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“Taking the government budget constraint seriously can overturn some widely held beliefs about policy effects.” (Leeper and Nason, 2005)

Estimating the effects of fiscal policy under the budget constraint

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\textbf{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 main finding is that fiscal consolidation has expansionary effects on output and inflation. Non-Keynesian effects also dominate when debt expands. The expectation of policy adjustments to guarantee fiscal sustainability by future tax rises or spending cuts contracts output today.

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1. INTRODUCTION

Drafting of the annual government budget suscites 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 and the stabilisation of the aggregate economy. Policymakers usually take it for granted that fiscal expansions have expansionary effects on the economy in the short run. Opinions about fiscal policy in the economics profession are rather diverse. Ricardian Equivalence is a reasonable starting point for the theoretical analysis of the effects of fiscal policy. Few economists would endorse it as a realistic description of fiscal policy, however. There are few definite conclusions on the economic effects of changes in fiscal policy. Different theoretical models can predict contrasting responses of the main economic aggregates.¹ However, all studies would agree that increased government spending or tax cuts have positive output effects in the short run.

Empirical evidence does not provide unanimous results either. 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 is the positive Keynesian effect of fiscal expansions on consumption, albeit the total multiplier effect is small. Studies that make a detailed historical study of policy decisions (Romer and Romer, 2007) or use information on large fiscal events that may be assumed exogenous (e.g. the timing of wars and defence spending) find small or negative effects of fiscal policy on consumption (Ramey and Shapiro, 1998).² However, all studies would agree that increased government spending or tax cuts have slightly positive output effects in the short run. In contrast, spending cuts or tax rises may have positive – non-Keynesian – effects on output if the government consolidates debt. Like VAR studies, these papers look at the effect of fiscal policy after a large fiscal event, in this case a successful consolidation (Alesina and Perotti, 1997).

How can we explain this puzzle? Should the VAR and event studies not give a similar answer? After all, large fiscal events should be similar in both types of studies, as both involve large shifts in spending or taxes. One can name several reasons for the lack of robust results of fiscal VARs,

¹ See Perotti (2007) for a comparison of the theoretical models, and their different empirical predictions.
² Studies that impose ‘sign’ restrictions on fiscal VARs do not find significant responses either (Mountford and Uhlig, 2005).
such as aggregation bias, anticipation, non-invertibility, etc.. One important aspect often ignored in the VAR analysis of fiscal policy is public debt. Fiscal policy is subject to a budget constraint. 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. In addition to the response of spending or taxes to economic variables 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. This has two serious consequences for the interpretation of fiscal VARs (Sims, 1998; Favero and Giavazzi, 2007). First, the 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. 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.

There are several ways to include the budget constraint in a VAR. Favero and Giavazzi (2007) directly include the stock of public debt in the VAR. Forni et al. (2007) include a fiscal tax rule that responds to debt in their calibrated VAR. Chung and Leeper (2007) impose the present value condition on an identified VAR. This indirectly imposes some restrictions on the parameters of the model (Hansen et al., 1991). In this paper, I propose to directly incorporate the budget constraint in the empirical VAR. Typical time series tests for fiscal sustainability are based on unit root behaviour of the fiscal surplus and debt, or the cointegrating properties of spending and revenues. I use this long-term relationship between spending and revenues to identify as a fiscal ‘solvency’ shock any deviation from long-term sustainability. Shocks to the budget constraint 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 three 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 is directly incorporated in the VAR. Second, I take an explicitly intertemporal view on the response of economic agents to the changes in fiscal policy. Finally, I reconcile the non-Keynesian and VAR literature by identifying large fiscal events with jumps in public debt.

Ravn et al. (2007) argues that event studies identify anticipated shocks. This explains and Perotti (2007) argues that some atypical events dominate the the findings of VAR and event studies differ because Roberds (1991) and Carriero et al. (2005) develop a test of fiscal sustainability based on the violation of these cross-equation coefficient restrictions on the joint DGP of debt and surplus. Other studies develop a test for the Fiscal Theory of the Price Level (Canzoneri et al., 2001).
I apply this common trends model on US quarterly data since 1965. The main finding is the expansionary effect on output and inflation of fiscal consolidations. The results can be read as a non-Keynesian effect of fiscal policy due to expectations of fiscal adjustments by future tax rises or spending cuts to restore fiscal sustainability (Gordon and Leeper, 2005). Monetary policy is modestly reinforcing the wealth effect. Supply side effects of fiscal policy are less important.

The plan of the paper is as follows. The econometric set up of this common trends model, its identification and the interpretation of the structural shocks are detailed in section 2. I then analyse the effects of fiscal policy in section 3. I conclude in section 4.

2. THE SVAR-COMMON TRENDS ANALYSIS

2.1. 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 $X_t^T$.

$$X_t = X_t^p + X_t^T.$$  (1)

Neither component in (1) is observable as such. We can give content to the permanent and transitory components in the following way (Stock and Watson, 1988). There are some stochastic trends underlying the data series that are driven 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 $k = n - r$ 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 $X_t$ is (2)

$$X_t = X_0 + F\tau_t + \Phi(L)\nu_t, \quad \nu_t \approx WN(0, I_n),$$  (2)

where (a) $\Phi(\lambda)$ is assumed finite for all $\lambda$ in and on the unit circle; (b) $X_0$ is stationary; and (c) $\nu_t$ 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 $\tau_t$ is a random walk (3) with drift $\mu$ and innovation $\phi_t$:

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

The common trends model then decomposes the time series $X_t$, as in (4):

$$\begin{cases}
X_t^T = X_0 + \Phi(L)\nu_t \\
X_{t+1}^p = F[\tau_0 + \mu t + \sum_{i=1}^r \phi_i] 
\end{cases}$$

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 $\beta$. 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 $m$ as in (5)

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

and if $X_t$ is cointegrated of order $(1,1)$ with $r$ cointegrating vectors, then we know from the Granger Representation Theorem that $\text{rank}[A(1)] = r$ and $A(1) = \alpha \beta'$, with $\alpha$ the loading matrix of adjustment coefficients on the $r$ cointegrating vectors. Equation (5) can then be rewritten as a VECM (6):

$$A'(L)\Delta X_t = \rho - \alpha \beta X_{t-1} + \varepsilon_t \quad \text{with } A'(\lambda) = I_n - \sum_{i=1}^{m-1} A^* \lambda_i, \ A^* = - \sum_{j=r+1}^m A_j.$$

If we rule out orders of integration larger than unity, there also exists a Wold VMA representation of (6) as in (7)

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

This can be rewritten as a common trends model similar to (4), 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(m) process:

$$\begin{cases}
X_t^T = X_0 + \tilde{C}(L)\varepsilon_t \\
X_{t+1}^p = C(1)[\xi_0 + \rho t + \sum_{i=1}^r \varepsilon_i]
\end{cases}$$
This is nothing else than the Beverdige-Nelson decomposition of $X_t$. In (8), $C(1)$ has reduced rank $k$ under the assumption of $r$ cointegrating vectors. I.e., only $k$ elements of $C(1)\xi_t$ result in independent permanent effects on $X_t$. As the loading matrix $F$ is not orthogonal to the elements of $C(1)\xi_t$, permanent shocks also have transitory effects on $X_t$. In contrast, transitory shocks have an effect on the temporary component only. 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 (6) and the Wold VMA representation as in (7).

Identification of the permanent and transitory shocks requires some assumptions on the long-term behaviour of the model. Call $\Gamma$ any regular matrix of dimension $n$ such that $\Gamma \Sigma \Gamma'$ is diagonal, and $R(1) \equiv C(1)\Gamma^{-1}$ the total impact matrix to innovations in the long term. Let $\eta_{i,t}$ be the i-th component of the vector $\Gamma \xi_t$. The matrix $\Gamma$ is said to identify the common trends model if: (a) $\Gamma$ is uniquely determined from the parameters in (6); (b) the covariance matrix of $\Gamma \xi_t$ is diagonal with non-zero diagonal elements; and (c) the total impact matrix $R(1) = [F \ 0]$. 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 (7) is equivalent to the structural model (9)

$$\Delta X_t = \delta + R(L)\eta_t,$$

where $\eta_t$ 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 $\beta$. 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

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5 Warne (1993) moreover derives the asymptotic impulse response functions under the hypothesis of a known finite
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,\(^6\) gives us a further $\frac{k(k+1)}{2}$ restrictions. The choice of $\pi$ 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 (6).

2.2. Specification and identification

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

$$\Delta X_t = \alpha \beta' X_{t-1} + A_1 \Delta X_{t-1} + \ldots + A_m \Delta X_{t-m} + \epsilon_t,$$

where $X_t = [y \ p \ i \ g \ t]'$. (10)

I assume three cointegration relations may be present in the DGP of $X_t$, each of which I associate with the temporary shocks in $\eta$. 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. I account for economic conditions by including output, inflation and interest rates. With three long-term relationships, there are two common stochastic trends.

The common real and nominal trend

Following the literature, I associate these two permanent components with a real and a nominal trend respectively (King et al., 1991). The real trend captures a technology shock with a

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\(^6\)That is, $\pi$ is lower triangular.
permanent effect on real variables. The nominal trend instead captures any (monetary) shock with a long-term effect on nominal variables.

With \( n = 5 \) variables and \( r = 3 \) cointegration relations, we already specify 15 parameters in the cointegration vectors. A further six restrictions come from the orthogonality of the three cointegration vectors to the \( k = 2 \) common trends. The choleski ordering of uncorrelated permanent shocks gives us three more restrictions. This leaves one further restriction to be imposed. I distinguish the real from the nominal trend by assuming that the nominal trend has no long-term effect on real output. This provides us with a decomposition of the series into permanent and transitory components that carry an economic interpretation within the model.

The shocks that have only a temporary impact on the economy come from the following three cointegration relations:

\[ \text{The fiscal 'solvency' shock } (g_t - t_t) \]

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 \( g_t \).

\[ b_t \equiv (1 + r_t) b_{t-1} - s_t \quad \text{where} \quad s_t = t_t - g_t. \quad (11) \]

Without loss of generality, one may express the fiscal variables in nominal or real terms, and we accordingly denote by \( r_t \) the appropriate nominal or real return on debt. Solving forward (11), one obtains the intertemporal budget constraint IGBC (12).

\[ b_t = \lim_{n \to \infty} E_t \left[ \prod_{j=1}^{n} \frac{b_{j+n}}{1 + r_{t+j}} \right] + \sum_{n=0}^{\infty} E_t \left[ \prod_{j=1}^{n} \frac{g_{j+n} - t_{j+n}}{1 + r_{t+j}} \right]. \quad (12) \]

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

\[ \lim_{n \to \infty} E_t \left[ \prod_{j=1}^{n} \frac{b_{j+n}}{1 + r_{t+j}} \right] = 0. \quad (13) \]

Some equivalent tests for fiscal sustainability can then be derived under various assumptions on this condition. Standard (time series) tests have been based on the unit root properties of debt (Hamilton and Flavin, 1986). 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.\footnote{These assumptions require that: (a) agents hold rational expectations; (b) utility of consumption follows a random walk; and (c) the covariance of marginal rates of substitution with fiscal variables is time-invariant. See Ahmed and Rogers (1995) for a proof.} It then follows that the total government deficit series is stationary when spending and revenues are cointegrated.

There are two alternative ways to include the budget constraint in (10). Theoretically, from (12), cointegration implies that the coefficients on spending and revenues are 1 and -1 respectively. Bounded sustainability – or a cointegrating vector $\beta' = [1, -1]$ with drift – implies that public debt is finite in the long run. An alternative to this theoretical condition is the empirical condition for sustainability in-sample. This weaker condition leaves the cointegration coefficients in $\beta$ unspecified (Quintos, 1995). 'Weak' sustainability implies the undiscounted debt process is exploding at a rate slower than the growth rate of the economy. I do not impose strong sustainability on the cointegrating relation between expenditures and revenues, but examine in-sample sustainability.

I incorporate this cointegration relation between government spending $g_t$ and revenues $t_t$ in the VAR model (10). The structural innovation to the budget constraint is the fiscal 'solveny' shock. These are spending or tax policies that make fiscal policy stray away from a sustainable policy. Consider a hike in public spending, for example. For a given level of tax revenues, this leads – \textit{ceteris paribus} – to a deficit and higher public debt. This requires a future rise in tax revenues (or an equivalent cut in spending) for the intertemporal government budget constraint to hold. The anticipation of these offsetting policies has an impact on current economic variables, and these in turn determine the dynamic path of fiscal variables. 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. Bohn (1991) examines the adjustment following spending cuts or tax increases in a VEC model that includes fiscal series only.

There might be other innovations to the economy than fiscal ones that make a given debt position unsustainable, however (Lambertini \textit{et al.}, 2007). First, consider a permanent shock to technology. These innovations determine the growth potential of an economy. A negative shock is partially absorbed by public debt. Lower potential growth reduces tax revenues permanently.
The ensuing deficits make a similar debt position less sustainable *ceteris paribus*. In my model, these effects work out via the common real trend that affects the dynamic growth path of the economy. Actually, the common trends model does even better. The balanced growth path of the economy may be affected by both government spending and taxation. For example, if the hike in public spending falls on (productive) government investment, this would likely have permanent effects. 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 again the future responses – in taxes or spending. Second, higher (real) interest rates imply higher interest payments on outstanding public debt. The debt stabilising surplus is affected by monetary policy. A control for the endogenous reaction of interest rates is necessary, given the crucial role of monetary policy in the determination of sustainable fiscal plans. This is the reason for including interest rates and inflation in (10). In conclusion, the common trends model identifies ‘true’ fiscal shocks on the budget constraint.

This fiscal shock is not directly comparable to the ones commonly identified in fiscal VARs for two reasons. First, fiscal VARs that examine the effect of unanticipated changes in discretionary policy disregard the ensuing adjustment in expenditures or revenues. I.e., no attention is paid to the financing of the budget. Future fiscal adjustments constrain the short-term effects of current changes in fiscal policy. Second, by definition, the fiscal ‘solvency’ shock cannot be attributed to spending or revenues. In contrast to the identification of discretionary changes in spending or taxes, I only identify deviations from long-term budget balance. The examples given above can be applied to a cut in taxes, for a given level of spending.

The identification of shocks to fiscal solvency in a VAR model is probably closest to the identification procedure of papers that study the ‘non-Keynesian’ effects of fiscal policy. The typical two-step approach is to first identify periods of strong fiscal consolidation (Alesina and Perotti, 1997). This fiscal adjustment is usually measured by a large improvement in the cyclically adjusted primary balance (or some other ‘fiscal impulse’). Consequently, one selects those periods that are a ‘success’ in reducing the debt/GDP ratio after some years. Finally, they determine the consequences of fiscal consolidations on economic variables. My identification procedure generalises the ‘non-Keynesian’ procedure in two ways. First, I directly identify the policies that put spending and revenues on a sustainable path, and examine their economic effects. Second, the VAR is symmetric. I look at fiscal consolidations, but also at unsustainable

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8 I assume that the interest rate approximates the discount factor on future surpluses, as in Canzoneri et al. (2001).
9 The identification of changes in taxation is still fraught with difficulty, however (Romer and Romer, 2007).
policies that violate the budget constraint.

The business cycle shock \((y_t + \theta_t)\)

Public debt absorbs the effects of a permanent technology shock. Similarly, some of the variation in the budget owes to cyclical variations in the short-term, mainly due to automatic stabilisers. I contribute the co-movement of government revenues and real output to business cycle shocks. This allows identifying temporary economic shocks.

The underlying assumption is that government spending absorbs a determinate share of output; I take account of secular growth in the government sector with a real trend. Revenues are adjusted to finance public spending, but fluctuate because of short-term variations in the economy. 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 trend.

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

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 such an interpretation. Unlike bivariate tests for cointegration between a nominal interest rate and inflation, I have sufficient restrictions in the other parts of the common trends model to identify a monetary shock. There are three reasons. First, the real rate is not a constant: the real common trend, which reflects changes in productivity, shifts the natural rate of interest. Second, the nominal trend models any longer term trends in monetary policy (Vlaar, 2004). Third, transitory business cycle shocks have been taken out.

2.3. Some caveats on the common trends model

In contrast to most other fiscal VARs, I characterise both the long-term growth and business cycle features of a variety of macroeconomic models in which the effects of fiscal policy are derived subject to the government budget constraint. The common trends model addresses some of the empirical problems with fiscal VARs. The incorporation of the budget constraint in
the VAR allows tracking the adjustment of both spending and taxes after an initial fiscal shock. I also control for cyclical variation in fiscal variables. As I include public debt in the model, I avoid the non-invertibility problem of small fiscal VARs (Chung and Leeper, 2007). Including the budget constraint pays off if fiscal policy has non-linear effects (Giavazzi et al., 2000). The common trends model allows for such non-linear behaviour around the long-term relations in the model. Moreover, I do not restrict contemporaneous links between fiscal and economic variables with timing or institutional restraints. A large number of identifying restrictions are thus consistent with the steady state properties of economic models. I do not need to draw upon some \textit{ad hoc} assumptions.

But there are two caveats. First; as in other fiscal VARs, there still is a problem of handling 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 does not give a solution.

Second, long-term restrictions come at a cost (Sarte, 1997). Even in large samples, substantial uncertainty surrounds the estimates of the long-term inverted MA representation in (7). Hence, we cannot construct asymptotically correct confidence intervals on $C(I)$ (Faust and Leeper, 1997). Specifying \textit{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. This seems a very robust restriction anyway – as King and Watson (1997) demonstrate. 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.\footnote{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).}

3. ESTIMATING THE EFFECTS OF FISCAL POLICY

3.1. Data
All data series for US fiscal and monetary policy are on a quarterly basis. The budget law is commonly drafted every parliamentary year. Nonetheless, discretionary in-year revisions are not unusual, and cyclical responses are automatic. Whereas the latter effect mainly influences government revenues, the former mainly involves modifications to spending (Van den Noord, 2002). 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 spending.

The central bank decides more regularly on interest rate revisions. The intermediate choice of data frequency reflects a trade-off in the joint modelling of fiscal and monetary policy. Given a high degree of interest rate smoothing, the quarterly measure is probably not too coarse. 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.

Finally, 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.$^{11}$

3.2. 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

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$^{11}$ Fiscal and GDP data are from the BEA (NIPA tables), public debt from the FRED II database, deflators from BLS and the Federal Funds rate from the Federal Reserve.
derailing of public finances, and also reflect an I(2) process.

**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 \( \beta \)**
the three cointegrating relations were estimated in the VECM (6) via maximum likelihood. *A priori*, four lags are included in the VAR. Seasonal dummies are included. The 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 \( F_y \). 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.

\[
F_0 = \begin{bmatrix}
1 & 0 & 0 & \hat{\theta} & -\hat{\theta} \\
0 & 1 & 1 & \hat{\omega}_g & \hat{\omega}_t
\end{bmatrix}
\]  

(14)

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.\(^{13}\)

---

\(^{12}\) I reject this hypothesis, hence debt converges to a finite value.

\(^{13}\) I plot the impulse responses over 40 quarters, with the 66% asymptotic confidence bounds.
3.3. The non-Keynesian effects of fiscal policy

All fiscal VAR studies agree on a positive effect of fiscal expansion on output. In contrast, the dynamic responses of the common trends model reverses this standard results. Figure 2 plots the impulse responses after a one standard deviation shock to the budget constraint. We cannot attribute the fiscal shock to a rise in government spending or a cut in taxes directly. Fiscal expansions that move fiscal policy from a sustainable path seem to be mainly due to expansions in government spending. Tax cuts play a smaller role. I find that an expansionary fiscal policy has contractionary effects on output. The contraction is quite stark in the first year after the shock, but is not very persistent. Output slightly expands in the second and third year but returns to equilibrium afterwards.

Fiscal VARs disagree on the effects of increases in government spending on private consumption. Several approaches have been used to identify fiscal shocks to government spending. 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 is the positive Keynesian effect of fiscal expansions on consumption, albeit the total multiplier effect is small. Studies that make a detailed historical study of policy decisions (Romer and Romer, 2007) or use information on large fiscal events that may be assumed exogenous (e.g. the timing of wars and defence spending)(Edelberg et al., 1999; Burnside et al., 2003) find small or insignificant positive effects of fiscal policy on consumption. Both results can be justified by 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 on 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.14

However, all studies would agree that increased government spending or tax cuts have positive

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14 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.
output effects in the short run. How can we explain this non-Keynesian output effect of a fiscal policy shock? The reason is that I identify shocks to solvency that are a temporary innovation to the budget constraint and not to government consumption (or taxes) separately. 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). These results are in line with the predictions of some partial equilibrium models of fiscal sustainability. A fiscal consolidation that cuts spending to restore fiscal sustainability has positive wealth effects (Giavazzi and Pagano, 1990). Consumers that perceive a consolidation as a permanent fall in public spending, anticipate a future reduction in the tax burden. This raises their lifetime disposable income permanently. This wealth effect predicts that consumption increases.

A similar point can be made for a tax increase. A tax hike signals a change in the fiscal regime. Consumers discount this tax adjustment and do not expect further tax increases in the future. If this change in fiscal regime convinces consumers, they will expect future policy changes in the opposite direction, i.e. tax cuts (Bertola and Drazen, 1993).

According to this view, consumers will react more strongly to a fiscal adjustment if it is more credible. Hence, non-Keynesian effects will be more likely if initial conditions are bad (Sutherland, 1997; Perotti, 1999). If there is some urgency in consolidating debt, like in a fiscal crisis, the commitment of the government will be more credible. Davig (2004) shows in a switching model how (beliefs) of changes in the US public debt regime are sufficient to induce economic effects.

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 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.

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.

I identify all shocks to US public debt that lead to an adjustment in spending or taxes over time. Usually, most studies on fiscal consolidation do no use time series for a single country, and identify only a few large consolidations. The consequence is that the result differs from the usual findings in the non-Keynesian literature in a few aspects. First, I generalise the result: fiscal consolidations may boost output, but fiscal expansions that raise debt may have negative output effects. Finally, I identify all shocks to the budget constraint, whether they are small or large. Few studies identify any strong fiscal consolidation in US fiscal policy, if they do at all. I detect more than a few isolated fiscal events but derive a time series of shocks to the budget constraint. I plot the structural innovations to the budget constraint in Figure 3. It may be too much to require an exact timing or an appropriate size of these fiscal shocks. Nevertheless, periods of strong deviations away from or towards solvency should be easily discernible. Large outliers are exceptional, but some periods of fiscal insolvency 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 have declined since the eighties (around 1985-1987) and have been at a low level since (Figure 4). 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. The supposedly symmetric responses on a series with few large shocks may be the reason for the less precisely estimated impulse responses.

The forecast error variance decomposition gives some additional insights in the common trends model (Figure 5). Fiscal policy shocks have a small role in explaining variability in interest rates or fiscal policy at very short-term horizons. Systematic fiscal policy is more important than discretionary shifts. Variability in government expenditure is largely explained by shocks to solvency. This is consistent with spending being a largely independent process, driven by factors exogenous to the model. Fiscal policy has only a minor role in the variability of other series, in particular government revenues. This is a rather disappointing result: a consolidation with fiscal instruments, as in Bohn (1991), seems 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.

3.4. The role of monetary policy
A complementary monetary policy may strengthen the non-Keynesian effects. First, in situations of fiscal stress, public debt may carry an interest rate premium due to inflation or default risk. A credible fiscal consolidation that brings debt below some threshold value may cause a fall in real interest rates (Alesina et al., 1992). This supports economic activity, and boosts the wealth effect. Second, and more generally, 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 its expansionary output effect. Linnemann and Schabert (2003) or Galí and Monacelli (2005) show in a New Keynesian model that the effects of fiscal policy can vary with the monetary regime. Fiscal and monetary policy would appear as strategic substitutes. Moreover, strong changes in interest rates further burden the government with higher interest payments on outstanding debt. By absorbing some of the effects of future fiscal policy changes, there is a muted response of macroeconomic variables to non-systematic fiscal shocks (the so-called ‘Sims conjecture’).

I find that 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.\(^{15}\) In my model, the short-term interest rate falls in the first quarters after the shock, as in Mountford and Uhlig (2005).\(^{16}\) 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. I can discard the hypothesis that the monetary policy regime is the reason of the output contraction.

The results of the monetary part of the model are 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 6a). 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

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15 As in Favero and Giavazzi (2007), I find that interest rate responses are much more muted.
16 The interest rate falls in the first quarters after the fiscal shock
rate cuts reduce the tax base for capital taxation; output growth instead translates into higher tax revenues. Monetary policy shocks have a small role in explaining variability in interest rates or fiscal policy at very short-term horizons (Figure 5). 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.

Finally, the business cycle shock temporarily raises output and growth dampens out after about two years (Figure 6b). 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.

### 3.5. Supply side effects of fiscal consolidation

In addition to the expectations and monetary channel, a second view emphasizes the composition of the fiscal adjustment in determining non-Keynesian effects (Alesina and Perotti, 1997). A fiscal consolidation that cuts government wages and social spending has positive labour market effects. Notice that in contrast to the predictions of both RBC and New Keynesian models, fiscal policy cuts in specific areas have positive supply-side effects. If budget consolidation mainly falls on investment, or is achieved by tax hikes instead, there will be no such positive supply side effects.

As the common trends model includes a technology shock with permanent output effects, I can indirectly infer on the importance of supply side effects of fiscal policy. The comparison of dynamic responses to the economic shocks generally conforms to economic priors. A positive supply shock has permanent effects on output (Figure 7a). As one might expect, it also exerts a temporary deflationary impact on the economy. I find that supply shocks cause government spending to permanently expand. This is not the consequence of an increase in the available budget: higher potential growth does not lead to similar increases in tax revenues. We even observe an initial decline in tax intake in the first two years after the supply shock. As revenue windfalls do not fuel a rise in spending, the common trends model suggests there are important supply side effects of government spending, but much less so of taxation. However, these are mingled with ‘other’ technology shocks. Given the identification of the VAR, I cannot attribute these effects to the composition of spending. The nominal trend has a persistent impact on
inflation and interest rates, but no impact on real variables. These permanent effects are rather imprecisely estimated, however (Figure 7b).

At short horizons, most of the variability in the various series is accounted for by the transitory components (Figure 5). 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.

3.6. Robustness checks

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

VAR estimates are sensitive to structural breaks. This may particularly be the case when analysing debt. The effects of fiscal policy may be highly non-linear. I.e. at high debt levels, the standard results on fiscal policy can reverse as a reversal in the policy regime is expected. Many papers find evidence for this non-linearity (Giavazzi et al., 2000). I perform stability tests on both the VAR-coefficients and the cointegration tests to detect possible breaks.

Following Bai et al. (1998), I first construct a sequential Quandt-Andrews likelihood ratio test for checking structural breaks in the joint DGP of $X_t$ (Table 3, Figure 8a-b). The (single) breakdate is the supremum value of this recursively generated series over the central 70% of the sample. I scale the latter series by the change in residual variance before and after the initially estimated breakdate. 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 the ‘Volcker-Greenspan’

\footnote{Applying this heteroskedasticity correction more correctly locates the break if the data series did not only change its first, but also its second moments (Stock and Watson, 2003).}
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 heteroskedastic 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 reduced form evidence, I cannot tell apart the importance of shifts in either fiscal or monetary policy. 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 regime of ‘passive’ fiscal policy and ‘active’ monetary policy is crucial in bringing about macroeconomic stability.

To get some more insight in the break tests of the unrestricted VAR, I subject the structural VEC model to some recursive stability tests. The common trends model is non-linear because of its decomposition in permanent and transitory components. A breakdown of cointegration indicates a faulty identification. Unfortunately, I cannot test for different responses to negative and positive solvency shocks, but only for time varying responses. I can only test whether the dynamic adjustment to shocks changes with the level of debt.

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 a constant rank is rejected, I continue to test for constancy of the cointegration parameters. I first concentrate out the short-term parameters $\alpha$ 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 in the cointegrating relations. 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 $\tau$-test of Hansen and Johansen (1999). 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 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 $\tau$-test on the sum of eigenvalues is also strongly in favour of parameter constancy (Figure 9d).

I can only square this parameter constancy in $\beta$ 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 $\tau$-test on all the parameters in $\alpha\beta'$. 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 decline gradually 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.

3.6.2. Model fit

There are several other ways to assess the robustness of the results from the common trends model. Following (8), the common trends model provides us with a decomposition of the series into (time-varying) permanent and transitory components. The permanent component can be directly compared to measures of potential output or cyclically adjusted government balances. We can also compute measures of core inflation and ‘natural’ real interest rates, as in Bagliano and Morana (2003).

The underlying real and the nominal stochastic trends that I extract from the model do make sense (Figure 10). The permanent component gives a reasonable characterisation of the trends 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

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18 I 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.
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 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 (Table 4).

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. There is no correspondence to more standard core measures of inflation. The ‘Goldilocks’ US economic boom with steady economic growth and low inflation has often been attributed to the accumulation of positive supply shocks.

3.6.3. Data definitions

The results are confirmed for other definitions of the fiscal series and output. One wonders if the protracted crowding out effect is not simply due to (national) accounting conventions. One might expect that the multiplier effect of fiscal policy is reinforced by including government spending in GDP. No significant difference results by using total instead of private output, however. I have also excluded government investment from the budget as it likely has long-term effects on output. Including public capital building does not affect the findings, however. Finally, I use total capital transactions instead of including only net interest payments, but again the main
result remains. In summary, alternative data definitions all confirm the contractionary effects following fiscal expansions, and even lead to more pronounced responses to fiscal shocks.

4. CONCLUSION

This paper contributes to the growing literature on fiscal VARs. I reconcile the use of VAR methods with the literature on non-Keynesian effects of fiscal consolidation. I identify innovations to sustainable fiscal policies by incorporating the long-term government budget constraint in a common fiscal VAR specification. The error correction structure of the common trends 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.

I find expansionary fiscal policy to have contractionary effects on output and inflation in the short run. These findings are in contrast to common results in the fiscal SVAR literature, but in line with the literature on non-Keynesian effects following fiscal consolidation. Expectations of an adjustment in spending or taxes that put fiscal policy on a sustainable path, may overturn widely held beliefs on policy effects (Leeper and Nason, 2005). 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, in contrast to what standard RBC or New Keynesian models predict (Gordon and Leeper, 2005). This expectational channel is slightly reinforced by a loosening of monetary policy, boosting the wealth effect of economic agents. Supply-side effects of fiscal policy are less important. An extension of the common trends model to analyse consumption and labour market responses to fiscal policy could validate the importance of the expectation or supply side effects.

The lesson we draw from this is that financing decisions of government policies matter, and may even outdo the effect of the initial change in fiscal policy. 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.

19 Results are not reported.
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### Table 1. Unit root tests.\(^{(a)}\)

<table>
<thead>
<tr>
<th>Test</th>
<th>level ADF(^{(b)})</th>
<th>level KPSS</th>
<th>first difference ADF(^{(b)})</th>
<th>first difference KPSS</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(\mu)</td>
<td>(\tau)</td>
<td>(\mu)</td>
<td>(\tau)</td>
</tr>
<tr>
<td>government expenditure</td>
<td><strong>-3.21</strong></td>
<td>-2.41</td>
<td>*<strong>3.50</strong></td>
<td>*<strong>0.66</strong></td>
</tr>
<tr>
<td>government revenues</td>
<td>-1.31</td>
<td>-2.44</td>
<td>***3.35</td>
<td>0.11</td>
</tr>
<tr>
<td>output</td>
<td>-1.08</td>
<td><strong>-3.88</strong></td>
<td>***3.57</td>
<td>*<strong>0.25</strong></td>
</tr>
<tr>
<td>inflation</td>
<td>-2.04</td>
<td>-2.44</td>
<td>***0.37</td>
<td>*<strong>0.31</strong></td>
</tr>
<tr>
<td>interest rate</td>
<td><strong>-2.91</strong></td>
<td>-2.88</td>
<td>*0.57</td>
<td>*<strong>0.57</strong></td>
</tr>
<tr>
<td>public debt</td>
<td>-0.47</td>
<td>-2.00</td>
<td>***3.59</td>
<td>*<strong>0.43</strong></td>
</tr>
<tr>
<td>net interest payments</td>
<td>-1.91</td>
<td>1.56</td>
<td>***3.40</td>
<td>*<strong>0.59</strong></td>
</tr>
<tr>
<td>total interest payments</td>
<td>-1.89</td>
<td>0.43</td>
<td>***3.40</td>
<td>*<strong>0.59</strong></td>
</tr>
</tbody>
</table>

**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) \(\mu\) for level stationarity, \(\tau\) for trend stationarity.

### Table 2. Trace tests for cointegration rank.\(^{(a)}\)

<table>
<thead>
<tr>
<th>null hypothesis</th>
<th>constant 5% critical value</th>
<th>constant and trend 5% critical value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Trace test</td>
<td></td>
<td></td>
</tr>
<tr>
<td>rank = 0</td>
<td>***167.83</td>
<td>76.81</td>
</tr>
<tr>
<td></td>
<td>***136.64</td>
<td>88.55</td>
</tr>
<tr>
<td>rank (\leq) 1</td>
<td>***68.32</td>
<td>53.94</td>
</tr>
<tr>
<td></td>
<td>***75.92</td>
<td>63.66</td>
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<tr>
<td>rank (\leq) 2</td>
<td>**38.08</td>
<td>35.07</td>
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<tr>
<td></td>
<td>*40.40</td>
<td>42.77</td>
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<tr>
<td>rank (\leq) 3</td>
<td>12.05</td>
<td>20.16</td>
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<tr>
<td></td>
<td>20.54</td>
<td>25.73</td>
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<tr>
<td>rank (\leq) 4</td>
<td>5.04</td>
<td>9.14</td>
</tr>
<tr>
<td></td>
<td>4.33</td>
<td>12.45</td>
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<tr>
<td>S&amp;L Test</td>
<td></td>
<td></td>
</tr>
<tr>
<td>rank = 0</td>
<td>***107.13</td>
<td>59.95</td>
</tr>
<tr>
<td></td>
<td>***110.81</td>
<td>66.13</td>
</tr>
<tr>
<td>rank (\leq) 1</td>
<td>**43.19</td>
<td>40.07</td>
</tr>
<tr>
<td></td>
<td>***58.53</td>
<td>45.32</td>
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<tr>
<td>rank (\leq) 2</td>
<td>13.32</td>
<td>24.16</td>
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<td></td>
<td>15.52</td>
<td>28.52</td>
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<tr>
<td>rank (\leq) 3</td>
<td>9.30</td>
<td>12.26</td>
</tr>
<tr>
<td></td>
<td>2.08</td>
<td>15.76</td>
</tr>
<tr>
<td>rank (\leq) 4</td>
<td>-</td>
<td>-</td>
</tr>
<tr>
<td></td>
<td>1.14</td>
<td>6.79</td>
</tr>
</tbody>
</table>

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

<table>
<thead>
<tr>
<th></th>
<th>QLR-test stat(^{(a)})</th>
<th>breakdate</th>
<th>confidence interval</th>
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</thead>
<tbody>
<tr>
<td>homoskedastic</td>
<td>38.38 (0.00)</td>
<td>1982:2</td>
<td>[1981:1; 1983:3]</td>
</tr>
<tr>
<td>heteroskedastic</td>
<td>50.24 (0.00)</td>
<td>1984:1</td>
<td>[1983:3; 1985:3]</td>
</tr>
</tbody>
</table>

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 4. Correlation of permanent and transitory components with common measures.

<table>
<thead>
<tr>
<th>temporary component of output</th>
<th>permanent component of output</th>
<th>permanent component of inflation</th>
</tr>
</thead>
<tbody>
<tr>
<td>potential NIPA</td>
<td>0.92 (0.00)</td>
<td>core inflation -0.29 (0.00)</td>
</tr>
<tr>
<td>HP-filter</td>
<td>0.91 (0.00)</td>
<td>HP filter -0.28 (0.00)</td>
</tr>
<tr>
<td>HP-filter</td>
<td>0.59 (0.00)</td>
<td>level, HP filter 0.44 (0.00)</td>
</tr>
<tr>
<td>linear trend</td>
<td>0.07 (0.87)</td>
<td>ratio, NIPA 0.38 (0.00)</td>
</tr>
<tr>
<td>quadratic trend</td>
<td>0.29 (0.00)</td>
<td></td>
</tr>
</tbody>
</table>
FIGURES

Figure 1. Data, United States, 1965:1-2003:4.

Figure 2. Impulse response function, 66% asymptotic confidence interval, response to a 1 standard deviation fiscal policy shock.
Figure 3. Fiscal policy shocks, standardised residual, from (9).

Figure 4. Volatility of fiscal policy shock, rolling window of 20 quarters.

Figure 5. Forecast error variance decomposition (FEVD).
Figure 6. Impulse response function, 66% asymptotic confidence interval.

(a) response to a 1 standard deviation monetary policy shock

(b) response to a 1 standard deviation business cycle shock
Figure 7. Impulse response function, 66% asymptotic confidence interval.

(a) response to a 1 standard deviation real supply shock

(b) response to a 1 standard deviation nominal shock
Figure 8a. Stability analysis for VAR – breakdate estimation.

(a) homoskedastic version

(b) heteroskedastic version

Figure 9. Stability tests.

(a) Recursive trace test for rank (Hansen and Johansen, 1999)

Note: scaled critical values: 1 is the significance level at 5%.
(b) VEC – recursive eigenvalues

(c) $\tau$-test on stability eigenvalues (Hansen and Johansen, 1999).
Figure 10. Series, decomposition in permanent and transitory components, from (8).

(a) series and the permanent component

(b) expenditure

(c) output

(d) interest rates

(e) revenues

(f) inflation

---

(d) $\tau$-test on sum eigenvalues

(e) $\tau$-test on sum eigenvalues

---

Figure 10. Series, decomposition in permanent and transitory components, from (8).

(a) series and the permanent component

(b) expenditure

(c) output

(d) interest rates

(e) revenues

(f) inflation
(b) the transitory component, from (8)