considerations is the *tax smoothing approach* advanced by Barro (1979) and Lucas and Stokey (1983). In a

nutshell, the theory predicts that budget deficits and surpluses should be used optimally to minimise the distortionary effects of taxation, given a certain path of public spending. A clear implication of the model is that tax rates should be kept constant over the business cycle and hence deficits should arise during recessions and surpluses during expansions.

In fact, the conduct of fiscal policy observed in most countries over the last three decades does not seem to be fully consistent with the predictions of the tax smoothing theory. Political economy models posit that in addition to strictly economic considerations, political and institutional factors affect the decisions of the policymaker, so that a *political bias* in fiscal policy formation arises. The literature has identified five basic sources of potential bias. This paper investigates the effective importance of each of these sources.[1]

The first potential source of bias goes under the name of *fiscal illusion* (Buchanan and Wagner, 1977). When voters overestimate the benefits of current spending and do not recognise the costs of future taxation, an opportunistic incumbent has the incentive to manipulate fiscal policy in order to rise his share of consensus when elections are approaching. The result is an increase in the size of fiscal deficit (higher expenditure and lower taxation) in pre-electoral periods. Evidence in support of this hypothesis is provided by Nordhaus (1989) and Corsetti and Roubini (1992), some counter-evidence is instead given by Alesina Roubini and Cohen (1997).

The second form of bias is the *ideological orientation of the policymaker*. Leftist governments are usually assumed to promote higher levels of spending and taxation compared to rightist ones. The overall impact on the size of the deficit is not clear *ex-ante*, albeit there is a rough consensus that a leftist ideology should be associated to more expansionary fiscal policy (see the discussion in Borrelli and Royed, 1995). As a matter of fact, Borrelli and Royed (1995)

[1] Alesina and Perotti (1994) survey the political economy literature on budget deficits. Persson and Tabellini (1997) and Alesina, Roubini and Cohen (1997) also provide detailed discussions of various related contributions.

and Roubini and Sachs (1989) do obtain that the ideology of the policymaker (measured on the basis of the ideologies expressed by the political parties that compose the ruling coalition) does matter in policy formation. But again, Alesina, Roubini and Cohen (1997) obtain that the size of the fiscal deficit is not significantly linked to the political colour of the incumbent.

The third source of bias is the myopic behaviour of the incumbent in countries characterised by *high government instability*. Faced with a positive probability of being replaced in office by a party with different policy preferences, the incumbent has an incentive to use fiscal policy strategically to tights the hands of his successors. In particular, large fiscal deficits are run by an incumbent that does not fully internalise the costs of debt accumulation (Alesina and Tabellini, 1990 and 1991). Grilli, Masciandaro and Tabellini (1991), Darby, Li and Muscatelli (1998), De Haan Sturm and Beekhuis (1999) and Petterson (1999) all provide evidence which is broadly consistent with the idea of a positive relationship between instability and size of the deficit (or level of consumption expenditure). However, in a systematic test of the Alesina and Tabellini' s model on a sample of OECD countries, Lambertini (1998) obtains non supportive evidence.

The fourth form of bias relates to the *dispersion of political power* within the ruling coalition. The idea is that when different political parties, holding different policy views, must all participate in the decision making process (as it is the case, for example, with coalition and/or minority governments), then a significant bias towards higher levels of government spending should be observed as a consequence of the attempt of each party to secure greater spending for its supporting constituency (Roubini and Sachs, 1989). Alesina and Dazen (1991) formalise a game-theoretic argument to show that coalition governments are more likely to generate debt growth. The empirical results in Roubini and Sachs (1989) and Perotti and Kontopoulos (1999) confirm the hypothesis that fragmentation increases fiscal deficit, whilst those in Franzese (1998) and De Haan and Sturm (1997) are not supportive of the theory.

Finally, political bias in fiscal policy formation might be due to the specific *institutional arrangements* prevailing in a country. Grilli, Masciandaro and Tabellini (1991) point out that there is a significant correlation between the type of electoral rule (proportional representation or plurality system) and the rate of debt accumulation. Von Hagen (1992) argues that budgetary procedures (such as the degree of centralisation of negotiations, the flexibility of budget execution, the existence of long-term planning constraints) contribute to determine the ability of a government to keep fiscal deficit growth under control. Hallerberg and Von Hagen (1997) suggest that electoral institutions determine the budgetary procedures at work in a given country at a given time. They also provide evidence to support the idea that these procedures affect the size of fiscal deficit in the EU countries. However, De Haan, Moessen and Volkerink (1999) find the effective impact of budgetary procedures to be rather small.

As it can be seen from the previous discussion, several issues concerning the effective importance of the various sources of political bias still represent open questions. This paper provides a comprehensive econometric analysis of these questions by making use of a newly constructed data-set to incorporate the key aspects of each potential source of bias. The econometric framework and the model specification for this analysis are defined in the next Section.

Section 2. Econometric set-up and model specification

The sample consists of annual observations on 13 western European coalition systems from 1960 to 1995.[2] I estimate a regression model of a linear form:

[2] Countries are: Austria, Belgium, Denmark, Finland, France, Germany, Iceland, Ireland, Italy, Luxembourg, the Netherlands, Norway, Sweden. All these are classified as *coalition systems* by political scientists.

(1) $y_{it} = \mathbf{X_{it}}\mathbf{b_1} + \mathbf{Z_{it}}\mathbf{b_2} + e_{it}$

where *i* denotes a generic country and *t* a generic year, $\mathbf{X}_{it}$ is a vector of observations on exogenous economic variables, $\mathbf{Z}_{it}$ is a vector of observations on exogenous political variables, $\mathbf{b}_1$ and $\mathbf{b}_2$ are two vectors of coefficients to be estimated and ε_{it} is a disturbance term.

As it is well known, the procedure for the estimation of model (1) depends upon the specific assumptions concerning the form of the disturbance term ε. If it is assumed that $\varepsilon_{it} \sim iid\ (0,\sigma^2)$ for all *i* and *t*, then a standard *pooled OLS estimator* can be used. If instead it is assumed that ε_{it} consists of an individual (country-specific) effect plus a random component, then an alternative *Random Effect* (RE) or *Fixed Effect* (FE) estimator must be used, depending on whether or not the individual effect is uncorrelated to the set of exogenous variables. The tables in the Appendix report for each specification of model (1) the results obtained with the estimator which is "favoured" by the Breusch and Pagan test of homoscedasticity of disturbances and by the Hausman test of orthogonality.[3] Furthermore, when the panel is unbalaced, the procedure designed by Verbeek and Nijman (1996) is adopted.

Three different measures of fiscal policy outcomes are considered as the dependent variable in model (1): the year-to-year change in the debt-to-GDP ratio (DEF) as a measure of fiscal deficit, the annual change in the ratio of government consumption expenditure to GDP (DG) and the annual change in the ratio of total tax revenues to GDP (DT).

[3] Breusch and Pagan (1979 and 1980) define a Lagrange Multiplier statistic to test the null hypothesis of homoscedaticity of disturbances. High values of this test-statistic thus favour the pooled OLS estimator against the RE and FE estimator. Hausman (1978) proposes a statistic to test the null hypothesis that the individual effect and the exogenous variables are uncorrelated. Thus, high values of this statistic favour the FE estimator against the RE estimator.

The set of economic regressors (that is, the variables in vector **X**) is designed to capture the key punch line of the tax smoothing argument, namely that fiscal deficits should be observed during recessions and surpluses (or smaller deficits) during expansions. Following Alesina, Roubini and Cohen (1997), and consistently with most of the empirical contributions in this field, four economic variables are considered: the lagged value of deficit[4] (DEF_{-1}), the change in the rate of unemployment (DU), the change in the rate of output growth (DY), the change in the difference between the real interest rate and the rate of GDP growth times the lagged debt-to-GDP ratio (Dr).[5] According to the theoretical argument, deficits should adjust slowly, so that the estimated coefficient on DEF_{-1} should be positive, but smaller than one. The fact that deficits have a counter-cyclical behaviour (high during recessions, low during expansions) implies that the coefficient on DU is expected to be positive, whilst the one on DY is expected to be negative. Finally, as the real rate of interest grows faster than output, the government might be tempted to compensate the increased burden of debt through fiscal deficits. Henceforth Dr should display a positive estimated coefficient. As it will be stressed in Section 3, the set of economic regressors has been further extended to include dummy variables for the oil shock and the end of the Bretton-Woods system, as well as, the lagged value of debt and the share o population aged over 65. However, all these additional variables do not appear to have any significant explanatory power.

When the dependent variable is either DG or DT, only DY and DU are included in vector **X**. The ex-ante expectations is that the coefficient on DU should be positive and the one on DY negative in the regression of DG. The opposite is true for the regression of DT.

[4] The inclusion of a lagged dependent variable might cause a problem of correlation with the disturbances when the model is estimated by RE or FE. In fact, given that my panel consists of a relatively long time dimension combined with a relatively short cross-section dimension, the Feasible GLS procedure I have adopted (see Judge et al., 1985) guarantees correct estimates. Moreover, for most of the specifications that include the lagged dependent variable, the favoured estimator (reported in the tables) is the pooled OLS, for which the problem encountered with FE and RE does not arise.

[5] The variable Dr is formally defined as $b_{-1}d(r\text{-}y)$, where b is the lagged value of the debt-to-GDP ratio, $(r\text{-}y)$ is the difference between the real interest rate r and the growth rate of GDP, y, and d denotes annual change.

Data on the economic variables are taken from the OECD-Statistical Compendium and from the Institutional Monetary Fund-International Financial Statistics and International Government Statistics.

The set of political regressors (that is, the variables in the vector **Z**) is specified to account for the five possible sources of political bias mentioned in Section 1. Here below is a short description (based on Carmignani, 1999) of the variables used in this analysis.

A. Indicators used to test the fiscal illusion argument

Two electoral dummies are defined. ELE takes value 1 in the pre-electoral year, if elections are held in the first or in the second quarter of the electoral year, and value 1 in the electoral year, if elections are held in the third or in the fourth term of the electoral year. ELX takes value 1 in the electoral year, independently from the quarter of the elections. According to the fiscal illusion argument, both dummies are expected to display positive coefficients in the regression of DEF (and DG): when elections are approaching the incumbent undertakes expansionary monetary policies to increase his share of consensus. In the regression of DT the opposite argument can be made and the estimated coefficients should be negative.

B. Indicators used to test the role of ideology of the policymaker

Given a policy scale that reports on a left-right continuum the cardinal location of political parties in a given country at a given time, the ideology of the policymaker can be computed as a weighted average of the ideological locations of the parties involved in the process of policy formation. Policy scales are made available in the political science literature for a number of modern democracies. They are constructed on the basis of the policy proposals stated in the party manifestos and hence they reflect *ex-ante* public policy positions. On these scales,

ideology is measured on a uni-dimensional space, where 1 represents the extreme left and 10 the extreme right. Henceforth, the ideology of a given party is a number x such that $1 \leq x \leq 10$ (see, for example, Huber and Inglehart, 1995). The key problem is then the one of finding an appropriate system of weights to obtain the aggregate measure of ideological orientation of the government (policymaker). Ideally, the weight attached to individual's party ideological locations should reflect the effective contribution that each party gives to the process of policy formation. In turn, the contribution of any party will depend upon the specific structure of the decision making process. Political scientists identify four possible structures of the decision making process in coalition systems (see, *inter alia*, Laver and Shepsle, 1994). For each of these four structures, a different system of weighs, and hence a different measure of the ideology of the policymaker, is defined. LOC1 is a measure where the location of each party in the ruling coalition is weighted by the share of cabinet portfolios hold by that party. The associated structure of decision making is one where policy formation is effectively a collective activity to which each individual minister contributes (a so called *cabinet government*). In the measure LOC2, the location of any party in the ruling coalition is weighted by the share of coalition seats held by that party. The underlying type of decision-making process is one where the executive is heavily constrained by party leaders and party agreements (a so called *party government*). LOC3 assigns a weight equal to 1 to the ideological location of the party holding the office of Prime Minister and 0 to the locations of all the other parties. This is the case of a *prime-ministerial government*, where decision making is essentially a monocratic exercise of a powerful prime-minister. LOC4 is computed by assigning to the ideological location of each coalition partner a weight equal to the share of *key* portfolios held by that party. Key portfolios are those whose jurisdiction extends over the most important areas of the policy space (key portfolios are identified following Laver and Hunt, 1992). The associated structure is one where decision making displays a strong

departmental character: the contents of the policies undertaken in a specific area are heavily constrained by the orientation of the party in control of the portfolio with jurisdiction over that area (*ministerial government*). A strong version of LOC4 will assign weight 1 only to the party in control of the key portfolio of finance.

The four measures thus obtained are cardinal in the sense that their estimated coefficients can be given partial derivative interpretations. However, besides these cardinal measures of ideology, a full set of dummies is constructed. ID is a dummy taking value 1 if in a given country in a given year the ideological location of the incumbent is higher than the threshold value 5.5. That is, ID is based on a bi-partition of the policy space and it is coded 1 for left-oriented governments. Of course, there are four dummies ID, each referring to a specific cardinal measure LOC. IDL takes value 1 if in a given country in a given year the corresponding measure LOC is below the threshold 4.6. IDC takes value 1 when LOC is above 4.6 but below 6.4. IDR takes value 1 when LOC is above 6.4. Thus, IDL, IDC and IDR refer to a tri-partition of the policy space (Left, Centre and Right). Again, since there are four measures LOC, four sets of dummies IDL, IDC and IDR are computed.

Since leftist policymakers are expected to promote greater spending and higher taxation, LOC should display a positive coefficient in the regression of both DG and DT. The same is true for IDL, whilst IDR should be associated to a negative coefficient. The coding of the dummy ID implies that its expected coefficient in the regression of DG and DT should be negative. Finally, it is not clear ex-ante whether fiscal deficits should be systematically larger with left-oriented governments. If this were indeed the case, then the ideology variables should display in the regression of DEF coefficients of the same sign as those in the regression of DG.

C. Variables used to role of fragmentation of political power.

The basic argument relating to the fragmentation of political power can be tested by mean of two dummies, COAL and MIN. COAL takes value 1 when the incumbent government is supported by a coalition of two or more parties.[6] MIN is coded 1 in the years when a minority government is in office. It is expected that both dummies exhibit negative coefficients in the regression of DEF and DG. A more sophisticated measure of fragmentation should take into account the effective degree of dispersion of the policy views of parties in the ruling coalition. To this purpose, the variable CI is defined as the variance of the cardinal locations of the parties supporting the government. Clearly, for single party governments, CI takes value 0. An alternative indicator of fragmentation is the *effective number of parties* (ENP) in the coalition (see Laasko and Taagepera, 1979 for a definition of the concept of effective number of parties): the higher this number, the more fragmented the coalition. The hypothesis of dispersion of the political power suggests that in the regression of DEF and DG both ENP and CI should display positive coefficients. The polarisation of the legislature as a whole is captured by the variable POL, which is defined following Powell (1982) as the share of support for *extremist* parties. As intra-parliamentary negotiations become more difficult in heavily polarised systems, the coefficient on POL in the regression of DEF and DG should again be positive.

D. Variables used to test the role of government instability

A rather common approach adopted in the literature is to use the average number of government transfers observed in the previous *n* years (where *n* could be eventually equal to one) as a proxy for instability. Often, minor transfers are separated from major changes (that is, transfers that involve a change in the composition of the ruling coalition and/or a shift of

[6] As noted in footnote 2, the sample does include several single-party governments. Thus the distinction between coalition and single-party governments is not meaningless.

the ideological orientation of the policymaker). Here I follow a different strategy. Through probit analysis I estimate the probability for a given government at a given time to collapse. According to the theory, instability leads to strategic accumulation of deficits to the extent that the incumbent is afraid to be replaced by a successor with significantly different preferences. Thus, I weight the measure of instability obtained from the probit model by the past (observed) alternation in office and portfolios volatility. Alternation in office (ALT) is defined following Powell (1982) as the sum of shares of parliamentary seats held by parties entering the government plus the shares of parties leaving the government. Total portfolios volatility (TPV) is the total number of portfolio transfers observed between two consecutive cabinets. Party portfolio volatility (PPV) is the number of transfers occurring between any two different parties in two consecutive cabinets. Ideological portfolios volatility (IPV) is the number of transfers between different parties weighed by the ideological distance between the parties.[7] Obviously, these measures of instability should all display positive coefficients in the regression of DEF (and DG).

E. Variables used to test the impact of electoral insitutions and budgetary procedures

In most of the coalition systems included in the sample, a proportional representation (PR) system is at work. However, whilst some countries adopt a purely PR rule, others combine the PR rule with various procedures for the allocation of a quota of seats. I therefore define a dummy PR taking value 1 for purely PR systems and zero otherwise. Notice that the group of non purely PR systems includes a variety of alternative rules, such as the single transferable vote in Ireland, the plurality system in France, the 2-Tier PR system plus transfers for the reminder in Austria, Denmark, Italy (prior to 1993), Germany and Sweden. As noted by Hallerberg and Von Hagen (1997), it is commonly argued that PR systems should run larger deficits and hence the coefficient on PR in the regression of DEF and DG should be positive.

[7] The measures of portfolios volatility are originally proposed by Huber (1998).

In terms of budgetary procedures, I follow Hallerberg and Von Hagen (1997) and divide the countries of the sample in three groups: those where fiscal powers are delegated to a strong Ministry of Finance, those where commitments to negotiated fiscal targets defined in contract agreed upon by coalition partners are usually taken and those where no such constraint is at work. Then, three dummies are used: DELEGATION takes value 1 for the countries in the first group (France and Germany), CONTRACT takes value 1 for the countries in the second group (Austria after 1985, Denmark after 1982, Finland, Ireland after 1987, Luxembourg, the Netherlands), UNCONSTRAINED takes value 1 for the countries in the third group (Belgium, Italy, Sweden).[8] The degree of tightness of fiscal policy should be minimum in the absence of constraint, intermediate with fiscal contracts, and maximum with delegation to a strong ministry of finance.

Section 3. Econometric results: political bias in fiscal policy formation.

The tables in the Appendix report the results of the estimation of various specifications of model (1). In each Table, the first row reports the number of the Column, the second row the dependent variable and the last row the estimators favoured by the statistical tests mentioned in Section 2. The full set of results obtained with the non-favoured estimators is available from the author upon request.

Before focusing on the importance of the five sources of political bias, it is worth considering the purely economic specification of Column 1 Table 1. All the estimated coefficients of the economic regressors are consistent with the *a priori* expectations based on the tax smoothing model. In particular, there is evidence that deficits adjust slowly over time and exhibit a

[8] For Norway and Iceland I could not obtain reliable information concerning the type of procedure and hence they are not included in the sample when the model specification includes the dummies for the type of budgetary process.

counter-cyclical behaviour. To check the robustness of this basic economic specification, dummies to account for the oil shocks and the end of the Bretton Woods system, the lagged value of debt-to-GDP ratio and the share of population aged over 65 have been added on the r.h.s. The coefficients on the four key variables do not change substantially, whilst those on the added variables are not statistically different from zero. Moreover, in order to control for the potential joint endogeneity of deficit and business cycle, DU and DY have been instrumented using their lagged values. Again, no significant changes in the estimated coefficients are observed. Thus, it appears that the basic economic specification in Column 1 is an appropriate benchmark for the analysis of political bias in fiscal policy formation.

3.1 Fiscal illusion

In Column 2 of Table 1 the electoral dummy ELE is added to the basic economic specification to see whether the size of fiscal deficits displays any systematic pre-electoral pattern. It turns out that this does not seem to be the case: budget deficits are not significantly higher in electoral or pre-electoral years. The same conclusion holds when the dummy ELX replaces ELE. In addition to that, ELE (and ELX) does not seem to play any relevant role in the determination of consumption expenditure and taxation. In the regression of DG, ELE does display a positive coefficient, as the theoretical argument suggests, but this coefficient is not significant at usual confidence levels. Similarly, in the regression of DT, the negative coefficient on the electoral dummy fails to pass a zero restriction test. Based on this evidence, I suggest that in the western European coalition systems, over the time 1960-1995, there is little support for the hypothesis of fiscal illusion as a relevant source of political bias in fiscal policy formation.

3.2 The ideological orientation of the policymaker

In Column 3 of Table 1, the ideological variable LOC is included. The estimates reported are those obtained when the specific definition LOC4 is used to measure the ideology of the policymaker using a system of weights based on the share of key portfolios controlled by each party. Even if the key findings are unchanged when the other definitions of location are used, it is worth stressing that LOC4 (and the associated dummies) does seem to have a superior econometric performance. That is, the level of significance of its estimated coefficients is higher and the general goodness of fit of the model is greater relative to the case where the other definitions of LOC are used. Furthermore, in some regressions, when the variable LOC4 (or one of the associated dummies) is entered jointly with any of the other ideological variables, the coefficient on LOC4 retains its size and significance, whilst the one of the other ideological variable becomes not different from zero. In this sense, it could be that using a system of weights based on the share of key ministers controlled by each party allows a better (more precise) econometric representation of the effective ideology of the policymaker. Further work on this point is thus desirable.

With the favoured estimator (OLS), there seems to be no statistically significant correlation between ideology and size of fiscal deficit, once controlling for the economic determinants of fiscal policy. The result is confirmed when the dummy ID (defined over a simple bi-partition of the policy space) replaces LOC. However, when the dummies DL and DC (defined over a tri-partition of the policy space) are added some interesting results, reported in Column 4, appear. The negative and significant coefficient on DL implies that leftist governments (located to the left of the threshold 4.6 on the ten points scale) sustain *smaller* deficits relative to rightist governments. This same difference shows up in a model where the three dummies

DC, DL and DR are added jointly (and the constant term dropped). In this case, DR displays a positive and significant coefficient, thus rightist governments appear to sustain significantly *larger* deficits relative to leftist (and possibly centrist) governments. Notice that this finding does not contradict the standard argument, commonly advanced in the political economy literature, that left oriented policymakers tend to favour more public spending. As the estimated coefficients on the dummy ID in Column 5 and 6 suggests, leftist governments effectively spend more, but they also tax more (recall that ID takes value 1 when the incumbent is ideologically shifted to the left of the policy scale). In other words, they appear to be more able than rightist governments to balance expenditures and revenues in aggregate. I can identify two possible, not mutually exclusive explanations for this result. The first one has to do with the fact that, at least in the sample considered in this analysis, leftist governments are characterised by a lower degree of internal fragmentation; that is, they are supported by a smaller number of parties and/or by parties with relatively homogeneous policy views. This higher internal cohesion makes negotiations over budget balancing significantly easier and hence it favours smaller deficits (see also Subsection 3.3). Evidence consistent with this explanation is provided in Column 7 of Table 1, where the estimated coefficients on the ideological dummies are found to be not statistically different from zero when the rough indicator of internal fragmentation COAL is added to the set of regressors.

The second explanation relies on the theory of "labour quiescence" developed by Cameron (1984). Following this argument, leftist governments face a less fierce opposition from the unions and this makes them intrinsically more stable. Carmignani (2000) finds that effectively, the probability for any government to collapse is smaller for left-oriented incumbents. As it will be discussed in Subsection 3.4, the ideological dummies significantly enter the probit specification used to estimate the instability of the government, thus

reinforcing the hypothesis that political colour affects fiscal policy outcomes through its impact on some structural characteristics of the policymaker.

3.3 The degree of internal cohesion, fragmentation of the decision making process and polarisation of the legislature.

The greater number of supporting constituencies imply that coalition governments should be less able than single-party ones to limit the size of fiscal deficit. The results in Column 1 of Table 2 are consistent with this hypothesis: after controlling for the key economic variables, coalition governments are found to sustain significantly larger deficits. In Column 2 an interesting finding is obtained: fiscal deficit is lower with minority governments. A possible interpretation for this finding is that minority governments are essentially single-party ones and hence they are internally more cohesive than most majority governments supported by coalitions of different parties. The fact that when entered jointly on the r.h.s. COAL and MIN loose any explanatory power suggests that the two variables might be collinear and hence that they are likely to capture the same effect of single vs. multi-party governments. Furthermore, when explicitly isolated, the impact of coalition-minority governments is to *increase* fiscal deficit, as it can be seen from the estimated coefficient on the interactive term COAL*MIN in Column 3. Thus, minority governments effectively generate lower deficits only when formed by a single-party. To put it differently, the minoritarian status of the incumbent is not a disadvantage, in terms of ability to control fiscal policy outcomes, as long as it is not associated with the coalition status. The implication of this argument is that for fiscal policy formation, negotiations within the cabinet might be more crucial than negotiations between the cabinet and the parliament. A result that, if confirmed by future research, would reverse the conclusion reached Edin and Ohlsson (1991). It is quite a disappointment to notice that

more sophisticated measures of fragmentation of the ruling coalition (such as CI and ENP) fail to play any significant role in the regression of DEF, not even when combined with the dummy for the majority status, as it is the case of the term MAJ*CI in Column 4. The inclusion of these more sophisticated measures however does not alter the previous findings on MIN and COAL*MIN.

Even if the above findings suggest that what really matters for fiscal policy formation is intra-coalition bargaining, the overall degree of polarisation of ideological preferences in the party system might be a possible source of inability for the incumbent to reduce spending and deficits. In particular, the effect of the degree of polarisation would arise in the presence of minority governments. The positive coefficient on the interactive term MIN*POL in Column 5 of Table 2 is consistent with this hypothesis.

3.4 Government instability

A standard approach in the political economy literature is to use the number of government changes in a given period as a proxy for government instability. Alternatively, the average duration of the cabinet or its survival rate (the ratio of duration to maximum time between two consecutive elections) can be entered the model specification to capture the impact of excessive turnover in office on the process of fiscal policy formation. In this paper, I adopt a different approach, that follows the one in Alesina et al. (1992). First, I compute the probability for a government to collapse in country *i* in year *t* through the estimation of a binary choice model. Then, this probability is entered as a proxy for government instability on the r.h.s. of model (1).

A systematic analysis of the determinants of government duration in western European coalition systems is undertaken in Carmignani (2000). Building on those results, I choose a

parsimonious specification for the binary choice model, that includes the political factors which are most likely to determine government collapse. The dependent variable of the model is a dummy TERM taking value 1 in years when a termination is observed. The model is thus specified as follows:

$$(2)\quad TERM_{it} = const + \gamma_1 TERM_{it-1} + \gamma_2 MAJ_{it} + \gamma_3 SING_{it} + \gamma_4 CI_{it} + \gamma_5 POL_{it} + \gamma_6 DL_{it} + \gamma_7 GR_{it} + \gamma_8 GR_{it-1} + e_{it}$$

where *const* is the intercept and *GR* the growth rate of GDP, γ's are the coefficient to be estimated and all the other variables are as described in Section 2.

The specification of model (2) incorporates the idea that government termination depends on the past history of terminations (a form of *duration contagion*), on several characteristics of the party system and of the ruling coalition (including its ideological location) and on the economic performance as captured by the growth rate of GDP. Both a probit and a logit version of model (2) are estimated. In addition to that, a basic specification where only the lagged value of TERM and the dummies MAJ and SING enter the r.h.s. Results are reported in Table 3.

From the set of estimates in Table 3, four different series of the probability for a government to collapse are obtained. These represent the proxies for the degree of government instability (INSTAB). In the models of Table 4, INSTAB is used as an explanatory variable in the regression of DEF and DG. Notice that whilst the table reports the estimates when INSTAB is computed from the probit version of model (2), nothing really changes when the other series of INSTAB are used instead.

The results in Column 1 of Table 4 are not supportive of the theoretical argument of strategic use of debt in countries with high instability: the coefficient on INSTAB does not pass a zero

restriction test. However, it might be the case that instability really matters only when government changes involve significant alternation in office and/or volatility of portfolios. As a matter of fact, not necessarily a termination implies a significant change in the composition of the government or of the ruling coalition or even in the allocation of portfolios. For example, in Italy prior to 1993 the cabinets death rate is very high, but the core of the ruling coalition has remained substantially unchanged for long spells. A similar situation has occurred in Finland and, to some extent, in France. The basic theoretical argument incorporated in the models of strategic use of deficit and debt claims that large deficits will be run by the incumbent only when he faces a positive probability of being replaced by a *different* party (or coalition of parties). Thus, for a more appropriate test of this hypothesis, the variable INSTAB should be weighted by a measure of the effective degree of change in office. The past observed values of the indicators of portfolios volatility PPV and of alternation in office ALT are here used as weights. The interactive terms INSTAB*ALT and INSTAB*PPV do display the expected positive and significant coefficients (Column 2 and 3 of Table 4), thus providing supporting evidence to the theory of strategic accumulation of debt in the presence of high government volatility (the same is true when INSTAB is weighted by the other two indicators of volatility TPV and IPV). A similar result is obtained in the regression of DG, where the interactive term TER*ALT is used instead of INSTAB*ALT. Notice also that all these results concerning the role of instability are robust to the inclusion of the ideological dummies, in spite of the fact that a dummy DL significantly enters model specification (2). This means that the incentive to spend more and to sustain larger deficits in periods of high instability exists for any incumbent policymaker independently from its ideological location.

An issue that deserves specific consideration when estimating the impact of instability on fiscal policy is the one of possible reverse causation. That is, it could be the case that larger fiscal deficits (as indicators of bad economic performance) induce higher cabinet instability. To test this hypothesis, the binary choice model (2) has been re-estimated adding DEF to the set of regressors. Interestingly, little changes are observed on the size and the standard errors of the estimated coefficients on the other regressors, whilst the estimated coefficient on DEF, albeit positive, remains statistically not significant at usual confidence levels. The same results are obtained when DEF is one period lagged. Thus, there is evidence that the size of the deficit does not affect the turnover in office. As a further check, the interactive terms INSTAB*ALT, INSTAB*TPV, INSTAB*PPV, INSTAB*IPV have been instrumented by their lagged values. In fact, the estimated coefficient on these variables display now larger standard errors, but still remain significant at usual confidence levels. Again, no relevant changes appear for the other regressors.[9] Future research in this area could further investigate the issue of joint endogeneity of instability and deficit by estimating appropriate systems of simultaneous equations, as Alesina et al. (1992) do for the case of economic growth and political instability.

3.5 Budgetary institutions

The dummies DELEGATION, CONTRACT and UNCONSTRAINED are used to investigate the role of different budgetary institutions.[10] The estimates in Column 1 of Table 5 show that a significant difference in the size of fiscal deficit can be traced back to the existence, in a given country at a given time, of some form of constraint (either a mechanism of delegation to

[9] All these results are available from the author upon request.
[10] The coding of any of the three dummies is constant for any given country (unit of observation). This implies that a fixed effect estimator cannot be computed. However, both OLS and random effect can be estimated, with OLS that always appears to be favoured by statistical tests.

a strong Ministry of Finance or a commitment to some negotiated fiscal target). In the absence of this constraint, fiscal deficit is larger. However, which of the two forms of constraint is effectively at work seems to matter very little for the size of the deficit. In a regression of DT, both UNCONSTRAINED and CONTRACT are found to have positive and significant coefficients. This implies that the adoption of a procedure such as the centralisation of the process in the hands of a strong Ministry of Finance effectively promotes lower spending and, more interestingly, less recourse to distortionary taxation.

Proportional representation systems do not appear to sustain systematically larger deficits (Column 2, Table 5). However, in these systems, the increase in government consumption expenditure appears to be significantly larger than in other countries (Column 3, Table 5). These latter results are robust to the inclusion of several other political variables (such as the measure of fragmentation of the party system, the index of cabinet instability and the dummies for the ideological location of the median voter) and of the dummies that account for cross-country differences in budgetary institutions. Therefore, the hypothesis that electoral institutions matter only because they contribute to determine the type of budgetary institutions at work in a given country might not be the end of the story. As suggested by Persson and Tabellini (1998), the analysis of comparative politics, relating the size and composition of government spending to the political system in general and the electoral system in particular is a promising area for future research.

3.6 The ideological location of the median voter

Theoretical models often assign the median voter a pivotal role in the determination of fiscal policy choices (see, inter alia, Persson and Tabellini, 1997). It is therefore interesting to

investigate to what extent the ideological orientation of the electorate (as summarised by the ideological orientation of the median voter) contributes to explain fiscal policy formation in coalition systems.

To address this issue, a measure of the ideological location of the median voter must be computed from the data available on the preferences expressed by voters in political elections and the location on the ideological scales of the parties competing in these elections. The approach I use is a straightforward extension of the one suggested by Laver and Schofield (1990) to compute the location of the median legislator. In particular, I order political parties from left to right. Each party is identified by a specific number indicating its location on a ten points left-right scale. Then, I add the share of votes received by each party, counting from left to right. Eventually, there will be a party whose share added to those of the parties on its left makes the cumulative sum of share larger than the threshold 50.1%. The ideological location associated to this specific party can be taken as an approximation of the ideological location of the median voter. Then, dummies can be constructed exactly in the same way as they are constructed to represent the ideological location of the policymaker (see Section 2 above). Thus, IDMV is coded 1 when the median voter is located to the right of 5.5 on the 10 points ideological continuum, IDMVL takes value 1 if the median voter is located to the left of 4.6, IDMVC is coded 1 if the median voter is located between 4.6 and 6.4 and IDMVR is coded 1 if the median voter is located to the right of 6.4.

The last three columns of Table 5 report the results concerning the role of the median voter. Before commenting on these figures, it is worth spending a few words on the relationship between the ideological location of the median voter and the one of the policymaker. Most theoretical models assume that these two coincide. A comparison between the dummies based on a bi-partition of the ideological space (ID4 and IDMV) yields that in 94 out of 305

observations the two are coded differently. That is, in 94 cases, the estimated location of the median voter was on the left whilst the one of the policymaker was on the right or vice-versa. Figures do not significantly change when other definitions of the location of the policymaker (ID1, ID2, ID3) are used instead of ID4. When instead the dummies based on a tri-partition of the space are considered, the number of cases where coding is different drops to 89 out 305. A simple probit regression of ID on a constant and IDMV (to test to what extent the location of the median voter is correlated to the one of the policymaker) yields an estimated coefficient on IDMV equal to 1.268971, with a standard error of .166803 and a p-value of .000. The restriction that the coefficient on IDMV is equal to 1 can be rejected at the 10% level of confidence. Without the constant, the coefficient on IDMV is .375643 and the p-value is .001, the restriction that the coefficient on IDMV is equal to 1 can be rejected at all confidence levels. Similar results are obtained when the probit regression is estimated for the dummies based on the tri-partition of the policy space. Overall, this evidence is supportive of the idea that, albeit positively correlated, the location of the median voter and the location of the policymaker do not necessarily coincide. This should not be surprising: in coalition systems, the continuous need for political bargaining makes it more likely that the preferences of the electorate are only imperfectly incorporated into the orientation of the policymaker.

In Column 4 of Table 5 both the location of the median voter and the location of the policymaker are included on the r.h.s. of a DEF regression. The large standard errors of the estimated coefficients on the ideological dummies might be a symptom of multicollinearity. More precise estimates are obtained in Column 5 and 6. In the regression of DG (Column 5), only the dummies reflecting the ideological location of the policymaker displays significant coefficients. Thus, it seems that, at least with respect to spending decisions, the orientation of the policymaker prevails over the one expressed by the voters (when the two are conflicting). Again, this result is not surprising since in a representative democracy the policy is decided

by the government and not directly by the electorate. Notice also that the ideological location of the median voter does not significantly affect the change in government consumption expenditure even after that the ideological location of the policymaker is dropped (Column 6). All in all, the evidence proposed suggests that the ideological preferences of the median voter (and by extension of the electorate as a whole) could be imperfectly incorporated into those of the policymaker and when this happens only the latter ones count, at least for fiscal spending decisions.

Section 4. Conclusions

The empirical analysis undertaken in this paper on a sample of coalition systems has achieved some interesting results concerning the role of political and institutional variables in fiscal policy formation. The fiscal illusion argument receives little support whilst significant effects on fiscal policy decisions can be traced back to the ideological orientation of the policymaker, the degree of cabinet instability in the country, differences in electoral and budgetary institutions. The dispersion of political power also matters to the extent that it requires negotiation within the ruling coalition. The ideological orientation of the median voter does not necessarily coincide with the one of the policymaker and when the two are conflicting, the latter prevails, at least with respect to fiscal spending decisions.

I believe that further investigation on the role of electoral institutions represents a major direction for future research in this area. There are various theoretical reasons to believe that different electoral institutions will produce different forms of bias in fiscal policy formation. The results proposed in Section 3 does provide some preliminary support to this hypothesis. However, both theoretical and empirical work is needed in order to identify the channels

through which this effect operates. The question takes particular relevance for the countries of the European Union, where the existence of the *Growth and Stability Pact* already poses significant constraints on the possibility for national governments to use fiscal policy discretionary (or strategically) and where electoral reforms are often at the top of the political agenda.

Appendix.

Table 1

	1	2	3	4	5	6	7
	DEF	DEF	DEF	DEF	DG	DT	DEF
DEF_{-1}	.575904	.574924	.593553	.577268			.577116
	(.01501)	(.05017)	(.05076)	(.05246)			(.05261)
DU	1.52256	1.51938	1.47000	1.43200	.190635	-.00214	1.41801
	(.19413)	(.19428)	(.19474)	(.19814)	(.02428)	(.27808)	(.19945)
DY	-.33584	-.33798	-.35748	-.34990	-.09772	.053946	-.35493
	(.07592)	(.07601)	(.07593)	(.07684)	(.00807)	(.09499)	(.07724)
Dr	.050460	.049912	.052229	.051303			.051246
	(.01261)	(.01263)	(.01253)	(.01265)			(.01268)
ELE		.003385					
		(.00403)					
LOC			.000784				
			(.00124)				
ID					-.001051	-.017962	
					(.00054)	(.006578)	
DL				-.009514			-.00720
				(.005280)			(.00590)
DC				-.001799			-.00221
				(.004789)			(.00485)
COAL							.005335
							(.00615)
favoured estimator	OLS	OLS	OLS	OLS	FE	FE	

Note: standard error in brackets. Estimates of the constant term (OLS and RE) are not reported. The measures of ideological location in the Table are those defined as ID4, DL4 and DC4.

Table 2

	1	2	3	4	5
	DEF	DEF	DEF	DEF	DEF
DEF_{-1}	.575721	.531739	.567737	.571221	.527991
	(.50015)	(.05196)	(.04987)	(.04988)	(.05171)
DU	1.48341	1.60917	1.47639	1.45264	1.58992
	(.19420)	(.19610)	(.19244)	(.19294)	(.19411)
DY	-.33952	-.29627	-.331738	-.34191	-.29378
	(.07539)	(.07639)	(.07512)	(.07548)	(.07554)
Dr	.050600	.050546	.050605	.050178	.050683
	(.01253)	(.01292)	(.01243)	(.01243)	(.01278)
COAL	.009181				
	(.00492)				
MIN		-.01978	-.01709	-.01577	-.03599
		(.00739)	(.00581)	(.00642)	(.01102)
COAL*MIN			.015988	.016099	
			(.00693)	(.00692)	
MAJ*CI				.001129	
				(.00282)	
MIN*POL					.090324
					(.03935)
favoured estimator	OLS	FE	OLS	OLS	FE

Note: standard error in brackets. Estimates of the constant term (OLS and RE) are not reported. The measures of ideological location in the Table are those defined as ID4, DL4 and DC4.

Table 3

	1	2	3	4
	TERM	TERM	TERM	TERM
Constant	.540417	.878644	.552551	.706075
$TERM_{-1}$	-.282381	-.444807	-.16260	-.060021
MAJ	-.465663	-.768122	-.507800	-.189512
SING	-.599685	-.979283	-.840032	-.312042
CI	-.068678	-.110429		
POL	1.08258	1.79327		
DL	-.240313	-.383687		
GR	-.010712	-.017187		
GR_{-1}	-.005428	-.008515		
model	PROBIT	LOGIT	PROBIT	LOGIT

Note: the measures of ideological location in the Table is that one defined as DL4.

Table 4

	1	2	3	4
	DEF	DEF	DEF	DG
DEF_{-1}	.574463	.560630	.548328	
	(.05019)	(.05016)	(.05158)	
DU	1.51488	1.50471	1.53621	.168422
	(.19447)	(.19271)	(.19709)	(.02448)
DY	-.33807	-.342839	-.32173	-.09873
	(.07601)	(.07536)	(.07654)	(.00822)
Dr	.051144	.050373	.052794	
	(.01264)	(.01251)	(.01297)	
INSTAB	.003981			
	(.00481)			
INSTAB*ALT		.028881		
		(.01247)		
INSTAB*PPV			.000826	
			(.00039)	
TER*ALT				.002037
				(.00115)
favoured estimator	OLS	OLS	OLS	RE

Note: standard error in brackets. Estimates of the constant term (OLS and RE) are not reported. The measure of instability (INSTAB) reported in the table is the one btained from Column 1 of Table 3.

Table 5

	1 DEF	2 DEF	3 DG	4 DEF	5 DG	6 DG
DEF_{-1}	.618123	.574031		.571893		
	(.05349)	(.05044)		(.05274)		
DU	1.46590	1.52706	.180424	1.39286	.201832	.197991
	(.18969)	(.19478)	(.02429)	(.20130)	(.02475)	(.02435)
DY	-.33246	-.33539	-.10074	-.34201	-.09894	-.09911
	(.07921)	(.07605)	(.00820)	(.77228)	(.08078)	(.00795)
Dr	.054783	.050199		.052715		
	(.01391)	(.01265)		(.01272)		
Delegation	.004107					
	(.00439)					
Contract	.002221					
	(.00269)					
Unconstrained	.006385					
	(.00336)					
PR		-.00143	.001535			
		(.00365)	(.00046)			
DR					-.001349	
					(.000741)	
DL				-.00759		
				(.00561)		
DC				.000069	-.001165	
				(.00506)	(.000582)	
IDMVL				-.011460		
				(.10266)		
IDMVR					-.001824	-.00092
					(.00130)	(.00145)
IDMVC				-.00849	.000923	.001295
				(.00767)	(.00081)	(.00103)
favoured estimator	OLS	OLS	OLS	OLS	OLS	OLS

Note: standard error in brackets. Estimates of the constant term (OLS and RE) are not reported. The measures of ideological location in the Table are those defined as ID4, DL4 and DC4.

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