Asymmetric and non-linear adjustment in the revenue-expenditure models

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Abstract

The purpose of this paper is to empirically analyse the revenue-expenditure models of public finance by considering the possibility of non-linear and asymmetric adjustment. A long-run relationship between general government expenditure and revenues is identified for Italy. Following system-wide shocks, the estimated relationship adjusts slowly to equilibrium, mainly due to complex administrative procedures that add to the sluggishness of tax collection and undermine the effective monitoring of public spending. Exogeneity of public expenditure implies that taxes rather than spending, carry the burden of short-run adjustment to correct budgetary disequilibria. Allowing for non-linear adjustment and the possibility of multiple equilibria, our findings show evidence of asymmetric adjustment around a unique equilibrium. In particular, we find that when government expenditure is too high, adjustment of taxes takes places at a faster rate than when it is too low. Further, there is evidence of a faster adjustment when deviations from the equilibrium level get larger, pointing to a Leviathan-style, revenue-maximiser government.

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1. Introduction

This paper analyzes the revenue-expenditure patterns of the Italian public finances, whose most remarkable feature is the presence of a sustained post-war growth of the general government expenditure, through deficit spending and the creation of debt. It was only after the Maastricht (1992) treaty that public spending control in Italy became an important objective of economic policy with the aim of gaining admission to the European Monetary Union (EMU).

Between the late 1950s and 1998, the Italian general government expenditure grew at an average nominal rate of 14.6%, raising the government share from 33% up to 57.6% of the national GDP in 1993. During the same years, public revenues grew at an average nominal rate of around 13%, which was not enough to reverse the budgetary imbalances, as evidenced by the constant presence of deficit spending and the enormous dimensions of the Italian public debt. The latter rose from around 41% of GDP in the late 1950s to around 124% of GDP in 1994, only to fall to 110% of GDP in 2000, representing the highest value for the EU countries. At the same time, fiscal pressure in Italy is higher than the OECD average; in particular, the taxes to GDP ratio is 43.3% for Italy compared with an OECD average of 37.3%. [1]

Given the importance of the above variables for economic policy purposes as well as for the attainment to the Maastricht and the European Stability and Growth Pact (ESGP) criteria, it is important to analyze more in detail the relationship between public expenditures and revenues. Indeed, examining the relationship between government expenditure and revenues can shed some light on the causes and the consequences of fiscal disequilibria.

Public finance theory provides three different models of revenue-expenditure. Within the public finance tradition, revenues and expenditures are simultaneously set by the government,

[1] In 2000, the taxes to GDP ratio for the European OECD countries stood at 39.9%.

to attain its economic policy objectives. Peacock and Wiseman (1979) discuss the spend-and-tax model according to which the government would raise the necessary resources to cover its spending, and therefore higher government expenditure would lead to higher taxes. In this case, a government spending restraint would be expected to lower the size of the government deficit. Friedman (1978) puts forward the tax-and-spend model according to which the government would spend all its revenues, and therefore higher taxes would lead to higher government expenditure. In this case, imposing higher taxes to restrict the size of the government deficit would raise it instead.

Taking into account the latest developments in the area of cointegration, recent empirical studies focus on the joint modelling of government expenditure and revenues. The popular empirical approach is to use either Engle and Granger's (1987) model or Johansen's (1988, 1995) methodology to determine the existence of a long-run relationship between government expenditure and revenues, and then run Granger's (1969) causality tests in order to see whether the tax-and-spend or the spend-and tax hypotheses (or both) hold. [2]

This paper examines the above hypotheses for Italy within a bivariate model of public expenditure and taxation. We apply the Johansen (1988, 1995) cointegration methodology, which allows for the long-run properties and the short-run dynamics of the variables in question to be jointly analysed. However, there are two main differences between our paper and earlier empirical studies in public finance. First, we use the persistence profile analysis (Pesaran and Shin, 1996) to investigate the speed with which deviations from the estimated cointegrating relation, resulting from system-wide shocks, are eliminated. Second, and much more important, we consider the case of asymmetric and non-linear adjustment back to equilibrium also testing for the possibility of multiple equilibria. We view this as an important

[2] There are numerous studies in the area. Among the most recent ones, Baghestani and McNown (1994) reject both the tax-and-spend and spend-and-tax hypotheses for the US. Using data for 9 industrialized countries, Koren and Stiassny (1998) find evidence in favor of the spend-and-tax hypothesis for Italy whereas Cheng (1999) rejects the spend-and-tax hypothesis for eight Latin American countries.

contribution to the public finance literature where non-linear error correction models have not been considered so far. [3] Non-linearities in the relationship between government spending and taxes may arise if for instance, the fiscal authorities react differently to positive and negative deviations from the long-run budgetary equilibrium level.

Our main findings point to a long-run relationship between general government expenditure and revenues for the case of Italy. We also find that the persistence profile of the cointegrating relationship converges to zero rather slowly. The sluggish rate of convergence towards equilibrium is explained by complex administrative procedures, problems in budget control and other rigidities in the functioning of the Italian public sector. Further, the structure of the estimated model suggests that the short-run adjustment to correct budgetary disequilibria is mainly done by changes in taxes rather than changes in government spending. In addition, short-run increases in government expenditure lead to higher taxes, which provides support for the spend-and-tax hypothesis. Allowing for non-linear effects in the short-run adjustment process provides evidence of asymmetric adjustment around a unique (at zero) equilibrium rather than multiple ones. In particular, when government expenditure rises above its equilibrium level, the fiscal authorities respond by raising taxes rapidly. On the other hand, when the expenditure falls below its long-run level, the fiscal authorities respond by lowering taxes slowly. Further, there is evidence of a faster adjustment when deviations from the equilibrium level get larger. These results point to the presence of a Leviathan-style, revenue-maximiser government.

The paper is organised as follows. Section 2 provides the theoretical background to our analysis by discussing the main revenue-expenditure models in public finance theory. Section 3 applies multivariate cointegration techniques to determine the existence of a long-run

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[3] Other studies explore (but not in an error correction context) the non-linear response of national saving to fiscal policy for a number of developed and developing countries, see e.g. Giavazzi *et al*. (2000) and the references cited therein. Sarno (2001) estimates a non-linear univariate model of the US debt-GDP ratio.

relationship between government spending and taxes for Italy, whereas Section 4 presents the short-run dynamics of the empirical model also allowing for asymmetric and non-linear adjustment. Finally, Section 5 offers some concluding remarks and provides some policy implications.

2. Public finance models of revenues-expenditure

Within public finance analysis, the traditional models consider both general government expenditure and revenues as simultaneously determined by a "benevolent government" in order to maximize a social welfare function.

A different stream of theories is concerned with the interdependencies between the two variables, originating the debate on the tax-and-spend or the spend-and-tax hypotheses. It is important to stress that this debate is not only theoretical, since the identification of the relevant revenue-expenditure pattern is fundamental to set the appropriate strategy of fiscal discipline.

Within Peacock and Wiseman analysis (1979), the government raises taxes to cover its expansionary spending, constrained by what citizens would perceive as a tolerable burden[4]. Within the revenue-expenditure theories, this is termed as the "spend-and-tax" hypothesis, with clear policy implications: a spending restraint is needed in order to reduce public deficits.

A different approach is adopted in Friedman's (1978) analysis: in the tax-and-spend model, the government would spend all its revenues, and therefore higher taxes would increase public expenditure. In this case, imposing higher taxes to restrict the size of the government deficit would raise it instead.

[4] Notice that our paper considers a post-war sample. Therefore, we do not discuss Peacock and Wiseman's displacement effect.

The traditional empirical testing of the above models has been based on Granger's (1969) causality testing; the spend-and-tax model requires statistical significance of the lagged values of the general government expenditure in the revenue autoregression. On the other hand, the tax-and-spend model requires statistical significance of the lagged values of taxes in the expenditure autoregression, whereas simultaneity of the models requires the presence of both causality patterns. Our approach is based instead on the Johansen (1988, 1995) cointegration methodology testing for the possibility of weak exogeneity of the fiscal policy components and allowing for asymmetric and non-linear adjustment towards the long-run equilibrium.

3. The long-run model

We use a set of $p = 2$ endogenous variables, $y = [G, T]'$. G refers to general government expenditure (measured as the consolidated expenditure of state and local agencies) and T refers to general government revenues that include taxes, social security contributions and other entries (with limited importance). We use annual Italian data from 1957 to 1998 in log form. Nominal values are used, as it is difficult to identify an appropriate deflator for expenditures and taxes. Further, the possibility of contamination of the data from the use of an inappropriate deflator is eliminated when the variables are expressed in current prices (the choice of data in current prices is also discussed in Joulfaian and Mookerjee, 1990). Expenditures and revenues are not considered as a share of GDP. The reason is that the Italian *GDP* appears to be an $I(2)$ series whereas recursive OLS estimation of a bivariate model involving G/GDP and T/GDP revealed serious parameter instability in terms of forecast Chow tests (Figures are available by the authors upon request). The data set is taken from ISTAT, *Annuario Statistico Italiano* (various issues). The use of annual data is chosen on economic grounds given that important decisions on government spending and taxation are outlined in the annual *Legge Finanziaria e di Bilancio* (Finance and Budget Law). We believe that the use of higher frequency data, although useful from an econometric perspective since it increases the regression degrees of freedom, could undermine the economic rationale of the fiscal policy analysis.

Following Johansen (1988, 1995), we write a p-dimensional Vector Error Correction Model (VECM) in first differences as:

$$\Delta y_t = \Gamma_1 \Delta y_{t-1} + ... + \Gamma_{k-1} \Delta y_{t-k+1} + \Pi y_{t-1} + \mu + \varepsilon_t, t = 1,...,T \quad (1)$$

where y_t is a (p x 1) vector of the $I(1)$ variables discussed above, $\varepsilon_t \sim niid(0,\Sigma)$, μ is a drift parameter, and Π is a (p x p) matrix of the form $\Pi = \alpha\beta'$, where α and β are (p x r) matrices of full rank, with β containing the r cointegrating vectors and α carrying the corresponding loadings in each of the r vectors.

The two series are plotted in Figure 1. Preliminary analysis of the data using the Augmented Dickey-Fuller (ADF) tests suggested that both series are $I(1)$ in levels. The levels of the two equations in the unrestricted VAR model (1) are estimated using a lag length of $k = 2$ and allowing for the intercept term, μ, to enter the VAR model (1) unrestrictedly. [5] The determination of the number of cointegrating vectors is based on the maximal eigenvalue (λ-*max*) and the trace (λ-*trace*) statistics (see Johansen, 1988 and 1995). Cointegration results are shown in Table 1, which reports the λ_i eigenvalues, the λ-*max* and the *trace* statistics. To account for our small sample, both tests use a small sample correction (for exact mathematical formulas, see e.g. Doornik and Hendry, 2000, p.282). The 95 percent critical values are taken from MacKinnon *et al*., 1999). [6] Both the λ-*max* and the *trace* statistics support the existence of $r = 1$ cointegrating vector. Normalizing on G, we get the following long-run relationship:

[5] The lag length is selected by the Akaike Information Criterion (AIC). The diagnostic tests (available on request) suggest that both equations are free from normality failures, autocorrelation (up to order 2) and ARCH (up to order 1) effects. Estimations are done in PCFIML (Hendry and Doornik, 1997).

[6] These critical values are more accurate than those reported in Osterwald-Lenum (1992) and Johansen (1995). The reason is that contrary to previous studies, MacKinnon *et al*. (1999) use a higher number of simulations as

$G =$	$1.059T$
	(0.010)

where the number in parenthesis is the standard error of the estimate on T. The elasticity of G with respect to T is estimated at 1.059. The adjustment coefficients on G and T are estimated at 0.496 (standard error = 0.170), and 0.525 (standard error = 0.125), respectively. The estimate of the adjustment coefficient on G has the wrong sign; however it turns out to be insignificantly different from zero suggesting that G is weakly exogenous for the long run parameters. Testing weak exogeneity of G is based on a Likelihood Ratio (LR) test. This is distributed as a $\chi^2(1)$ under the null, giving a value of 6.95, which is just insignificant at 1 percent (p-value = 0.01). Imposing the above restriction, yields: $G = 1.040\ T$. Figure 2 plots the deviations from the estimated relationship.

We have also tested weak exogeneity of T. This is strongly rejected, giving a $\chi^2(1)$ value of 12.20 (p-value = 0.00). Noticing that the elasticity on T is close to one, we have also tested proportionality between G and T. This is rejected, giving a $\chi^2(1)$ value of 10.48 (p-value = 0.00). [7] Although a tax elasticity that exceeds one provides evidence against a balanced budget, the presence of cointegration points to a stable long-run co-movement of expenditures and revenues in the sense that G and T do not deviate too much from each other. This cointegrating equilibrium does not provide evidence in favor of a fiscal discipline in the EGSP sense; taxes will be lower than expenditures, resulting in a positive long-run deficit.

Further, we have the intriguing result that G is weakly exogenous; the relevant loading is insignificantly different from zero. In statistical terms, weak exogeneity of G implies that we

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well as response surfaces rather than discrete random walk approximations of the quantiles of the asymptotic distribution (for a detailed technical discussion see MacKinnon *et al.*, 1999).

[7] This result is in general consistent with Koren and Stiassny (1998) and Garcia and Hénin (1999) who examine the tax and spend relationship for a number of industrialized countries (including Italy) over the period 1953-1992 (annual data) and 1960-1996 (semi-annual data), respectively. In the case of Italy, they report non-stationarity of the $(G - T)$ to GDP ratio.

can proceed by estimating a short-run equation for taxes conditioning on government expenditure without any loss of information. In economic terms, weak exogeneity of government expenditure indicates the presence of asymmetries in the structure of the estimated model; it suggests that changes in taxes rather than changes in government expenditure equilibrate the system. Therefore, the short-run adjustment to correct budgetary disequilibria is mainly done by changes in the tax policy. This has to do with public expenditure rigidities not only due to multi-annual contracts and planning, but also due to strong resistance against expenditure reductions arising not only from the demand-side but also from bureaucratic powers (see for example the discussion in OECD, 1997, and Legrenzi and Milas, 2002). On the other hand, tax changes (tax increases in particular) can be supported by the presence of fiscal illusion, arising from the complexity of the fiscal system therefore resulting in several indirect taxes (see also the discussion in Puviani 1903, and Wagner 1976). [8] Given that changes in taxes equilibrate the system, the next section explores further the role of asymmetries only along the tax dimension; in particular, we examine whether positive versus negative and large versus small disequilibrium deviations have different effects on the short-run behavior of taxes.

Next, we discuss the persistence profile of the estimated cointegrating vector. The persistence profile analysis (see e.g. Pesaran and Shin, 1996), sheds some light on the speed of convergence of the estimated vector towards its long-run equilibrium following system-wide shocks. In particular, a system-wide shock refers to a shock drawn from the multivariate distribution of the error term in the VAR model (1) defined above, rather than a shock drawn from the distribution of the error term of a particular equation. If there is cointegration among a set of $I(1)$ variables, the impact of a system shock is only transitory and eventually disappears as the economy moves back to its steady state. By contrast, if there is no cointegration among these

[8] Using data for Greece, Hondroyiannis and Papapetrou (1996) also find exogeneity of government expenditure in a bivariate model of government spending and revenues.

variables, then the effect of a system-wide shock persists forever. In that sense, the persistence profile analysis provides complementary evidence that the estimated vector is indeed a cointegrating relationship. Furthermore, the persistence profile approach has the advantage of being invariant to the way shocks in the underlying VAR model are orthogonalized, and therefore provides an important extension to the traditional impulse response analysis, which is sensitive to the ordering of the variables in the VAR (see e.g. Lütkepohl, 1991).

Figure 3 shows the persistence profile for the estimated cointegrating relationship following system-wide shocks (these are implemented in Microfit 4.0, see Pesaran and Pesaran, 1997). The persistence profile of the cointegrating relationship converges to zero rather slowly. In particular, around 90 percent of the adjustment is made after around 3.5 years, and there is a very small impact of shocks on the cointegrating relationship even after 7-8 years. This sluggish rate of convergence towards equilibrium has to do with complex administrative procedures, problems in budget control, multi-annual contracting and planning and other rigidities in the functioning of the Italian government sector. These include a plethora of (often) uncoordinated laws that add to the sluggishness in the collection of taxes and undermine the tasks of public administration as well as an inadequate monitoring system of public spending; management accountability is underdeveloped therefore hampering the assessment of economic results. [9] In addition, frequent general elections have also undermined the control of public spending for electoral purposes, at least until the early 1990s when more prudent policies were put in place in order to meet the Maastricht (1992) convergence criteria (see e.g. OECD, 1997 and 2001)[10].

[9] As the OECD (1997) study for Italy points out, the number of laws in the early 1990s was estimated at between 100,000 and 150,000 for Italy, compared to 7,000 in France and 6,000 in Germany.

[10] It is interesting to notice that in the time span considered here, Italian political life saw 43 governments and 12 general elections.

4. The short-run model

4.1 Linear adjustment

Table 2 reports the OLS estimates of the error correction model for ΔT conditioning on ΔG which was found to be weakly exogenous. Lagged ΔG and ΔT regressors are dropped as they are insignificant (LR test of over-identifying restrictions: $\chi^2(2) = 0.32$ (p-value=0.85)). The estimated coefficient on the cointegrating vector (i.e. CV_{t-1} expressed in mean corrected form) suggests that approximately 33 percent of the disequilibrium error is corrected within a year by changes in taxes. Further, taxes are affected positively by current short-run increases in government expenditure. Overall, the positive effect from the disequilibrium error and government expenditure growth in the short-run equation for taxes provides empirical support for the spend-and-tax model according to which increased government spending leads to higher taxes. [11] The diagnostic tests in Table 2 show no evidence of misspecification for the estimated ΔT equation. To assess the impact of the Maastricht Treaty on the revenue-expenditure patterns, we also added in our model a Maastricht dummy variable, taking the value of 1 from 1993 onwards and 0 elsewhere. This turned out to be statistically insignificant, meaning that the adjustment of the Italian public finances did not take place through changes in the revenue-expenditure patterns. [12]

The stability of the model is analyzed using recursive OLS estimates. The plots (in Figure 4a) of the 1-step residuals $\pm$ 2*standard errors (SE) and the break point F-tests (N$\uparrow$ step Chow

[11] As our results suggest a feedback from ΔG to the ΔT model, we avoid reference to Granger causality running from government expenditure on taxes. The latter would suggest a feedback from lags of ΔG rather than ΔG itself on taxes.

[12] Italy qualified to the European Monetary Union although did not meet the debt criterion. The adjustment of the Italian economic policy variables was mainly achieved via monetary adjustments, although privatisations and a "European tax" were imposed on households.

tests) together with their 5 percent critical values provide evidence of parameter non-constancy (for a detailed discussion of these tests see Hendry and Doornik, 1997).

4.2 Asymmetric and non-linear adjustment

Considering that government expenditure is weakly exogenous, non-linearities along the tax dimension may arise if the fiscal authorities react differently to deviations of the government expenditure from its equilibrium level. For instance, the authorities may be more willing to raise taxes rapidly when the government expenditure is above its equilibrium rather than lowering taxes rapidly when government spending falls below its long-run level.

Various authors have examined non-linearities in the behavior of error correction models (see Escribano, 1986; Granger and Lee, 1989; Granger and Teräsvirta, 1993; Escribano and Granger, 1998; Escribano and Pfann, 1998; and Escribano and Aparicio, 1999, among others). For instance, Granger and Lee (1989) partition the error correction term into its positive and negative components, and feed them back into the short-run dynamic equation, whereas Escribano and Granger (1998) and Escribano and Aparicio (1999) use a cubic error correction term.

The second and third panel of Table 2 report the error correction equation based on different types of non-linear adjustment. First, as in Granger and Lee (1989), we take the deviations of CV_{t-1} around its mean value, and partition them into their positive and negative components (denoted by CV^{+}_{t-1} and CV^{-}_{t-1}, respectively). The results in the second panel of Table 2 indicate that the speed of adjustment varies depending on whether the estimated relationship is above or below its *unique* (at the zero point) equilibrium. The results suggest that when government expenditure is higher than it should be, taxes increase rapidly (i.e. coefficient on $CV^{+}_{t-1} = 0.490$; t-ratio = 3.554). On the other hand, when government expenditure is below than it should be, taxes fall slowly (i.e. coefficient on $CV^{-}_{t-1} = 0.172$; t-ratio = 1.245). Given the relatively high number of regressors in our model and the small size of our sample, we test equality of the coefficients on CV^{+}_{t-1} and CV^{-}_{t-1} based on a bootstrapped *p*-value rather than an asymptotic *p*-value although our results are not sensitive to the above choice. [13] The bootstrapped *p*-value of the F test is equal to 0.14, which in statistical terms shows very weak evidence of different effects from CV^{+}_{t-1} and CV^{-}_{t-1}. Hence, in economic terms our results points to downward inflexibility of taxes; the fiscal authorities react to an increase in government expenditure above its equilibrium level by raising taxes quickly. On the other hand, they are less willing to lower taxes rapidly when government expenditure is below its equilibrium level. Nevertheless, in terms of statistical tests, these results are less clear.

In the formulation of asymmetric error correction models it is important to identify the number of equilibria. Assuming a unique equilibrium (around zero) in our analysis above, may be too restrictive. In order to test this, we estimate a cubic error correction model as suggested by Escribano and Granger (1998) and Escribano and Aparicio (1999). More specifically, we allow for CV^{2}_{t-1} and CV^{3}_{t-1} to enter the short-run equation. This type of non-linear adjustment allows for a faster adjustment when deviations from the equilibrium level get larger. Further, it allows for the possibility of more than one equilibrium points when the additional regressors, that is, CV^{2}_{t-1} and CV^{3}_{t-1} are statistically significant. In that sense, the cubic error correction model is more flexible than the Granger and Lee (1989) type of asymmetric adjustment. The bootstrapped *p*-value of the F test (using 10,000 replications based on the method discussed above) for the joint significance of the CV^{2}_{t-1} and CV^{3}_{t-1} regressors is equal to 0.08, indicating evidence of this type of non-linear adjustment at the 10% level of statistical significance.

[13] Bootstrapping was done in Gauss. To compute the bootstrapped *p*-values of the Lagrange Multiplier F tests reported in the second and third panel of Table 2, we implemented the following procedure. First, we estimated the linear error correction model for taxes. Then, draws were taken from the residuals and 10,000 artificial data series were constructed. For each of these artificial series, F statistics were constructed and then compared to the corresponding statistic from the actual data. The bootstrapped *p*-values were derived as the number of times the F statistics from the artificial data exceeded the corresponding statistics from the actual data, divided by 10,000.

The estimates of all other regressors in the asymmetric and non-linear models are very similar to those of the linear model in the first panel of Table 2. The error variance ratio of both the asymmetric and non-linear models relative to the linear one (i.e. σ^2_{AS}/σ^2_L and σ^2_{NL}/σ^2_L) is less than one, indicating that the asymmetric and non-linear models have a better fit. In particular, the ratio shows a reduction in the residual variance of the asymmetric compared to the linear model by 4 percent and by around 12 percent for the non-linear model. Further, a comparison of Figure 4a with Figure 4b suggests that there are significant gains in terms of parameter constancy tests for the non-linear error correction model relative to the linear one (this is also the case for the asymmetric model). All other diagnostic tests (i.e. autocorrelation, ARCH and normality tests) are almost identical to those given in the first panel of Table 2 (and for this reason not reported here).

Figure 5a and Figure 5b plot the asymmetric and non-linear adjustments against the cointegrating vector, respectively. Figure 5a provides evidence of asymmetric adjustment as the cross-plot is rather far from being a straight line. Given the weak evidence of additional explanatory power from the CV^2_{t-1} and CV^3_{t-1} regressors, the plot of the non-linear cubic polynomial in Figure 5b also suggests that a unique equilibrium point (at zero) exists for the ΔT equation. Therefore, the non-linear adjustment results support the findings of asymmetric adjustment around a unique equilibrium. [14]

5. Conclusions

In this paper we looked at the relationship between government expenditure and revenues in the case of Italy. Using the Johansen (1988, 1995) cointegration methodology, a long-run relationship between government expenditure and revenues is shown to exist. The persistence

[14] Figures 5a,b are upward sloping, whereas the Figures reported in Escribano and Granger (1998) and Escribano and Aparicio (1999) are downward sloping. This is simply due to the fact that in our equation the error correction term has a positive coefficient because of the normalization we adopted.

profile analysis suggests a sluggish rate of convergence towards equilibrium mainly due to complex administrative procedures and the existence of a complicated legal system that adds to the sluggishness in the collection of taxes and undermines the effective monitoring of public spending. The structure of the estimated model suggests that the short-run adjustment to correct budgetary disequilibria is mainly done by changes in taxes rather than changes in government expenditure. In addition, short-run increases in government expenditure lead to higher taxes, which provides support for the spend-and-tax hypothesis. Finally, we considered the case of asymmetric and non-linear adjustment also testing for the possibility of multiple equilibria. According to our results, there is evidence in favor of asymmetric adjustment around a unique equilibrium (at zero) rather than multiple ones, suggesting that when government expenditure is too high, adjustment of taxes takes places at a faster rate than when government expenditure is too low. Further, there is evidence of a faster adjustment when deviations from the equilibrium level get larger, pointing to a Leviathan style revenue-maximiser government.

Taking into account that taxation carries the burden of correcting budgetary disequilibria, a simplified structure of the tax system would increase the speed with which deviations from the estimated long-run spending - revenues relation are eliminated, although it would off-set the fiscal illusion. However, progress towards fiscal consolidation cannot rely solely on tax pressure due to possible distortionary effects from the latter. Tighter government spending underpinned by a more effective monitoring system of government activities would also reduce the persistency of fiscal imbalances.

Bearing in mind that the class of non-linear models is infinite, we view the empirical results discussed in this paper as a promising step towards investigating the role of non-linearities in the area of public finance. It is notable that commenting on Ericsson *et al.*'s (1998) UK money demand model, Teräsvirta (1998) pointed out that non-linear models with quadratic and cubic error correction terms are first-order approximations to smooth transition autoregression models (STAR; see e.g. Granger and Teräsvirta, 1993), where the transition mechanism between different regimes is driven by the disequilibrium error. In the context of our public finance bivariate model, it would be interesting to formally introduce and estimate a two-regime STAR model where adjustment takes place in every period but the speed of the adjustment (as well as the impact of the lagged values of expenditure and taxes) vary conditional on whether disequilibrium deviations from the expenditure/taxes relationship are large (call this regime 1) or small (call this regime 2). [15] We intend to address these issues in future research.

[15] Testing for unit root behavior in the US public debt, Sarno (2001) models the growth of the debt-GDP ratio conditional on large and small values of the level of the (lagged) debt-GDP ratio.

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TABLE 1
Eigenvalues, test statistics, and critical values

λ_I	*λ-max*				*λ-trace*			
	H_0	H_1	Stat.	95%	H_0	H_1	Stat.	95%
0.338	$r = 0$	$r = 1$	16.51	15.0	$r = 0$	$r \geq 1$	20.38	18.1
0.092	$r \leq 1$	$r = 2$	3.8	8.2	$r \leq 1$	$r = 2$	3.8	8.2

Notes: r denotes the number of cointegration vectors. The λ-*max* and *trace* statistics use a small sample correction. The critical values are taken from MacKinnon *et al*. (1999).

TABLE 2

OLS estimates of the error correction model

Variable	ΔT equation Coeff.	*t*-ratio
Constant	0.070	5.125
ΔG_t	0.457	5.356
CV_{t-1}	0.331	4.142
Diagnostics		
σ_L	0.052	
$\overline{R_L^2}$	0.669	
F*ar* (2, 35)	0.33[0.72]	
F*arch* (1, 35)	0.92[0.34]	
$\chi^2 nd$ (2)	3.51[0.17]	

Asymmetric adjustment

Variable	ΔT equation Coeff.	*t*-ratio
CV^+_{t-1}	0.490	3.554
CV^-_{t-1}	0.172	1.245

$\sigma_{AS} = 0.051$, $\sigma^2_{AS}/\sigma^2_L = 0.960$

$\overline{R_{AS}^2}$ 0.678

F (1, 36) test of equal effects from CV^+ and CV^-:

1.94 [0.14]*

Non-linear adjustment

Variable	ΔT equation Coeff.	*t*-ratio
CV_{t-1}	0.127	0.993
CV^2_{t-1}	0.855	2.094
CV^3_{t-1}	3.662	2.052

$\sigma_{NL} = 0.049$, $\sigma^2_{NL}/\sigma^2_L = 0.880$

$\overline{R_{NL}^2}$ 0.698

F (2, 35) test of zero effects from CV^2 and CV^3:

2.74 [0.08]*

Notes: F*ar* is the Lagrange Multiplier F-test for 2nd order residual serial correlation. F*arch* is the 1st order ARCH F-test. $\chi^2 nd$ is a Chi-square test for normality. Numbers in parentheses are the degrees of freedom of the tests. Numbers in square brackets are the *p*-values of the tests. $\overline{R^2}$ is the adjusted coefficient of determination and σ refers to the standard error of the regression.

CV = $G - 1.04\,T$, in mean corrected form.

*Bootstrapped *p*-values in square brackets. The *p*-values are derived from bootstrapping with 10,000 replications.

Figure 1: Plots of the levels of the series

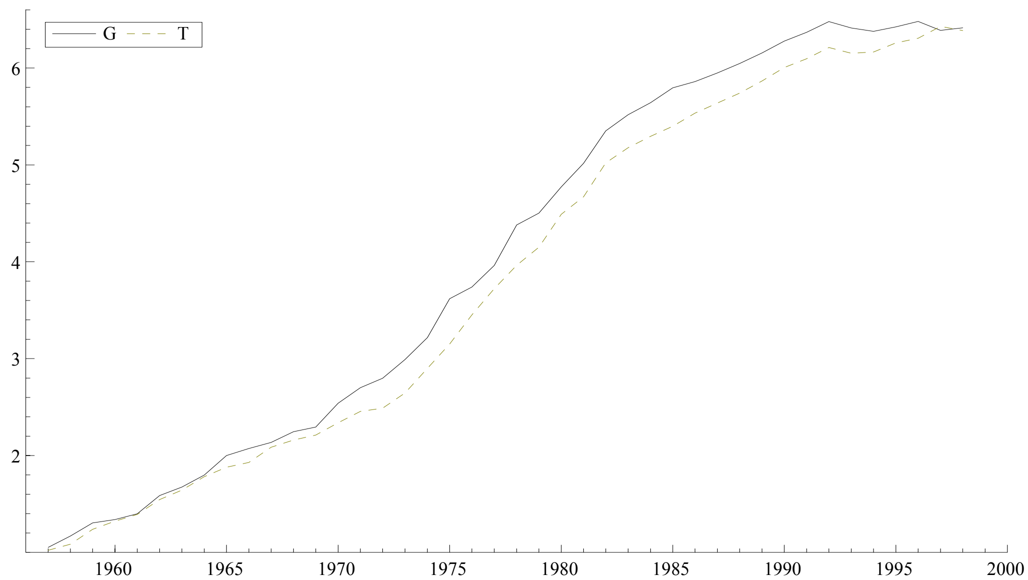


Figure 2: Long-run relationship: CV = $G - 1.040\, T$

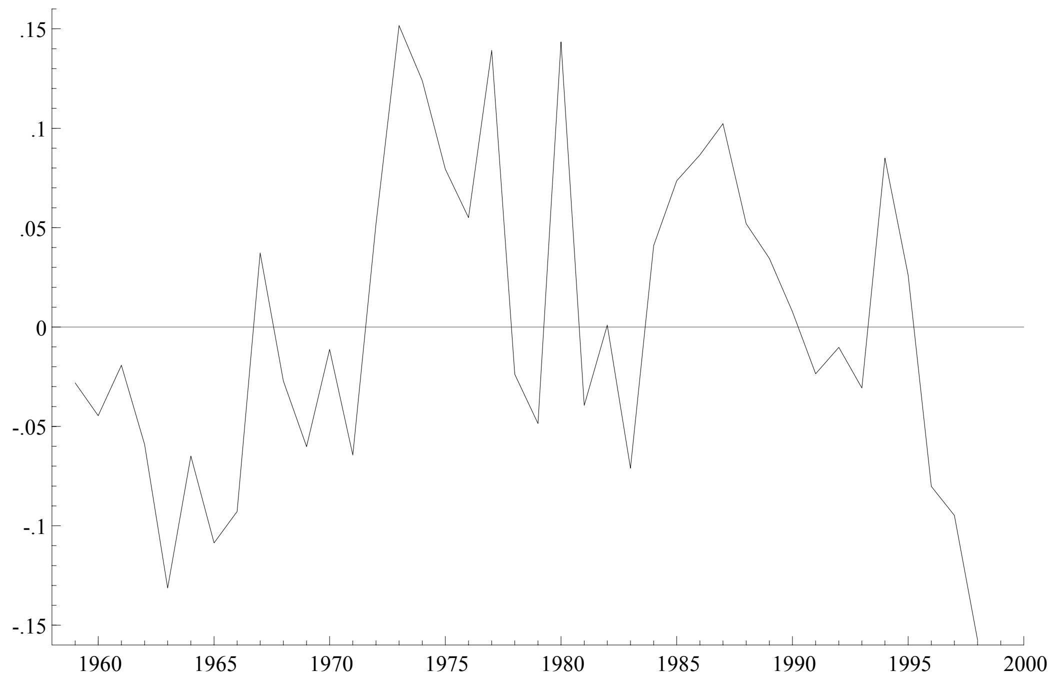


Figure 3: Persistence profile of the cointegrating vector to system-wide shocks

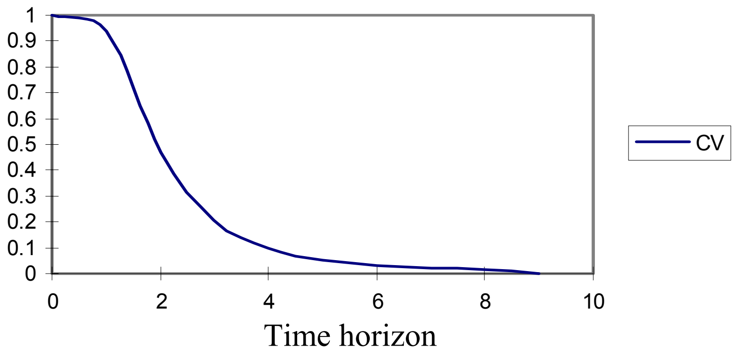
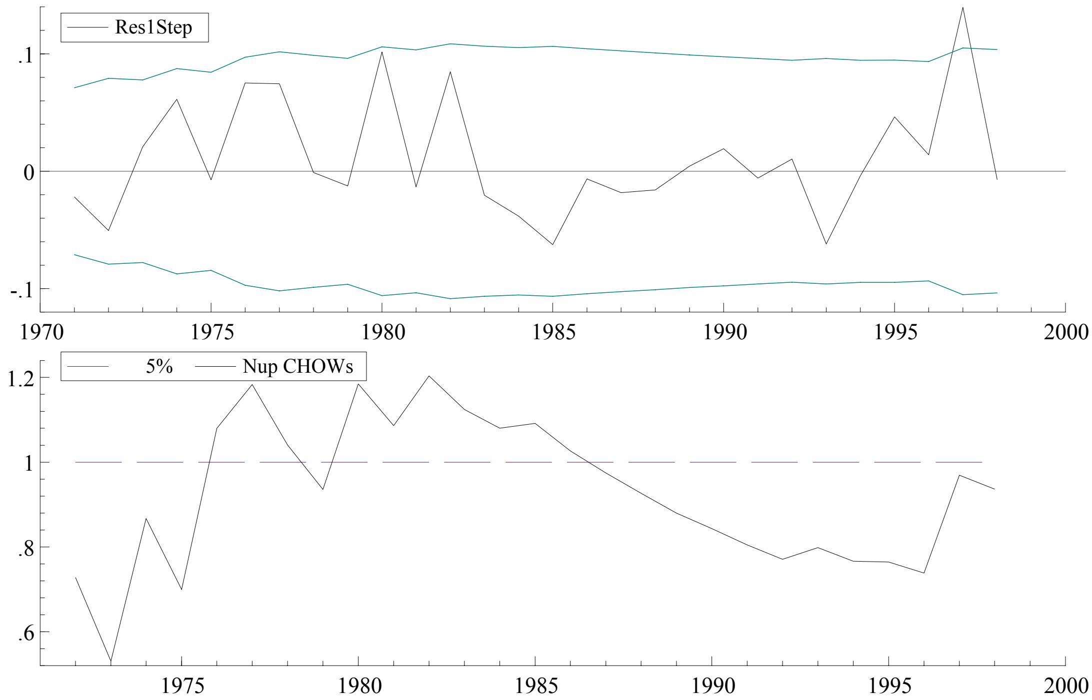
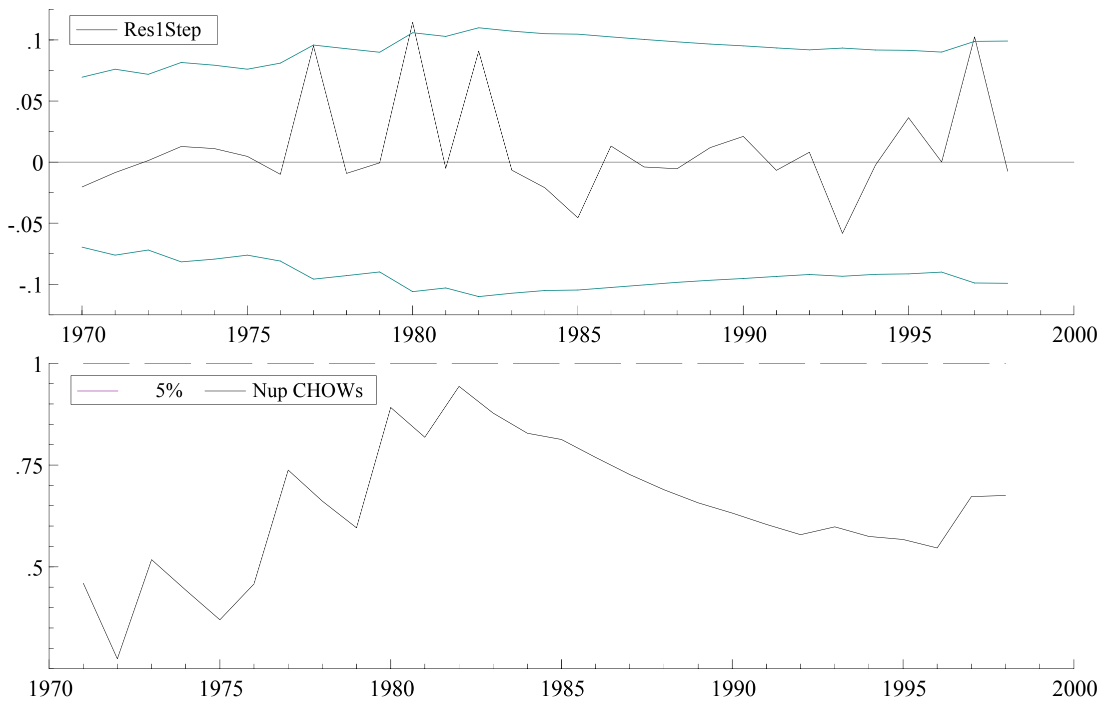


Figure 4: Parameter constancy tests for the error correction models

(a) Linear error correction model



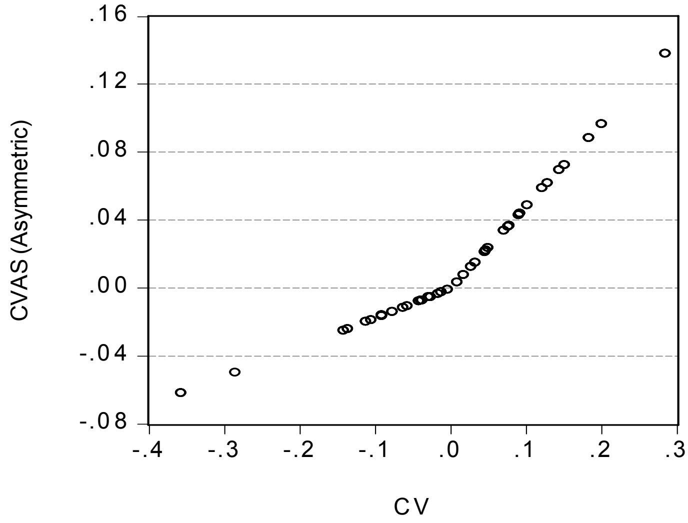
(b) Non-linear error correction model



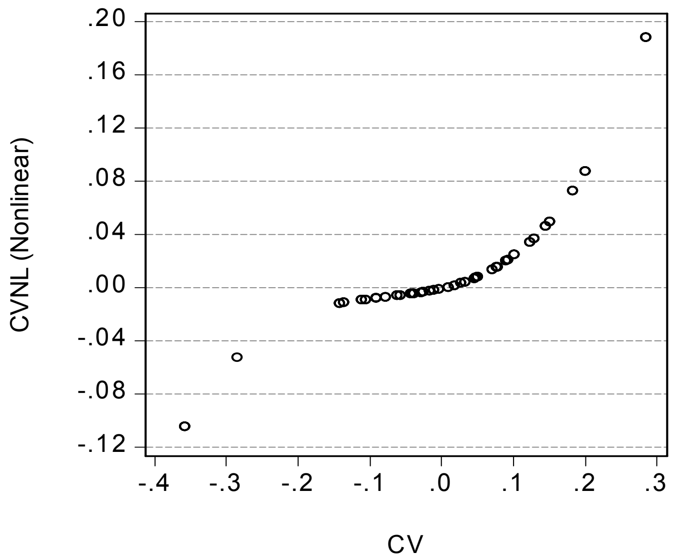
Notes: Res1Step : 1-step residuals ± 2 standard errors
Nup CHOWs: Forecast Chow tests and 5% critical value

Figure 5: Asymmetric and Non-linear adjustment

(a)



(b)



Notes: $CVAS = 0.490CV^{+} + 0.172CV^{-}$, $CVNL = 0.127CV + 0.855CV^{2} + 3.662CV^{3}$