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Cyclical Expenditure Policy, Output Volatility, and Economic Growth

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Abstract: This paper provides an empirical assessment of the relation between the cyclical policy of fiscal expenditure policy, output volatility, and economic growth, using a cross-section of 88 countries over the period 1960 to 2004. Identification of the effects of (endogenous) cyclical expenditure policy is achieved by exploiting the exogeneity of countries' political and institutional characteristics, which we find to be relevant determinants of the cyclical policy of expenditures. There are three main results: First, both pro- and countercyclical expenditure policy amplify output volatility, much in a way like pure fiscal shocks. Second, output volatility, due to variations in cyclical and discretionary fiscal policy, is negatively associated with economic growth. Third, there is no direct effect of cyclical policy on economic growth other than through output volatility.

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I. Introduction

Does fiscal policy affect economic growth? This is clearly one of the most fundamental and policy relevant macroeconomic questions. According to Easterly (2005) there is no robust evidence for a relation between macroeconomic policies (including fiscal policy) and economic growth, once institutions are controlled for. In contrast, Caballero (2008, p. 1) argues that “Good macroeconomic policy helps growth ... I do not think this view is in any dispute in the applied and policy world.”

Notwithstanding the wide agreement that macroeconomic policies can influence economic performance, it remains a challenge for both theory and empirics to identify the channels through which economic policy affects growth. The emergence of new endogenous growth theory, overcoming the traditional dichotomy between business cycle and growth theory, has laid the ground for such an analysis. It is hardly questioned that economic policy affects economic activity in the short run. If business cycle volatility and economic growth are related as suggested by endogenous growth theory, economic policy can indirectly affect growth through its effect on volatility. Such a finding would also lead to a reassessment of macroeconomic priorities: The welfare cost of volatility *per se* are widely regarded as negligible since Lucas (1987). But if volatility turns out to have a negative effect on economic growth, its costs – or equivalently, the gains from stabilization – will be substantial (Barlevy, 2004).

Regarding the role of fiscal policy, Fatas and Mihov (2003) suggest introducing fiscal rules as a means to reduce the use of discretionary fiscal policy, defined as fiscal policy unrelated to the business cycle, based on their finding for a large cross-section of 91 countries that aggressive use of discretionary policy lowers growth by increasing output volatility.

This paper highlights the role of another important element of fiscal policy, namely cyclical fiscal policy. So far, there are hardly studies investigating the effects of fiscal cyclicity on economic growth. One notable exception is Aghion and Marinescu (2007), who consider an (unbalanced) panel of annual data for 19 OECD countries from 1960 to 2007. Regressing growth on alternative cyclicity measures (and standard controls for economic growth regressions), they find a positive effect of the ‘countercyclicity’ of fiscal policy on economic growth.

The main goal of the present paper is to shed more light on the role of cyclical fiscal policy and its transmission channels, considering both its effect on output volatility, and – in a second step – its effect on economic growth. Other than previous studies we consider a large cross-section of 88 countries covering the period 1960 to 2004, which is motivated by the use of (de facto) time-invariant variables on the countries’ political and institutional characteristics to identify the causal effect of (endogenous) cyclicity on output volatility. The use of institutional variables as instruments for fiscal policy was first suggested by Fatas

and Mihov (2003). The present study extends their analysis in considering the role of cyclical fiscal policy (as well as that of discretionary fiscal policy). We demonstrate that institutional variables (such as political of constraints and the average number of elections) provide considerable information on the variation in fiscal cyclicalities across countries, and we use this exogenous variation to identify the causal effect of cyclicalities on output volatility and economic growth.

We find that cyclical fiscal (expenditure) policy has a destabilizing effect on the economy, no matter whether it is pro- or countercyclical. In fact, it amplifies output volatility much the same way as discretionary fiscal policy. This adds to the widespread scepticism against the usefulness of fiscal policy as a fine-tuning instrument. We also find that output volatility, induced by variations in cyclical or fiscal policy, negatively affects economic growth. Taken together this has an important policy implication: Economic growth could be enhanced by introducing fiscal rules, designed to restrict both the use of discretionary fiscal policy (Fatas and Mihov, 2003) as well as the use of cyclical fiscal policy.

The remainder of the paper is organized as follows. Section II constructs measures of fiscal cyclicalities and the aggressiveness of discretionary fiscal policy for a large cross-section of 88 countries. Section III motivates the identification strategy and provides evidence on the relation between cyclical (and discretionary) fiscal policy and output volatility. Section IV considers the effect of cyclicalities on economic growth. Section V summarizes the results and concludes.

II. Constructing Measures of Cyclical and Discretionary Fiscal Policy

We use government consumption as indicator of fiscal policy. This choice is dictated by data availability, since there are no internationally comparable data for other measures of fiscal policy for our large cross-section of countries. On the one hand, this limits the generality of our results. On the other hand, an advantage is that government expenditures – compared with revenues – are less responsive to the cycle through stabilizers ‘built-in’ the fiscal system and can be changed with relative ease. As a consequence, expenditures are more indicative of a government’s intentional cyclical policy than revenues, whose cyclical behaviour is driven by automatic stabilizers to a much larger extent. Moreover, previous studies suggest that the cyclicalities of government consumption reflects the cyclicalities of overall government expenditures reasonably well.¹

¹ In Lane (2003), for example, who studies the determinants of cyclical fiscal policy using a sample of 22 OECD countries, the correlation between the cyclicalities of government consumption and that of total government expenditures is 0.71.

We follow the standard approach in the literature and estimate cyclical parameters (χ) by regressing growth of real government consumption (G) on the growth of real GDP (Y), correcting for serial correlation in the error term:

$$\Delta \ln G_{i,t} = \alpha_i + \chi_i \Delta \ln Y_{i,t} + \eta_{i,t}, \quad (1a)$$

$$\eta_{i,t} = \rho_i \eta_{i,t-1} + \varepsilon_{i,t}. \quad (1b)$$

Equation (1) is estimated separately for each of the $i = 1, \dots, 88$ countries, which is the largest set of countries for which the key variables required in the present study are available. The time dimension t ranges from 1960 to 2004; for some countries, a slightly shorter time span had to be used (see Appendix A1).

Equation (1) is considered as reduced form equation for government consumption and estimated by ordinary least squares. As a result we obtain a decomposition of the growth of government consumption into a cyclical and a discretionary component. The time series of country i 's cyclical fiscal policy is given by $\hat{\chi}_i \Delta \ln Y_{i,t}$; and the estimate of the (structural) residual of equation (1), i.e., $\hat{\varepsilon}_{i,t}$, is interpreted as series of discretionary fiscal policy shocks. By least squares algebra these two series are orthogonal.

In the following, we will estimate the effects of cyclical fiscal policy on volatility and growth from a cross-section regression, where we will use the estimates of the parameter χ_i as country-specific indicators of the average cyclical policy. A positive (negative) value of χ_i is associated with procyclical (countercyclical) fiscal behaviour. Our results indicate substantial cross-country variation in the cyclical parameters, whose estimates range from -0.835 to 2.698 . Most of the countries show procyclical fiscal expenditure policy; only 11 of the 88 coefficients are negative (see Appendix A1). In line with Fatas and Mihov (2003), we use as a measure of discretionary fiscal policy ($DISCR$) the standard deviation of the residuals from equation (1), i.e., $DISCR_i = \text{sd}(\hat{\varepsilon}_i)$.

Regarding the relevance of cyclical versus discretionary fiscal policy, the R^2 in equation (1), averaged over all 88 countries, amounts to 0.201. This means that roughly one-fifth of the total variation in fiscal policy is due to cyclical fiscal policy. Our estimates are well in line with previous studies. Comparing our cyclical coefficients $\hat{\chi}_i$ (for period 1960 to 2004) with those of Lane (2003) for 22 OECD countries (for 1960 to 1998), the correlation is 0.872. Our measure of discretionary fiscal policy is very close to that of Fatas and Mihov (for 1960 to 2000); their correlation is 0.887.

III. Cyclical Policy and Output Volatility

1. The Empirical Model

Our basic empirical framework builds on Fatas and Mihov (2003); as a novel feature the cyclical policy is included as explanatory variable for output volatility:

$$\ln \sigma_i^y = \gamma_0 + \gamma_1 \ln CYC_i + \mathbf{x}_i \boldsymbol{\gamma} + u_i. \quad (2a)$$

The dependent variable is output volatility (σ^y), defined as standard deviation of the growth rate of (real) output per capita; CYC is our measure of the cyclical policy, which we construct from the estimates of equation (1) as will be outlined more in detail below; \mathbf{x}_i is a vector of control variables, and u is a stochastic error term. The cross-section dimension (i) comprises 88 countries, the largest sample for which the required data are available. Unless mentioned otherwise, all data are averages over the period 1960 to 2004 (see Appendix A1).

As it is standard in skedastic regressions, we choose a logarithmic specification to avoid negative predicted values; it is then natural to use the cyclical policy measure in log form as well, such that the parameter of our main interest (γ_1) measures the relative change of output volatility with respect to relative changes in cyclical policy.²

We define cyclical policy (CYC) as absolute value of $\hat{\chi}$ to allow for negative values of the cyclical policy coefficients $\hat{\chi}_i$ in the logarithmic specification (2a). Obviously, the variable $CYC = |\hat{\chi}|$ then measures only the cyclical responsiveness of fiscal policy, but not whether it is pro- or countercyclical. This could be addressed by properly signing $\ln CYC$ for the respective observations. But this would assume that – if procyclical policy amplifies business cycles – countercyclical policy smoothes business cycles. This is an assumption we wish to test rather than impose right from the beginning, given the widespread scepticism against the effectiveness of fiscal policy as fine-tuning instrument. Countercyclical fiscal policy might actually turn out destabilizing due to lags in (recognition, implementation, and) materialization, a point prominently made by Friedman (1953).

Consequently, we do not impose any assumption about the relation between the effects of pro- and countercyclical policy. Instead, we define CYC_i as absolute value of $\hat{\chi}_i$ ($CYC_i = |\hat{\chi}_i|$) and allow for different parameters of $\ln CYC_i$, depending on whether $\hat{\chi}_i$ is positive or negative for the respective observation i :

$$\ln \sigma_i^y = \gamma_0 + \gamma_1 \ln CYC_i + \gamma_1^{counter} D_i^{counter} \ln CYC_i + \mathbf{x}_i \boldsymbol{\gamma} + u_i, \quad (2b)$$

where $D^{counter}$ is a dummy variable taking a value of 1 for ‘countercyclical observations’, i.e., $D_i^{counter} = 1$ for all i where $\hat{\chi}_i < 0$ and 0 otherwise. The parameter $\gamma_1^{counter}$ then measures the

² While the logarithmic specification yields a slightly better fit, it is not crucial for the results.

difference between the effect of countercyclical fiscal policy and the effect of procyclical fiscal policy on output volatility (γ_1).

We start from a simple regression of output volatility on cyclical policy (CYC) and then add three standard controls: Government size ($GSIZE$), openness ($OPEN$), measured as imports plus exports as a share of GDP, and real GDP per capita ($GDPPC$), i.e., $\mathbf{x}_i = [GSIZE_i, OPEN_i, \ln GDPPC_i]$. Regarding government size, it has been argued that more volatile economies may have an incentive to set up larger governments as a means to reduce macroeconomic volatility (Rodrik, 1998). As a consequence, $GSIZE$ might be endogenous in equation (2). In line with Fatas and Mihov (2003) we use the standard approach and instrument $GSIZE$ by the (log of) population (POP), the urbanization rate ($URBAN$), and the dependency ratio (DEP).

In a final step, we will include the aggressiveness of discretionary fiscal policy as defined in section II ($DISCR$), yielding our most comprehensive model:

$$\ln \sigma_i^y = \gamma_0 + \gamma_1 \ln CYC_i + \gamma_2 \ln DISCR_i + \mathbf{x}_i \gamma + u_i. \quad (3)$$

Before turning to the estimation of models (2) and (3), two issues warrant discussion. First, our variable of main interest, the cyclical policy (CYC), is endogenous with respect to output volatility as a result of reverse causality. Talvi and Vegh (2005) show in a political economy model that high output volatility tends to generate procyclical fiscal behaviour.³ This would introduce an upward bias in the estimated effect of cyclical policy on output volatility. In addition, the cyclical policy measure (CYC) is an estimate of its true value and might thus be subject to classical measurement error, causing an attenuation bias.

Another issue, related to the fact that our country-specific cyclical policy measures (χ_i) are generated by model (1), is that the observations on our variable CYC_i are estimated with different precision. This is addressed by using a weighted (two stages) least squares procedure, using the inverse of the variance of $\hat{\chi}_i$ as weights.⁴ This implies that observations, for which the variable CYC is measured more precisely, are assigned a higher weight in the regression. As we show below the weighting improves the fit but it is not crucial for the results.

2. Identification, First Stage Regressions, and Instrument Quality: Political and Institutional Characteristics as Determinants of Cyclical Policy

³ They also provide evidence from a large cross-section of countries that the degree of procyclicality in government consumption is positively correlated with output volatility. Lane (2003) obtains a similar results for a sample of 22 OECD countries.

⁴ The choice of the weight is not affected by the logarithmic transformation. By the delta method, $\sigma_{\ln CYC}^2 = 1/\mu_{CYC} \sigma_{CYC}^2$, i.e. the variances of the level and log of CYC are equal up to a rescaling by the mean of CYC (μ_{CYC}).

In order to identify the causal effect of cyclical volatility on output volatility, we use countries' political and institutional characteristics as instruments. We hypothesize that countries' political and institutional characteristics are not only relevant determinants of discretionary fiscal policy (as shown by Fatas and Mihov (2003)) but also of cyclical fiscal policy. In particular, we consider four institutional variables: the average number of elections (*NELEC*), an index of political constraints (*POLCON*) by Henisz (2000), a dummy for majoritarian systems (*MAJ*), and a dummy for presidential regimes (*PRES*).

To motivate the choice of instruments, notice that the variable *CYC*, defined as absolute value of cyclical ($CYC = |\hat{x}|$), in the first place measures the aggressiveness (but not the direction) of cyclical fiscal policy. As a consequence, part of the discussion by Fatas and Mihov (2003) motivating the use of the institutional variables (*NELEC*, *POLCON*, *MAJ*, *PRES*) as instruments for the aggressiveness of discretionary policy directly carries over to our measure of cyclical fiscal policy (*CYC*).

The extent of political constraints (*POLCON*) is the instrument with the strongest theoretical motivation. According to the 'voracity hypothesis' (Tornell and Lane, 1998), power diffusion among more agents induces procyclicality, since fiscal competition by multiple power groups for fiscal revenues increases (decreases) in booms (recessions). On the other hand, governments less constrained in implementing their policy can respond more flexibly to the business cycle and will thus be better able to translate their 'cyclical preferences' into actual policy.

Regarding electoral characteristics, the frequency and timing of elections (*NELEC*) and the induced electoral cycles will not be systematically related to the business cycle. As a consequence, the observed pattern of fiscal policy will show a smaller association with the business cycle, the larger the number of elections, i.e., the more the responsiveness of fiscal policy to the business cycle is diluted by fiscal measures related to the electoral cycle. A similar point can be made for *MAJ* in light of the argument by Persson and Tabellini (2001) that majoritarian systems will have more pronounced electoral cycles. Finally, it could be the case that presidential regimes will not only be associated with a more aggressive use of discretionary fiscal policy as argued by Fatas and Mihov (2003), but also with a more active conduct of cyclical fiscal policy.

While our choice of the institutional variables *NELEC*, *POLCON*, *MAJ*, and *PRES* is well motivated, the ultimate question is whether they are also relevant in our empirical model. Table 1 reports the results of a regression of the log of *CYC* on the four institutional variables separately (columns (1a) to (1d)) and simultaneously (column (2a)). The number of elections (*NELEC*) and political constraints (*POLCON*) turn out to have the strongest effect; they are significant both in a simple regression (columns (1a) and (1b)) of Table 1 and in a multiple regression on all four political variables (column (2a)). The sign of the coefficient of *NELEC*

is negative as expected. The variable *POLCON* also enters with a negative sign; this does not necessarily reject the voracity hypothesis but suggests that – among the various ways through which political constraints affect the cyclicity of fiscal (expenditure) policy – the voracity effect does not appear to be the most dominant force.⁵ The variables *MAJ* and *PRES* are insignificant or only weakly significantly and would thus weaken the quality of our set of instruments (compare the F-statistic in columns (2a) and (2b)). Hence, we will use only *NELEC* and *POLCON* as instruments for *CYC* in the two stages least squares regressions below.

< Table 1 >

Notice that column (2b) corresponds to the first stage regression for equation (2a) only in the most parsimonious specification without additional explanatory variables for output volatility. The first stage regression for the most comprehensive model including all controls – *GSIZE* (instrumented by *POP*, *URBAN*, and *DEP*), *OPEN*, and *GDPPC* – is given in column (3). An important result is that the variables *NELEC* and *POLCON* remain significant, both individually and jointly.

Column (4) shows the corresponding first stage regression for discretionary fiscal policy (*DISCR*). Results are in line with Fatas and Mihov (2003). Notice that – in contrast to the first stage regression for *CYC* – the two variables *MAJ* and *PRES* turn out significant at the five and one percent level. At least from an empirical perspective, this suggest that the variation in the variables *MAJ* and *PRES* may be helpful to identify the separate effect of discretionary policy in model (3), where both *CYC* and *DISCR* are included.

Overall, the results reveal interesting links between institutions and cyclicity. Exploring these links more in detail is beyond the scope of this paper. For the purpose of the present study, the most relevant message from the results in Table 1 is that the two variables *NELEC* and *POLCON* are relevant instruments for cyclicity (*CYC*); this is not the case for the variables *MAJ* and *PRES*, which are, however, strongly associated with the aggressiveness of discretionary fiscal policy (*DISCR*).

3. Estimation Results

We start from the most parsimonious specification of model (2), which includes only cyclicity (*CYC*) as explanatory variable.⁶ Columns (1a) and (1b) show the weighted least

⁵ Lane (2003) also finds little support for the voracity hypothesis in his study of the cyclicity of expenditures in a sample of 22 OECD countries; in particular, the effect of political constraints on cyclicity is often insignificant or shows the wrong sign.

⁶ The weighting accounts for the fact that *CYC* is a generated regressor, not for the presence of heteroskedasticity in the error term in models (2) or (3) (which is also confirmed by standard tests). Hence, we use robust standard errors for inference throughout.

squares estimates (WLS) of equation (2a), which allows the effects of pro- and countercyclical fiscal policy to differ by including an interaction between *CYC* and a dummy for countercyclical policy ($D^{counter}$). The estimated elasticity of output volatility with respect to procyclical fiscal policy is 0.184; the effect of countercyclical policy appears to be even larger (0.278), but the difference is insignificant with a p-value of 0.300. This conclusion holds up when the model is estimated by weighted two stages least squares (WTLS), using the average number of elections (*NELEC*) and the index of political constraints (*POLCON*) as instruments for *CYC* (column (1b)). In that case the elasticities with respect to pro- and countercyclical fiscal policy are 0.595 and 0.501 respectively, but again the difference is insignificant (p-value: 0.618).

< Table 2 >

In light of this result we proceed with a restricted model, imposing parameter equality for pro- and countercyclical fiscal policy. Columns (2a) and (2b) show the WLS and WTLS estimates of the simple regression of volatility on cyclical policy. For comparison, columns (3a) and (3b) show the results of the unweighted LS and TSLS estimates. Notice first, that endogeneity of *CYC* is indeed pronounced: The difference between the (W)LS and (W)TSLS coefficients is sizeable. More formally, a Hausman test rejects that *CYC* is exogenous at the one percent level in all specifications. It is interesting to note that both the weighted and unweighted LS estimates of the effect of *CYC* on volatility show a strong attenuation bias towards zero. This suggests that measurement error is the dominant source of endogeneity rather than reverse causality (causing an upward bias). This view is also supported by a comparison of the weighted and unweighted estimates. In the weighted regressions, less precise estimates are assigned a lower weight, rendering the role of measurement error less relevant. As a consequence, the attenuation bias is less pronounced in the WLS regression, yielding coefficients that are closer to the WTLS estimates.

While we postpone a more comprehensive sensitivity analysis to below, we emphasize that the weighting is not crucial for the results: A comparison of columns (2b) and (3b) shows that the weighted and unweighted TSLS estimates are virtually identical, pointing to an elasticity of output volatility with respect to cyclical policy of around 0.6. The choice of the logarithmic form of *CYC* is not essential for the qualitative conclusions as well: The corresponding results for the specification in levels (columns (4a) and (4b)) are in line with the logarithmic specification (columns (1b) and (2b)).

A final observation is that the tests for overidentifying restrictions reject the null hypothesis of valid instruments in some specifications. This is not too surprising, given that several important variables have been omitted from the regression so far. Results for a more comprehensive model, including *Gsize*, *OPEN* and *GDPPC* as controls are given in Table 3. As already discussed above, government size is likely to be endogenous with respect to

volatility, which is addressed by using population (*POP*), the urbanization rate (*URBAN*), and the dependency ratio (*DEP*) as instruments for *G**SIZE*.

< Table 3 >

Columns (1a) and (1b) show the unweighted LS and TSLS estimates, whereas column (1c) gives the results of the WTSLS estimation. As expected the estimated elasticity of volatility with respect to cyclical policy becomes smaller in magnitude when the control variables are added (around 0.3) but remains significant. Exogeneity of *CYC* is still clearly rejected in all models; we thus focus on the (W)TSLS results. The OID tests are insignificant in most specifications, suggesting that the institutional variables *NELEC* and *POLCON* (as well as *POP*, *DEP*, *URBAN*) are valid instruments. To reinforce the finding of our parsimonious specification, we repeat the test for parameter equality between pro- and countercyclical policy (see columns (2a) and (2b)). The conclusion is the same as before: There is no evidence for a stabilizing effect of countercyclical fiscal policy. In contrast, it adds to output volatility, in a way not significantly different from that of procyclical fiscal policy.

Of course, the results regarding the role of countercyclical policy should be interpreted with care. The number of countries which pursued countercyclical fiscal policy on average is rather small (11 out of the 88). While this might be too little variation to yield a significant difference in the estimated effect, it does not explain that the effect of countercyclical policy on volatility is always positive, a finding that is extremely robust. We also emphasize that our results should be interpreted as averages over countries and time. One cannot rule out that a highly effective government, which is aware of the relevant lag structures and able to respond very quickly, might be successful in its fiscal efforts to smooth business cycles. What our evidence suggest, however, is that such a constellation is rather the exception than the rule.

Finally, we consider the results for model (3), which includes both cyclical policy (*CYC*) and discretionary fiscal policy (*DISCR*). It is a subtle question, whether discretionary fiscal policy should be controlled for. On the one hand, discretionary fiscal policy is certainly a relevant determinant of output volatility (Fatas and Mihov (2003), Badinger (2008)). Moreover, while the time series of discretionary and cyclical fiscal policy measures for a single country are orthogonal, this does not carry over the cross-country variation in cyclical and discretionary policy (averaged over time): Countries that are more responsive to the cycle might also more actively engage in discretionary fiscal policy.

Under these two assumptions – *DISCR* matters for volatility and is related to *CYC* – the estimated effect of *CYC* in model (2) will be upward biased due to the omission of *DISCR*. In that case, however, we would also expect the OID test to reject instrument validity, since the instruments used for *CYC* (i.e., *POLCON*, *NELEC*) are also related to discretionary policy (see section III, subsection 2). But this is not the case in any of the specifications, suggesting that these two elements of fiscal policy could be (close to) orthogonal in the cross-section.

One could still argue that the OID test has small power and *CYC* and *DISCR* should be regarded as related for theoretical reasons. Even in that case the question remains, whether discretionary policy should be controlled for: A main reason for a possible association between *DISCR* and *CYC* is that an active conduct of cyclical fiscal policy might partly result in (unintentional) discretionary policy (unrelated to the cycle) as a result of lags in implementation and materialization (again, on average over countries and time). Since these unintentional consequences of cyclical fiscal policy can hardly be ruled out by policy makers in practice, it might be reasonable to let the parameter of cyclical policy in model (2) also capture its indirect effects on volatility through its relation to discretionary fiscal policy.

Notwithstanding these arguments that might favour model (2) over model (3), we now turn to the results when *DISCR* is included (see column (3a)). As expected the coefficient of *CYC* becomes smaller, pointing to an elasticity of volatility with respect to cyclical policy of around 0.163, but it remains significant at the 10 percent level. The elasticity with respect to discretionary fiscal policy is 0.454 but insignificant with a p-value of 0.103. This is not too surprising; since both variables are instrumented using mainly the same set of institutional variables (only *MAJ* and *PRES* are added as additional instruments for *DISCR*), the predicted values for *CYC* and *DISCR* from the first stage regressions will be strongly correlated, causing a multicollinearity problem in the second stage regression. This is aggravated by the presence of a third endogenous variable (*Gsize*).⁷

To address this weak instruments problem, we reestimate the model by limited information maximum likelihood (LIML), which is superior to TSLS estimation in the presence of weak instruments (Stock and Yogo, 2004). In that case *CYC* and *DISCR* turn both out significant with elasticities of 0.182 and 0.417 respectively (see column (3b)). Another route would be to use a compound measure of discretionary and cyclical fiscal policy ($\ln CYC + \ln DISCR$), which could be justified in light of the fact that the hypothesis of parameter equality cannot be rejected (F-statistic: 0.704, p-value: 0.403). In the restricted model, the compound measure of fiscal policy turns out highly significant with a coefficient of 0.238 (column (4)). The economic interpretation of this restricted model with equal parameters for cyclical and discretionary fiscal policy carries our finding regarding the irrelevance of the direction of cyclical policy one step further: Not only has countercyclical policy the same effect on volatility as procyclical fiscal policy. It also implies that cyclical fiscal policy (*CYC*) has the same amplifying effect on output volatility as ‘random’ discretionary fiscal policy shocks, suggesting that the effects of intentional cyclical policy measures – due

⁷ If government size is treated exogenous and population is included as instrument, *CYC* and *DISCR* turn out significant at the five and one percent level respectively. However, since there are strong theoretical arguments to regard government size as endogenous, and since the theoretical motivation for using country size (population) as instrument for *CYC* and *DISCR* is weak (despite the fact that is highly significant in the first stage regression), we pursue the more conservative approach here and treat *Gsize* as endogenous.

to poor timing and lags in implementation and materialization – spread over time in a way such that the implied average outcome is random with respect to the cycle.⁸ A sensitivity analysis of the results is given in Appendix A3.

IV. Fiscal Cyclicity, Volatility, and Economic Growth

From a theoretical perspective, the relation between output volatility and economic growth is ambiguous (see Imbs, 2007). Empirically, Ramey and Ramey (1995) found a negative effect of output volatility on economic growth, and – though there is no consensus so far – the evidence that has emerged since then tends to support this finding. While a number of studies have considered the effect of output volatility on growth, there is hardly evidence on the relation between the cyclicity of fiscal policy and growth. One notable exception is Aghion and Marinescu (2007), who find a positive effect of countercyclical measures in a growth regression, using an (unbalanced) panel of annual data for 19 OECD countries from 1960 to 2007. Moreover, no previous study has considered the relations between cyclicity, volatility and growth in a joint empirical framework.

We first consider the effects of cyclicity on growth directly, running a cross-section regression of average growth of real GDP per worker over $(\Delta \ln GDPPW)$ on the cyclicity of fiscal policy (CYC), again testing for differences in the effect of pro- and countercyclical policy:

$$\overline{\Delta \ln GDPPW}_i = \delta_0 + \delta_1 \ln CYC_i + (\delta_1^{counter} D_i^{counter} \ln CYC_i) + \mathbf{w}_i \boldsymbol{\delta} + \mathcal{E}_i \quad (4)$$

The following standard controls (\mathbf{w}_i) are included in our cross-country growth regression: the (log of the) initial level of real GDP per worker ($GDPPW^{in}$), the average level of human capital in terms of educational attainment, i.e., the fraction of males above 25 with primary schooling (HC^{prim}) and secondary schooling (HC^{sec}). Model (4) refers to the time period from 1960 to 2004 again; the cross-section dimension comprises 80 rather than 88 countries due to missing human capital data.

As in section III, the endogeneity of CYC is addressed by using the institutional variables $NELEC$ and $POLCON$ as instruments; and the fact that CYC is calculated from fitted values of model (1) is accounted for by a weighted least squares approach, using the inverse of the variance of CYC as weight.

⁸This argument was already made by Friedman in his informal essay on fiscal policy: “In fiscal as in monetary policy, all political considerations aside, we simply do not know enough to be able to use deliberate changes in taxation and or expenditures as a sensitive stabilization mechanism. In the process of trying to do so, we almost surely make matters worse... by introducing a largely random disturbance that is simply added to other disturbances.” (Friedman, 1962, p. 78).

< Table 4 >

Columns (1a) and (1b) in Table 4 report the unweighted LS and TSLS estimates, allowing the effect of pro- and countercyclicality to differ. Columns (2a) and (2b) report the respective weighted estimates. In all specifications we find a significantly negative effect of cyclicalities on economic growth. As in section II, only the magnitude of cyclicalities seems to matter: We find a negative effect of both pro- and countercyclicality on economic growth, and while the coefficient of countercyclical fiscal policy is smaller in magnitude, the difference in the coefficients is not significantly different from zero. This holds true for both the unweighted and the weighted estimates. According to the Hausman test there is no strong evidence for endogeneity of *CYC*, though the (W)LS estimates of the parameter of *CYC* are always smaller in magnitude than the (W)TSLS estimates.

The results in Table 4 suggest a negative relation between the cyclicalities of fiscal policy and economic growth. In the corresponding model (4), which omits output volatility (and further controls), the parameter of *CYC* (δ_1) captures all effects of cyclicalities on economic growth, both through its relation with output volatility (direct and indirect through *DISCR*), through its relation with other variables affecting growth, as well as ‘direct’ effects of cyclicalities on economic growth (if any). From an economic policy perspective, this might in fact be the most relevant question.

Nevertheless, we would like to provide a more detailed picture of the relation between cyclicalities and growth. In particular, we also wish to answer the question whether *CYC* affects growth only through output volatility or also directly.⁹ The empirical framework employed to address this question is sketched by Figure 1, which illustrates the interrelationships between the key variables in our empirical models. Potential endogeneity is indicated by reversed arrows (though the source of endogeneity is not necessarily simultaneity); relations between variables of the same equation (such as *CYC* and *DISCR*) are omitted for simplicity here.

< Figure 1 >

To test for a direct effect of cyclicalities on growth, we first consider a model relating growth to output volatility (and controls). We then add cyclicalities as additional regressor (again considering potentially different effects of pro- and countercyclical policy):

$$\overline{\Delta \ln GDPPW}_i = \varphi_0 + \varphi_1 \ln \sigma^y + \mathbf{w}_i \boldsymbol{\varphi} + (\delta_1 \ln CYC_i + \delta_1^{counter} D_i^{counter} \ln CYC_i) + \zeta_i \quad (5)$$

⁹ Such a direct link could be motivated through the model by Aghion et al. (2005). They argue that credit constrained firms have a borrowing capacity that depends on current earnings, which are reduced in recessions. Hence, countercyclicality may foster productivity growth by reducing the magnitude of the output loss induced by market failures (as credit market imperfections) in a recession, and one could argue that such an effect should also hold for a given output volatility.

Column (1a) in Table 5 reports the least squares estimates of equation (5), column (1b) the TSLS estimates using the full set of institutional variables (*NELEC*, *POLCON*, *MAJ*, *PRES*) as instruments for volatility.¹⁰ Results of the TSLS estimates point to a negative effect of volatility on growth, consistent with our finding that cyclical increases volatility (models (2) and (3)) and reduces growth (model (4)). The coefficient of output volatility (-2.968) suggests that an increase in volatility by one percent reduces average growth by some 0.3 percentage points. This is close to Fatas and Mihov (2003), who obtain a coefficient of -3.371 in a similar regression. We emphasize that results are very similar when *MAJ* and *PRES* are excluded from the set of instruments, or when *CYC* or *DISCR* are used as instruments directly. This is supportive of the finding in section III that the transmission mechanisms from political institutions to output volatility through cyclical and discretionary fiscal policy are very similar.

< Table 5 >

Although the OID tests, reported at the bottom of Table 5, suggest that there is nothing wrong with our instruments, we use, as a robustness check, an alternative instrument for output volatility (Z^σ), which is entirely unrelated to the respective country's characteristics. In particular, we use for each country i , the trade share weighted output volatility of all other countries $j (\neq i)$ in the sample. To ensure exogeneity two modifications are made: actual trade shares are replaced by predicted values from a bilateral gravity model including geographical variables only; actual output volatility is replaced by predicted values from a regression of output volatility on (a constant and) the four institutional variables (*NELEC*, *POLCON*, *MAJ*, *PRES*) (see Badinger (2010)). The variable Z^σ also turns out to be a relevant instrument for output volatility besides the four institutional variables (*NELEC*, *POLCON*, *MAJ*, *PRES*). In the first stage regression for model (5), including all four institutional variables, Z^σ is significant at the one percent level. In fact, Z^σ turns out to be the strongest instrument for output volatility besides political constraints (*POLCON*). This is an important result, since it implies that Z^σ adds variation to identify the effect of output volatility in the growth regression (5) with *CYC* included as regressor, in addition to the identifying variation, which comes from the effects of institutions on volatility through on *DISCR* (which might be too closely associated with the effects of *CYC* on volatility) (see Figure 1).

Column (1c) shows the TSLS estimates of model (4), using only Z^σ as instrument; the estimated effect of volatility is even larger in magnitude, though the estimates are less precise. Nevertheless, output volatility remains significant with a p-value of 0.087. The fit of the model improves, when the four institutional variables are included as additional instruments (see column (1d)). Overall, the estimates of model (5) in columns (1b) to (1d), which differ

¹⁰ Results for specifications without *CYC* are unweighted estimates.

only by the set of instruments used, are very similar. This is a reassuring result. In the following, we will use as instruments for volatility all four institutional variables (*NELEC*, *POLCON*, *MAJ*, *PRES*) as well as exogenous volatility spillovers (Z^σ) to provide additional identifying variation.

We next consider the robustness of the results with respect to controlling for institutional quality, which we measure using the government antidiversion policy (*GADP*) index by Hall and Jones (1999). Since institutional quality might be endogenous as a result of reverse causality, we follow Hall and Jones (1999) and use proxies for Western influence as instruments for *GADP*: distance from equator, the fraction of a country's population speaking English as mother tongue, and the fraction of a country's population speaking one of the five European languages (English, French, German, Portuguese, Spanish) as mother tongue. Results in column (2) show that output volatility remains significant and negatively related to economic growth, even if institutional quality in terms of *GADP* is controlled for.

We now test for a direct effect of cyclical policy on growth. This means, that the variable *CYC* is included in the main equation (and instrumented by itself), whereas output volatility is instrumented by the institutional variables (*NELEC*, *POLCON*, *MAJ*, *PRES*) and exogenous volatility spillover (Z^σ). Columns (3a), (3b), and (3c) show the LS, TSLS, and WTSLS estimates of equation (5) including institutional quality as control.¹¹ The estimates show no evidence for a direct effect. This holds up for the unweighted and least squares regressions or when we allow the effect of pro- and countercyclical policy to differ (see column (4)). We also explored the subsample stability with respect to the cross-country dimension (rich countries, excluding countries with large volatility or cyclical policy coefficients) and the time dimension, considering the more recent time span 1980 to 2004. In none of the specifications, we could identify a direct effect of cyclical policy besides output volatility, irrespective of whether we allow the effect of countercyclical policy to differ from that of procyclical policy or not.

Overall, the findings in sections III and section IV provide a consistent picture. Cyclical as well as discretionary fiscal policy amplify output volatility (Tables 2 to 4), which is in turn negatively related to economic growth (Table 5). This also shows up in direct estimates of growth on cyclical policy (Table 4). The effects found for rich countries are of the same order of

¹¹ Notice that, since *CYC* is included in the main model, the independent variation to identify the effect of volatility on growth comes from the instrument Z^σ as well as from the effect of institutions on volatility through discretionary fiscal policy (see Figure 2). The latter is also apparent from the fact, that in a regression with output volatility as dependent variable, the political variables (*MAJ*, *PRES*, *NELEC*, *POLCON*) are jointly significant determinants of output volatility when *CYC* is controlled for, but they become insignificant if *DISCR* is added to the regressions as well. This suggests an alternative approach, using *DISCR* as instrument for σ^y directly (instead of the instead of the institutional variables); results turned out almost identical and are not shown here for brevity.

magnitude but estimated less precisely, rendering the effects insignificant in some specifications with p-values slightly above the 10 percent level.

V. Conclusions

Previous studies found that discretionary fiscal policy, defined as policy unrelated to the business cycle, lowers output growth by increasing output volatility. Using a large cross-section of 88 countries over the period 1960 to 2004, the present paper provides comprehensive empirical evidence that this is also true for cyclical fiscal policy.

We estimate simple average measures of cyclicalities of government consumption over the period 1960 to 2004 for each of the 88 countries of our sample. We find that cyclical fiscal policy constitutes a non-negligible share of overall fiscal policy, accounting for roughly one fifth in the total variation of government consumption in our sample. We then demonstrate that institutional variables (such as political constraints and the average number of elections) contain considerable information about the cross-country variation of fiscal cyclicalities, similar as for discretionary fiscal policy (Fatas and Mihov, 2003).

Using institutional variables as instruments, we identify a destabilizing effect of cyclical fiscal policy on economic activity, irrespective of whether fiscal policy is pro- or countercyclical. This not only confirms the scepticisms against the usefulness of countercyclical fiscal policy as fine tuning instrument to smooth business cycles. It also implies that countercyclical policy has the same amplifying effect on volatility as procyclical fiscal policy; in fact, we find some support for the hypothesis that cyclical fiscal policy affects volatility much in the same way as pure fiscal shocks. According to this result, the way towards stabilization does not lead over more active countercyclical fiscal policy, but less cyclical fiscal policy at all.

The gains from this (passive) stabilization policy could be substantial in light of our finding that aggressive use of cyclical (as well as discretionary fiscal) policy has a negative effect on economic growth. This result is obtained both in a direct regression of growth on cyclicalities and in a two stages least squares regression of growth on volatility, using the fiscal policy-related institutional variables as instrumental variables in the first stage regression.

Overall, our results have an important policy implication: Economic growth could be enhanced by introducing fiscal rules, designed to limit the use of discretionary fiscal policy on the one hand (as already argued by Fatas and Mihov, 2003) but also the use cyclical fiscal expenditure policy on the other hand. Notwithstanding the robustness of our results with respect to subsample stability over the cross-country and time dimension, It should be added, however, that our cross-section estimates should be interpreted as averages over countries and time, not as economic laws that apply to every government at any time. Moreover, the use of government consumption as measure of fiscal policy– a choice required to obtain a relatively

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3 large sample of countries – limits the generality of the results. As a consequence, it would be
4 interesting to investigate the effect of fiscal policy on volatility and growth for smaller groups
5 of countries with more comprehensive, more detailed and higher frequency data on fiscal
6 policy. Another question that remains to be addressed in future research is how existing fiscal
7 rules affect the cyclical responsiveness of governments. Additional evidence on these issues
8 would be informative about the channels, through which fiscal policy affects economic
9 growth, and the optimal design of fiscal rules.
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Appendix

A1. Sample Description

The largest set of countries for which the required key variables are available comprises 88 countries. The list of countries is reported in Table A1, along with our estimates of the cyclical parameter χ and the aggressiveness of discretionary fiscal policy (*DISCR*) from equation (1).

Table A1. *Sample and Data on Output Volatility, Cyclical and Discretionary Fiscal Policy*

country	time period (obs)	$\sigma^{\Delta \ln y}$		Cyclical ($\hat{\chi}$)		<i>DISCR</i>	
		1960-2004	1980-2004	1960-2004	1980-2004	1960-2004	1980-2004
Argentina	1960-2004 (45)	5.753	6.378	1.540	2.190	29.741	37.765
Australia	1960-2004 (45)	1.840	1.807	-0.375	0.009	3.086	1.965
Austria	1960-2004 (45)	1.777	1.162	0.264	0.493	2.035	1.602
Burundi	1960-2004 (45)	5.798	4.399	1.070	0.875	16.535	17.037
Belgium	1960-2004 (45)	1.833	1.333	-0.080	-0.031	2.396	1.841
Benin	1960-2004 (45)	3.099	3.058	-0.034	0.038	8.456	9.478
Burkina Faso	1960-2004 (45)	3.101	3.338	0.740	1.211	13.896	11.118
Bangladesh	1960-2004 (45)	4.133	1.473	2.698	1.834	12.215	7.613
Bolivia	1970-2004 (35)	3.616	2.650	2.049	2.394	9.886	9.686
Brazil	1960-2004 (45)	3.877	3.453	0.480	0.596	8.490	9.000
Centr. Afr. Rep.	1960-2002 (43)	4.034	4.678	0.891	1.095	11.129	12.311
Canada	1965-2004 (40)	1.982	2.122	-0.464	-0.283	1.871	1.767
Switzerland	1960-2004 (45)	2.247	1.572	0.440	0.446	2.065	1.873
Chile	1960-2004 (45)	4.879	4.565	0.824	0.428	8.088	5.652
Cote d'Ivoire	1960-2004 (45)	5.214	4.104	1.270	1.059	9.886	10.647
Cameroon	1965-2004 (40)	5.818	5.306	0.988	1.004	8.558	10.020
Congo, Rep.	1960-2004 (45)	5.500	6.002	0.461	0.448	17.857	22.703
Colombia	1960-2004 (45)	2.084	2.095	0.686	0.693	9.275	9.074
Costa Rica	1960-2004 (45)	3.241	3.585	1.016	1.177	5.458	6.202
Germany	1960-2004 (45)	1.880	1.441	0.135	0.414	2.443	1.525
Denmark	1960-2004 (45)	2.151	1.624	0.215	-0.342	2.526	1.804
Dom. Republic	1960-2004 (45)	5.080	3.624	0.728	3.369	21.349	19.347
Algeria	1960-2004 (45)	7.147	2.591	0.931	0.551	9.207	7.538
Ecuador	1960-2004 (45)	3.448	3.162	1.365	0.826	11.162	10.418
Egypt	1960-2004 (45)	2.726	1.758	0.662	-0.202	9.407	6.722
Spain	1960-2004 (45)	2.326	1.574	0.656	0.430	2.659	2.223
Finland	1960-2004 (45)	2.838	2.859	-0.054	-0.074	2.958	2.857
Fiji	1960-2000 (41)	4.616	4.594	0.628	0.441	7.567	8.167
France	1960-2004 (45)	1.585	1.138	-0.281	-0.432	1.411	1.186
Gabon	1960-2000 (41)	9.397	5.968	0.616	0.683	20.247	17.257
United Kingdom	1960-2004 (45)	1.735	1.754	-0.262	-0.273	2.320	1.914
Ghana	1960-2004 (45)	4.592	3.603	1.776	3.716	14.358	14.671
Greece	1960-2004 (45)	3.736	2.321	0.437	0.771	5.451	5.355
Guatemala	1960-2004 (45)	2.458	2.178	1.121	2.170	8.607	9.144
Honduras	1960-2004 (45)	2.886	2.423	-0.110	0.596	7.057	7.434
Haiti	1967-2003 (37)	4.469	4.146	1.146	1.134	10.656	10.947
Indonesia	1960-2004 (45)	3.981	4.444	1.560	1.453	14.998	8.224
India	1960-2004 (45)	3.038	1.817	0.518	0.340	5.047	3.237
Ireland	1960-2004 (45)	2.595	2.930	0.602	0.610	3.561	3.373
Iceland	1960-2004 (45)	3.701	2.914	0.860	0.669	3.904	3.244

Table A1 (continued). *Sample and Data on Output Volatility, Cyclical and Discretionary Fiscal Policy*

country	time period (obs)	$\sigma^{\Delta \ln y}$ in percent		Cyclicalit y ($\hat{\chi}$)		DISCR in percent	
		1960-2004	1980-2004	1960-2004	1980-2004	1960-2004	1980-2004
Israel	1960-2004 (45)	3.331	2.118	1.181	1.334	9.928	6.283
Italy	1960-2004 (45)	2.148	1.275	0.059	0.293	2.848	2.578
Jamaica	1966-2004 (39)	4.596	3.167	1.150	1.427	9.612	10.284
Japan	1960-2004 (45)	3.453	1.791	0.128	0.123	2.205	0.879
Kenya	1960-2004 (45)	4.562	1.982	1.202	1.789	5.977	4.584
Korea, Rep.	1960-2004 (45)	3.441	3.782	-0.281	-0.121	6.844	3.137
Sri Lanka	1960-2004 (45)	1.901	1.722	0.319	1.089	8.646	9.695
Lesotho	1960-2004 (45)	6.383	3.844	0.239	-0.350	15.132	9.970
Morocco	1960-2004 (45)	4.449	4.796	0.700	0.549	7.568	4.719
Madagascar	1960-2004 (45)	4.266	4.782	1.240	1.220	9.628	10.979
Mexico	1960-2004 (45)	3.312	3.619	1.660	1.760	5.305	5.752
Mali	1967-2004 (38)	5.140	5.154	0.660	1.296	16.133	15.721
Mauritania	1960-2004 (45)	5.890	3.270	0.635	1.900	17.809	18.398
Mauritius	1980-2004 (25)	1.535	1.535	0.583	0.583	3.751	3.751
Malawi	1960-2004 (45)	5.423	5.721	-0.835	-2.066	15.870	16.817
Malaysia	1960-2004 (45)	3.385	3.971	0.229	0.548	8.451	8.308
Niger	1960-2004 (45)	6.253	5.542	0.762	1.270	11.605	11.165
Nigeria	1960-2004 (45)	7.186	5.100	0.429	0.814	19.854	22.460
Nicaragua	1960-2004 (45)	6.895	4.200	0.277	1.633	19.112	22.562
Netherlands	1960-2004 (45)	1.833	1.503	0.350	-0.027	2.177	1.770
Norway	1960-2004 (45)	1.554	1.618	0.631	0.023	3.675	3.982
New Zealand	1971-2004 (34)	2.835	1.946	0.360	0.159	4.133	2.658
Pakistan	1960-2004 (45)	2.187	1.793	0.975	1.308	8.104	7.276
Panama	1960-2004 (45)	4.197	4.832	1.168	1.044	6.571	6.125
Peru	1960-2004 (45)	5.038	6.201	1.248	1.381	9.718	7.948
Philippines	1960-2004 (45)	3.027	3.629	1.261	1.490	5.344	5.665
Pap. New Guinea	1961-1999 (39)	4.610	5.118	1.042	0.893	8.776	9.588
Portugal	1960-2004 (45)	3.295	2.521	0.750	1.261	3.549	2.571
Paraguay	1960-2004 (45)	3.793	3.768	0.469	0.988	9.713	10.104
Rwanda	1960-2004 (45)	11.997	14.623	1.171	1.303	16.300	16.653
Senegal	1960-2004 (45)	4.173	3.777	0.827	0.033	18.002	5.998
Singapore	1960-2004 (45)	4.175	3.832	0.043	-0.339	6.823	7.359
El Salvador	1960-2004 (45)	3.834	4.215	0.561	0.319	5.802	5.548
Sweden	1960-2004 (45)	1.925	1.837	0.032	0.050	2.252	1.828
Syr. Arab Rep.	1960-2004 (45)	7.809	5.298	0.569	-0.005	9.749	9.548
Chad	1960-2004 (45)	8.186	8.773	0.064	0.027	14.689	19.314
Togo	1960-2004 (45)	6.131	6.234	0.174	0.410	17.114	13.173
Thailand	1960-2004 (45)	3.635	4.436	0.519	0.446	4.761	3.667
Trin. and Tobago	1960-2004 (45)	4.749	5.666	1.195	1.269	11.929	13.900
Tunisia	1961-2004 (44)	3.343	2.557	0.443	-0.019	5.614	2.266
Turkey	1968-2004 (37)	4.147	4.637	0.629	0.602	7.737	8.357
Uruguay	1960-2004 (45)	4.744	5.812	0.989	0.974	10.328	5.996
United States	1960-2004 (45)	1.906	1.820	-0.048	-0.091	1.967	1.310
Venezuela	1960-2004 (45)	4.949	5.938	1.616	1.913	16.543	20.776
South Africa	1960-2004 (45)	2.525	2.658	0.892	0.664	4.724	3.700
Congo, Dem. Rep.	1960-2004 (45)	6.079	5.363	1.570	1.425	26.618	32.791
Zambia	1960-2004 (45)	4.703	3.919	1.485	1.928	21.520	24.135
Zimbabwe	1960-2004 (45)	5.828	5.811	0.335	0.339	12.693	15.324
Correlation		0.842		0.766		0.935	

A2. Variable Definitions and Data Sources

<i>DEP</i>	dependency ratio, defined as ratio of people younger than 15 and older than 64 to working age population (people from 15 to 64). Source: World Development Indicators (WDI).
<i>GC</i>	real general government consumption in national currency. Source: WDI.
<i>GDPPC</i>	real GDP per capita in PPP\$. Source: Penn World Tables (PWT) 6.2.
<i>GDPPW</i>	real GDP per worker in PPP\$. Source: PWT 6.2.
<i>GSIZE</i>	ratio of government consumption to GDP.
<i>HC^{prim}</i>	primary educational attainment, defined as fraction of males above 25 with primary schooling. Source: Barro and Lee (2002).
<i>HC^{sec}</i>	secondary educational attainment, defined as fraction of males above 25 with secondary schooling. Source: Barro and Lee (2002).
<i>MAJ</i>	zero-one dummy for electoral system (1 for majoritarian, 0 for proportional). Sources: Person and Tabellini (2001), Database of Political Institutions.
<i>NELEC</i>	average number of elections. Sources: Database of Political Institutions.
<i>OPEN</i>	ratio of imports plus exports to GDP. Source: PWT 6.2.
<i>POLCON</i>	index of political constraints. Source: Henisz (2000).
<i>POP</i>	population in million persons. Source: WDI.
<i>PRES</i>	zero-one dummy for political regime (1 for presidential, 0 for parliamentary). Source: Person and Tabellini (2001), Database of Political Institutions.
<i>URB</i>	ratio of urban population to total population. Source: WDI.
<i>y</i>	real GDP in national currency per capita; $y = Y/POP$. Source: WDI.
<i>Y</i>	real GDP in national currency. Source: WDI.

Data source for the geographical variables (such as bilateral distance, area, common border dummies) is the CEPII database (see Mayer and Zignago, 2006).

A3. Robustness

Table A2 shows that the logarithmic transformation of *CYC* is not crucial for the results (see columns (1a) and (1b)). However, the fit is worse than for the specification in log form. We next consider subsample stability of the results, excluding countries with ‘large’ output volatility or ‘large’ cyclicalities from the sample. Columns (2a) and (2b) give the results for models (2) and (3), excluding countries whose output volatility exceeds the sample average by more than one standard deviation. The same exercise is repeated in columns (3a) and (3b), excluding countries whose cyclicalities coefficients exceed the sample average by more than one standard deviation. Overall, the results for the full sample hold up, confirming that our results are not driven by a few outlying observations.

< Table A2 >

In columns (4a) and (4b) we focus on a subsample of 28 ‘rich’ countries, constituting the upper third of the income distribution of our sample in terms of GDP per capita. In both models (2) and (3) the variable *CYC* becomes insignificant with p-values of 0.245 and 0.193, respectively. However, if the level of development (*GDPPC*) is excluded, which appears to be justifiable for a group of countries with a similar level of development (in particular in model (3) where *GDPPC* is insignificant), *CYC* becomes significant again in models (2) and (3) at the 10 and 5 percent level respectively.

In a final step we consider the results for the more recent period 1980 to 2004. First note that our estimates of *CYC* and *DISCR* for the full period from 1960 to 2004 and for the period of 1980 to 2004 are fairly similar (see Appendix A1). Many of the countries that pursued countercyclical fiscal policy from 1960 to 2004 on average did so as well in the period from 1980 to 2004. The correlation between *CYC* (*DISCR*) for the two time periods is 0.766 (0.935). This also favours the interpretation that (de facto) time invariant institutional features of countries have a strong impact on the cross-country variation in fiscal policy. A slight difference to the estimation results for period 1960 to 2004 is that the effect of countercyclicality is significantly different from that of procyclical policy in the logarithmic specification (see column (5a)), though its effect on output volatility is still clearly positive. One might argue that countercyclical policy has become more effective (and thus overall less destabilizing). But this appears to be overstressing results a bit, given that there is no significant difference between the effects of pro- and countercyclical policy if the level rather than the log of *CYC* is used; column (5b) shows the (restricted) model using the level of *CYC*.

The estimates of model (3) for the period 1980 to 2004 (columns (6a) and (6b)), where *DISCR* is included along with *CYC*, are in line with the results for the full period 1960 to 2000, in particular when the model is estimated using LIML (column (6b)). As before the hypothesis of parameter equality of *CYC* and *DISCR* cannot be rejected. Finally, the subsample stability with respect to the cross-country dimension for the period 1980 to 2004 is qualitatively very similar to that for the full period.

Table 1. *Political Determinants of Cyclicality and First Stage Regressions*

Dependent variable is:	(1a) ln CYC	(1b) ln CYC	(1c) ln CYC	(1d) ln CYC	(2a) ln CYC	(2b) ln CYC	(3) ln CYC	(4) ln σ^{Discr}
<i>NELEC</i>	-2.674*** (0.774)				-2.566** (1.000)	-2.391*** (0.660)	-2.305*** (0.767)	-0.513 (0.355)
<i>POLCON</i>		-1.905*** (0.334)			-1.704*** (0.569)	-1.822*** (0.313)	-1.489** (0.658)	-0.470** (0.236)
<i>MAJ</i>			0.007 (0.233)		-0.036 (0.197)			-0.139** (0.068)
<i>PRES</i>				0.460* (0.242)	0.112 (0.415)			0.408*** (0.143)
ln <i>POP</i>							-0.196*** (0.064)	-0.082** (0.024)
<i>URBAN</i>							-0.800 (0.753)	0.004 (0.259)
<i>DEP</i>							3.003* (1.549)	0.477 (0.568)
<i>OPEN</i>							-0.577** (0.299)	0.091 (0.102)
ln <i>GDPPC</i>							0.609* (0.359)	-0.283** (0.122)
F-stat. ¹⁾	11.934***	32.589***	0.001	3.594*	12.312***	25.142***	6.587**	6.289***
F-stat. ²⁾							8.472***	8.663***
<i>R</i> ²	0.122 (0.007)	0.275 (0.146)	0.000 (0.017)	0.040 (0.018)	0.372 (0.091)	0.372 (0.089)	0.502 (0.13)	0.872 (0.674)
Observations	88	88	88	88	88	88	88	88

Notes: A constant is included in all models. *, **, *** denote significance at 10, 5, and 1 percent level respectively; standard errors in parenthesis. Table reports weighted least squares estimates, using the inverse of the variance of *CYC* as weight. *R*² refers to weighted model. ¹⁾ F-test on excluding the institutional variables (*NELEC*, *POLCON*, *MAJ*, *PRES*) from first stage regression for ln *CYC* and ln σ^{Discr} . ²⁾ F-Test on excluding all instruments (institutional variables, ln *POP*, *URBAN*, *DEP*) from regression.

Table 2. *Cyclical Policy and Output Volatility – Basic Model*

	(1a)	(1b)	(2a)	(2b)	(3a)	(3b)	(4a)	(4b)
	WLS	WTLS	WLS	WTLS	LS	TSLS	WTLS	WTLS
$\ln CYC$	0.184** (0.082)	0.595** (0.232)	0.239*** (0.060)	0.600*** (0.207)	0.150*** (0.044)	0.742*** (0.185)		
$D^{counter} \ln CYC$	0.094 (0.083)	-0.094 (0.208)						
CYC							1.432*** (0.311)	1.762*** (0.332)
$D^{counter} CYC$							-0.528 (0.517)	
Hausman ¹⁾ (p-val.)		(0.000)		(0.000)		(0.000)	(0.000)	(0.000)
OID ²⁾ (p-val.)		(0.128)		(0.236)		(0.227)	(0.036)	(0.887)
R^2	0.128	0.096	0.108	0.108	0.108	0.108	0.105	0.099
SEE	0.448	0.606	0.454	0.626	0.431	0.731	0.768	0.915
Observations	88	88	88	88	88	88	88	88

Notes: Dependent variable is $\ln \sigma^y$. A constant is included in all models. *, **, *** denote significance at 10, 5, and 1 percent level respectively; robust standard errors in parenthesis. WLS denotes weighted least squares estimates, using the inverse of the variance of CYC as weight. (W)TSLS denotes (weighted) two stages least squares, using $NELEC$ and $POLCON$ as instruments for CYC . R^2 refers to unweighted model, calculated as squared correlation between actual and predicted values. ¹⁾ Heteroskedasticity-robust test for exogeneity; H_0 : CYC is exogenous. ²⁾ Heteroskedasticity-robust test of overidentifying restrictions; H_0 : $NELEC$, $POLCON$ are valid instruments.

Table 3. *Cyclical Policy and Output Volatility – Extended Model*

	(1a)	(1b)	(1c)	(2a)	(2b)	(3a)	(3b)	(4)
	LS	TSLs	WTSLs	TSLs	WTSLs	WTSLs	LIML	WTSLs
$\ln CYC$	0.079* (0.040)	0.676** (0.270)	0.292** (0.119)	0.318* (0.166)	0.324* (0.188)	0.163* (0.090)	0.182** (0.083)	0.238*** (0.068)
$D^{counter} \ln CYC$				-0.067 (0.133)	-0.084 (0.153)			
$\ln DISCR$						0.454 (0.276)	0.417** (0.204)	restricted
$Gsize$	-0.045 (0.620)	1.840 (4.894)	-3.406 (3.903)	0.635 (0.260)	-6.354 (4.654)	-2.709 (2.854)	-1.299 (0.994)	-3.431 (2.724)
$OPEN$	0.176** (0.080)	0.447* (0.257)	-0.003 (0.194)	0.293* (0.166)	0.077 (0.196)	-0.118 (0.129)	-0.108 (0.109)	-0.061 (0.158)
$\ln GDPPC$	-0.239*** (0.041)	-0.096 (0.125)	-0.206 (0.127)	-0.189*** (0.071)	-0.166 (0.158)	-0.004 (0.107)	-0.013 (0.099)	-0.088 (0.114)
Hausman ¹⁾ (p-val.)		(0.000)	(0.000)	(0.179)	(0.000)	(0.016)		(0.000)
OID ²⁾ (p-val.)		(0.893)	(0.683)	(0.057)	(0.554)	(0.817)		(0.913)
R^2	0.404	0.159	0.279	0.292	0.220	0.479	0.534	0.376
SEE	0.358	0.658	0.507	0.418	0.583	0.469	0.472	0.485
Observations	88	88	88	88	88	88	88	88

Notes: Dependent variable is $\ln \sigma^y$. A constant is included in all models. *, **, *** denote significance at 10, 5, and 1 percent level respectively; robust standard errors in parenthesis. WLS denotes weighted least squares estimates, using the inverse of the variance of CYC as weight. (W)TSLs denotes (weighted) two stages least squares, using $NELEC$ and $POLCON$ as instruments for CYC and using $NELEC$, $POLCON$, MAJ , and $PRES$ as instruments for CYC and $DISCR$ in columns (3) and (4). ¹⁾ Heteroskedasticity-robust Hausman test for exogeneity; H_0 : $\ln CYC$, $Gsize$ (and $\ln DISCR$) are exogenous. ²⁾ Heteroskedasticity-robust test of overidentifying restrictions; H_0 : $NELEC$, $POLCON$ (MAJ , $PRES$) and $\ln POP$, DEP , $URBAN$ are valid instruments.

Table 4. *Cyclical Policy and Economic Growth*

	(1a)	(1b)	(2a)	(2b)	(3a)	(3b)	(3c)
	LS	TSLS	WLS	WTSLs	LS	TSLS	WTSLs
$\ln CYC$	-0.362** (0.161)	-0.767* (0.441)	-0.341* (0.188)	-0.606* (0.344)	-0.297** (0.118)	-1.149** (0.514)	-0.439* (0.226)
$D^{counter} \ln CYC$	0.151 (0.157)	0.379 (0.331)	0.090 (0.152)	0.173 (0.215)			
$\ln GDPPW^{in}$	-0.748*** (0.158)	-0.798*** (0.166)	-1.046*** (0.195)	-1.016*** (0.120)	-0.748*** (0.158)	-0.895*** (0.191)	-1.087*** (0.200)
$\ln HC^{prim}$	0.524** (0.218)	0.568** (0.244)	1.134*** (0.212)	1.180*** (0.236)	0.532 (0.217)	0.688** (0.318)	1.163*** (0.242)
$\ln HC^{sec}$	0.724*** (0.136)	0.646*** (0.185)	0.716*** (0.177)	0.567** (0.279)	0.726 (0.132)	0.515** (0.223)	0.663*** (0.240)
Hausman ¹⁾ (p-val.)		(0.467)		(0.182)		(0.015)	(0.418)
OID ²⁾ (p-val.)		(0.028)		(0.335)		(0.929)	(0.250)
R^2	0.400	0.363	0.370	0.358	0.395	0.268	0.365
SEE	1.051	1.100	1.203	1.121	1.048	1.332	1.107
Observations	80	80	80	80	80	80	80

Notes: Dependent variable is $\Delta \ln GDPPW$. A constant is included in all models. *, **, *** denote significance at 10, 5, and 1 percent level respectively; robust standard errors in parenthesis. WLS denotes weighted least squares estimates, using the inverse of the variance of CYC as weight. (W)TSLs denotes (weighted) two stages least squares, using *NELEC* and *POLCON* as instruments for CYC . ¹⁾ Heteroskedasticity-robust Hausman test for exogeneity; H_0 : $\ln CYC$ is exogenous. ²⁾ Heteroskedasticity-robust test of overidentifying restrictions; H_0 : *NELEC*, *POLCON* are valid instruments.

Table 5. Cyclical Policy, Output Volatility, and Economic Growth

	(1a)	(1b)	(1c)	(1d)	(2)	(3a)	(3b)	(3c)	(4)
	LS	TSLS	TSLS	TSLS	TSLS	LS	TSLS	WTSLS	WTSLS
$\ln \sigma^y$	-0.994** (0.302)	-2.968*** (0.107)	-3.689* (2.128)	-2.993*** (0.955)	-2.174* (1.196)	-0.662** (0.292)	-2.101* (1.187)	-1.173* (0.083)	-1.416** (0.611)
$\ln CYC$						-0.100 (0.109)	-0.071 (0.158)	-0.194 (0.133)	-0.309 (0.232)
$D^{counter} \ln CYC$									0.182 (0.169)
$\ln GDPPW^{in}$	-0.821*** (0.157)	-1.066*** (0.223)	-1.155*** (0.236)	-1.069*** (0.213)	-1.065*** (0.176)	-0.971*** (0.198)	-1.058*** (0.175)	-1.412*** (0.249)	-1.364*** (0.252)
$\ln HC^{prim}$	0.551** (0.203)	0.697*** (0.230)	0.750*** (0.274)	0.699*** (0.225)	-0.569* (0.292)	0.401*** (0.146)	0.587* (0.297)	1.090*** (0.0186)	1.127*** (0.190)
$\ln HC^{sec}$	0.580*** (0.148)	0.144** (0.305)	-0.015** (0.491)	0.138 (0.270)	0.446 (0.241)	0.598*** (0.001)	0.341 (0.242)	0.671*** (0.149)	0.061*** (0.147)
$GADP$					1.007 (1.562)	2.067 (0.650)	0.861 (1.642)	-0.314 (1.610)	-0.786 (1.463)
Hausman ¹⁾ (p-val.)		(0.023)	(0.026)	(0.011)	(0.251)		(0.302)	(0.187)	(0.041)
OID ²⁾ (p-val.)		(0.760)	-	(0.771)	(0.644)		(0.607)	(0.692)	(0.673)
R^2	0.409	0.316	0.278	0.314	0.430	0.537	0.463	0.417	0.406
SEE	1.036	1.222	1.363	1.227	1.053	0.935	0.484	1.138	1.172
Observations	80	80	80	80	78	78	78	78	78

Notes: Dependent variable is $\Delta \ln GDPPW$. A constant is included in all models. *, **, *** denote significance at 10, 5, and 1 percent level respectively; robust standard errors in parenthesis. WLS denotes weighted least squares estimates, using the inverse of the variance of CYC as weight. TSLS denotes (weighted) two stages least squares, using the following instruments for σ^y : $NELEC$, $POLCON$, MAJ , $PRES$ in (1b), Z^σ in (1c), and both the institutional variables and Z^σ in all other columns. ¹⁾ Heteroskedasticity-robust Hausman test for exogeneity; H_0 : $\ln CYC$, $Gsize$ (and $\ln DISCR$) are exogenous. ²⁾ Heteroskedasticity-robust test of overidentifying restrictions; H_0 : $NELEC$, $POLCON$ (MAJ , $PRES$) and $\ln POP$, DEP , $URBAN$.

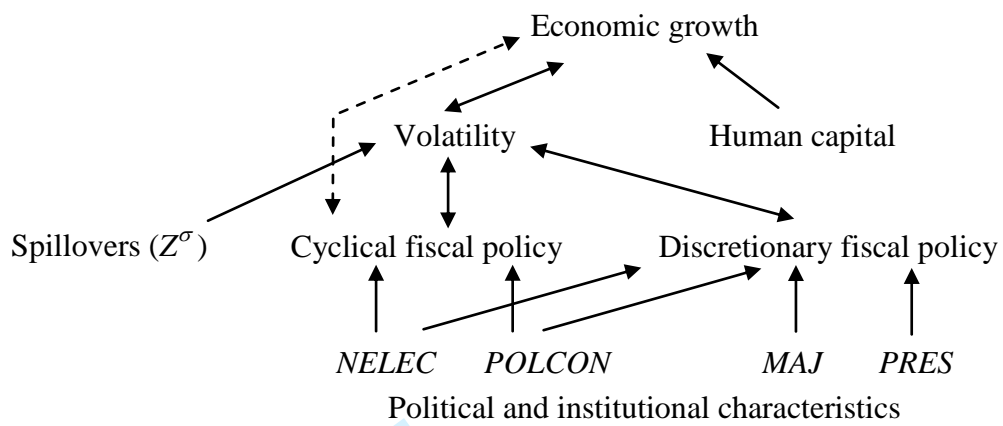
Figure 1. *Relationships between Key Variables in Empirical Models*

Table A2. Cyclicalities of Fiscal Policy and Output Volatility – Robustness

	1960-2004								1980-2004			
	(1a)	(1b)	(2a)	(2b)	(3a)	(3b)	(4a)	(4b)	(5a)	(5b)	(6a)	(6b)
	level of <i>CYC</i>		exl. large volatility		exl. large <i>CYC</i>		Rich		Level of <i>CYC</i>			
	equ. (2b)	equ. (2a)	equ. (2b)	equ. (3)	equ. (2a)	equ. (3)	equ. (2a)	equ. (3)	equ. (2b)	equ. (2a)	equ. (3)	equ. (3)
	WTSLs	WTSLs	WTSLs	WTSLs	WTSLs	WTSLs	WTSLs	WTSLs	WTSLs	WTSLs	WTSLs	LIML
<i>ln CYC</i>			0.130	0.071*	0.282**	0.129**	0.131	0.063	0.231***		0.078	0.273*
			(0.087)	(0.042)	(0.116)	(0.059)	(0.110)	(0.047)	(0.066)		(0.056)	(0.153)
<i>D^{counter} ln CYC</i>									-0.100**			
									(0.043)			
<i>CYC</i>	0.825***	1.034**								0.946**		
	(0.314)	(0.317)								(0.459)		
<i>D^{counter} CYC</i>	-0.026											
	(0.630)											
<i>ln DISCR</i>				0.505**		0.554**		0.488**			0.405***	0.408**
				(0.170)		(0.212)		(0.235)			(0.072)	(0.164)
<i>Gsize</i>	-4.899	-4.988	-2.389	-2.345	-2.310	-2.199	-2.122	-2.547	-1.369	-4.708	-3.606**	-6.977**
	(5.222)	(4.202)	(3.300)	(2.164)	(3.598)	(0.283)	(2.844)	(1.936)	(1.782)	(3.681)	(1.676)	(3.155)
<i>OPEN</i>	0.037	0.023	0.0063	-0.076	-0.029	-0.157	-0.021	-0.116	-0.039	0.208	-0.089	0.092
	(0.144)	(0.149)	(0.128)	(0.088)	(0.210)	(0.116)	(0.171)	(0.110)	(0.137)	(0.231)	(0.072)	(0.216)
<i>ln GDPPC</i>	-0.145	-0.095	-0.187**	0.063	-0.237**	0.025	-0.502*	-0.172	-0.388***	-0.164	-0.105	0.065
	(0.150)	(0.145)	(0.093)	(0.090)	(0.114)	(0.090)	(0.277)	(0.314)	(0.065)	(0.194)	(0.098)	(0.186)
Hausman (p-val.)	(0.009)	(0.000)	(0.100)	(0.218)	(0.000)	(0.028)	(0.109)	(0.187)	(0.000)	(0.023)	(0.138)	
OID (p-val.)	(0.696)	(0.949)	(0.075)	(0.122)	(0.544)	(0.609)	(0.336)	(0.247)	(0.206)	(0.309)	(0.200)	
<i>R</i> ²	0.204	0.167	0.344	0.520	0.309	0.507	0.309	0.651	0.296	0.175	0.463	0.317
SEE	0.642	0.713	0.335	0.366	0.479	0.433	0.399	0.342	0.654	1.008	0.643	0.673
Observations	89	89	77	77	79	79	28	28	88	88	88	88

Notes: Dependent variable is $\ln \sigma^y$. A constant is included in all models. *, **, *** denote significance at 10, 5, and 1 percent level respectively; robust standard errors in parenthesis. See also Tables 2 and 3.