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Tax and Spend or Spend and Tax?  
An International Study

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1. Introduction

Traditionally, economists and political experts presume that the dynamics of budgetary developments are determined by government expenditure, assigning to revenues a rather passive and accommodating role. This assessment derives from the assumption that the long-run optimal supply of public goods and services is comparatively more autonomous than the sources of its financing and, as a consequence, policy makers in a first step determine government expenditures, while securing the flow of revenues sufficient to meet the public sector’s intertemporal budget constraint only in a second step.

This traditional view regarding the causal dependency between expenditure and receipts has been challenged by various government officials and experts.1 They argue that it is the development of revenues rather than of outlays which dominates the budgetary decision making process. From the perspective of policy makers, it is claimed, receipts are comparatively more autonomous than outlays so that the latter are adjusted to the former and not the other way round.

Although, at first glance, this change in interpretation of budgetary interdependence may seem to be of minor importance, upon closer examination it turns out to be highly relevant for the interpretation of the policy mechanisms determining fiscal policy.

First of all, this alternative view changes the common interpretation of the long-term increase in the share of government activities in GDP that we have experienced over the last decades. This increase is not seen to be caused by a permanently higher income elasticity of the demand for public goods compared with private goods as argued by traditional public finance theory (Wagner’s Law), but to originate in too high an elasticity of government revenues arousing sectional demands from various pressure groups and lobbyists. As a consequence, an administration anxious to reduce the share of government activities in GDP has to start by cutting receipts; any attempt to cut spending without previously cutting receipts is doomed to failure.

Secondly, this alternative view also changes the interpretation of budget deficits. Deficits are no longer understood as the result of economically or politically induced expenditure shocks compensated for by subsequent shifts in revenues (Peacock-Wiseman 1979) nor as the outcome of a tax smoothing process that minimises dead weight losses of financing public services (Barro 1979). On the contrary, they are interpreted as the result of revenue shocks, which must be followed, if governments want to remain within their intertemporal budget constraints, by expenditure adjustments.

The advocates of this view, therefore, do not see any inconsistencies in simultaneously calling for deficit reductions and tax relief (Evans 1981, p.103; Hall 1982,p.225). Moreover, analogously to the intensively discussed problem of how to implement credible monetary policy strategies (see Giavazzi-Pagano 1988), they even interpret this proposition as a fiscal policy device. Only credible signals of a “dry” tax policy can show the unwillingness of an administration to finance current deficits through future tax increases, i.e. signal a hard budget constraint that reduces demands from pressure groups and lobbies.

Over the last years, the question of which causal relation between government outlays and receipts prevails in practice has been met with increasing interest in the economic profession. Various attempts have been made to discriminate empirically between them (Anderson et.al. 1986, Manage et.al. 1986, von Furstenberg et.al. 1986, Ram 1988a,b, Miller-Russek 1990, Bohn 1991, Koren-Stiassny 1992) but the evidence remains inconclusive.

In this paper we propose an empirical testing procedure to discriminate between the competing "spend and tax" and "tax and spend" hypotheses using data for nine industrialised countries at federal level. The countries have been chosen with a view to encompassing a broad variety of tax systems, underlying redistributional objectives and general economic and political attitudes that may be relevant for the issue under consideration.

The estimation procedure is implemented by using a tri-variate, structural VAR model. We compute impulse responses to investigate the response of future taxes and spending to shocks in current taxes and spending. We further examine the frequency domain properties of the estimated VARs. The stronger the causal dependency of taxes on expenditures and the more periods of budgetary consolidation are characterised by tax increases, the more we will tend the see the "spend and tax" hypothesis confirmed. The more the inverse relationship holds, the more we will be willing to accept the alternative "tax and spend" hypothesis.

In section two of this paper we try, in line with standard public choice reasoning, to give some explanations for the causality direction that can be derived from a country's specific political and economic characteristics. Section three presents in some detail the methodology used for our empirical analysis. The empirical results are presented in section four. Section five summarises the main findings and draws some final conclusions.

2. Can Higher Revenues Increase Expenditures and the Deficit?

Traditional public finance theory interprets the process of political decision making as the action of a benevolent dictator. Outlays and revenues are thought to be determined simultaneously so as to maximise a society's intertemporal social welfare function (Musgrave 1985). However, simultaneity does not mean that the budget must be in balance at all times. Differing time profiles of expenditure and revenue flows can easily bring the government budget into transitory deficit or surplus. But anticipated tax smoothing strategies (Barro 1979) do not give rise to exclusive or one-sided relationships between revenues and expenditures.

Although rigorous public finance theory can hardly derive a specific causality direction between the two sides of the budget, there seems to prevail a wide-spread belief among economists that government expenditures determine budgetary developments. As already mentioned, an early and highly plausible justification for this view has been offered by Wagner's Law. More general explanations of the historical development of the public sector stressing the importance of changes in a society's voting rules (extension of the franchise) and in the distribution of income among the electorate (Meltzer-Richard 1981), also seem to be in accordance with the spend and tax hypothesis.

The long-term and significant increase of budget deficits and government debts in many industrialised countries over the last two decades has initiated a wide body of research analysing the political and economic causes and consequences of debt financing. Many of these approaches have been very successful in explaining differences in the development of budget deficits in various countries. Some of the indicators and explanatory factors most often cited in these studies also tend to support the traditionally expected direction of causality between the budget aggregates.

The concept of fiscal illusion (Buchanan-Wagner 1977, Rowley 1987), for instance, strongly supports the view that expenditures dominate the budget. This results from the basic assumption of a short time horizon of both voters and policy makers. Since expenditure programs can be manipulated more easily than complex tax legislations, we would expect the expenditure dominance to increase with the shortening of the planning horizon.
A similar conclusion can be drawn from those studies which, following the pioneering work of Alesina-Tabellini (1987,1990), stress the importance of *time inconsistencies* resulting from disagreements between current and future majorities in the constituencies, and thus the stability of the political system as the major factor explaining the magnitude and development of government deficits. Under short-lived governments and sharp political swings associated with every change of administration, societies are presumed to accept higher deficits than under stable administrations with a longer average tenure and minor ideological discrepancies. Incumbent governments want to force their successors to implement unpopular consolidation measures and to impede new public programs that are incompatible with their own political preferences. Likewise, multi-party coalition governments with sharply diverging preferences and a strong need for internal compromise are expected to run significantly higher deficits than single party governments.

Another important factor often seen as being of central relevance for the level of public debt and deficits is the degree of central bank independence. Central bank independence differs significantly among the OECD countries (Bean 1992, Alesina-Summers 1993). As a consequence, government access to central bank profits or to non-interest bearing central bank funds varies. Given their short time horizon, policy makers tend to underestimate the effective costs of financing public debt by seniorage and inflation taxation. Governments are therefore often presumed to find deficit financing the more attractive the lower the degree of effective central bank independence (Burdekin-Wihlborg-Willett 1992). Again, we would expect that such a softening of the government's budget constraint may favour *spend and tax*.

The alternative hypothesis, i.e. the assumption of government outlays being dependent on receipts, has found far less attention. Commonly it has been seen relevant only for subnational and local governments which are endowed with very limited tax raising powers. As, in addition, their borrowing capacities are often constrained by constitutional law and/or limited by effective credit rationing, these entities are highly dependent on central government grants, i.e. a source which is exogenous from their point of view. The same outcome can be expected from strict revenue appropriations. Whenever specific tax revenues or contributions are earmarked for specific categories of outlays, there is good reason to expect causality running from taxes to spending.

During the prosperous years of the 1960s, some authors surmised that the progressivity of the tax system would cause the built-in elasticity of government revenues to exceed the dynamics of government outlays (Heller 1967). As it was assumed in addition that political decision makers were reluctant to run permanent budget surpluses and that they were not interested in cutting taxes either, a revenue-driven long-term increase of the government's share in total GDP was expected. It is under dispute whether steps towards economic and monetary union in Europe will strengthen budgetary discipline or will, on the contrary, enhance tendencies towards overborrowing (Buiter 2002).

---

2 Already two centuries ago, Adam Smith feared that the availability of credit facilities to the government would induce a (suboptimal) increase of public sector spending above the level that would prevail in a world where outlays have to be financed out of current revenues (Smith 1776, V). Keynesian models of demand management commonly are also primarily interested in analysing expenditure multipliers while revenues are treated as implicit budget constraint (see e.g. Blinder-Solow 1978).

3 Roubini-Sachs 1989, on the contrary, expect that a change in seniorage does not alter the magnitude of the deficit but only the composition of its financing.

4 This coincides with the findings of some recent studies aiming to demonstrate that budget deficits and public debt are significantly lower in countries with a high degree of fiscal federalism compared to those with a high degree of centralisation (Moesen-Van Rompuy 1990).

5 Meanwhile, however, the built-in elasticity of most countries' tax systems is lower than the dynamic of outlays, so that in fact there is no tendency towards budgetary surpluses anymore (Mckee-Visser-Saunders 1986). It remains unclear whether the logic of Heller's argumentation also holds in this case.
Many experts argue that under credible non-bailout arrangements efficient credit markets will charge high-risk premia on uncontrolled credit demands, thereby preventing individual countries from debt hazard and overborrowing. Others, most notably many central bankers, doubt the credibility of such non-bailout arrangements and, thus, the effectiveness of credit markets in penalising overborrowing. It appears, however, that if economic integration affects fiscal policies, this could also have an impact on the causality direction of the budgetary process. This seems plausible because the regions and countries within the integration area are forced to compete with regard to their attractiveness as an economic and industrial location, this attractiveness being dependent, among other things, on comparative tax advantages. National tax autonomy is therefore significantly reduced, so that it becomes increasingly difficult for national authorities to finance current deficits out of higher future taxes (Streissler 1989).

A further consequence of international integration is the increasing concern addressed to the externalities of budget deficits. This is especially true for the budget deficits of major economies, as the international community feels adversely affected because of crowding out effects induced by an increased level of world interest rates. These international pressures to reduce budget deficits clearly limit national budget autonomy (De Grauwe 1992), and may thus have some effects on the causality direction.

3. The Methodology: Identification, impulse response functions, frequency domain techniques and error correction forms

To discriminate between the two hypotheses elaborated above we will first of all examine the causal relation between the development of government taxes and spending. In addition, as we assume that the government's intertemporal budget constraint is strictly met, we can investigate whether budgetary imbalances, which are deemed intolerable in the long run, are reduced predominantly by expenditure restraint or by tax increases.

Commonly, impulse response functions of bi-variate VAR models are used to analyse causality relations between time series (Bohn 1991). However, such a representation is insufficient for our purposes as the variables in question are highly dependent on the development of aggregate income. Neglecting this joint dependency might easily lead to biased results. We therefore estimated a tri-variate, structural VAR model using nominal GDP as an additional explanatory variable. This approach allows us to distinguish between the direct causality relation between spending and taxes and the indirect causality effects via GDP.

We interpret the development of the net deficit as an error correction process, defining any deviation of the development of the net deficit from its long-run constraint as an error that the government must correct by adjusting revenues (increasing taxes) and/or outlays (reducing spending). We thus investigate which side of the budget in practice bears the burden of adjustment.

To account for a broad variety of socio-economic frameworks, estimates have been conducted with respect to nine industrialised countries. These countries are listed in Table 1. Following standard OECD classifications, five of these countries can be labelled as "major" countries, four as "small" countries. The main economic and political indicators that, following the considerations in the last section, might prove to be of relevance for the issue at hand are also listed in Table 1.

6 Note, that the often used Granger causality tests (Ram 1988a,b, Miller, Russek 1990), besides having other serious drawbacks, do not permit such a distinction.
The "openness" of the various countries, interpreted as the share of exports in GDP, ranges from 9% (USA) to 54% (Netherlands).

Central bank independence also differs markedly. Following standard indicators (see Alesina-Grilli 1991, Burdekin-Whilborg-Willett 1992), the central banks of Switzerland and Germany are endowed with the highest degree of independence, those of Italy, France and Sweden with the lowest. As Table 1 indicates, central bank independence seems inversely correlated with the average inflation rate.

Indicators of political stability (Roubini-Sachs 1989) show a significant North-South disparity. While average government tenure is longest in Germany, Austria and Great Britain, it is shortest in Italy. In addition, while the Northern countries are predominantly characterised by small (two party) coalitions or single-party governments, multi-party governments are common in the South. One should be aware, however, that such an indicator of "political stability" remains superficial.

Indicators of the fiscal regime also show significant differences among the countries under consideration. Central government budget deficits as a share of GDP have on average been highest in Italy (7.1%) and lowest in Switzerland (0.9%). Fiscal federalism seems most pronounced in Switzerland, the USA and Germany, with the share of central government revenues as a percentage of general government revenues achieving only 30% to 40%.

Table 1. Main economic and political indicators

<table>
<thead>
<tr>
<th>Exports as % of GDP</th>
<th>Central bank independence</th>
<th>Average inflation rate 1955-1992</th>
<th>Average tenure of government (Years)</th>
<th>Predominant political regime</th>
<th>Average fiscal deficit as % of GDP</th>
<th>Fiscal federalism</th>
</tr>
</thead>
<tbody>
<tr>
<td>Austria</td>
<td>39.7</td>
<td>3.5</td>
<td>4.0</td>
<td>6.9</td>
<td>single party, two party coalition</td>
<td>2.65</td>
</tr>
<tr>
<td>France</td>
<td>23.8</td>
<td>2.0</td>
<td>6.2</td>
<td>2.8</td>
<td>presidential with coalition government</td>
<td>2.34</td>
</tr>
<tr>
<td>FRG</td>
<td>30.5</td>
<td>4.0</td>
<td>3.2</td>
<td>6.9</td>
<td>two party coalition</td>
<td>1.80</td>
</tr>
<tr>
<td>Italy</td>
<td>20.5</td>
<td>1.8</td>
<td>7.9</td>
<td>1.7</td>
<td>multi party coalition</td>
<td>7.10</td>
</tr>
<tr>
<td>Netherlands</td>
<td>54.5</td>
<td>2.5</td>
<td>4.3</td>
<td>3.6</td>
<td>multi party coalitions</td>
<td>3.18</td>
</tr>
<tr>
<td>Sweden</td>
<td>30.8</td>
<td>2.0</td>
<td>6.4</td>
<td>6.0</td>
<td>small coalition</td>
<td>1.53</td>
</tr>
<tr>
<td>Switzerland</td>
<td>41.7</td>
<td>4.0</td>
<td>4.5</td>
<td>constitution al coalition</td>
<td>0.90</td>
<td>28.4</td>
</tr>
<tr>
<td>UK</td>
<td>24.1</td>
<td>2.0</td>
<td>7.1</td>
<td>5.7</td>
<td>single party</td>
<td>1.92</td>
</tr>
<tr>
<td>USA</td>
<td>9.4</td>
<td>3.5</td>
<td>4.5</td>
<td>5.0</td>
<td>presidential with divided legislation</td>
<td>2.04</td>
</tr>
</tbody>
</table>

1) IFS 1990 values; 2) See: Alesina,Grilli 1991; Burdekin, Whilborg, Willett 1992. Average of political and economic independence measured in terms of statutory independence, government representation on policy board, legislated objectives etc. The higher the index number the greater independence. 3) IFS, annual change of consumer prices. 4) Average tenure between 1945 and 1992 except France where only the 4th Republic has been
taken into account. 5) IFS data. 6) OECD Revenue Statistics; Central Government tax revenues as percentage of General Government tax revenues.

By its very nature, budget consolidation is a long-term task. In addition, VAR-estimations are data intensive procedures. Some authors therefore employed very long time series, some of them dating back to the 18th century, for testing fiscal dependencies (Bohm 1991). Compatible international time series of government finance are, however, not available for such a long time period. But a series of thirty-five to forty fiscal years, i.e. observations, should be sufficient for both methodological requirements and the requested adherence to the long-term budget constraint.

We used the data of the Government Finance Statistics of the IMF and chose a sample period from 1953 to 1992. Taking into account the required differencing and lagging of the variables, the effective estimation period begins with 1956. As we estimated a tri-variate VAR, the following variables were needed:

- $GDP$ (log (nominal GDP))
- $EXPR$ (log (expenditure ratio) = log (federal expenditures minus debt repayments) - GDP)
- $REVR$ (log (revenue ratio) = log (federal revenues) - GDP)
- $NDEFR$ (net deficit ratio = $EXPR - REVR$

Whenever possible the net deficit ratio was used as an error correction term. The figures for government expenditures, tax revenues and the net deficit are expressed as GDP ratios. There are several reasons for this: first of all, political and public discussions are usually in terms of GDP ratios so that estimation in ratios is an almost natural choice. Further, this transformation proved effective in reducing the degree of integration and in decreasing the dependence on nominal income. Moreover, as all variables exhibit a very strong trend, estimation in levels would have led to a severe multicollinearity problem.

All variables have been transformed into log-values to take account of the nonlinearity introduced by taking GDP ratios. Logarithms also facilitate imposing identifying restrictions, as these restrictions are now expressed in terms of elasticities.

For a meaningful interpretation of a VAR, identification of the model has to be assured. If this is not the case, impulse responses and causality spectra are not unambiguously determined, even if the parameterisation of the VAR is such that the error terms are uncorrelated. The reason for this is that, in general, many parameterisations exist which yield a diagonal covariance matrix of the VAR errors. To illustrate this consider the following VAR:

$$A(L)x_t = u_t,$$  \hspace{1cm} \text{(I)}

with $A(L) = A_0 + A_1L + \ldots + A_pL^p$, $A_0 = I$.

Let $\Sigma_u$ be the covariance matrix of the error terms, which in general is not diagonal. Now we can find a matrix $W$, such that $WAW' = \Sigma_u$ and $\Lambda$ is diagonal. The decomposition matrix $W$ could for

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7 Earlier fiscal data might be problematic for some countries as currency reform and debt rescheduling and international aid after World War II distort the data. Data for general government was not available on a comparable basis.

8 In using nominal GDP ratios, we implicitly assume that changes in real GDP and in the price level affect expenditures and revenues in the same way. Especially for expenditures this is clearly a problematic assumption, but after having weighed the costs and benefits of introducing a fourth variable we decided to make this simplification.

9 The term causality spectrum, i.e. causality coherence and causality phase was originally introduced by Granger 1969. We shall discuss a similar concept later in this section.
instance be obtained by the well known Choleski decomposition. In that case \( W \) is lower triangular. Premultiplying equation (1) by \( W^{-1} \) yields:

\[
W^{-1} A(L)x_t = W^{-1} u_t = v_t,
\]

or:

\[
A_0 x_t + W^{-1} A_1 L x_t + \ldots + W^{-1} A_p L^p x_t = v_t,
\]

with \( A_0 = W^{-1} \).

The covariance matrix of the new errors \( v_t, \Sigma_{vv} \), is now equal to:

\[
W^{-1} \Sigma_{ww} W^{-1} = W^{-1} W \Lambda W W^{-1} = \Lambda,
\]

which is diagonal by definition. Unfortunately, in general there exist many matrices \( W \) such that \( W A W' = \Sigma_{ww} \). Therefore a diagonal covariance matrix is not sufficient for identification. Further restrictions must therefore be imposed. These restrictions, which should be based on economic theory, are necessary to find a decomposition matrix \( W \) such that the corresponding new innovations \( v_t = W^{-1} u_t \) can be interpreted in an economically useful way.

Essentially, there are two procedures to accomplish this. The first is to impose restrictions on the lag structure of the model, especially on long-term effects, such that \( A_0 \) is unambiguously determined. This procedure has been applied with success to identify demand and supply shocks in VARs\(^{10}\). For our problem, i.e. identification of expenditure and revenue shocks, this method is not well suited because our aim is to estimate long-term effects rather than presuppose them. For that reason we used the second method, originally proposed by Bernanke 1986 and Sims 1986. In applying this procedure we directly place restrictions on the matrix \( A_0 = W^{-1} \), which represents the contemporaneous correlations of the VAR innovations in equation (1).

In our specification an additional problem arises. Because we are estimating expenditures and revenues in GDP ratios, shocks in these ratios can also be due to shocks in GDP. This is clearly a complication for identification purposes.\(^{11}\) To overcome this problem we applied a simple trick and premultiplied the estimated VAR (eq. (1)) by the following matrix \( B \) to get the residuals \( e_t \).

\[
[1 \ 0 \ 0] \quad B = \begin{bmatrix} 1 & 0 & 0 \\ 1 & 1 & 0 \\ 1 & 0 & 1 \end{bmatrix}
\]

\Rightarrow BA(L)x_t = Bu_t = e_t.

(3)

This transformation converts, after estimation of the VAR, ratios back to levels. The residuals \( e_t \) can now be interpreted as shocks in the levels of expenditures and revenues. Identification of the model can then be achieved along the lines discussed above: we must find a matrix \( A_0 \) such that \( A_0 \Sigma_{
abla A_0} \) is diagonal. We therefore postulated the following innovation model (note that lower case letters now indicate innovations of the VAR whereas upper case letters denote the variables themselves):

\[
A_0 Bu_t = v_t:
\]

\[
[1 \ \alpha \ \beta] \begin{bmatrix} 1 & 0 & 0 \\ \gamma & 1 & \delta \\ \varepsilon & \zeta & 1 \end{bmatrix} \begin{bmatrix} gdpt \\ expr \\ revt \end{bmatrix} = \begin{bmatrix} v_{u_t} \\ v_{u_t} \\ v_{u_t} \end{bmatrix}
\]

(4)

The innovations \( v_{1t}, v_{2t}, \) and \( v_{3t} \) are assumed to be uncorrelated and are interpreted as primitive shocks in the sense of Bernanke. Using the estimated covariance matrix \( \Sigma_{uu} \) of the VAR residuals the

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\(^{10}\) See for instance Blanchard, Quah 1989.

\(^{11}\) This problem is commonly not addressed in the literature, for instance in Bohn 1991. von Furstenberg et al. 1986 implicitly reduced it in using potential GDP ratios.
parameters $\alpha$ to $\zeta$ can be determined. This can be accomplished by applying the "method of moments": calculating the covariance matrix of $\nu_t, \Sigma_{\nu\nu}$, equation (4) implies:

$$\Sigma_{\nu\nu} = A_0 B \Sigma_{\epsilon\nu} B^\prime A_0'. $$

(5)

The right hand side of (5) depends on the six parameters $\alpha$ to $\zeta$. To calculate them, we need six equations. As $\Sigma_{\nu\nu}$ is diagonal by hypothesis all upper-right off-diagonal elements of the right hand side of (5) must be zero. This leads to the first three required restrictions. Hence, we must impose three additional restrictions, for instance upon $A_0$, to be able to determine all parameters.

We used the following three restrictions:

1. We assume $\gamma$ (the contemporaneous reaction of expenditures on GDP shocks) to be 0.1. That means that an unpredictable positive shock in GDP lowers expenditures to some extent, for instance because the government has to pay less unemployment benefits in that case. Likewise, one could also suppose $\gamma$ to be zero (or to be 0.2) as this does not alter our results in any noticeable way.12

2. We restrict $\varepsilon$ (response of revenues on GDP shocks) to an estimate of the income tax elasticity for each country times minus one. These estimates were obtained simply by regressing $\Delta REV$ on $\Delta GDP$. The estimated elasticities range from 0.89 to 1.15, with a mean value of approximately 1.05.

3. Finding a third restriction was somewhat more cumbersome. As we were especially interested in the contemporaneous effects of expenditures on revenues and vice versa, we proceeded as follows: We investigated three variants. In the first one we restrict $\delta$ to zero (instantaneous causality running only from $EXP$ to $REV$) and estimated $\zeta$ as well as $\alpha$ and $\beta$ by the method of moments, considering the above two restrictions. In the second variant, we restricted $\zeta$ to zero and estimated $\delta, \alpha$ and $\beta$. Finally, in the third variant we assume $\delta$ to be equal to $\zeta$ (symmetric effects, instantaneous causality running in both directions) and estimated $\delta, \zeta, \alpha$ and $\beta$. We thus cover a wide range of possible values for $\delta$ and $\zeta$. Fortunately, it turned out that for most countries the implied results were very similar, if not identical, in all three variants.13 Only for Austria did the variant with $\zeta = 0$ lead to a very different outcome. But we could easily dismiss this variant for Austria as it implied an extremely unreasonable result concerning the effects of expenditure shocks on nominal GDP. In the following section, favouring somewhat the case of bi-directional causality but bearing in mind that the other two variants lead to very similar results, we report the symmetric case.

Having achieved identification of the model, estimation of impulse responses is a standard procedure in applied econometrics. However, point estimates of impulse response functions contain no information about the statistical reliability of the effects detected. In our case this is especially true for the effects of shocks in expenditures on the revenue ratio and vice versa. We therefore implemented Monte Carlo simulations to assess whether the impulse responses are statistically significant at a 90% confidence level.

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12 We have to rely on this crude guess for $\gamma$ because estimation of $\gamma$ by simply regressing $\Delta EXP$ on $\Delta GDP$ is not helpful here, as only (unpredictable) shocks in GDP are relevant in this case.

13 This might be due to the fact that the instantaneous correlations between shocks in $EXPR$ and $REVR$ were comparatively small in most cases, especially if one takes into account that by definition both ratios depend on variations in GDP.
VAR estimates also provide some insights into the frequency domain properties of the relationships between the variables. To analyse these properties we employed a somewhat modified version of a procedure originally proposed by Geweke, who defined linear feedback measures and their decomposition by frequency (Geweke 1982, 1984, 1986). Let, for illustrative purposes,

\[ A(L)x_t = v_t, \]  

be an identified (structural) AR-process with the covariance matrix of \( v_t \), \( \Sigma_{v_t} \), being a diagonal matrix.

If \( A(\lambda) \) is invertible at all frequencies \( \lambda \in [0, \pi] \), the following MA-representation exists:

\[
x_t = C(L)v_t,
\]

where \( C(L) = \begin{bmatrix} c_{11}(L) & c_{12}(L) & \ldots \\ c_{21}(L) & c_{22}(L) & \ldots \\ \vdots & \vdots & \ddots \end{bmatrix} = A(L)^{-1}. \]

The cross spectral density matrix of \( x \) therefore is:

\[ S_{xx}(\lambda) = C(e^{-i\lambda})\Sigma_{v_t}C(e^{i\lambda})'. \]  

If we have a tri-variate system, the spectrum of the first variable \( x_1 \) is equal to

\[ S_{x_1}(\lambda) = \sigma_1^2 \tilde{c}_{11}(\lambda) + \sigma_2^2 \tilde{c}_{12}(\lambda) + \sigma_3^2 \tilde{c}_{13}(\lambda), \]

where \( \sigma_1^2 \) is the variance of the \( v_1 \), \( \sigma_2^2 \) is the variance of the \( v_2 \), \( \tilde{c}_{11}(\lambda) \) represents \( c_{11}(e^{-i\lambda})c_{11}(e^{i\lambda}) \), and so on. The spectrum of \( x_1 \) is therefore decomposed into three independent sources, originating from independent shocks in the three variables. To determine the influence of say \( v_2 \) (innovations in \( x_2 \) on \( x_1 \) one can define the causality spectrum

\[ S_{x_1 \rightarrow x_2}(\lambda) = \frac{\sigma_2^2 \tilde{c}_{12}(\lambda)}{S_{x_1}(\lambda)}, \]

which represents the portion of the spectrum of \( x_1 \) at frequency \( \lambda \) that can be attributed to shocks in \( x_2 \).\(^{14, 15}\) The similarity of this measure with the coherence spectrum is evident, although only one direction of causality is taken into account here. In this sense the term causality spectrum seems justified, bearing in mind the fact that a causal interpretation requires identifying restrictions.

\(^{14}\) For the bi-variate case, the definition of \( S_{x_1 \rightarrow x_2}(\lambda) \) is equivalent to \( f_{x_1 \rightarrow x_2}(\lambda) \) in Geweke 1982, considering Pierce's (1982) comment on this contribution. It is further similar to the causality coherence in Granger 1968.

For the multivariate case, this is an extension of Geweke 1982 which, compared to Geweke 1984, has the advantage that indirect causality chains are also taken into account.

\(^{15}\) There is an interesting connection between \( S_{x_1 \rightarrow x_2}(\lambda) \) and common forecast error variance decomposition techniques for VARs. The variance decomposition for \( t \rightarrow \infty \) is equal to

\[
\frac{1}{2\pi} \int_{-\pi}^{\pi} S_{x_1 \rightarrow x_2}(\lambda) \frac{S_{x_2}(\lambda)}{\sigma^2_{x_2}} d\lambda,
\]

which is a weighted average of \( S_{x_1 \rightarrow x_2}(\lambda) \), given the spectrum of \( x_1 \) as the weighting function. Therefore, these spectra are to some extent the frequency domain counterparts of conventional error variance decompositions. For a proof of this see Stiassny 1994.
If $A(\lambda)$ is not invertible at $\lambda=0$ because of a unit-root in the vector $x$, problems in deriving $S_{xx}(\lambda)$ can arise. Fortunately, this does not matter for the definition of $S_{x\rightarrow x}(\lambda)$, as the non-stationary parts of the AR-polynomial cancel. If the variables are I(1) processes, the spectra of $x$ and $\Delta x$ differ by the terms $(1-e^{\pm i\omega\lambda})^{-1}$ only. And these terms obviously cancel out, taking into account the ratio in the definition of $S_{x\rightarrow x}(\lambda)$.

VAR estimates are generally of comparatively low precision if only few observations are available. To ease this problem one can try to add more structure to the model at hand, for instance by estimating an error correction form imposing a cointegration restriction upon the VAR. As we assume that the public sector meets its long-term budget constraint such that the net deficit ratio cannot rise indefinitely, the idea of an error correction formulation occurs almost naturally (Bohn 1991). Unusually high deficit ratios must induce the administration to lower expenditures and/or to raise revenues. Since an error correction model requires the variables to be non-stationary and the error correction term to be stationary, a pre-testing procedure is necessary to examine the time series properties of the variables. We employed the augmented Dickey-Fuller test (ADF) to investigate for each country whether the four time series used in our models contain unit roots.

Table 2 reports the test statistics. To allow for the proper alternative we used the test variant with time trend whenever the variable indicated a significant trend component in a regression on a constant and a time variable. In almost all cases an autoregressive correction factor of 1 was sufficient to render the Ljung-Box Q-statistic insignificant at the 10% level.

Table 2. Unit Root Tests - (Augmented Dickey-Fuller Test)

<table>
<thead>
<tr>
<th>Country</th>
<th>$\Delta GNP$</th>
<th>$\Delta GDP$</th>
<th>$\Delta E X P R$</th>
<th>$\Delta E X P R$</th>
<th>$\Delta R E V R$</th>
<th>$\Delta R E V R$</th>
<th>$N D E F R$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Austria</td>
<td>-0.69*</td>
<td>-2.90**</td>
<td>-7.72**</td>
<td>-2.41**</td>
<td>-6.57**</td>
<td>-1.76**</td>
<td>-4.48***</td>
</tr>
<tr>
<td>France</td>
<td>-0.62*</td>
<td>-0.88</td>
<td>-5.08**</td>
<td>-2.17**</td>
<td>-3.43**</td>
<td>-1.81**</td>
<td>-3.57***</td>
</tr>
<tr>
<td>FRG</td>
<td>-1.51*</td>
<td>-3.10*</td>
<td>-5.80**</td>
<td>-0.53</td>
<td>-2.96**</td>
<td>-2.54**</td>
<td>-3.80***</td>
</tr>
<tr>
<td>Italy</td>
<td>-2.11*</td>
<td>-1.84*</td>
<td>-7.54**</td>
<td>-2.10*</td>
<td>-5.22**</td>
<td>-1.53*</td>
<td>-3.29*</td>
</tr>
<tr>
<td>Netherlands</td>
<td>-0.49*</td>
<td>-1.91</td>
<td>-6.74**</td>
<td>-1.40*</td>
<td>-4.49*</td>
<td>-1.13*</td>
<td>-3.64**</td>
</tr>
<tr>
<td>Sweden</td>
<td>-2.35*</td>
<td>-3.40*</td>
<td>-6.51**</td>
<td>-1.26*</td>
<td>-2.86*</td>
<td>-4.25**</td>
<td>-5.91***</td>
</tr>
<tr>
<td>Switzerland</td>
<td>-2.00*</td>
<td>-2.61</td>
<td>-4.12*</td>
<td>-3.00*</td>
<td>-5.19**</td>
<td>-0.85</td>
<td>-3.92*</td>
</tr>
<tr>
<td>UK</td>
<td>-2.04*</td>
<td>-2.21*</td>
<td>-5.79**</td>
<td>-2.65*</td>
<td>-4.61**</td>
<td>-2.98*</td>
<td>-4.62*</td>
</tr>
<tr>
<td>USA</td>
<td>-1.87*</td>
<td>-2.76*</td>
<td>-7.35**</td>
<td>-3.63*</td>
<td>-5.33**</td>
<td>-4.17**</td>
<td>-4.75**</td>
</tr>
</tbody>
</table>

* ... significant at 10% level, ** ... significant at 5% level, *** ... significant at 1% level.
+ ... time trend included.

Table 2 clearly indicates that for most countries the expenditure ratio as well as the revenue ratio are likely to be integrated of degree one. The exceptions are the USA (both expenditure and revenue ratios are stationary) and Sweden (revenue ratio is stationary). According to our tests, an I(2) property for nominal GDP cannot be ruled out for France, Netherlands, Italy and the UK. The net deficit ratios seem to be borderline cases for some countries. We nevertheless assume $N D E F R$ to be trend stationary for all countries but Sweden.

The results of these unit root tests suggest the following specification of a tri-variate error correction model for all countries except Sweden and the USA:
\[
\begin{bmatrix}
\Delta^{(2)}GDP_r \\
\Delta EXPR_r \\
\Delta REVR_r
\end{bmatrix}
= f\left[\Delta^{(2)}GDP_{r,t-1}, \Delta EXPR_{r,t-1}, \Delta REVR_{r,t-1}, NDEFR_{r,t-1}(t)\right].
\] (9)

Because an I(2) property of GDP could not be ruled out for France, Netherlands, Italy and the UK, \( \Delta^{(2)}GDP \) was included instead of \( \Delta GDP \) for those countries. We allowed for a time trend whenever it was significant at a 10% level in any of the three equation. Because of the stationary Swedish revenue ratio we included \( REVR_{r,1} \) instead of \( NDEFR_{r,1} \) in the Swedish model. In the US model we used both \( REVR_{r,1} \) and \( EXPR_{r,1} \) (both are stationary) instead of \( NDEFR_{r,1} \).

A final remark has to be made on specifying the lag length of the error correction model. Since we chose the last thirty-six years as our estimation period in a tri-variate model structure, a lag length of three seems to be the maximum. It turned out, however, that a lag length of two was sufficient to render the Q-statistics insignificant in almost all cases.

4. Empirical Results

The key parameters of the estimated error correction model described in the previous section are reported in Table 3 for each of the countries under consideration. The first two columns of Table 3 report the t-values of the error correction terms (the net deficit ratio) in the expenditure equation and in the revenue equation, respectively. The higher the t-value of this term in the respective equation, the more statistically significant the contribution of expenditures or revenues in reducing an unsustainable deficit pattern. As can be seen, the "expenditure dominance", defined as a statistically reliable relation between the development of revenues and the deficit, is most pronounced in Austria with a t-value of 7.2. The inverse case of a "revenue dominance" can be found in the case of FRG and UK. In these countries, revenues seem to be comparatively autonomous so that in fact expenditures have to bear the primary burden of budgetary adjustment. For the other countries, the net deficit ratio is insignificant at a 5%-level in both equations.
Table 3. Some Key-Parameters of the estimated equations

<table>
<thead>
<tr>
<th>Country</th>
<th>t-value of NDEFR _EXPR-equation</th>
<th>t-value of NDEFR _REVR-equation</th>
<th>$R_D^2$ _EXPR-equation</th>
<th>$R_D^2$ _REVR-equation</th>
<th>time trend</th>
</tr>
</thead>
<tbody>
<tr>
<td>Austria</td>
<td>-1.88</td>
<td>7.20</td>
<td>0.66</td>
<td>0.70</td>
<td>yes</td>
</tr>
<tr>
<td>France</td>
<td>-0.97</td>
<td>1.30</td>
<td>0.12</td>
<td>0.26</td>
<td>yes</td>
</tr>
<tr>
<td>FRG</td>
<td>-2.78</td>
<td>1.28</td>
<td>0.52</td>
<td>0.31</td>
<td>yes</td>
</tr>
<tr>
<td>Italy</td>
<td>0.18</td>
<td>1.24</td>
<td>0.10</td>
<td>0.21</td>
<td>yes</td>
</tr>
<tr>
<td>Netherlands</td>
<td>-1.51</td>
<td>-0.19</td>
<td>0.20</td>
<td>0.14</td>
<td>no</td>
</tr>
<tr>
<td>Sweden</td>
<td>-</td>
<td>-</td>
<td>0.46</td>
<td>0.43</td>
<td>yes</td>
</tr>
<tr>
<td>Switzerland</td>
<td>-1.6</td>
<td>1.44</td>
<td>0.61</td>
<td>0.88</td>
<td>yes</td>
</tr>
<tr>
<td>UK</td>
<td>-2.36</td>
<td>0.32</td>
<td>0.51</td>
<td>0.28</td>
<td>no</td>
</tr>
<tr>
<td>USA</td>
<td>-</td>
<td>-</td>
<td>0.71</td>
<td>0.40</td>
<td>yes</td>
</tr>
</tbody>
</table>

Figure 1 summarises the effects that are induced by shocks in revenues on the expenditure ratio, and vice versa. For that purpose, the first row of each country's representation demonstrates the estimated impulse responses with the corresponding 90% confidence intervals. These impulse responses show how future values of the variables react to current shocks. The second row depicts the share that these shocks contribute to the respective spectra. The results for the individual countries are ranked according to the detected degree of estimated revenue dominance. As a measure of revenue dominance we used an indicator based on our estimated causality spectra. One has to be aware, however, that this measure is in no way an indisputable indicator. But as it represents an average value of the mutual dependency between the budgetary aggregates across the whole frequency band it seems to be an appropriate indicator for our purpose. Others could, however, argue that only the dependencies at a specific point (e.g. at $\lambda=0$, the long run) are relevant.

$R_D^2$ is a measure of goodness-of-fit suited for non-stationary time series. Thereby the residual sum of squares of the estimated equations is compared to the residual sum of squares obtained by a pure random walk model with drift. A value of 0.2 therefore means that the models fit is 20% better than the fit of a pure random walk model.

We used the area of the $EXP \rightarrow REVR$ causality spectrum, weighted by the spectrum of the revenue ratio and subtracted the area of the $REV \rightarrow EXPR$ causality spectrum weighted by the spectrum of the expenditure ratio. According to Stixatny 1994, this is equal to the difference of the forecast error variance decompositions for an infinite horizon. The estimated values of this measure are: UK: -0.9, NL: -0.78, FRG: -0.21, USA: -0.27, SWITZ: -0.05, SWE: 0.04, FRA: 0.24, AUT: 0.57, ITAL: 0.9. A negative value implies revenue dominance, a positive expenditure dominance.
Fig. 1. Impulse Responses and Spectral Decompositions

UK:

\( \text{EXP} \rightarrow \text{EXPR} \)

\( \text{EXP} \rightarrow \text{REVR} \)

\( \text{REV} \rightarrow \text{EXPR} \)

\( \text{REV} \rightarrow \text{REVR} \)

NL:

\( \text{EXP} \rightarrow \text{EXPR} \)

\( \text{EXP} \rightarrow \text{REVR} \)

\( \text{REV} \rightarrow \text{EXPR} \)

\( \text{REV} \rightarrow \text{REVR} \)
Fig. 1. Continued

FRG:

USA:

14
Fig. 1. Continued

SWITZ:

\[
\begin{align*}
\text{EXP} &\rightarrow \text{EXPR} \\
\text{EXP} &\rightarrow \text{REVR} \\
\text{REV} &\rightarrow \text{EXPR} \\
\text{REV} &\rightarrow \text{REVR}
\end{align*}
\]

SWE:

\[
\begin{align*}
\text{EXP} &\rightarrow \text{EXPR} \\
\text{EXP} &\rightarrow \text{REVR} \\
\text{REV} &\rightarrow \text{EXPR} \\
\text{REV} &\rightarrow \text{REVR}
\end{align*}
\]
Fig. 1. Continued

FRA:

\[
\text{EXP} \rightarrow \text{EXPR} \quad \quad \quad \text{EXP} \rightarrow \text{REVR} \quad \quad \quad \text{REV} \rightarrow \text{EXPR} \quad \quad \quad \text{REV} \rightarrow \text{REVR}
\]

AUT:

\[
\text{EXP} \rightarrow \text{EXPR} \quad \quad \quad \text{EXP} \rightarrow \text{REVR} \quad \quad \quad \text{REV} \rightarrow \text{EXPR} \quad \quad \quad \text{REV} \rightarrow \text{REVR}
\]
As Figure 1 shows, it is the UK that shows the most significant revenue dominance. According to the estimated impulse responses as plotted in the first row, within the first ten periods, shocks in expenditures do not have statistically relevant effects on the expenditure ratio or on the revenue ratio. On the other hand, shocks in revenues have a statistically significant and permanent effect on both the revenue and the expenditure ratio in Great Britain. The pattern of the estimated spectral decomposition, as shown in the second row, leads to the same conclusion: only at high frequencies (e.g. $\lambda = \pi$, in the short run) do shocks in expenditures contribute in a noticeable way to the variance of the revenue ratio. There are, however, no detectable effects at business cycle frequencies or in the long run. Shocks in revenues, on the other hand, nearly exclusively explain the long-term variation of the expenditure ratio. There is also a strong connection at business cycle frequencies ($\lambda = \pi/2$, cycles of approximately 4 years). Further, apart from the very short term, the variance of the revenue ratio can be attributed nearly exclusively to its own shocks. Both results thus support the view that Britain’s fiscal policies can be better explained by the tax and spend hypothesis than by the spend and tax hypothesis.

Our estimations with regard to the Netherlands and the FRG also give strong support to the tax and spend hypothesis. The results for the USA, although less pronounced, indicate a dominance of revenues as well. In the case of Switzerland, the estimated impulse responses only slightly favour revenue dominance, whereas for Sweden our impulse responses cannot detect any significant cross effects. In addition, the spectral decompositions do not support any of the competing hypotheses.

For the remaining countries, France, Italy and Austria, the results tend to support the spend and tax hypotheses. While this dominance remains rather weak in the case of France it seems far more pronounced in the case of Italy and Austria. In Austria, for instance, we detect a statistically significant effect of expenditure shocks on the revenue ratio (first row) while the effects of revenue shocks on the expenditure ratio remain negligible. It is further noticeable that shocks in expenditures lead to a permanent effect in the expenditure ratio, whereas shocks in revenues do not have
statistically significant permanent effects on the revenue ratio. This again suggests that, in Austria, revenues rather than expenditures are adjusted in order to meet the government's intertemporal budget constraint. The spectral decompositions (second row) show a very pronounced dependency of revenues on expenditures at longer business cycle frequencies and in the long run. Across the whole frequency band no effects of revenues on expenditures are detectable.

Our results thus show that the dependencies between the budgetary aggregates diverge significantly among the countries under consideration. Obviously, the political-economic traditions in the various countries regarding budget decision making seem to differ markedly. As already mentioned, such discrepancies cannot be interpreted as contradicting or supporting those approaches of public finance theory that presume the simultaneity of revenue and outlay planning. Our computations estimate the dependencies between actual revenues and outlays. They cannot identify whether these flows result from a rational, long-term budget decision of a "benevolent dictator" or are the outcome of short-sighted interventions of a self-interested government aiming to win the next election or to prejudge its successors.

However, we tried to find some indicators which could help us to explain systematically the detected differences by attributing them to a well-defined set of political and economic factors characteristic for each country. We therefore estimated linear regressions of our measure of revenue dominance on the indicators summarised in Table 1.

The share of exports in GDP, the inflation rate, the measure of central bank independence, the average government tenure and the degree of fiscal federalism did not prove significant in these regressions, although all the estimated parameters had the sign which could be expected from the considerations in section two. With regard to the monetary data, this insignificance may result from the fact that our country sample is in fact too homogenous to allow different patterns to be effectively detected. Central bank independence and the inflation rate are probably better suited to discriminating between industrialised and developing countries than between the developed countries themselves. The political indicators, on the other hand, might be just too simple to model the complex structures of political decision making in advanced societies.

The only variable in our regressions which turned out to be marginally significant at a 15% level is the average deficit ratio, indicating that the probability of spend and tax behaviour increases with a country's tendency towards higher deficits. This could be interpreted such that societies which favour tax and spend, e.g. whose tax decisions are more autonomous than their spending decisions, tend to run lower overall deficits. One could, however, also argue that this relationship is the result of successful signalling strategies. Countries whose governments have gained credibility in their efforts to limit deficits by instrumentalising taxes to reduce the demands from various lobbyists for government spending programs have in fact been successful. Our estimations do not allow us to discriminate between these two alternative explanations.
Our paper aims to improve the understanding of the political economy determining the interrelation between government spending and taxes. For that purpose we estimated a tri-variate, structural VAR model, implementing impulse-response functions and frequency domain techniques to identify the causal relationship between government outlays and receipts. Since we interpreted the budgeting process as an error correction model, we were also able to investigate to what extent revenues and outlays bear the burden of adjustment in cases of deviations from long-term budget restrictions. To take account of a wide variety of socio-economic factors estimations were carried out for nine industrialised countries.

The empirical results indicate significant differences in the political economy of budget decision making in the various countries. While for some countries (most prominently the United Kingdom, in a less pronounced way for the Netherlands, Germany and the USA) the dynamics of tax revenues clearly dominate the development of outlays and budgetary adjustments, the opposite is true for other countries (e.g. France, Austria and Italy). For Switzerland and Sweden no dominance of either revenues or expenditures were detectable. These results are only partially in line with earlier empirical findings on this issue.

In the case of the USA, v. Furstenberg et. al. 1986 and Bohn 1991 - both using VAR techniques without, however, addressing the identification problem, which is a central issue in our study - found a significant, though weak dominance of expenditures over revenues. Miller - Russo 1990 detected a one-way Granger causality from taxes to spending using yearly data at the federal level. Using quarterly data they found a bi-directional causality. Ram 1988a concluded that taxes Granger-cause spending at the federal level.

In the case of Austria, an earlier finding derived for different budget concepts (Koren-Stiassny 1992) led to the same results as detected here. To the best of our knowledge, Rati 1988b is the only comparable study so far that has been implemented for other countries at the federal level. For the UK he found bi-directional Granger causality, while for Austria, Netherlands, Switzerland and the USA he detected no causality at a 5% significance level. The differences in the results of these studies compared to our work might be attributed to the use of different sample periods (which is by far longer in Bohn 1991 and covers only the stable post-war period ending 1982 in the case of v. Furstenberg 1986), a broader budget concept (including off-budget items) and different estimation techniques (structural VAR compared to ordinary VAR techniques or Granger causality tests).

We tried to find systematic explanations for the detected differences among the countries under consideration. We therefore used various variables that are often mentioned as relevant indicators of a society's political-economic system to explain our results. All of these indicators, except the budget deficit pattern, proved insignificant, although we always got the expected sign in our regressions. The significance of the average deficit provides some evidence of a correlation between a country's propensity to "tax and spend" and its deficit level. It may well be that societies with a pronounced public awareness of the tax burden and significant problems regarding tax evasion are confronted with a strict budget constraint that in fact leads to lower overall deficits. It could, on the other hand, as well be that these lower deficits are the result of successfully implemented signalling strategies. Governments that signalled a consequent hard budget constraint by intentionally provoking a situation of budget crisis have been successful in reducing demands for public goods and services.

We are, however, aware that the causes for the differences detected in this paper are probably more complex than these simple considerations suggest. We must leave the task of finding more appropriate explanations to further research.
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