Estimating the effects of fiscal policy under the budget constraint

Peter Claeys[a]
Universitat de Barcelona, Research group AQR-IREA[b]

ABSTRACT: I reconsider the short-term effects of fiscal policy when both government spending and taxes are allowed to respond to the level of public debt. I embed the long-term government budget constraint in a VAR, and apply this common trends model to US quarterly data. The results overturn some widely held beliefs on fiscal policy effects. The main finding is that expansionary fiscal policy has contractionary effects on output and inflation. Ricardian effects may dominate when fiscal expansions are expected to be adjusted by future tax rises or spending cuts. The evidence supports RBC models with distortionary taxation. We can discard some alternative interpretations that are based on monetary policy reactions or supply-side effects.

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1. INTRODUCTION

Drafting of the annual government budget suscite lively debates both in parliament and the public at large. The reason is that shifts in fiscal policy have considerable implications for the allocation of public goods, the redistribution of resources between different groups and the stabilisation of the aggregate economy. There are also widely different opinions about fiscal policy in the economics profession. There are few definite conclusions on the economic effects of changes in fiscal policy. Different theoretical models not only predict a different persistence or magnitude, but also result in a different sign of the economic responses.[1] Empirical evidence does not provide unanimous results either. The lack of an empirical consensus owes to the problems in the identification of exogenous shifts in fiscal policy. There are basically four empirical problems with fiscal VARs. First, the government has available plenty of different fiscal instruments. The aggregation of the scattered use of many spending and tax tools in a summary fiscal shock mutes the economic responses. Second, the public can easily anticipate policy changes because of the long time lags between the announcement and the implementation of fiscal packages. This makes it hard to pin down the exact timing of a fiscal shock. Third, fiscal variables move with the economic cycle, mainly due to automatic stabilisers. Finally, fiscal policy is subject to a budget constraint. Hence, in addition to the response of spending or taxes to economic variables (output, interest rates, inflation) there is a feedback effect of public debt. Spending will need to be cut or taxes raised if the government is to keep public debt under control.

The reason for the widely different empirical effects of fiscal policy mainly owe to different assumptions in the identification of the fiscal shocks. Empirical fiscal SVAR studies have typically ignored public debt. This can have quite dire consequences, especially in models that incorporate forward-looking economic behaviour (Christ, 1979; Sims, 1998). Constraints that are believed to work out in the long-term also have their repercussions for expected adjustments in fiscal policy in the short-term. These may overturn the perceived wisdom about the economic effects of fiscal policy (Gordon and Leeper, 2005). In this paper, I propose to directly incorporate the budget constraint in the empirical model. The structural vector error correction model captures the long-term relationship between

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[1] See Perotti (2007) for a comparison of the different theoretical models, and their empirical predictions.

spending and revenues: shocks to this equilibrium and the ensuing adjustment are responsible for short-term fluctuations around some common trends, which are driven by the shocks with permanent effects. This approach has two benefits. First, the budget constraint implies some identifying restrictions on the behaviour of fiscal policy that are directly taken from theory. I.e., the fiscal adjustment on the spending or revenue side to keep the budget in balance over time are directly incorporated in the VAR. Second, I take an explicitly intertemporal view on the response of economic agents to the changes in fiscal policy.

I apply the common trends model on US quarterly data since 1965. The main finding is the contractionary effect on output and inflation of fiscal expansions in the short run. This is consistent with a Real Business Cycle model of fiscal policy that displays period-by-period budget balance, with distortionary taxation. The results can alternatively be read as a negative Ricardian effect of fiscal policy due to expectations of fiscal adjustments by future tax rises or spending cuts that restore fiscal sustainability (Gordon and Leeper, 2005). Policy interaction might give an additional argument for reversing the usual crowding in effects of fiscal policy (Linnemann and Schabert, 2003). I indeed find a contractionary response of the Federal Reserve to fiscal expansions of the Treasury. The slight contraction of moneary policy is unlikely to explain the large size of the output contraction, however. Conversely, monetary policy shocks do not elicit a direct fiscal response. I also find government spending to have substantial supply-side effects in the long term.

The plan of the paper is as follows. In section 2, I review the findings of the fiscal VAR literature in the light of the sustainability of public debt. This implies also a discussion of the interaction of fiscal variables with monetary policy. The econometric set up of this common trends model, its identification and the interpretation of the structural shocks are detailed in section 3. I then analyse the effects of fiscal policy in section 4. I conclude in section 5 and offer some suggestions for further research.

2. THE EFFECTS OF FISCAL POLICY, AND PUBLIC DEBT

Ricardian Equivalence is a reasonable starting point for the theoretical analysis of the effects of fiscal policy (Barro, 1974). Few economists would endorse it as a realistic description of fiscal policy, however. The view that private savings fully offset the change

in public savings is not based on a firm empirical rejection of Ricardian Equivalence, for this hypothesis is not directly testable. Many empirical studies have therefore examined the alternative hypothesis as to whether fiscal policy has any real economic effects. Results widely differ in the size and sign of the effects of government spending and taxes, but all studies would agree that the financing decisions of government are not neutral for the real economy. Although public debt is crucial to the Ricardian model, fiscal VARs have typically not paid much attention to the government budget constraint.

The period-by-period dynamics of total debt b_t are given by the accumulation of interest payments on past fiscal imbalances and the current primary surplus s_t, which is the difference of government revenues t_t and government spending g_t.

$$b_t \equiv (1 + r_t)b_{t-1} - s_t \qquad \text{where } s_t = t_t - g_t. \qquad (1)$$

Solving forward (1), one obtains the intertemporal budget constraint IGBC (2).

$$b_t = \lim_{n \to \infty} E_t \left[\prod_{j=1}^{n} (\frac{b_{t+n}}{1 + r_{t+j}}) \right] + \sum_{n=0}^{\infty} E_t \left[\prod_{j=1}^{n} (\frac{g_{t+n} - t_{t+n}}{1 + r_{t+j}}) \right]. \qquad (2)$$

The sustainability condition is met when the public sector does not leave any public assets or liabilities with positive probability. For this to hold, the transversality condition needs to be satisfied, i.e. when

$$\lim_{n \to \infty} E_t \left[\prod_{j=1}^{n} (\frac{b_{t+n}}{1 + r_{t+j}}) \right] = 0. \qquad (3)$$

Some equivalent tests for fiscal sustainability can then be derived under various assumptions on this condition. Cointegration between non-stationary government expenditures g_t and revenues t_t can be shown – under some weak economic assumptions – to be a necessary condition for the IGBC to hold. Conversely, this cointegration relation is also a sufficient condition under weak assumptions on the DGP of debt, viz. that it is an I(1) process (Ahmed and Rogers, 1995). It then follows that the total government deficit series is stationary when spending and revenues are cointegrated.[2] Alternatively, cointegration between the primary surplus s_t and public debt b_t also implies that public debt is sustainable. This relation can also be rewritten in terms of a reaction function (4):

$$s_t = \rho b_t + \mu_t \qquad\qquad\qquad\qquad (4)$$

Bohn (1998) proofs that a strictly positive reaction of the primary surplus to public debt in this 'fiscal rule' is a robust sufficient condition for the sustainability of public finances. Most of the empirical tests of fiscal sustainability have been based on the (co)integration implications of the transversality condition, or on tests of fiscal rules.

The analysis of public debt in fiscal VARs has commonly been limited to an analysis of the joint VAR process of surplus and debt. Roberds (1991) and Carriero *et al.* (2005) develop a test of fiscal sustainability based on the violation of the cross-equation coefficient restrictions that the IGBC imposes on the joint DGP of debt and surplus (Hansen *et al.*, 1991). Other studies develop a test for the Fiscal Theory of the Price Level. According to this theory, the IGBC is not a constraint but rather a value equation. If the government fails to take actions to ensure its budget constraint is satisfied, equilibria are possible where fiscal – rather than monetary – policy determines the price level (Leeper, 1991; Sims, 1994). The possibility of such a non-Ricardian regime arises as government solvency eventually has to be ensured in real terms. If fiscal policy is sufficiently reactive to debt, however, then the IGBC will be satisfied for all price paths. Monetary policy retains the ability to control prices in this Ricardian environment. Canzoneri *et al.* (2001) distinguish Ricardian from non-Ricardian regimes in fiscal policy by examining the sign of serial correlations in deficits following shocks to debt in a bivariate VAR.[3]

In contrast, VAR studies that are interested in the economic effects of fiscal policy have typically ignored the path of public debt. This has two serious consequences (Sims, 1998; Favero and Giavazzi, 2007). First, the VAR coefficients are typically biased. The economic shocks that have been identified as structural still include a response to the future changes in spending and taxation that are needed to keep the level of the public debt in check.[4] Second, due to this bias, the interpretation of the results of usual fiscal VARs in terms of a corresponding DSGE model could be invalid. The paths of government spending and taxes may imply unstable dynamics for the public debt ratio.

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[2] Such 'bounded' sustainability implies that undiscounted public debt is finite in the long run. With an unspecified cointegrating vector, only 'weak' sustainability can be said to hold (Quintos, 1995).
[3] Sala (2004) finds evidence for different regimes in which the FTPL is relevant. He compares the responses to a tax shock in an extended FTPL-model with the impulse responses from a VAR.

Depending on the time path of financing, economic effects will wind out over shorter or longer periods, but as Ricardian Equivalence applies, all variables eventually return to steady state in the long term.[5] The economic interpretation of fiscal policies that lead to maintaining sustainability over time has governments adjust spending or taxes (or both) so as not to let debt explode. I incorporate the long-term budget constraint in a standard fiscal VAR, and employ a structural vector error correction model to estimate the economic effects and the debt stabilising (error correction) changes in fiscal policy.[6] I account for economic conditions by including output and inflation. A control for the endogenous reaction of interest rates is necessary, given the crucial role of monetary policy in the joint determination of sustainable fiscal plans. Omission of either fiscal or monetary policy when interaction matters may bias the findings of analysing fiscal or monetary policy in isolation.[7],[8]

3. THE SVAR-COMMON TRENDS ANALYSIS

3.1. The common trends methodology

A common trends model decomposes time series into trends and stationary variables. Denote by X_t a vector of n integrated time series, and decompose these series in some permanent trends X_t^P and a stationary residual component X_t^T.

$$X_t \equiv X_t^P + X_t^T. \tag{5}$$

Neither component in (5) is observable as such. We can give content to the permanent and transitory components in the following way (Stock and Watson, 1988).[9] There are

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[4] Favero and Giavazzi (2007) show how including the stock of public debt produces coefficient estimates that are signficantly different from a usual SVAR.
[5] Public debt can be non-neutral in the short-term in DSGE models if government bonds serve transaction purposes (Schabert, 2004).
[6] Bohn (1991) considers spending cuts or tax increases in a VEC model that includes fiscal series only.
[7] Different theories of policy interaction highlight the role of interest rates and inflation for the path of public debt (Sims, 1998; Dixit and Lambertini, 2003).
[8] Time series tests for sustainability based on the transversality condition are not very informative about which fiscal policies are sustainable (Bohn, 2007). We need economic assumptions that imply more stringent bounds on the feasible path of fiscal variables, such that the usual necessity conditions for cointegration apply. Embedding the budget constraint in an economic model is an example of an 'economic' test of fiscal sustainability.
[9] King *et al.* (1991) apply the common trends model to analyse the long and short-term properties of economic growth in the US. Other applications are in the monetary field (Bagliano *et al.*, 2002; Englund *et al.*, 1994; Jacobson *et al.*, 2002) or in the transmission of shocks in open economies (Villani and Warne, 2003; Corsetti and Konstantinou, 2005).

stochastic trends underlying the data series that are drive by some shocks with persistent effects. As there are usually fewer of these trends than variables in the model, the steady state relationships determine some stationary fluctuations around the stochastic trends. In formal notation, the existence of r cointegrating relationships between the n variables in X_t allows us to extract the permanent components, which are driven by $k \equiv n - r$ common stochastic trends. The stationary fluctuations around these trends are given by the r cointegrating relationships. The permanent shocks do not only determine the long-term properties of the series, but can have transitory effects as well.

An equivalent way to write X_t then is (6)

$$X_t = X_0 + F\tau_t + \Phi(L)v_t \qquad v_t \approx WN(0, I_n), \qquad (6)$$

where (a) $\Phi(\lambda)$ is assumed finite for all λ in and on the unit circle; (b) X_0 is stationary; and (c) v_t is white noise. The trend component is then described by $F\tau_t$, where F is the loading matrix of dimension $n \times k$ and rank k on the permanent components. The trend τ_t is a random walk with drift μ and innovation φ_t (7):

$$\tau_t = \mu + \tau_{t-1} + \varphi_t \qquad \varphi_t \approx WN(0, I_k). \qquad (7)$$

The common trends model then decomposes the time series X_t, as in (8):

$$\begin{cases} X_t^T = X_0 + \Phi(L)v_t \\ X_t^P = F[\tau_0 + \mu t + \sum_{i=1}^{t} \varphi_i] \end{cases} \qquad (8)$$

When the number of k common trends is less than the number n of variables, there are exactly $r = n - k$ linearly independent cointegrating vectors. Let these vectors be collected in $\in$. As these are orthogonal to the loading matrix F, the process $\beta'X_t$ is jointly stationary.

In particular, when X_t is generated by a VAR with lag length p as in (9)

$$A(L)X_t = \rho + \varepsilon_t \qquad \varepsilon_t \approx WN(0, \Sigma), \qquad (9)$$

and if X_t is cointegrated of order (1,1) with r cointegrating vectors, then we know from the Granger Representation Theorem that $rank[A(1)] = r$ and $A(1) = \alpha\beta'$, with α the

loading matrix of adjustment coefficients on the r cointegrating vectors. Equation (9) can then be rewritten as a VECM (10):

$$A^*(L)\Delta X_t = \rho - \alpha\beta'X_{t-1} + \varepsilon_t \qquad \text{with} A^*(\lambda) = I_n - \sum_{i=1}^{p-1} A_i^* \lambda_i, A_i^* = -\sum_{j=i+1}^{p} Aj. \qquad (10)$$

If we rule out orders of integration larger than unity, there also exists a Wold VMA representation of (10) as in (11)

$$\Delta X_t = \delta + C(L)\varepsilon_t, \qquad (11)$$

This can be rewritten as a common trends model similar to (8), if we let $C(\lambda) = C(1) + (1-\lambda)\tilde{C}(\lambda)$ and assume the sum of coefficients $C_i^* = \sum_{j=i+1}^{\infty} C_j$ $(i \geq 0)$ in $C^*(\lambda) = \sum_{i=0}^{\infty} C_i^* \lambda_i$ to be absolutely summable. In that case we obtain the following common trends representation of the VAR(p) process:

$$\begin{cases} X_t^T = X_0 + \tilde{C}(L)\varepsilon_t \\ X_t^P = C(1)[\xi_0 + \rho t + \sum_{i=1}^{t} \varepsilon_i] \end{cases} \qquad (12)$$

This is nothing else than the Beverdige-Nelson decomposition of X_t. In (12), $C(1)$ has reduced rank k under the assumption of r cointegrating vectors. I.e., only k elements of $C(1)\varepsilon_t$ result in independent permanent effects on X_t. As the loading matrix F is not orthogonal to the elements of $C(1)\varepsilon_t$, permanent shocks also have transitory effects on X_t. In contrast, transitory shocks have an effect on the temporary component only. Hence, short-term dynamics are determined by all innovations in the system. Warne (1993) derived a general estimation strategy of the common trends model based on (10) and the Wold VMA representation as in (11).[10]

Identification of the permanent and transitory shocks requires some assumptions on the long-term behaviour of the model. Call Γ any regular matrix of dimension n such that $\Gamma\Sigma\Gamma'$ is diagonal. Then $R(1) \equiv C(1)\Gamma^{-1}$ is the total impact matrix to innovations in the long term. Let $\eta_{i,t}$ be the i-th component of the vector $\Gamma\varepsilon_t$. The matrix Γ is said to

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[10] Warne (1993) moreover derives the asymptotic impulse response functions under the hypothesis of a

identify the common trends model if: (a) Γ is uniquely determined from the parameters in (10); (b) the covariance matrix of $\Gamma \varepsilon_t$ is diagonal with non-zero diagonal elements; and (c) the total impact matrix $R(1) = [\ F\ \vdots\ \mathbf{0}\ ]$. The innovation is categorised as transitory (*permanent*) if column $i \in \{1,...,n\}$ of $R(1)$ is (non-)zero. Permanent innovations are thus associated to the k common trends in F. Hence, the reduced form VMA representation in expression (11) is equivalent to the structural model (13)

$$\Delta X_t = \delta + R(L)\eta_t, \tag{13}$$

where η_t contains the serially uncorrelated structural disturbances with mean zero and with covariance matrix I_n. This then allows retrieving the permanent and transitory component of each series in X_t.

In practice, after establishing the cointegrating rank r we determine the cointegrating vectors β. These may either be obtained via the usual ML estimation techniques or can be directly imposed from the steady state properties of some economic theory. This associates the cointegrating vectors to the r transitory innovations and thus imposes $n \times r$ identifying restrictions. In a second step, we calculate the matrix of common trend parameters using the orthogonality of the cointegrating vectors to the permanent components. Following King *et al.* (1991), we may write $F = F_0 \pi$, where F_0 is an $n \times k$ matrix chosen so as to have $\beta' F_0 = 0$. The latter restriction results in a further $r \times k$ restrictions. However, these restrictions do not attribute any particular economic meaning to the k trends. We therefore need $k \times k$ additional assumptions to isolate k unique (and economically interpretable) trends. Assuming that the permanent shocks are uncorrelated and satisfy a (choleski) ordering on their reciprocal influence,[11] gives us a further $\frac{k(k+1)}{2}$ restrictions. The choice of π does not imply recursiveness for the total loading matrix F, as F_0 actually determines how the trends influence X_t. Finally, at least $\frac{k(k-1)}{2}$ additional constraints on the effect of permanent shocks on the variables included in the model need be motivated by economic theory. Estimation then proceeds on the VECM (10).

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known finite upper bound p on the lag order of the VAR.
[11]That is, π is lower triangular.

3.2. Specification and identification

I specify a VAR in the (log) levels of real GDP (y), inflation (p), a short-term nominal interest rate (i) and the (log) levels of real government expenditure (g_t) and revenues (t_t). The model for the DGP can be written in VECM form as follows, with m the number of lags:

$$\Delta X_t = \alpha \beta' X_{t-1} + A_1^* \Delta X_{t-1} + ... + A_s^* \Delta X_{t-m} + \varepsilon_t .$$
$$\text{where} \quad X_t = [\, y \ \ p \ \ i \ \ g \ \ t \,]'$$

(14)

I assume three cointegration relations may be present in the DGP of X_t, each of which I associate with the temporary shocks in η_t. With $n = 5$ variables and $k = 3$ cointegration relations, we thus specify 15 parameters in the cointegration vectors. A further six restrictions come from the orthogonality of the three cointegration vectors to the $r = 2$ common trends. The choleski ordering of uncorrelated permanent shocks gives us three more restrictions. This leaves one further restriction to be imposed. In accordance with the literature, I interpret the two permanent components as a real and a nominal trend respectively (King *et al.*, 1991). I distinguish both by assuming that the nominal trend has no long-term effect on real output. The real trend captures a technology shock with a permanent effect on real variables. The nominal trend instead captures any shock with long-term effect on nominal variables.[12] This provides us with a decomposition of the series into permanent and transitory components that carry an economic interpretation within the model.[13]

The fiscal 'solvency' shock ($g_t - t_t$)

I impose the cointegration relation between government spending g_t and revenues t_t as a necessary and sufficient condition for the intertemporal government budget constraint to hold. The structural innovation to the budget constraint is our fiscal shock. I do not impose strong sustainability on the cointegrating relation between expenditures and revenues, but instead leave the cointegration coefficients unspecified. Also, I include interest payments

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[12] This models any longer term shifts in monetary policy. No effect of the nominal trend on either spending or taxes is assumed.

in government spending and thus implicitly impose a unity coefficient on this variable.

This fiscal shock is not directly comparable to the ones commonly identified in fiscal VARs for various reasons. First, these studies examine the effect of unanticipated changes in discretionary policy, but disregard the ensuing adjustment in expenditures or revenues to maintain solvency. I.e., no attention is paid to the financing of the budget. It can be shown in basic RBC models that the effects of fiscal policy can be reversed depending on deficit or tax financing of budget expansions (Baxter and King, 1993). In my model, future fiscal adjustments constrain the short-term effects of current changes in fiscal policy.[14] Consider a hike in public spending, for example. For a given level of tax revenues, this leads – *ceteris paribus* – to a deficit and higher public debt. This requires a future rise in tax revenues (or a cut in spending). The anticipation of these offsetting policies has an impact on current economic variables, and these in turn determine the dynamic path of the fiscal variables. Second, the fiscal 'solvency' shock cannot be attributed to spending or revenues. The same example can be applied to a cut in taxes, for a given level of spending. I only identify deviations from the long-term budget balance. Finally, the fiscal shock is supposed to have no long-term effect on other variables. This is consistent with standard theories of fiscal policy and public debt where all variables eventually return to steady state. In practice, both government spending and taxation can have permanent long-term economic effects. In my model, these effects work out via the common trends and affect the dynamic path of the economy. For example, if the hike in public spending falls on (productive) government investment, this would likely boost long-term economic growth. Stronger potential growth enlarges tax bases and thus boosts revenues in the long term. This makes consolidation of public debt feasible without strong adjustments – and hence modifies future responses –in taxes or spending.

The business cycle shock ($y_t + \theta t_t$)

some of the variation in the budget owes to endogenous cyclical variations in the short-term, mainly due to automatic stabilisers. Revenues are adjusted to finance public spending, but fluctuate because of short-term variations in the economy. The underlying assumption is that government spending absorbs a determinate share of output; I take

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[13] These components can be directly compared to statistical measures of potential output, cyclically adjusted government balances, core inflation and base real interest rates.

account of secular growth in the government sector with a real trend. I contribute the co-movement of government revenues and real output to business cycle shocks. This allows identifying temporary economic shocks.[15] In contrast to Blanchard and Perotti (2002), I remove the cyclical fluctuations in fiscal policy from a real trend that is directly derived within the model as a permanent component.

The monetary shock: Fisher relation ($i - p$)

nominal interest rates and inflation are cointegrated. I may identify the relationship between interest rates and inflation either as a short-term Taylor or a long-term Fisher relation (Benhabib *et al.*, 2001). I argue that we can identify a short-term shock to the real interest rate as a monetary policy shock. Sufficient variation in economic variables – other than inflation – is necessary for this interpretation. Unlike a bivariate relation between interest and inflation, I have sufficient restrictions in the other parts of the common trends model to identify the monetary shock. The real rate is not a constant: the real common trend, which reflects changes in productivity, shifts the natural rate of interest. The nominal trend models any longer term trends in monetary policy (Vlaar, 2004). Transitory business cycle shocks have been taken out.

3.3. Some caveats on the common trends model

The incorporation of the budget constraint in the VAR allows tracking the adjustment of spending and taxes after an initial fiscal shock. I control also for cyclical variation of the budget variables. But there are two other identification problems that fiscal VARs need to tackle. Foremost, how do we handle the anticipation effects of fiscal policy? There are long time lags between the announcement, the lengthy and visible budget negotiations and the implementation of a fiscal package. As fiscal shocks need not affect fiscal variables first, the shock can be non-fundamental (Lippi and Reichlin, 1994). The common trends model has some advantages in this regard. The long-term restrictions do not restrict contemporaneous links and should therefore catch anticipation effects.[16] I characterise both the long-term growth and business cycle features of a variety of macroeconomic

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[14] Blanchard and Perotti (2002) note that cointegration does not affect their findings for spending and tax shocks, but do not attempt to identify the cointegration relation.
[15] Given the identification of an aggregate supply shocks with one common trend, temporary shocks might be labelled as aggregate demand shocks.
[16] In particular if price variables are added to the system.

models in which the effects of fiscal policy are derived subject to the government budget constraint. A large number of identifying restrictions are thus consistent with the steady state properties of these models. I do not need to draw upon *ad hoc* assumptions.

Long-term restrictions come at a cost though (Sarte, 1999). Even in large samples, substantial uncertainty surrounds the estimates of the long-term inverted MA representation in (11). Hence, we cannot construct asymptotically correct confidence intervals on $C(1)$ (Faust and Leeper, 1997). Specifying *a priori* the lag length of the VAR, or choosing the horizon at which the long run effect nullifies can solve this problem. I do not test the robustness of the long-term restriction of the nominal trend on output. As this seems a very robust restriction anyway – as King and Watson (1997) demonstrate – inferential problems may be minimal. How realistic is it to test the implication of the IGBC that spending and taxes are cointegrated? After all, there may simply not be enough memory in the DGP of fiscal data to infer upon a transversality condition that holds at infinite horizons (Canzoneri *et al.*, 2001). Even assuming ergodicity of the DGP underlying fiscal variables, in-sample identification is only indicative: a rejection of sustainability would be consistent with either unsound fiscal management or with a data set that is insufficient for identification. Unit root and cointegration tests have well known small sample bias.[17]

There is a potentially greater problem with the linear aggregation of budget data in total spending and taxes. I extract a limited number of shocks from the possibly large set of underlying shocks. Only if each shock affects the economy in qualitatively the same way will the shocks not be commingled. Different budget variables may have very diverse – perhaps even non-linear – effects (Giavazzi *et al.*, 2000). I am interested only in the effect of deviations from budget balance. Nonetheless, I perform stability tests on both the VAR-coefficients and the cointegration tests to detect structural breaks.

4. ESTIMATING THE EFFECTS OF FISCAL POLICY

4.1. Data

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[17] The omission of cointegration in fiscal VARs is usually justified by the fact that standard asymptotic tests remain valid in the presence of unit roots (Sims *et al.*, 1990).

All data series for US fiscal and monetary policy are on a quarterly basis. The intermediate choice of data frequency reflects a trade-off in the joint modelling of fiscal and monetary policy. The budget law is commonly drafted every parliamentary year. Discretionary in-year revisions are not unusual, and cyclical responses are automatic. Whereas the latter effect mainly influences government revenues, the former mainly involves modiifications to spending (Van den Noord, 2002). In contrast, the central bank decides more regularly on interest rate revisions. Given a high degree of interest rate smoothing, the quarterly measure is probably not too coarse.

Essential to the identification procedure is the construction of the fiscal series (Figure 1). All cyclically sensitive budget items are collected in government revenues. As in Blanchard and Perotti (2002), tax revenues are net of government transfers. On the other hand, capital expenditure – which is mainly related to interest payments on outstanding debt – is added to government consumption. The monetary policy instrument is the Federal Funds rate. This rate became the tool for monetary policy actions since 1965, and is also the start of the data sample. I choose the GDP deflator for the inflation rate. Total output is net of government spending. All data are deflated by GDP and in log levels; inflation and nominal interest rates are in percentages.[18]

4.2. Preliminary analysis

Prior to the estimation and the dynamic analysis of the common trends model, we need to establish the following results:

Unit root properties of the data series (Table 1)

I find that all series in X_t can be considered as non-stationary. Government spending and interest rates are level stationary according to the ADF test, but the KPSS test suggests that spending is an I(2) process. A similar finding results from a test for trend stationarity. In that case, output or government revenues might be trend stationary, however. Some classical tests of sustainability check for a unit root in public debt. Sustainability does not seem to be guaranteed in-sample. Debt is rather characterised as an I(2) series. Net or total interest payments follow this derailing of public finances, and also reflect an I(2) process.

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[18] Fiscal and GDP data are from the BEA (NIPA tables), public debt from the FRED II database, deflators from

Cointegrating rank of VAR (Table 2)

I include a constant in the long-term part of the model. As a consequence, public debt converges to a bounded value in the long-term. For the real interest rate, I include a drift so as to allow a linear trend over time. Alternatively, I include a trend in the short-term part of the model next to the constant term in the cointegration part. Both models give nearly identical results. At a 5% significance level, the trace test favours a cointegrating rank of 3. When a trend is included, this hypothesis is accepted at the 10% level. Instead, the Saikkonen and Lütkepohl test (S&L) finds evidence for only 2 cointegrating relations.

Estimation of the cointegrating vectors β

the three cointegrating relations were estimated in the VECM (10) via maximum likelihood. *A priori*, four lags are included in the VAR. No seasonal dummies are included. The cointegration coefficient estimates are used as input for the common trends model.

Imposition of the long-term restrictions

identification of the permanent innovations requires the imposition of parameter values on the common trend parameters in F_0. In order to identify the real trend in F, I take the permanent (supply) shock on real output to result in a corresponding long-term increase in spending. The coefficient then $\hat{\theta}$ refers to this estimated long run equilibrium coefficient between y_t and g_t. From cointegration between spending and revenues follows the relation between permanent output and revenues $(-\hat{\theta})$. Similarly, the nominal trend is assumed to affect both spending and revenues. The coefficients $\hat{\omega}_G$ and $\hat{\omega}_T$ are obtained from a regression of g_t and t_t on inflation respectively.

$$F_0^{'} = \begin{bmatrix} 1 & 0 & 0 & \hat{\theta} & -\hat{\theta} \\ 0 & 1 & 1 & \hat{\omega}_G & \hat{\omega}_T \end{bmatrix} \tag{15}$$

After these initial steps, I estimate the common trends model, compute the Wold VMA representation and calculate the impulse responses and forecast error variance decomposition – and their respective asymptotic standard errors.[19]

BLS and the Federal Funds rate from the Federal Reserve.
[19] We plot the impulse responses over 20 periods, with their 66% asymptotic confidence bounds.

4.3. Results

Fiscal VARs disagree on the effects of increases in government spending on private consumption. Empirical models either find insignificant or weakly positive responses. All studies agree on the positive effect of fiscal expansion on output, however. The dynamic responses of the common trends model reverse some of these standard results. I find that an expansionary fiscal policy has contractionary effects on output (Figure 2a). There is a stark contraction in the first year after the shock. This response is not very persistent. Afterwards, output slightly expands in the second and third year but returns to a steady value afterwards. At the same time, a fiscal shock has a particularly pronounced and persistently negative effect on inflation. Inflation remains significantly below baseline till three years after the shock. Consequently, there is also a slight monetary contraction. This interest rate response is not very significant nor very persistent, however.[20] Fiscal expansions that stray away from sustainability seem to be mainly due to expansions in government spending. Tax cuts play a smaller role. The model does not give a key on how a future fiscal adjustment will be brought about. I find a very gradual return of spending to the baseline value. For taxes, the adjustment is much less persistent. But in contrast to the results in Bohn (1991), I do not find any significant reversal of spending or an increase in tax revenues to offset the rise in public debt.

The results of the monetary part of the model is more in line with standard findings in the literature. A loosening of monetary policy leads to a moderate increase in output, and a stronger and rather persistent rise in inflation (Figure 2b). Monetary policy shocks do not trigger a direct response of fiscal variables. There is a small indirect impact, as the budget responds to the output effects of changes in monetary policy. This is the effect of two opposing forces. Interest rate cuts reduce the tax base for capital taxation; output growth instead translates into higher tax revenues.

The comparison of dynamic responses to the economic shocks generally confirms to economic priors (Figures 2c-e). The business cycle shock temporarily raises output and growth dampens out after about two years. Interest rates react in accordance: monetary policy reacts in a non-accommodative way to growth. This causes an initial decline in inflation that gradually rises over the economic boom. Owing to automatic stabilisers, tax

revenues rise in step with the increase in output. A positive supply shock has permanent effects on output. As one might expect, it also exerts a temporary deflationary impact on the economy. The permanent supply shock also leads to a sustained rise in government spending. This is not the consequence of an increase in the available budget: higher potential growth does not imply higher tax revenues. We even observe an initial decline in tax intake in the first two years after the supply shock. The nominal shock has a persistent impact on inflation and interest rates, but no impact on real variables. These permanent effects are rather imprecisely estimated, however.

The forecast error variance decomposition gives some additional insights in the common trends model (Figure 3). At short horizons, most of the variability in the various series is accounted for by the transitory components. In the short- to medium-term, aggregate demand shocks are the main source of output variability whilst both supply-side and nominal shocks are so in the long-term. This translates in a consequent important role for business cycle and supply side fluctuations in government revenues. This obviously owes much to automatic stabilisation. Aggregate demand shocks are also driving interest rates in the short-term, but the major part is accounted for by nominal factors. The latter also largely explain movements in inflation, even if the contribution of monetary policy is rather large at short horizons. Monetary policy shocks have a small role in explaining variability in interest rates or fiscal policy at very short-term horizons. Interest rates do play an increasing role in the variability of expenditure. This could be consistent with a stronger reaction to interest payments over time. As in most other studies, the contribution to output is negligible at all horizons. The same holds for fiscal policy: systematic fiscal policy is more important than discretionary shifts. Variability in government expenditure is largely explained by its own past behaviour. This is consistent with spending being a largely independent process, driven by factors exogenous to the model. Spending has only a minor role in the variability of other series, in particular government revenues. This is a rather disappointing result : consolidatory adjustment with fiscal instruments, as in Bohn (1991), seem to occur only on the spending side of the budget. There are of course other channels (higher inflation, higher potential growth) that could render the debt position sustainable.

4.4. The Ricardian effects of fiscal policy

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[20] As in Favero and Giavazzi (2007), I find that interest rate responses are much more muted.

The contractionary effects of fiscal policy shocks detected here stand in stark contrast to most previous results in the fiscal VAR literature. Several approaches have been used to identify fiscal shocks. Blanchard and Perotti (2002) make assumptions about the sluggish reaction of some variables to fiscal policy and use detailed institutional information on the tax system. The main finding of their semi-structural VAR, which has been repeated by other authors for different samples, is the positive Keynesian effect of fiscal expansions on consumption, albeit the total mulitplier effect is small. Studies that make a detailed historical study of policy decisions (Romer, 1994; Romer and Romer, 2007) or use information on large fiscal events that may be assumed exogeneous (e.g. the timing of wars and defense spending)(Edelberg *et al.*, 1998; Burnside *et al.*, 2003) find small or insignificant positive effects of fiscal policy on consumption.[21] Howver, all studies would agree that increased government spending or tax cuts have positive output effects in the short run.

These results are in line with the predictions of baseline DSGE models of fiscal policy. The basic mechanism behind the expansion of output in RBC models is the wealth effect (Baxter and King, 1993). An increase in government spending implies higher (lump sum) taxes in the future. As a consequence, consumption and leisure are substituted for an increase in the labour supply. Models that extend the baseline neo-classical model with some nominal rigidities also find positive effects of output. The expansionary effect in these New Keynesian models basically owes to a shift in labour demand and an increase in the real wage that induces higher consumption, via a substitution effect or because of some fraction of agents that are credit constrained.[22] What may explain the contractionary output response after fiscal expansions in the common trends model then?

A first reason might be monetary policy. Responses of the central bank to changes in fiscal policy might reverse its usual crowding out effects. An aggressive anti-inflationary reaction of the central bank that raises interest rates in the wake of a fiscal shock could actually reverse the expansionary output effect. Linnemann and Schabert (2003) or Gali and Monacelli (2005) show in a New Keynesian model that the effects of fiscal policy can vary

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[21] Studies that impose 'sign' restrictions on fiscal VARs do not find significant responses either (Mountford and Uhlig, 2005).

with the monetary regime. Moreover, strong changes in interest rates further burden the government with higher interest payments on outstanding debt.[23] In fiscal VARs, monetary policy has been treated in a rather incomplete way. It is usually considered sufficient to include some price variables. By absorbing some of the effects of future fiscal policy changes, these augmented VARs reveal muted responses of macroeconomic variables to non-systematic fiscal shocks (the so-called 'Sims conjecture' (Perotti, 2005)).[24] SVAR studies of this type find more moderate but still significantly positive output effects though. Mountford and Uhlig (2005) is the only study to explicitly identify both fiscal and monetary policy. They do not find significant output responses to fiscal shocks, but no contraction either. In my model, the short-term interest rate falls in the first quarters after the shock.[25] However, this fall is much smaller than the drop in inflation. We would have anticipated at least some accommodating reaction of the central bank. Hence, the fiscal shock implies some monetary contraction. It is not unlikely that the Federal Reserve sets its policy as a strategic substitute to expansionary budgets of the Treasury. Nonetheless, it seems unlikely that the slight contraction of monetary policy explains the large size of the output contraction. Hence, the monetary policy reaction is genuinely meant to dampen the adverse economic effects of the fiscal shock. We can discard the hypothesis that the monetary policy regime is the reason of the output contraction.[26]

Not only do I condition on monetary policy, the empirical model also includes a technology shock with permanent output effects. Both RBC and New Keynesian models allow for cyclical fluctuations around some steady-state trending growth path. In RBC models, fiscal policy would only have the supply-side effect of boosting labour supply; New Keynesian models display additional transitory output deviations due to nominal rigidities. Consequently, conditional on the supply shifts, there are either zero or positive output variations. I find that supply shocks cause government spending to permanently expand. Supply shocks do not lead to similar increases in tax revenues. Hence, revenue windfalls do not fuel a rise in spending. Hence, the common trends model picks up important supply

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[22] The real proof of the RBC or the (New) Keynesian model lies in the response of consumption, which goes in opposite directions. The large contraction in output must be due to a contraction in consumption, and hence contradicts both models.

[23] Favero and Giavazzi (2004) illustrate how interest rate hikes lead to a snowballing debt service.

[24] Moreover, the omission of interest rates may lead to aggregation of monetary and fiscal shocks if there are important systematic contemporaneous relationships between both policies.

[25] The interest rate falls in the first quarters after the fiscal shock in Mountford and Uhlig (2005).

[26] Note that the lack of a monetary response also excludes that the public expects monetary policy to resolve the fiscal problems.

side effects of government spending, which it mingles with 'other' technology shocks. This would still not explain the temporary contractionary response to the fiscal shock, however.

The finding is consistent though with a specific RBC model of fiscal policy. If the government keeps the budget balance period-by-period, output declines if spending shocks are offset with proportionate increases in distortionary taxation. Tax hikes make agents smooth intertemporally, reducing temporarily both current labour supply and investment. As tax bases decline, the rise in tax rates needs to be more than proportional to cover the additional spending. Consequently, the contractionary effect of taxes more than offset the positive output effects of increased spending.

The common trends model identifies exactly this type of fiscal shock: it is a temporary innovation to the budget constraint. In order to account for a negative output effect, taxation needs to be distortionary in the RBC model. In practice, distortionary taxation is definitely more common than lump sum taxation. The evidence can thus account for the features of a RBC model that is closer to reality.

Gordon and Leeper (2005) suggest an additional explanation for the negative output effect. Policies that increase government indebtedness raise future debt obligations. To the extent that fiscal expansions are *expected* to be brought back by a mix of tax rises or spending cuts in the future, negative Ricardian effects may dominate in the short run. DSGE models that account for this expectational channel indeed find opposite effects of fiscal policy, in contrast to what standard RBC or New Keynesian models predict (Gordon and Leeper, 2005). Fiscal expansions have negative output effects by affecting intertemporal saving choices.[27],[28] Taking the government budget constraint seriously can overturn widely held beliefs on policy effects (Leeper and Nason, 2005). In contrast to other fiscal VARs, the common trends model indeed explicitly incorporates the intertemporal government budget constraint. The VAR-counterparts to DSGE-models usually ignore the dynamic adjustment of government spending and taxes after a fiscal shock that menaces debt sustainability. The means of financing and the adjustment in spending and taxes wrap the empirically relevant role of expectations effects in the fiscal policy shock, without considering the path of public debt. By ignoring the effects of the

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[27] Their simulations with US data demonstrate that such expectations are sufficient to create business cycle like variations in macroeconomic variables.

budget constraint, the responses to fiscal shocks are biased. The common trends model explicitly embodies expectations of changes in fiscal policy, and thereby incorporates the forward looking behaviour of economic agents. This makes us recover the typical 'Ricardian' effect of fiscal expansions.

4.5. ROBUSTNESS CHECKS

4.5.1. Data definitions

The results are confirmed for other definitions of the fiscal series and output. Figures 4a-c display the results for the fiscal shock for some alternative specifications. One wonders if the protracted crowding out effect is not simply due to (national) accounting conventions. Figure 4a plots the response of total instead of private output. No significant difference results. I have excluded government investment from the budget as it might have long-term effects on output. Including public capital building does not affect the findings, however (Figure 4b). Finally, I use total capital transactions instead of including only net interest payments, but again the main result remains (Figure 4c). To sum up, alternative data definitions usually lead to more pronounced responses to fiscal shocks, but all confirm the contractionary effects following fiscal expansions.

4.5.2. Model fit

There are several other ways to assess the robustness of the results from the common trends model. Following (12), the common trends model provides us with a decomposition of the series into (time-varying) permanent and transitory components (Figure 5a-b). The permanent component can be directly compared to measures of potential output, cyclically adjusted government balances, core inflation and base real interest rates as calculated by international organisations, or as obtained by some mechanical data filters (Table 3).

The underlying real and the nominal stochastic trends that I extract from the model do make sense. The permanent component gives a reasonable characterisation of the trends

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[28] There are some more elaborate models of fiscal policy that specify a role for debt also challenge the conventional view (Linnemann and Schabert, 2005). The identification of these theoretical models would

underlying the series. This is perhaps least clear for the fiscal series. Actual and permanent spending coincide only over the period 1972-1985, but afterwards transitory fluctuations outdo the permanent component. Similarly, actual revenues are constantly larger than permanent tax income. As the budget constraint needs to be satisfied in-sample, the unusual budgetary events after 2001 cause the permanent component to be shifted down. The dramatic decline in revenues in 2002 accordingly attributes a larger part of fiscal series to the transitory component. In addition, a large part of the transitory increase in revenues is also due to the positive output gap over the nineties. The corresponding permanent net lending ratio nevertheless tracks the (primary) net lending ratio quite well, and the correlation with the primary deficit – in ratios or levels – is positive and significant.

The tendency for economic fluctuations to become less volatile and prolonged is clear from Figure 5b, and has also been documented elsewhere (Stock and Watson, 2003). The permanent output component tracks quite well other measures of potential output. The economic cycles correspond to periods of economic boom and bust. Both the permanent and transitory output series are correlated with measures provided by NIPA or a HP-filter.

A similar tendency for transitory fluctuations to become more regular and prolonged is visible for interest rates. The actual series oscillates around the permanent interest rate, reflecting the gradual loosening and tightening of policy. The trend decline in interest rates is also rather well captured. Interestingly, the strong fiscal expansion in recent years seems to be accompanied by a positive jump in the real interest rate. Fiscal consolidation in the nineties has been as important as *'good'* monetary policy in bringing down interest rates. In order to close the gap with the natural level of interest rates, the Federal Reserve will need to tighten monetary stance rather strongly. The model seems less successful in describing the behaviour of inflation. Permanent inflation is much lower than the actual series, and indicates deflation over the nineties. Nevertheless, the 'Goldilocks' US economic boom with steady economic growth and low inflation has often been attributed to persistenly positive supply shocks with possibly falling prices. There is no correspondence to more standard core measures of inflation.

4.5.3. Analysing structural shocks

require a more detailed empirical structure, however.

Since the criticism of Rudebusch (1998), we know that inference on the structural shocks is a dubious exercise as the residual series we obtain may depend on the particular specification of the empirical model. It may be too much to require an exact timing or an appropriate size of policy shocks. Nevertheless, periods of strong deviations away from or towards solvency should be easily discernible. Figure 6 displays the structural fiscal policy shock from the base common trends model. Large outliers are exceptional but some periods appear more evident. The tax cut of 1975 and the Carter-Reagan expansion of the early eighties are visible. The Bush tax cut of 2001 is less evident. Note that the volatility of discretionary fiscal interventions on sustainability have declined since the eighties (around 1985-1987) and have been at a low level since (Figure 7). This underlines the switch to a more consolidatory fiscal stance in the nineties. The recent rise announces another shift to a period of debt accumulation.

4.5.4. Structural breaks: regime shifts in fiscal and monetary policy

I first check for structural breaks in the joint DGP of X_t (Table 4, Figure 8a-b). I follow Bai *et al.* (1998) in constructing a sequential Quandt-Andrews likelihood ratio test based for a VAR. The (single) breakdate is the supremum value of this recursively generated series over the central 70% of the sample. Following Stock and Watson (2003), I scale the series by the change in residual variance before and after the initially estimated breakdate.[29] The confidence interval around this break indicates (Bai, 1997) how precise the structural break can be located. I set lag length of the VAR model to four, which should be sufficient for quarterly data.

The tests indicate a quite precisely estimated structural break in the first half of the eighties. The break occurred between the first quarter of 1981 (for the homoskedastic version) and the third quarter of 1985 (for the heteroskedastic version). This probably reflects 'Volcker-Greenspan' shift to an actively inflation combating Federal Reserve (Clarida *et al.*, 2000). Nevertheless, there is no reason to fancy a particular break date. A plot of the heteroskedatic test stats shows that instability is spread out over the entire eighties. A gradual decline only sets in after 1990, but never returns to its pre-1980 level. As this is

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[29] Applying this heteroskedasticity correction more correctly locates the break if the data series did not only change its first, but also its second moments.

reduced form evidence, I cannot tell apart the importance of shifts in either fiscal or monetary policy. A regime of *'passive'* fiscal policy and *'active'* monetary policy is crucial in bringing about stability. The strong reduction in public debt during the Clinton Administration follows upon a period in which the Federal Reserve kept inflation under control. Favero and Monacelli (2005) notice that fiscal policy aids in tracking inflation developments when the Treasury disregards debt consolidation. A debt stabilising regime started around 1987.

To get some more insight in the break tests of the general unrestricted VAR, I subject the structural VEC model to some recursive stability tests The decomposition into the long and short-term parts of the common trends model tells us whether the equilibrium relations or the speed of dynamic adjustment have changed over time. I recursively add data to the base period 1965:1-1976:1 to update the estimates of the VEC model. I scale each statistic by the 90% quantile of the distribution in Figure 9a. The time path of the recursively estimated trace statistic should be upward sloping for the significant rank values, but downward for the non-significant ranks. Only at the beginning of the sample is there clear evidence of three cointegrating relations. There is a strong fall in the significance of the trace parameters around 1980, and at most two cointegration relations can be said to exist till at least the mid-nineties.

As the hypothesis of constant rank is rejected, I continue to test for constancy of the cointegration parameters. I first concentrate out the short-term parameters α from the likelihood function, and estimate only the long-term part of the VEC. The recursively estimated eigenvalues give a visual test of parameter constancy. The normalised non-zero eigenvalues are plotted together with their approximate 95% asymptotic error bounds. Figure 9b shows the time paths of the three non-zero eigenvalues. In line with the results of the rank test, all eigenvalues are significant in the first part of the sample. Most of the instability occurs in the late seventies (for the largest eigenvalue), and early eighties (for the second eigenvalue). Towards the end of the sample, we observe a gradual decline. A formal test of constancy of the eigenvalues is given by the fluctuation τ-test of Hansen and Johansen (1999).[30] Only for the largest eigenvalue do I reject the stability-null (Figure 9c). The non-constancy of the cointegrating vector occurs over the period 1981-1992. This

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[30] We plot the sample path of the eigenvalues together with the 5% critical value of its supremum value. The null hypothesis is stability of the eigenvalues of the VEC model, but does not consider a specific alternative.

again suggests the importance of the Volcker shift in monetary policy, in conjunction with a gradual turnaround of fiscal policy to control public debt. As the other eigenvalues are not rejected to be constant, the overall fluctuation τ-test on the sum of eigenvalues is also strongly in favour of parameter constancy (Figure 9d).

I can only square this parameter constancy in β with the finding of instability in the VAR model, by locating changes in the short-term adjustment part of the model. I run the overall c-test on all the parameters in $\alpha\beta'$ as given by the VEC model. The recursive plot of fluctuation test statistics in Figure 9e indeed indicates instability in $\alpha\beta'$ over the first half of the sample up till 1980. Afterwards, a gradually decline sets in over time, confirming previous evidence of stability since around 1987 (except for a marginal rejection in 1995). I thus find that the long-term relations itself are stable, but that the persistence of the response to shocks underwent structural changes.

4.5.5. Alternative specifications of the VEC model

I now look in more detail into the role of public debt and interest payments and report selected results for some alternative specifications of the VEC model. So far, I did not use the theoretical cointegration relation between spending and revenues or the nominal interest rate and inflation. First, the data definitions for fiscal series do not completely match the theoretical concept. Second, we may expect debt to converge to some positive bounded value. The effect of imposing the cointegration vector $\beta' = [1, -1]$ on spending and revenues is to make the fiscal shock, and the consequent responses of output and inflation much more persistent relative to the base case (Figure 10a). An interest rate hike occurs over the first quarters. Again, there is no significant adjustment in revenues taking place. The decomposition in transitory and permanent components makes clear this is not a well fitting specification though (Figure 10b).

Instead of deriving the evolution of public debt from the budget constraint, a direct control for the accumulation of fiscal imbalances over time might be more straightforward. Several possibilities are open. A first alternative is to control for the initial stock of debt, and impose it as an exogenous process on the VEC model. The cointegrating relation between total spending and revenues can alternatively be rephrased as holding between current

spending, total revenues and interest payments.[31] Governments' fiscal policies then satisfy the intertemporal budget constraint if the vector $\beta' = [1,-1,1]$. Reformulating this with respect to debt, gives the following cointegrating vector: $\beta' = [1,-1,\kappa]$, where κ is now the coefficient on total public debt. I experiment also with imposing debt as an exogenous variable in the former case.

Table 5 gathers the cointegration rank test for these different specifications. Only if debt is imposed exogenously I do not reject a cointegrating rank of three. In all other specifications, at most a rank of two can be accepted. I test the stability of these alternative VAR models, but these do not give contrary results to the basic model for the homoskedastic version (Table 6). Controlling for heteroskedasticity, the models with interest payments locate the break about 1990 instead. Despite the gradual decline in interest rates, the repayment of public debt probably gathered pace, explaining the accelerated rise in interest payments.

The results for these models are rather disappointing: none of the impulse responses is really significant, nor does the common trends model fit the series well. This is clearly the case when debt is included as an exogenous variable.

Additional identifying restrictions are necessary in case public debt or interest payments are made endogenous in a VAR[6]. I continue imposing the three cointegration relations. I accordingly need additional identifying restrictions. I do so by distinguishing a third common trend. As before, I retrieve a real trend from a supply shock with having permanent effects on output, and a nominal trend with permanent effects on prices. I consider shocks to the real interest rate as causing a third common trend underlying all variables. Neither of these common trends models gives a precise indication on the effects of fiscal policy. The decomposition of the model tracks in both cases the variables rather well, and gives transitory fluctuations around the permanent trends that make economic sense (Figure 11a-b). Only core inflation remains well below zero. In contrast to the base model, the specification with interest payments attributes most of the shocks in fiscal policy to the period of monetary turbulence around 1980. A plot of the standardised residuals for both specifications shows that shocks in other periods are much less

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[31] We include alternatively net payments on outstanding debt, and overall capital non-investment transactions

pronounced, with the exception of the increase in the Vietnam War spending around 1967, or the 1975 tax cut (Figure 12a-b). Most of the volatility in fiscal policy is accordingly located in the early eighties. This corroborates the interpretation of the discipline imposed by much tighter monetary policy on fiscal behaviour through rising interest payments, albeit with some lag in its implementation. This is also consistent with the other findings of a break in fiscal policy happening in the late eighties (Favero and Monacelli, 2005).

5. CONCLUSION

This paper contributes to the growing literature on fiscal VARs. I proposes the common trends to shed some new light on the economic effects of fiscal policy, and resolve some of the empirical puzzles in fiscal VARs. The SVAR model embeds the long-term government budget constraint constraint in an analysis of short-term stabilisation policies. The error correction structure of the model permits analysing the short-term dynamics following shocks to the long-term equilibrium around some common trends, which are driven by shocks with permanent effects. The model is fully consistent with the theoretical properties of recent DSGE models of fiscal policy.

In contrast to common results in the fiscal SVAR literature, I find expansionary fiscal policy to have contractionary effects on output and inflation in the short run. As we control for monetary policy, we can discard the hypothesis that a tough anti-inflationary stance of the central bank causes this effect. Supply-side effects of fiscal policy are not responsible either. The explanation for this 'puzzle' lies in the identification of fiscal policy. I consider as a fiscal shock the deviation from the government budget constraint. In RBC models, fiscal expansions have negative output effects when the government keeps budget balance period-by-period, but only if it uses distortionary taxes. Taking the government budget constraint seriously can overturn usual policy effects also in a second way. To the extent that fiscal expansions are expected to be brought back by future tax rises or spending cuts, negative Ricardian effects may dominate. DSGE models that account for this expectational channel indeed confirm opposite effects of fiscal policy, as to what standard RBC or New Keynesian models predict (Gordon and Leeper, 2005).

The lesson we draw from this is that financing decisions of government policies matter,

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on the budget.

and may even outdo the effect of the initial change in fiscal policy.[32] RBC models may be less awkward than what recent extensions with nominal rigidities and other frictions seem to suggest. Attention should primarily focus on the friction of distortionarity taxation.

The implication for empirical research with fiscal VARs is that considering both government spending and taxes pays off with a better insight in the properties of fiscal policy models. There is also an urgent need to better understand the effect of changes in taxation. The identification of changes in taxation is fraught with enormous dfifficulties, however.

This paper provides a comprehensive framework to deal with both short- and long-term effects of economic policies. The common trends model of fiscal policy can be extended in at least two interesting directions. First, the analysis of consumption and investment responses to fiscal policy under a common trends identification may validate the importance of the expectation channel in RBC or New Keynesian models. Second, a recent literature combines the insights from finance and macroeconomic models to examine monetary policy (Rudebusch and Wu, 2003). A more elaborate specification of the interest rate structure, may allow for a feedback between the term structure factors and monetary or fiscal macro variables.

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[32] *" … [ ] The financing decisión is quantitatively more important than the resource cost of changes in government parchases."* (Baxter and King, 1993).

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TABLES

Table 1. Unit root tests.[a]

Test	level				first difference			
	ADF[b]		KPSS		ADF[b]		KPSS	
	μ	τ	μ	τ	μ	τ	μ	τ
government expenditure	**-3.21	-2.41	***3.50	***0.66	***-3.90	***-4.86	***1.33	*0.13
government revenues	-1.31	-2.44	***3.35	0.11	***-4.01	***-4.09	0.12	0.06
output	-1.08	**-3.88	***3.57	***0.25	***-4.30	***-4.36	0.05	0.03
inflation	-2.04	-2.44	***0.37	***0.31	***-4.15	***-4.40	0.12	0.04
interest rate	**-2.91	-2.88	**0.57	***0.57	***-3.89	***-4.13	0.19	0.03
public debt	-0.47	-2.00	***3.59	***0.43	-2.19	-2.18	**0.70	***0.70
net interest payments	-1.91	1.56	***3.40	***0.59	-2.07	-2.91	***1.26	***0.56
total interest payments	-1.89	0.43	***3.40	***0.59	-0.61	-1.33	***1.29	***0.58

Notes: (a) table entries are test statistics; */ **/ *** indicates the variable is I(1) at 10, 5 and 1% significance level respectively; the optimal lag order is determined by the Bayesian Information Criterion (BIC), with a maximum of 8 lags; (b) μ for level stationarity, τ for trend stationarity.

Table 2. Trace tests for cointegration rank.[a]

null hypothesis	constant	5% critical value	constant and trend	5% critical value
	Trace test			
rank = 0	***167.83	76.81	***136.64	88.55
rank $\diamondsuit$ 1	***68.32	53.94	***75.92	63.66
rank $\diamondsuit$ 2	**38.08	35.07	*40.40	42.77
rank $\diamondsuit$ 3	12.05	20.16	20.54	25.73
rank $\diamondsuit$ 4	5.04	9.14	4.33	12.45
	S&L Test			
rank = 0	***107.13	59.95	***110.81	66.13
rank $\diamondsuit$ 1	**43.19	40.07	***58.53	45.32
rank $\diamondsuit$ 2	13.32	24.16	15.52	28.52
rank $\diamondsuit$ 3	9.30	12.26	2.08	15.76
rank $\diamondsuit$ 4	-	-	1.14	6.79

Notes: (a) trace or S&L test statistics, critical values are from Doornik (1998) and Trenkler (2004).

Table 3. Correlation of permanent and transitory components with common measures.

permanent component of output		permanent component of inflation	
potential NIPA	0.92 (0.00)	core inflation	-0.29 (0.00)
HP-filter	0.91 (0.00)	HP filter	-0.28 (0.00)
cyclical component of output		permanent primary deficit	
HP-filter	0.59 (0.00)	level, HP filter	0.44 (0.00)
linear trend	0.07 (0.87)	ratio, NIPA	0.38 (0.00)
quadratic trend	0.29 (0.00)		

Table 4. Stability tests: breakdate in VAR[X_t], 4 lags, 1965:1-2003:4.

	QLR-test stat[a]	breakdate	confidence interval
homoskedastic	38.38 (0.00)	1982:2	[1981:1; 1983:3]
heteroskedastic	50.24 (0.00)	1984:1	[1983:3; 1985:3]

Notes: (a) the test statistic is the Quandt-Andrews LR version, with p-value in brackets. The breakdate is the heteroskedasticity corrected sup Quandt-Andrews breakdate (Stock and Watson, 2003). The years in brackets are the confidence interval at 33% (Bai, 1997); [-; -] indicates this interval exceeds the sample.

Table 5. Trace tests for cointegration rank, constant in cointegration relation.[a]

Null	basic model	+ debt exogenous	+ net interest payments	+ total interest payments	+ net interest payments + debt exogenous	+ total interest payments + debt exogenous	+ debt endogenous
rank = 0	***136.64	***113.02	***142.24	***135.41	***160.37	***158.80	***229.54
rank $\leq$ 1	***75.92	**51.69	**77.83	**77.07	**86.31	**93.67	***106.46
rank $\leq$ 2	*40.40	*31.63	44.66	44.62	48.85	48.07	50.66
rank $\leq$ 3	20.54	17.12	25.27	25.58	27.28	25.61	30.67
rank $\leq$ 4	4.33	6.67	12.52	12.36	14.86	13.99	16.26
rank $\leq$ 5	-	-	4.37	4.69	5.76	5.81	5.82

Notes: (a) entries are trace tests; */ **/ *** shows the 10, 5 and 1% significance level respectively.

Table 6. Breakdate test on VAR, 4 lags, 1960:1-2003:4.[a]

VAR	QLR-test stat	breakdate	confidence interval
Basic model			
homoskedastic	38.38 (0.00)	1982:2	[1981:1; 1983:3]
heteroskedastic	50.24 (0.00)	1984:1	[1983:3; 1985:3]
Basic model, with exogenous debt			
homoskedastic	40.61 (0.00)	1982:2	[1981:2; 1983:2]
heteroskedastic	52.36 (0.00)	1984:1	[1983:3; 1985:3]
Basic model, with net interest payments			
homoskedastic	43.46 (0.00)	1982:2	[1981:2; 1983:2]
heteroskedastic	60.58 (0.00)	1990:4	[1990:2; 1992:2]
Basic model, with exogenous debt and net interest payments			
homoskedastic	37.87 (0.00)	1982:2	[1981:1; 1983:3]
heteroskedastic	50.07 (0.00)	1990:4	[1990:2; 1992:3]

Notes: see table 4.

FIGURES

Figure 1. Data, United States, 1965:1-2003:4.
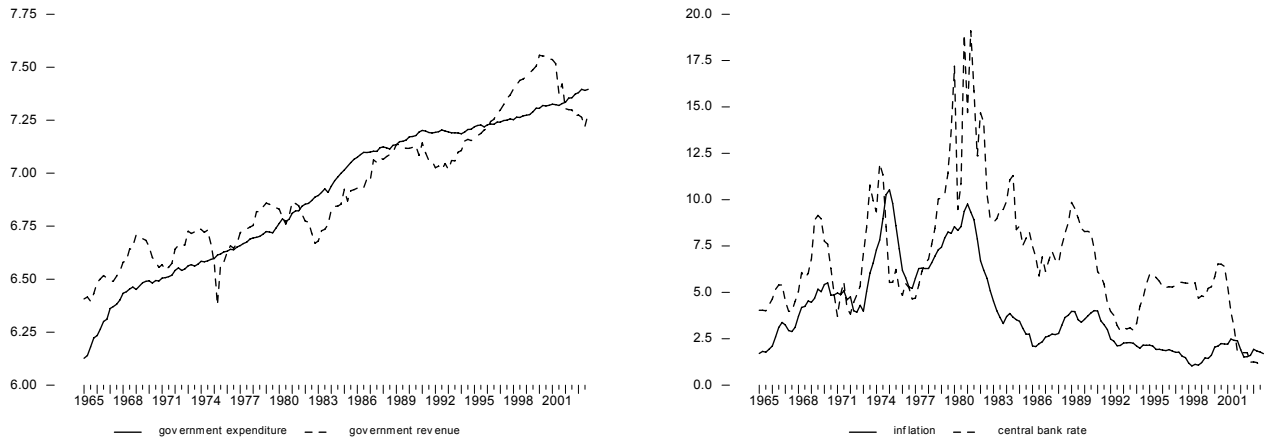
Figure 2a. Impulse response function, 66% asymptotic confidence interval, response to a 1 standard deviation fiscal policy shock.
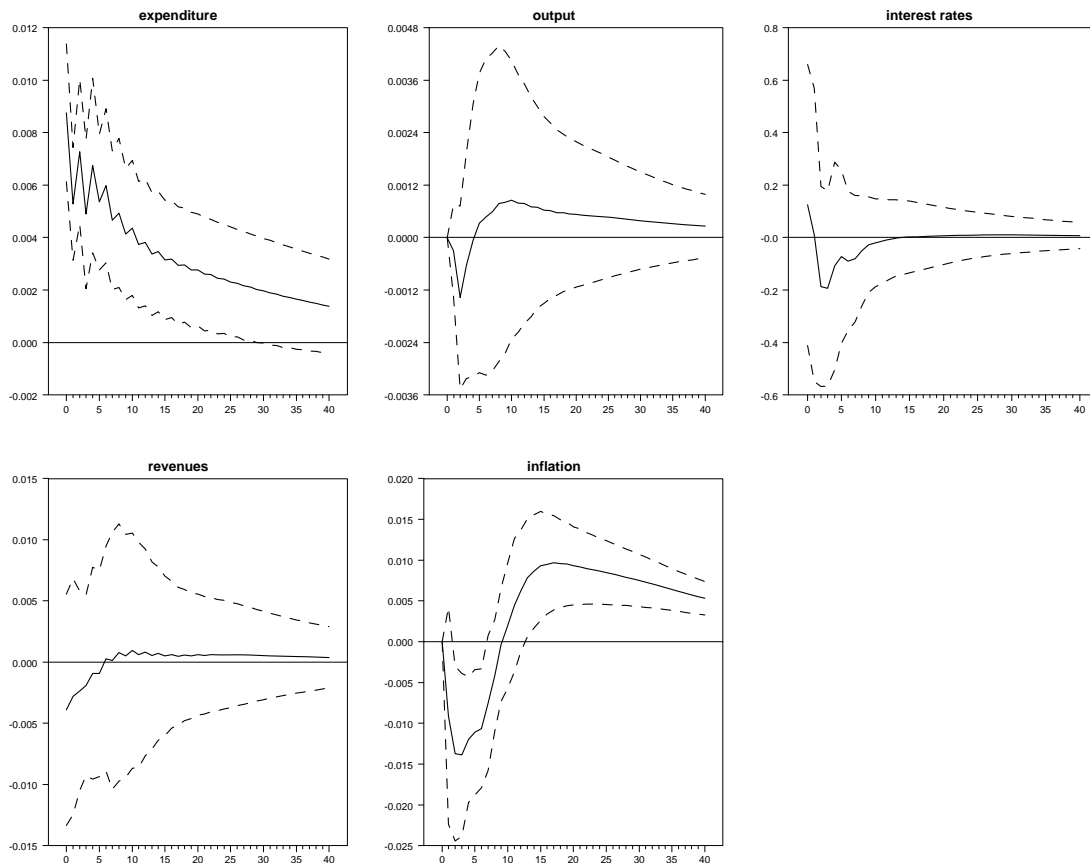
Figure 2b. Impulse response function, 66% asymptotic confidence interval, response to a 1 standard deviation monetary policy shock.
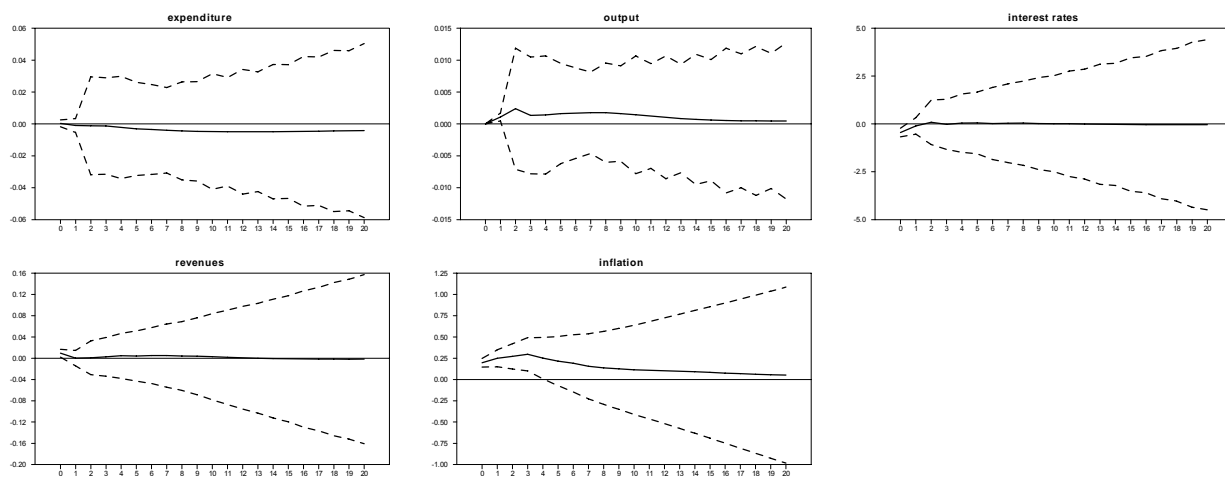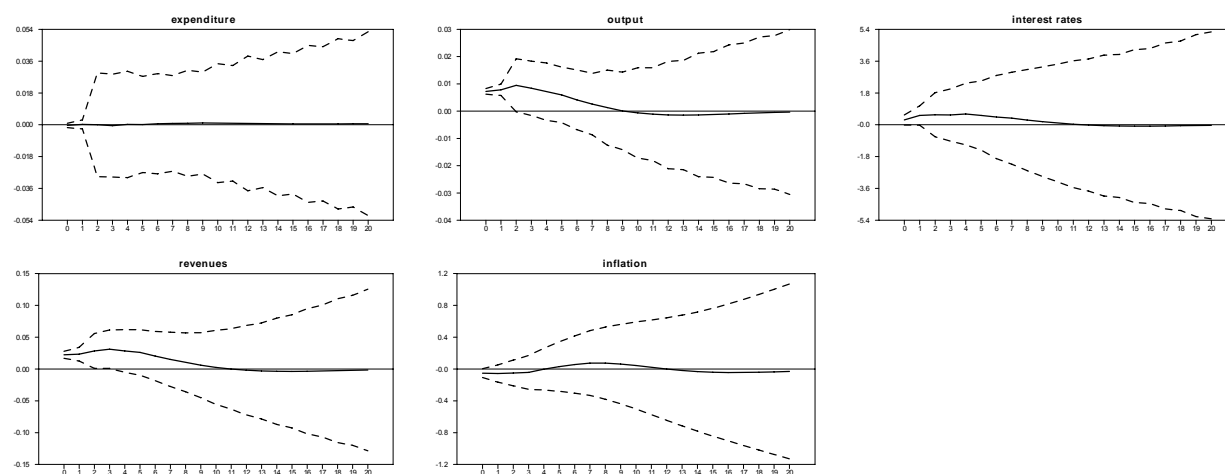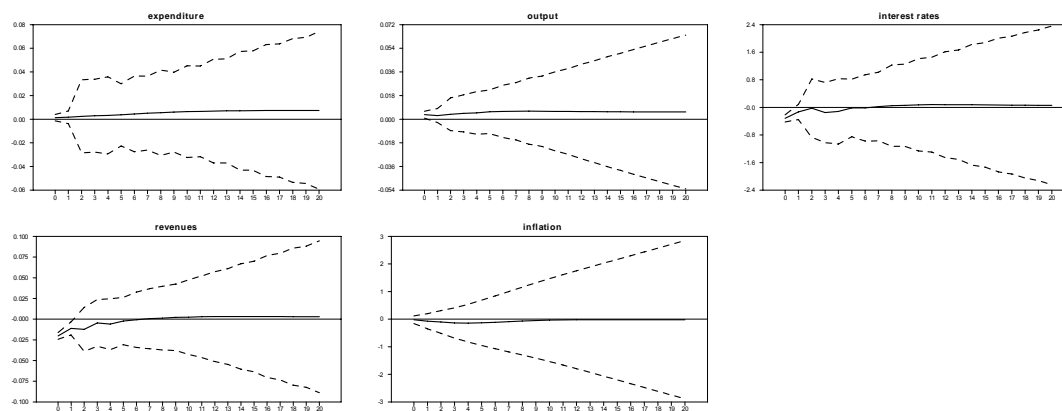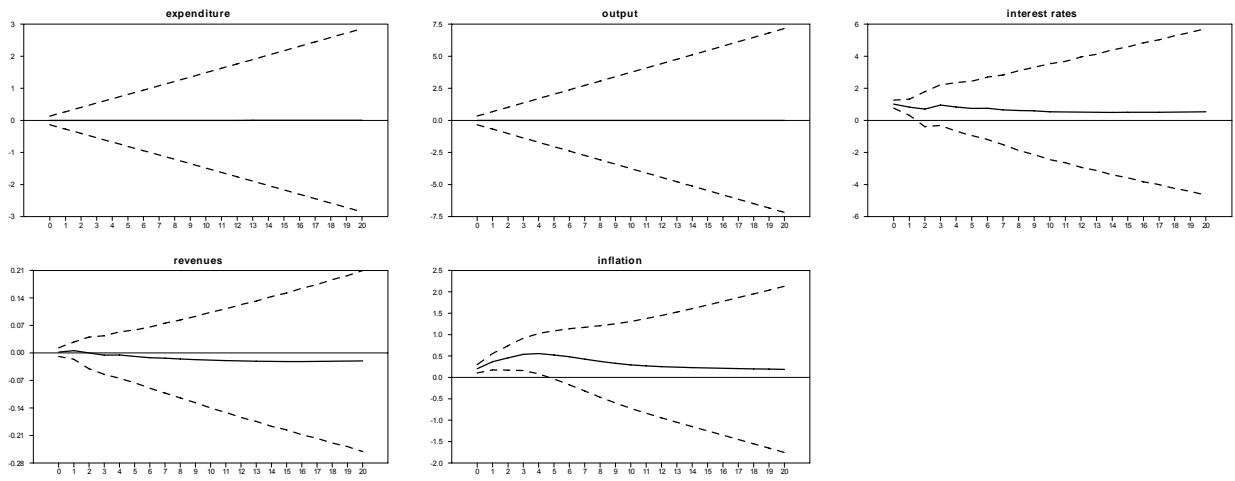

Figure 2c. Impulse response function, 66% asymptotic confidence interval, response to a 1 standard deviation business cycle shock.

Figure 2d. Impulse response function, 66% asymptotic confidence interval, response to a 1 standard deviation real supply shock.

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Figure 2e. Impulse response function, 66% asymptotic confidence interval, response to a 1 standard deviation nominal shock.

Figure 3. Forecast error variance decomposition (FEVD).

■ real ■ nominal ■ fiscal ■ monetary ■ business cycle
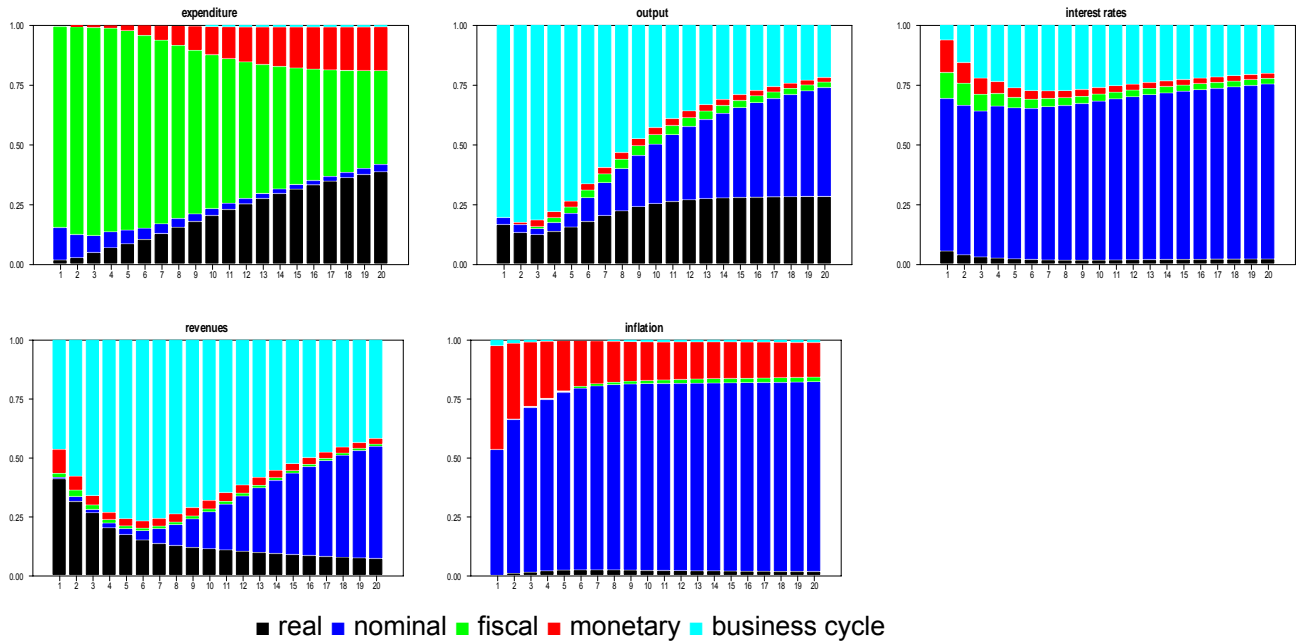
Figure 4a. Impulse response function, response to a 1 standard deviation fiscal shock.
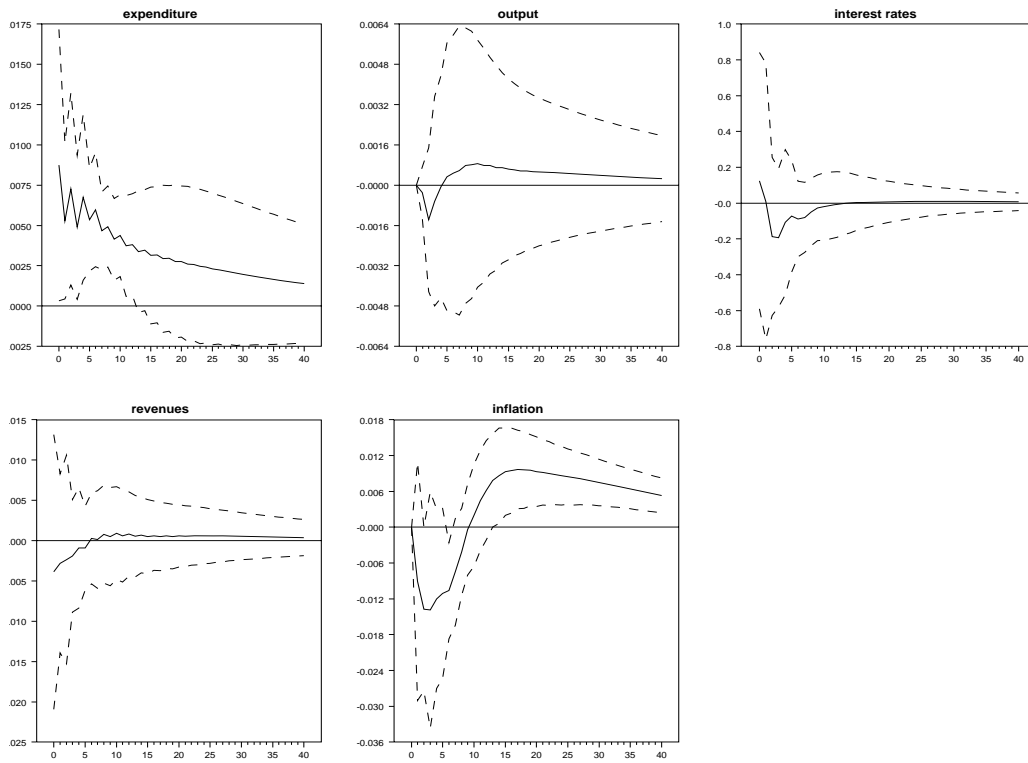
Figure 4b. Impulse response function, response to a 1 standard deviation fiscal shock, government expenditure = consumption + government investment + net interest payments.
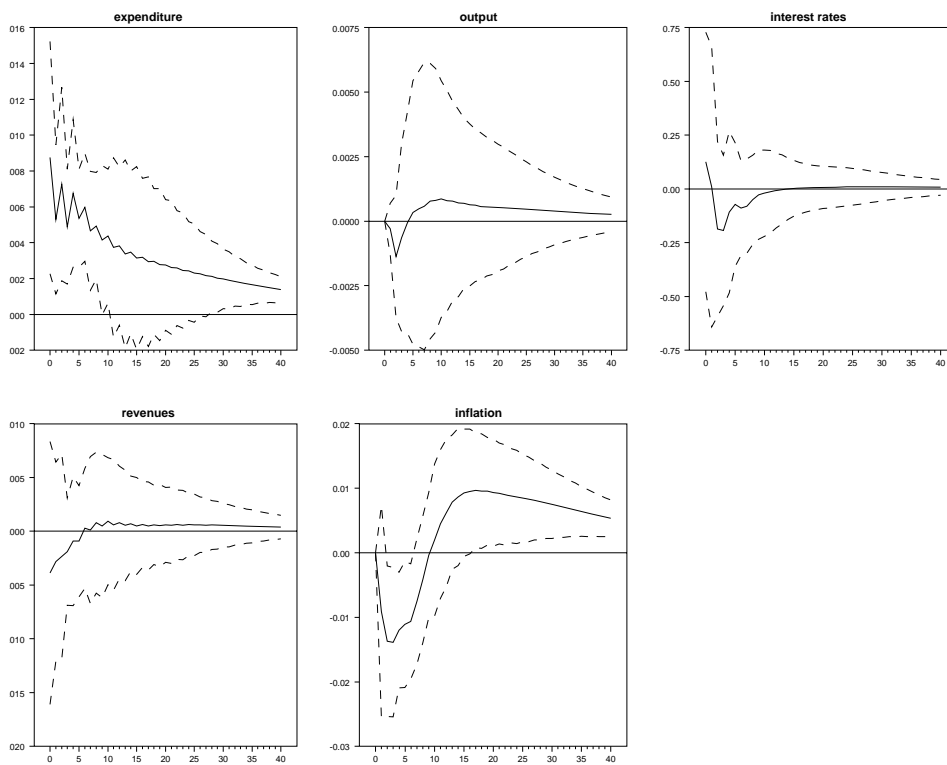
Figure 4c. Impulse response function, response to a 1 standard deviation fiscal shock, government expenditure = consumption + total interest payments.
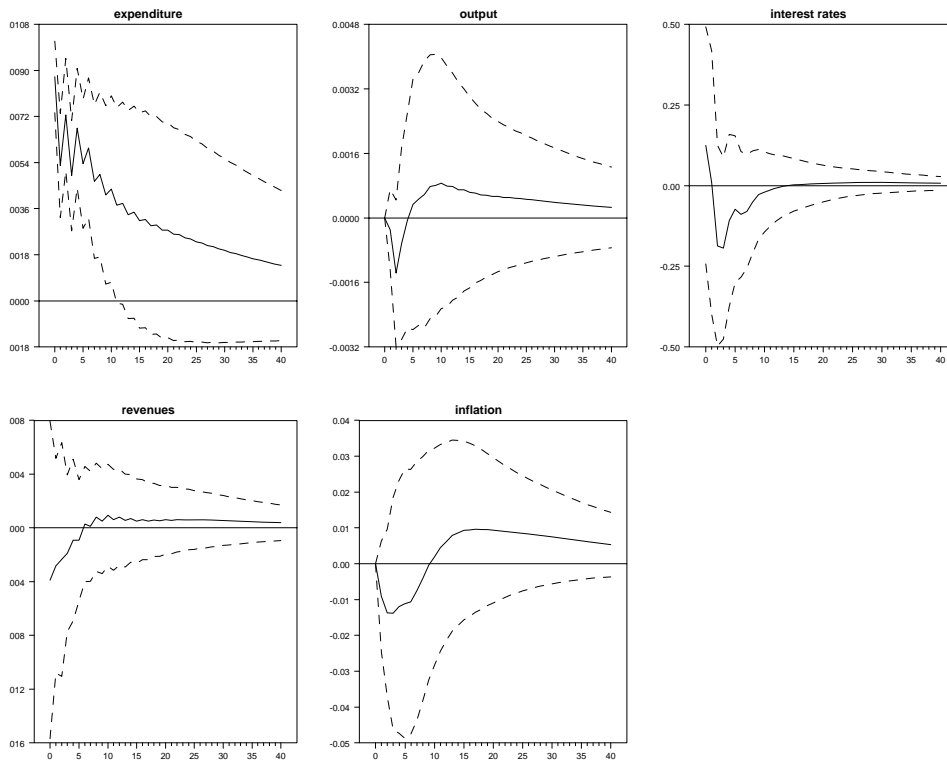
Figure 5a. Series and permanent component, from (12).

—— series —— permanent component
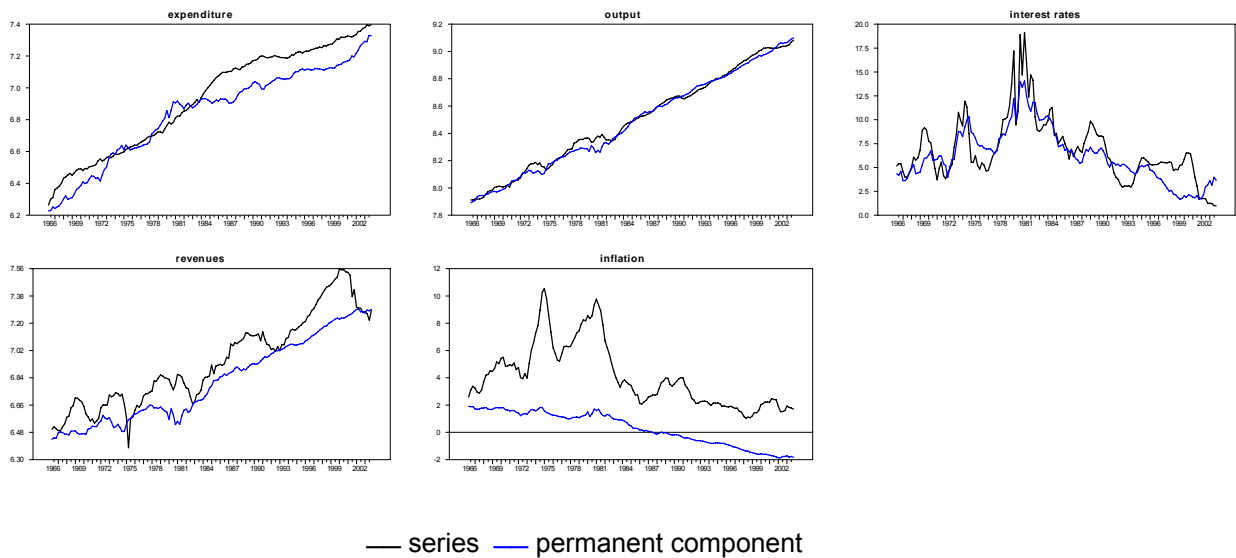
Figure 5b. Series and transitory component, from (12).
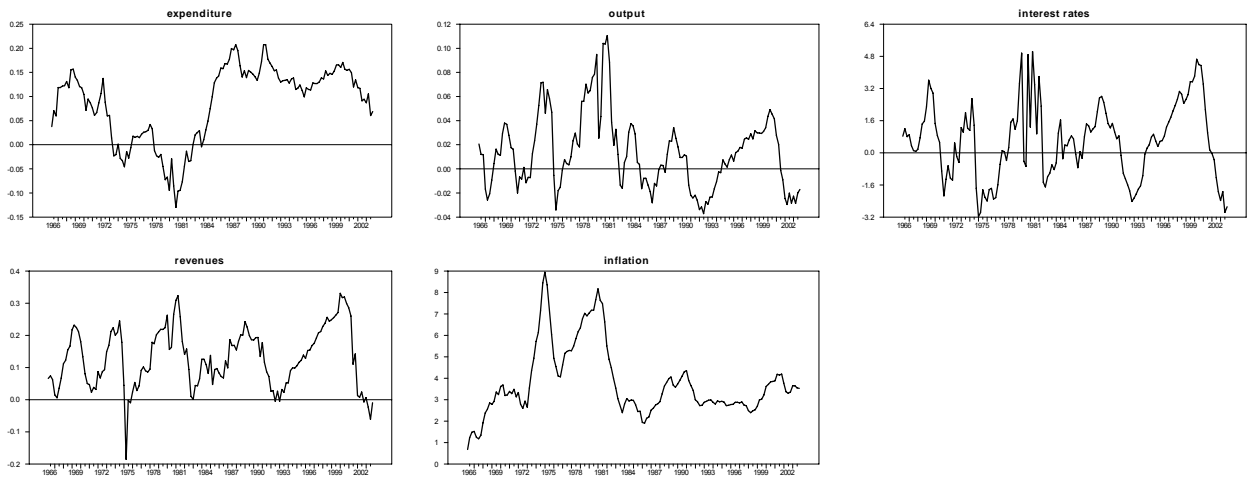
Figure 6. Fiscal policy shocks, standardised residual, from (13).
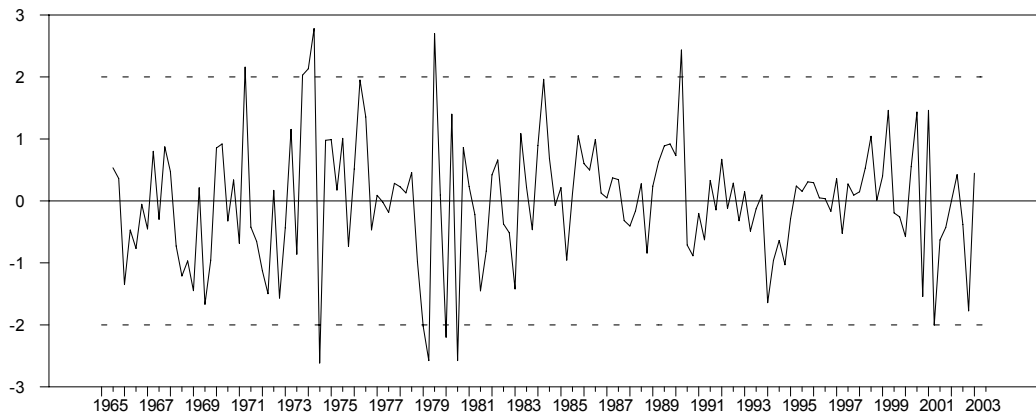
Figure 7. Volatility of fiscal policy shock, rolling window of 20 quarters.
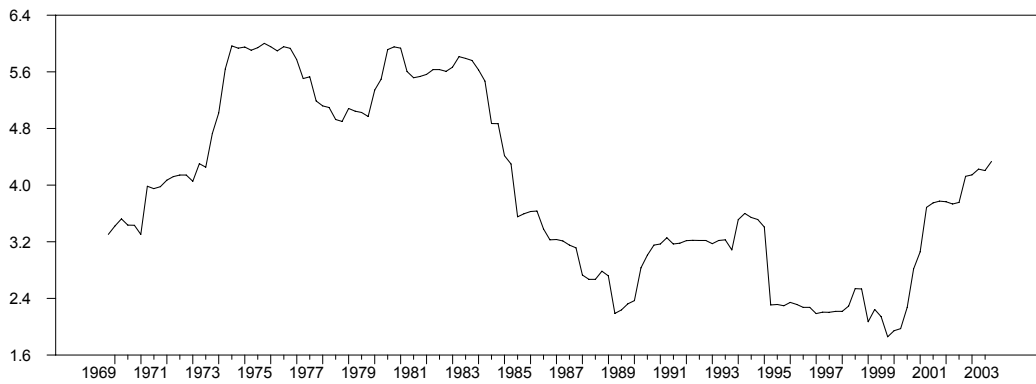
Figure 8a. Stability analysis for VAR – breakdate estimation (homoskedastic version).

Figure 8b. Stability analysis for VAR – breakdate estimation (heteroskedastic version).

Figure 9a. Stability test - Recursive trace test for rank (Hansen and Johansen, 1999).[a]
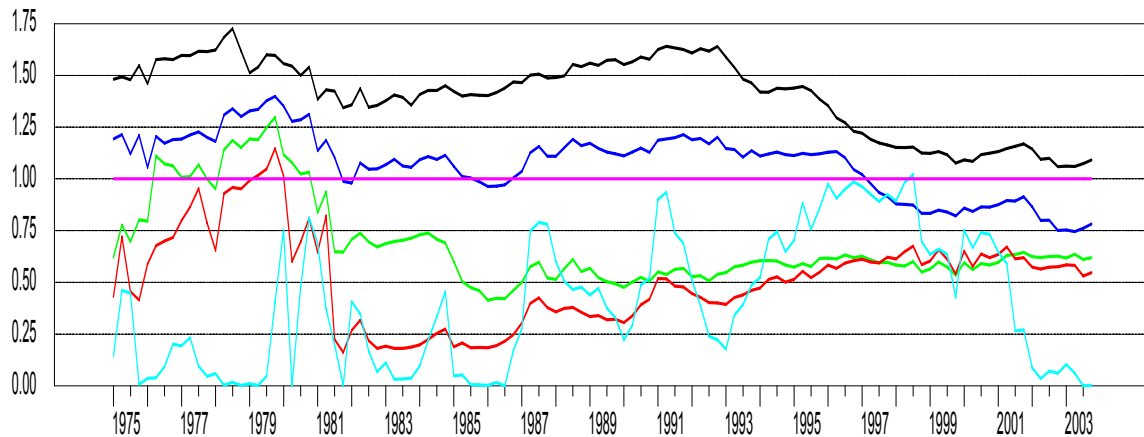
Note: (a) scaled critical values: 1 is the significance level at 5%.

Figure 9b. Stability analysis on VEC – recursive eigenvalues.

recursive eigenvalue 1

recursive eigenvalue 2

recursive eigenvalue 3
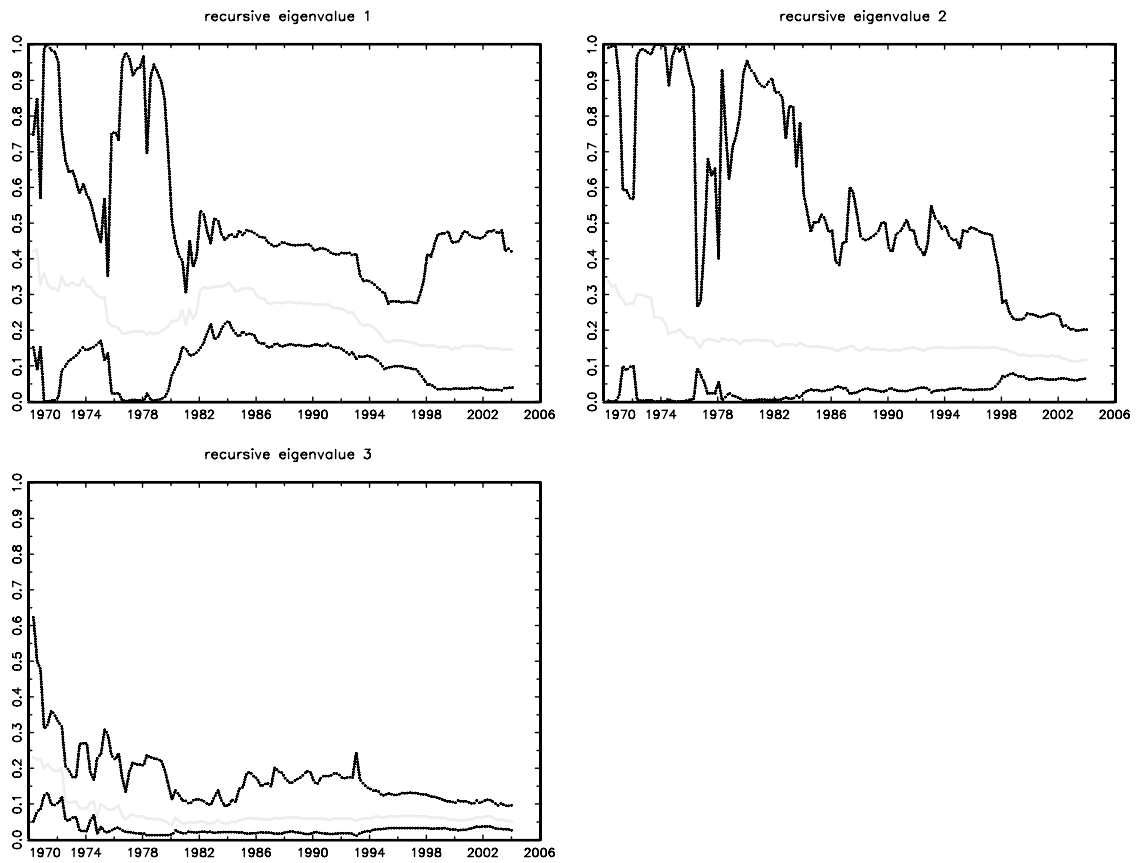
Figure 9c. Stability analysis– c-test on stability eigenvalues (Hansen and Johansen, 1999).
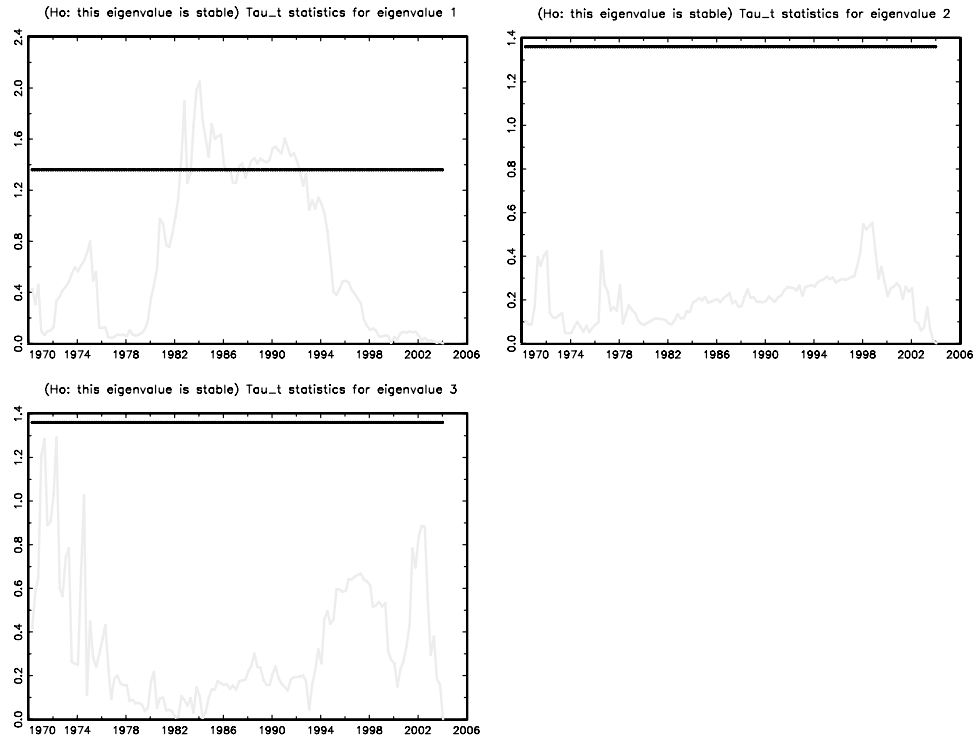
Figure 9d. Stability analysis – c-test on sum eigenvalues (Hansen and Johansen, 1999).
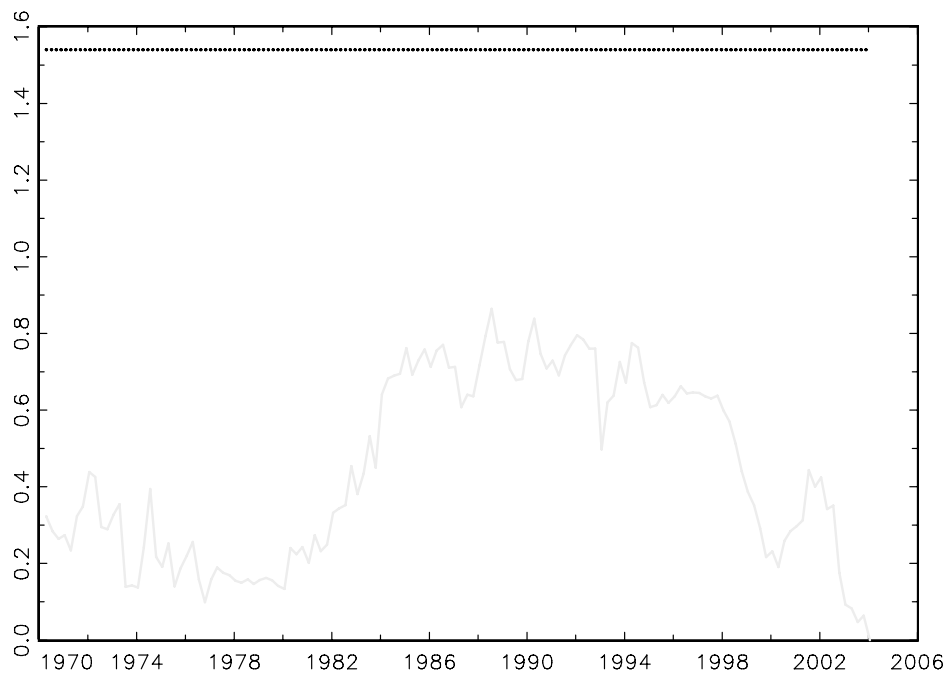
Figure 9e. Stability analysis – c-test on sum eigenvalues (Hansen and Johansen, 1999).
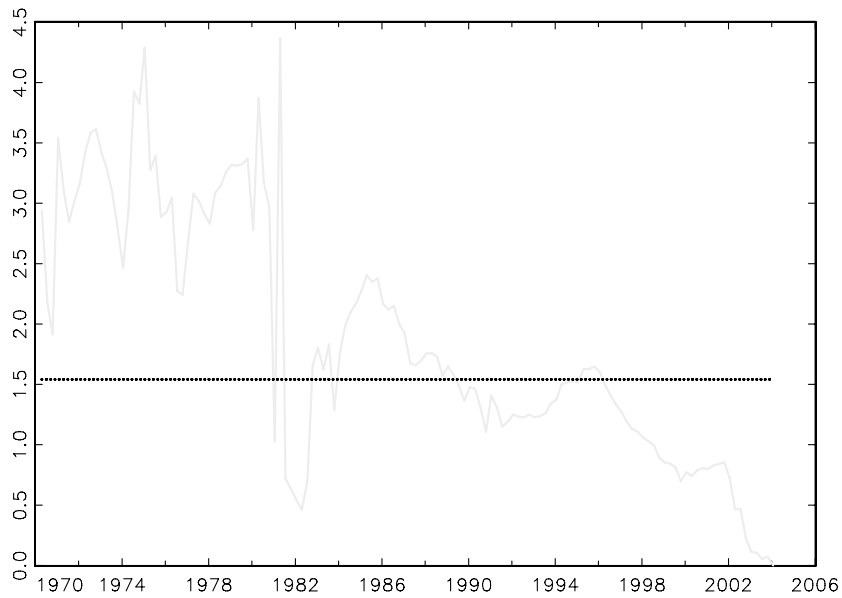
Figure 10a. Impulse response function, 66% asymptotic confidence interval, response to a 1 standard deviation fiscal policy shock, theoretical cointegration vectors.
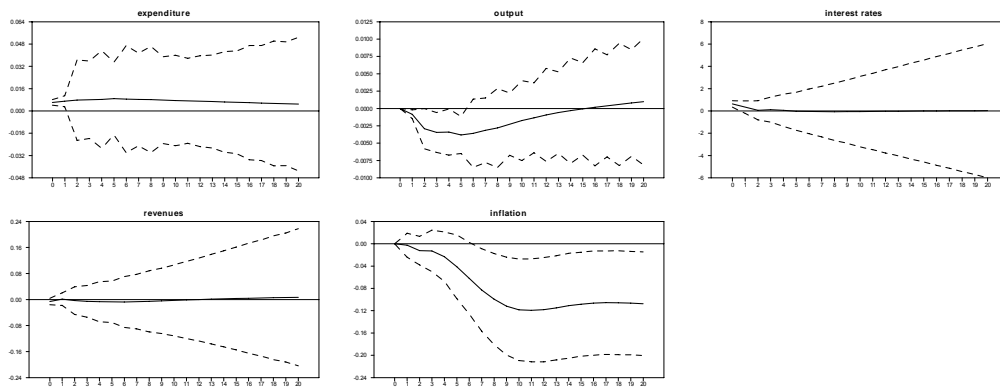
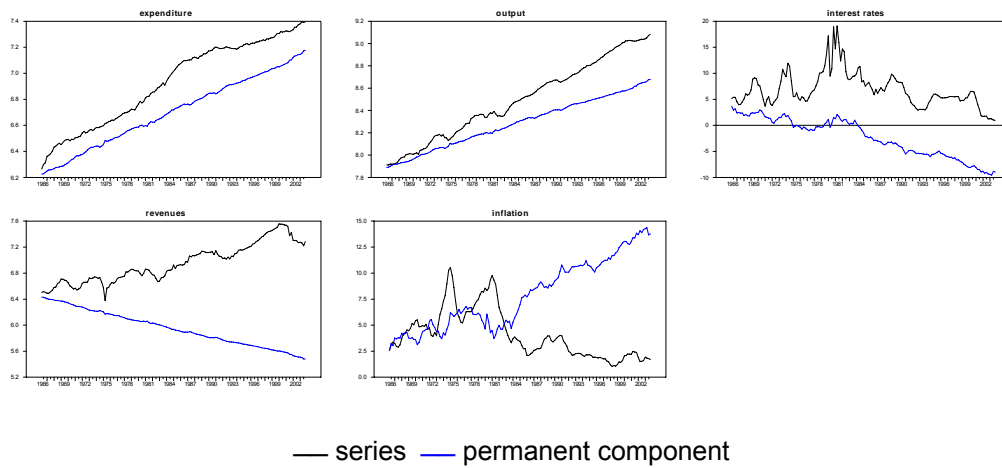**Figure 10b. Series and permanent component: theoretical cointegration vectors.**

series — permanent component

Figure 11a. Series and permanent component: interest payments in cointegration relation.

series — permanent component
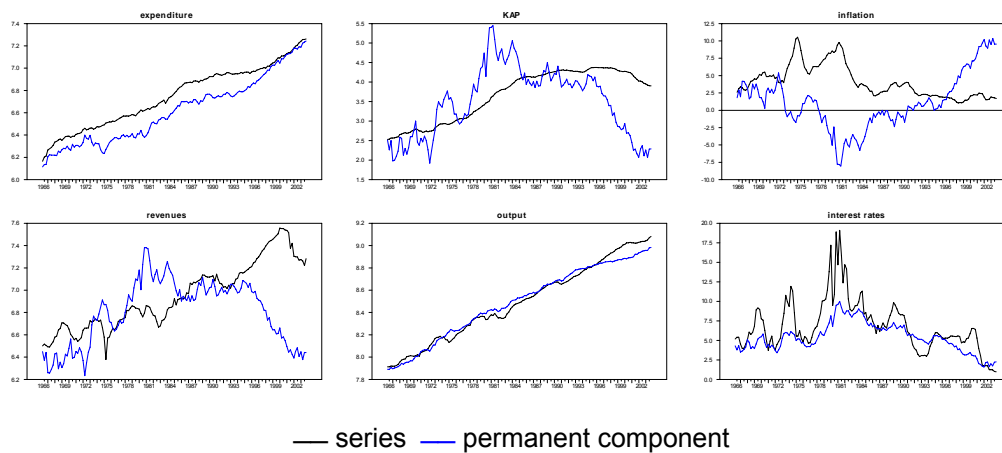
Figure 11b. Series and permanent component: debt in cointegration relation.
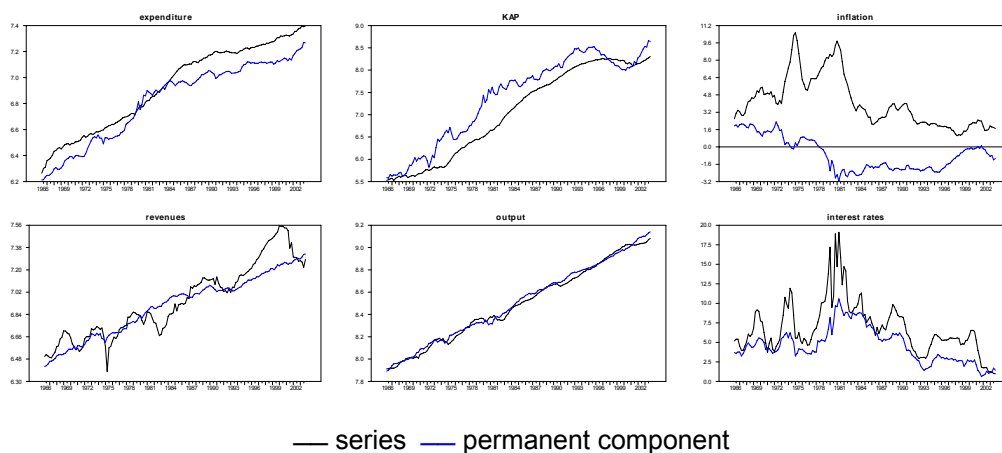

series — permanent component

Figure 12a. Fiscal policy shocks, standardised residual: interest payments in cointegration relation.
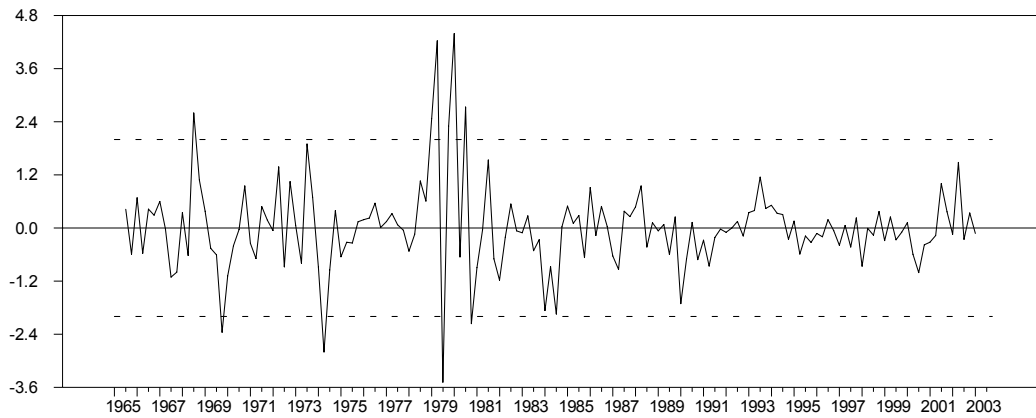

Figure 12b. Fiscal policy shocks, standardised residual: debt in cointegration relation.