ON THE MACROECONOMIC IMPACT OF FISCAL POLICY IN GERMANY: PRELIMINARY RESULTS OF A SVAR APPROACH

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1. Introduction

Although the effects of fiscal policy belong to the most extensively discussed issues in theoretical macroeconomics, not much is known about the actual impact of changes in government revenue or expenditure on the economy (the reader is referred to Perotti (2000) for a – somewhat disillusioning – overview on this issue). Against the background of debates on stabilisation policy dating back to the sixties, this seems to be surprising. In the traditional Keynesian approach to macroeconomic analysis, active fiscal policy was assigned a powerful and beneficial role as a potential macroeconomic stabiliser. Practical problems with the conduct of active stabilisation policy and new theoretical advances however, have led to a markedly more modest assessment of the potential and the benefits of active fiscal stabilisation policy. Practically, active fiscal stabilisation policy often turned out to be pro cyclical – and hence destabilising – rather than stabilising. Operational time lags in the conduct of fiscal policy and policy co-ordination failures between central government and local authorities – which may control a substantial amount of general government resources – are the perhaps most prominent hindrances to a successful active fiscal stabilisation policy.

New classical macroeconomics claimed that fiscal (and monetary) stabilisation policy is effective only if it surprises economic agents. According to this paradigm, *systematic* responses to macroeconomic shocks would be ineffective since by definition they can be expected and reactions by economic agents would in fact counteract discretionary policy. Additionally, real business cycle theorists interpreted macroeconomic shocks mainly as technology shocks and business cycles as consequences of welfare maximising choices of optimising economic agents adapting to such shocks. In this theoretical view, too, macroeconomic fine-tuning through fiscal policy is not advised since it potentially decreases welfare by

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interfering with the optimal choices of economic agents. In addition, a more recent challenge to the standard Keynesian view has emerged from the detection of so called Non-Keynesian effects of fiscal policy. If public debt has already reached high levels, economic agents would not have much trust in the sustainability of increases in expenditures. These may rather nurture expectations that government has to cut spending in the near future in order to observe the long run budget constraint. It is even possible that expenditure cuts have short-run expansionary effects by stabilising expectations and confidence.

Today, the view that discretionary fiscal policy should be concentrated on the provision of public goods and the continuos improvement of favourable supply side conditions rather than on reactive responses to cyclical fluctuations seems to constitute a widely accepted view among economists. Neo-Keynesians would advocate monetary rather than fiscal policy as the more flexible and hence more appropriate tool for active macroeconomic stabilisation[1] and New Classical economists remain generally sceptical against any discretionary stabilisation policy.

However, it could nevertheless be premature to declare fiscal stabilisation policy dead in view of the economic policy experience and the theoretical insights of the last two decades. A more prominent role for active fiscal stabilisation policy has recently been demanded especially for EMU in which monetary policy is centralised under the control of the ECB. The course of European monetary policy must necessarily be aimed at the macroeconomic conditions of the Euro area as a whole and may thus be inappropriate for individual countries hit by an asymmetric macroeconomic shock. Since the ECB would probably not adjust monetary policy in order to respond upon such idiosyncratic shocks, fiscal policy – so the argument goes – remains as the only tool which can should be used for active economic stabilisation. The demand for active fiscal policy has became more audible in the political arena, especially during and after the most recent economic slowdown in 2001. To conclude, it appears that the issue of active fiscal stabilisation remains on the agenda, especially under the conditions of EMU.

In this paper, short run impacts of fiscal policy in Germany on the macroeconomic environment in a small structural vector auto-regressive (SVAR) model are investigated. The results presented so far in this paper

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1 See, e.g., Taylor (2000).

are preliminary and need to be validated and qualified further. Furthermore, the views and results presented here reflect the opinion of the author and not the opinion of the institution the author belongs to.

Vector auto-regressive models have been used for the analysis of monetary policy since about 20 years now. It is only recently that their potential for the analysis of fiscal policy has been investigated. Blanchard and Perotti (2002) represents one of the early examples of a VAR analysis of fiscal policy.[2] Recent contributions to VAR analysis of fiscal policy shocks, on which this paper draws very much, are Mihov and Fatas (2000), Bruneau and De Bandt (1999) and Hoeppner (2001).

2. Data

The four series included in the analysis are GDP (Y), private consumption (CP), total government receipts (BR) and total government expenditure (BE). All variables are in real terms. The series were deflated with the respective price indices, except total government receipts which was deflated with the GDP deflator, and all series were seasonally adjusted. The data are based on semi-annual German national accounts[3] from 1970:1 to 2000:2. Total expenditures were adjusted for proceeds from UMTS licences.

Due to German unification the series exhibit a structural break in 1991:1. The series were adjusted for this break by applying an extended HP-Filter which allows for an endogenous estimation of deterministic structural breaks, given the period in which the break supposedly occurs. It is assumed thereby that the break is reflected in levels and not in trends.[4] The adjusted series was obtained by subtracting the estimated deterministic shifts in levels from the original series. Finally, all series were transformed into logarithms. Following standard tests, the adjusted and transformed series can be regarded as integrated of order 1.

Let Z_t be the 4x1 vector of the four endogenous variables Y, CP, BR and BE at time t. The VAR in reduced form is given by the equation:

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2 The working paper version of Blanchard and Perotti (2002) dates back to 1999.

3 Quarterly series of total government receipts and expenditure are not yet available.

4 See Appendix A.2 for the derivation of the extended HP-Filter.

$$Z_t = A(L)Z_{t-1} + u_t, \tag{2.1}$$

A(L) denotes a matrix polynomial in the lag operator L ($A(L) = A_0 + A_1L^1 + A_2L^2 + ... + A_pL^p$) and ε_t denotes a vector of white noise residuals with zero mean and covariance $E(u_t u_t^{'}) = \Sigma_u^2$. Hence, the residuals are assumed not to be auto-correlated but can and will in most cases be contemporaneously cross-correlated. The auto-regressive model in Equation (2.1) can be estimated by OLS, since the right hand side of the equation is predetermined and thus exogenous with respect to the vector Y_t. Hence, the ε_τ are in fact vectors of correlated one step prediction errors.

As a next step, the co-integrating relationships of the time series are to be analysed. As is well known, the reduced form in Equation (2.1) can be transformed into the error correction form:

$$\Delta Z_t = A_0 + \sum_{i=1}^{p} \pi_i \Delta Z_{t-i} + \Pi\, Z_{t-p} + u_t \tag{2.2}$$

If co-integrating relationships between the variables in Z exist, the n×n matrix Π can be decomposed as $\Pi = \alpha\beta'$, with α and β as n×r matrices, 0 < r < n. In this way, the vector of the endogenous variables, Z, can be described by r long run (or: co-integrating) relationships as specified by the vectors of β. The vectors in α and β can be estimated by rank regression and the number of co-integrating vectors in the β-matrix, r, can be tested using Johansen's trace test.[5]

The test depicts two co-integrating vectors at the 5% significance level.[6] Possible candidates for co-integrating vectors would be revenue and expenditure on the one hand and real private consumption and real GDP on the other. This can be tested by imposing identifying restrictions on the cointegrating vectors. With the ordering of the variables as Y, CP, BR, BE and C, (C denotes the constant in the co-integration space), the restrictions

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5 The was estimated with a lag of 2 periods. The lag length has been determined by information criteria. The Hannan-Quinn criterion was decisive in the choice of the maximum lag here. See Reimers (1993) for a discussion of information criteria in VARs.

6 The test has been performed under the assumption that both, the I(1) and the I(0) model contain constants and no trends, i.e. the deterministic component has been restricted to a non zero means in the cointegrating equations. This implies that $A_0 = \alpha\beta_0^{'}$, with β_0 as an n×r vector.

Table 2.1

Trace Tests for the co-integration rank (r)

r	Statistic	50%	80%	90%	95%	97.5%	99%
0	72.72	38.84	45.65	49.65	53.12	56.06	60.16
1	38.26	23.28	28.75	32.00	34.91	37.61	41.07
2	10.29	11.25	15.25	17.85	19.96	22.05	24.60
3	3.86	3.40	5.91	7.52	9.24	10.80	12.97

can be specified as zero restrictions on the 5×r Matrix β (with co-integration rank $r = 2$): $\beta_1' = [1, \beta_{21}, 0, 0, \beta_{51}]$ and $\beta_2' = [0, 0, 1, \beta_{42}, \beta_{52}]$. Furthermore, the co-integration vectors are normalised for Y and BR, respectively. The restrictions imply a stable long run relationship between private consumption and GDP, $Y + \beta_{21}CP + \beta_{51}C = 0$, and a stable long run relationship between expenditure and revenue, $BR + b_{42}BE + b_{52}C = 0$, respectively. The latter can be justified by the long run government budget constraint: For the long run budget constraint to hold, expenditure and revenues should not follow different trends.[7] The parameters b_{ij} are estimated freely. We would expect that $\beta_{21}, \beta_{42} < 0$ and, more specifically, $\beta_{21}, \beta_{42} \approx -1$.[8] The co-integration vectors are actually estimated as $\beta_1' = [1, \underset{(0.070)}{-1.30}, 0, 0, \underset{(0.450)}{1.63}]$ and $\beta_2' = [0, 0, 1, \underset{(0.123)}{-0.60}, \underset{(0.770)}{-2.92}]$, with associated standard errors in parentheses. Due to the restrictions imposed, the co-integration vectors are overidentified and the restrictions can thus be tested.

If the restrictions are correct, the log-likelihood of the restricted model should not differ much from that of the unrestricted model. This is the case as can be inferred seen from Table 2.2. The log-likelihood is

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7 Trehan and Walsh (1988) show that the inter-temporal budget constraints implies that the deficit be stationary. Actually, they employ this implication for a test of sustainability of the deficit.

8 Note that the variables are in logarithms.

Table 2.2

Test of overidentifying restrictions on the co-integrating vectors (weakly restricted)

LOG-LIKELIHOOD UNDER H0	1273.44104
LOG-LIKELIHOOD UNDER HA	1274.41926
NUMBER OF DEGREES OF FREEDOM	2
CHI SQUARE TEST	1.95646
SIGNIFICANCE LEVEL	0.38

Table 2.3

Test of overidentifying restrictions on the co-integrating vectors (strongly restricted)

LOG-LIKELIHOOD UNDER H0	1272.00356
LOG-LIKELIHOOD UNDER HA	1274.41926
NUMBER OF DEGREES OF FREEDOM	3
CHI SQUARE TEST	4.83142
SIGNIFICANCE LEVEL	0.18457

χ^2_2-distributed and the restrictions imposed cannot be rejected. (The significance level against the H0 is about 38%).

Although the signs of the freely estimated parameters come out as expected, the absolute estimated values, however, are largely different from unity. Thus, an additional test was performed with a more binding restriction on the second co-integrating vector such that the long run budget restriction of government is now explicitly enforced by imposing

$\beta_2^{'} = [0,0,1,-1,\beta_{52}]$. The co-integrating vectors are estimated as $\beta_1^{'} = [1,-\underset{(0.011)}{1.05},0,0,-\underset{(0.073)}{0.13}]$ and $\beta_2^{'} = [0,0,1,-1,-\underset{(0.02)}{0.16}]$ under the stronger restriction. Furthermore, the estimated relationship between consumption and GDP turns out as almost proportional.[9]

3. Results of the SVAR Analysis

Since all variables belong to the co-integration space, the VAR model can be estimated in levels (i.e.: as in Equation 2.1) rather than in first differences. The latter approach would be appropriate if no co-integration existed among the variables. Alternatively, the VAR could also be estimated in VEC-form (as in Equation 2.2). In this paper, only the results for the VAR in levels are presented. They are qualitatively comparable to those derived from the VEC-form albeit actual figures differ somewhat.

Estimating Equation (2.1) by OLS yields estimates of the one step prediction errors, u_t. However, the prediction errors do not constitute independent shocks which can be interpreted economically since they are mutually correlated. Rather, the "pure" shocks which can be exclusively assigned to certain variables, have to be derived from the estimated prediction errors. Let the pure structural shocks, ε_t, be implicitly defined by the VAR in structural form:

$$B\,Y_t = \Gamma_0 + \Gamma(L)Y_{t-1} + \varepsilon_t \tag{3.3}$$

such that $A_0 = B^{-1}\Gamma_0$, $A(L) = B^{-1}\Gamma(L)$, and:

$$u_t = B^{-1}\varepsilon_t \tag{3.4}$$

B is a Matrix with ones on the main diagonal. Note that the structural shocks are by definition mutually uncorrelated. Without loss of generality, the equation system (3.3) can be normalised such that the variances of the structural shocks are equal to one. Thus, the ε_t are assumed

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9 Likewise, the first co-integrating vector could be restricted thereby enforcing the stable CP/Y relationship. In this way, the inter-temporal budget restriction can be tested. As a result not reported here, the restriction cannot be rejected and the parameter β_{24} turns to be close to –1 in this case.

to be distributed as N(0, I_4) whereby I_4 denotes the 4×4 identity matrix. It follows that:

$$B\Sigma_u^2 B' = I_4 \tag{3.5}$$

If the 4×4 matrix B could somehow be derived from the covariance matrix Σ_u^2, which can be estimated from the reduced form, the structural shocks ε_t can be identified. Unfortunately, this is not possible without additional assumptions, since the estimated covariance matrix Σ_u^2 delivers n(n+1)/2 parameters whereas the n^2 entries of B are to be derived. Thus, it is necessary to impose at least n(n–1)/2 – in this case: 6 – additional restrictions on the matrix B. Actually, in this case it turns out that 7 restrictions are necessary.[10] The set of restrictions imposed on the contemporaneous responses (i.e.: the responses in the same period) of the endogenous variables upon structural shocks are described by Equation (3.6):

$$\begin{bmatrix} 1 & b_{12} & b_{13} & b_{14} \\ 0 & 1 & b_{23} & b_{24} \\ -0.85 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} u_Y \\ u_{CP} \\ u_{BR} \\ u_{BE} \end{bmatrix} = \begin{bmatrix} \varepsilon_Y \\ \varepsilon_{CP} \\ \varepsilon_{BR} \\ \varepsilon_{BE} \end{bmatrix} \tag{3.6}$$

The set of restrictions is very similar to the identification schemes in Blanchard and Perotti (1999) and Hoeppner (2001). The response of Y to any shock – the first line in (3.6) – is unrestricted. The second line implies that private consumption does not react contemporaneously to a GDP shock,[11] ε_Y. In the third line, the elasticity of total government revenue with respect to output is fixed to 0.85. This value has been taken from Bouthevillain *et al.* (2001). In addition, it is assumed that government revenue are not contemporaneously affected by government expenditure. The last line copies the identification scheme in Blanchard and Perotti (2002): First, government expenditure are assumed not to react immediately upon revenue shocks and, second, government is assumed not

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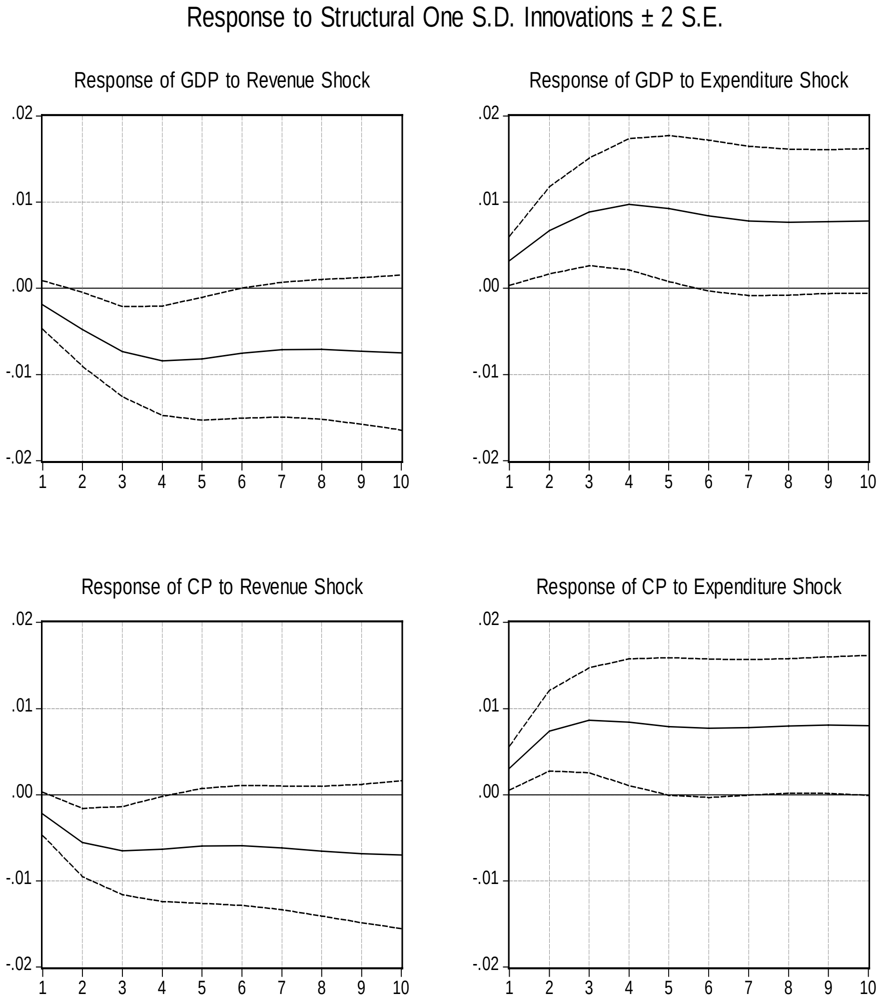
10 Note that Equation (3.5) defines non-linear relationships between the entries in B. Thus, by pure counting of equations and unknown parameters only a necessary, not a sufficient condition for exact identification can be derived.

11 Fixing the immediate response of CP to GDP to 0.5 – which comes close to 0.85×CP/Y, i.e.: setting b_{21} = –0.5, does not change the results below significantly.

Fig. 3.1

Response of Y and CP to a revenue and an expenditure shock



to react upon a macroeconomic shock by adjusting expenditure in the same period. By Equations (3.6) and (3.5), the B matrix and thus the structural shocks are now identified.

After estimating the VAR in reduced form (Equation (2.1)) the B matrix can be estimated and an impulse-response analysis can be performed. The complete set of impulse-response graphs is shown in the appendix. The responses of Y and CP to revenue and expenditure shocks – normalised to one standard deviation – are shown in Figure 3.1.Since the VAR has been estimated in levels, the response variables do not necessarily return to the zero line (which represents the state of the system in the absence of shocks). They rather adjust to new equilibrium levels induced by the impulses. As to be expected, a positive expenditure (revenue) shock causes GDP and private consumption to increase (decrease). The graphs exhibit the responses to a shock normalised to one standard deviation (together with the 5% confidence band). As all series are in logarithms, the results can easily be translated into elasticities by dividing the responses through the estimated standard deviations of the shocks. It follows that private consumption decreases by about 0.4% after two years following a one percent revenue shock and increases by about 0.35% after one year and a half following a one percent expenditure shock. GDP reacts with a decline of about –0.5% (0.4%) within two years after a one percent revenue (expenditure) shock. However, these results are preliminary and have to be interpreted with caution. The identification scheme, which is still arbitrary, has to be validated better in theoretical and methodological respect.

APPENDIX

A.1 THE EXTENDED HP-FILTER

Define a trend series with a deterministic break as:

$$x_t^* = a_t + s_t . \tag{A1}$$

Here, a denotes the smooth trend without break. The break occurs at an exogenously given period τ so that the series s is defined as follows:

$$s_t = \begin{cases} 0, & t < \tau \\ \bar{s}, & t \geq \tau \end{cases} \tag{A2}$$

Denote the original series to be filtered by x. The break, i.e. the parameter $\bar{s}$, can be determined endogenously by solving

$$\underset{S_t, a_t^*}{\mathrm{Min}} \sum_{t=1}^{T} (x_t - s_t - a_t)^2 + \lambda \sum_{t=3}^{T} \left(\Delta^2 a_t\right)^2 \tag{A3}$$

with **s** restricted by (A1). The simple HP Filter is implied by the condition $\bar{s} = 0$. For the general case, the solution of (A3) gives the following first order conditions:

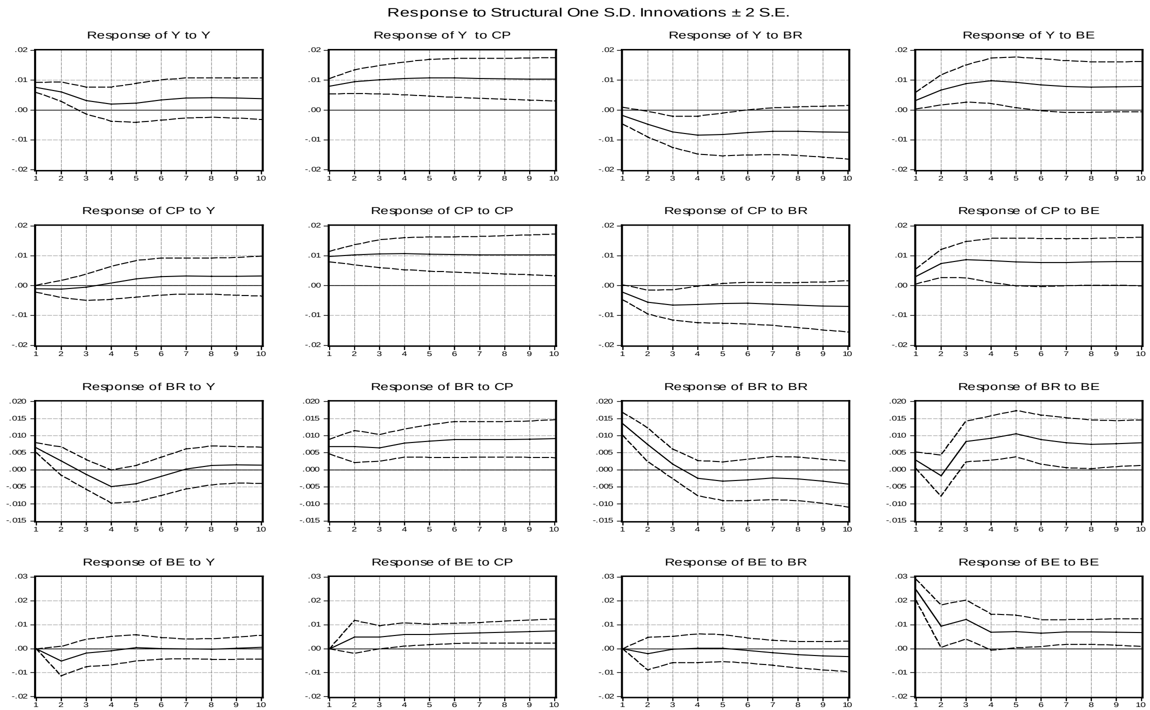
$$\begin{array}{ll}
t = 1: & x_1 - s_1 = a_1 + \lambda(a_1 - 2a_2 + a_3) \\
t = 2: & x_2 - s_2 = a_2 + \lambda(-2a_1 + 5a_2 - 4a_3 + a_4) \\
2 < t < T-1: & x_t - s_t = a_t + \lambda(a_{t-2} - 4a_{t-1} + 6a_t - 4a_{t+1} + a_{t+2}) \\
t = T-1: & x_{T-1} - s_{T-1} = a_{T-1} + \lambda(a_T - 4a_{T-2} + 5a_{T-1} - 2a_T) \; . \\
t = T: & x_1 - s_T = a_T + \lambda(a_{T-2} - 2a_{T-1} + a_T) \\
s_t = & \begin{cases} \dfrac{1}{T - \tau + 1} \displaystyle\sum_{t=\tau}^{T} \left(x_t - a_t\right) = \bar{s}, & t \geq \tau \\ 0, & t < \tau \end{cases}
\end{array} \tag{A4}$$

The intuition beyond (A4) is straightforward. The first 5 lines define the well-known first order conditions of the simple HP filter applied on the break-adjusted series, $x_t - s_t$. The last line, finally, determines the structural break endogenously as the average deviation of the smooth trend a from the original series being filtered after period τ. It must be kept in mind, however, that a and s are determined simultaneously by the equation system (A4).

A.2 IMPULSE-RESPONSE GRAPHS OF THE SVAR APPROACH

Fig. A.1

Set of Impulse-Response Graphs



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