An Analysis of Austrian Output Growth at a Sectoral Level

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Forschungsbericht/
Research Memorandum No. 327

July 1993

* Department of Economics. I would like to gratefully acknowledge comments by H. Pesaran, R. Alt, R. Kunst and T. Url.

Zusammenfassung

Pesaran, Pierse & Lee (1992) schlugen einen multisektoralen Ansatz zur Schätzung der Persistenz im BIP vor: Die Vermutung, anhand einer Vektorautoregression der BIP-Sektoren könnten genauere Persistenzschätzungen als mit univariaten Verfahren erzielt werden, wurde für US-Daten bestätigt. In der vorliegenden Arbeit wird der multisektorale Ansatz für österreichische Daten mit univariaten ARIMA-Modellen verglichen.

Weiters werden die Effekte verschiedener exogener Schocks auf die österreichischen Sektoren geschätzt. Innovationen im BIP europäischer OECD-Länder zeigen eine hohe Persistenz, während Innovationen im Dreimonatszinssatz nur transitorischen Effekt ausüben. Signifikante Effekte der Schocks im Ölpreis und im effektiven realen Wechselkurs konnten weder für einzelne Sektoren noch für das aggregierte BIP nachgewiesen werden. Die Reaktion der einzelnen Sektoren auf die exogenen Schocks ist höchst unterschiedlich, wobei die Unterschiede im Einklang mit informellen Beschreibungen stehen.

Summary

Recently, Pesaran, Pierse & Lee (1992) argued that estimation of aggregate output persistence may be more accurate when based on disaggregated data. They proposed a multisectoral approach based on a VAR of sectoral output growth. This work applies their methodology to Austrian output data with particular emphasis on the long-run impact of foreign shocks. I compare persistence estimates based on a VAR-model of sectoral output growth rates with those obtained from univariate ARIMA models of aggregate output.

Furthermore sectoral response to innovations in several exogenous variables is estimated. Innovations in European OECD GNP appear to have high persistence, whereas long-run influence of 3-month interest is not significant. Persistence estimates differ remarkably across sectors in accordance with informal characterizations of sectoral response. Influence of oil price and exchange rate innovations is statistically insignificant.

1 Introduction

Following the work of Nelson & Plosser (1982) it has now become a widely accepted view that aggregate output has a unit root and thus is best represented by a difference-stationary process. The extent to which the effects of shocks to economy persist over time has been the subject of extensive investigation over the past few years. Persistence estimates however vary considerably depending on the estimation procedure used. Among others Christiano & Eichenbaum (1989) found that the selection of lag length in univariate ARIMA-estimates influences persistence estimates considerably. They conclude that "data simply do not discriminate between the trend stationary and difference stationary views of real U.S. GNP."

The difficulties in determining the long run properties of time series based on a relatively small number of observations are not surprising. Recent approaches use error decompositions in multivariate models (e.g. Blanchard & Quah, 1989) based on information contained in the interaction of GDP with a stationary macroeconomic variable. It seems that these lead to more reliable estimates. [1]

Another approach was taken by Pesaran, Pierse & Lee (PPL; 1993) and Lee, Pesaran & Pierse (LPP; 1992). They argued that extra information obtained by a disaggregated analysis of sectoral output data may lead to a better representation of the stochastic process underlying aggregate output growth and thus may improve persistence estimates. Based on a multivariate approach they introduced a multisectoral persistence measure. They also showed how persistence effects may be partially lead back on exogenous macro variables.

The present paper applies PPL's methodology to Austrian output data. Six Austrian sectors are included in the analysis. After a short description of the multisectoral approach Dickey-Fuller-tests and a cointegration analysis of sectoral output growth are performed. Subsequently multisectoral persistence estimates are reported and compared with estimates based on univariate ARIMA models. Finally long-run impacts of unexpected changes in oil prices, interest rates, GNP of European OECD countries and effective exchange rate on sectoral outputs are estimated.

[1] In particular the results of Blanchard & Quah (1989) are insensitive against different VAR and VARMA-specifications (Rünstler & Hofer, 1992) and the choice of the stationary variable (Url & Wehinger, 1990).

2 Multisectoral Persistence Measures

Consider the multivariate linear difference stationary process

$$(1-L)y_t = \Delta y_t = \mu + A(L)\epsilon_t \tag{1}$$

where the m x 1 vector $\Delta y_t = (\Delta y_{1,t}, \ldots, \Delta y_{m,t})$ denotes sectoral output growth, μ is a vector of constants and ϵ_t is m-dimensional white noise with covariance matrix Σ. L denotes the lag operator. The moving average polynomial

$$A(L) = \sum_{i=0}^{\infty} A_i L^i$$

is assumed to be absolutely summable, i. e. $\sum_{i=0}^{\infty} \|A_i\|$, where $\|A_i\| = tr(A_i A_i')$. As in the univariate case the extent to which equation (1) deviates from a trend-stationary process is captured by the spectral density $f_{\Delta y}(\omega)$ of Δy_t at frequency zero. It is given by (see PPL)

$$2\pi f_{\Delta y}(0) = A(1)\Sigma A(1)'. \tag{2}$$

Let w be a constant m x 1 vector. For the linear combination $w'\Delta y_t$ the spectral density at frequency zero has the value $2\pi f_{w'\Delta y}(0) = w'A(1)\Sigma A(1)'w$. Scaling by the conditional variance $V(w'\Delta y_t | \epsilon_{t-1}, \ldots) = w'\Sigma w$ of $w'\Delta y$ given past innovations leads to the persistence measure

$$P_{w'y} = \left(\frac{w'A(1)\Sigma A(1)'w}{w'\Sigma w} \right)^{1/2}. \tag{3}$$

that may be regarded as a direct extension of the Campbell & Mankiw (1987) approach (see PPL).[2]

[2]For the univariate model $\Delta y_t = \mu + a(L)\epsilon_t$ the spectral density at frequency zero is given by $2\pi f_{\Delta y}(0) = a^2(1)\sigma_\epsilon^2$. Thus the Campbell & Mankiw (1987) persistence measure $P_y = a(1)$ can be regarded as the square root of $2\pi f_{\Delta y}(0)$ scaled by σ_ϵ.
The equivalence to the multisectoral approach can be best seen by Cholesky-decomposition $\Sigma = SS'$ with an arbitrary triangular matrix S. Let e_i be a m x 1 selection vector with unity at its i^{th} element and zero elsewhere. For the transformed process $\Delta y_t = A(L)S\xi_t = C(L)\xi_t$ with $\xi_t = S^{-1}\epsilon_t$, persistence of sector $y_{i,t}$ then simplifies to

$$P_i^2 = e_i'C(1)C(1)'e_i / e_i'e_i = \sum_{k=1}^{m} (c_{ik}(1))^2$$

Persistence P_{ik} of a particular shock $\xi_{k,t}$ on sector $y_{i,t}$ then could be defined as $P_{ik} = c_{ik}(1)$. However, since orthogonalization is arbitrary, Choleski decomposition does not provide unique values for P_{ik}.

It is important to note that (3) takes both direct effects - induced by $A(L)$ - and indirect effects - through contemporaneous residual covariances Σ - into account.

We can construct a matrix $P = (P_{ij})$ of cross-sectional persistence measures. The long-run-effects of shocks to sector j in sector i are given by (see Lee & Pesaran, 1993)

$$P_{ij} = \frac{e_i' A(1)\Sigma A(1)' e_j}{(e_i'\Sigma e_i)^{1/2}(e_j'\Sigma e_j)^{1/2}}. \tag{4}$$

Persistence of sector i is given by $P_i = \sqrt{P_{ii}}$ according to (3). A measure for persistence of aggregate output growth $\Delta Y_t = \sum_{i=1}^{m} y_{i,t}$ is obtained by using the selection vector $w = (1, 1, \ldots, 1)$ in (3).

Furthermore, as shown in PPL, possible cointegrating relationships are taken properly into account. The existence of cointegrating relationships leads to a rank reduction of the matrix P. Each cointegrating vector α fulfills the condition $\alpha' A(1) = 0$. This is equivalent to $\alpha' f_{\Delta y}(0) = 0$ and finally to $\alpha' P = 0$.

Persistence of Macro Shocks

Consider an extension of equation (1) by including a p x 1 vector of innovations $\nu_t = (\nu_{1,t}, \ldots, \nu_{p,t})'$ on the right hand side of (1). Innovation $\nu_{j,t}$ in a certain macroeconomic variable $x_{j,t}$ is given by the equation

$$x_{j,t} = z_{j,t}'\gamma_j + \nu_{j,t}, \tag{5}$$

where $z_{j,t}$ is a vector of predetermined variables. For the extended model

$$\Delta y_t = \mu + D(L)\nu_t + A(L)\epsilon_t \tag{6}$$

to be identified one must assume ϵ_t and ν_t to be uncorrelated. The polynomial $D(L)$ is defined by $D(L) = \sum_{i=0}^{\infty} D_i L^i$, where the D_i are m x p matrices. The p x p covariance matrix of ν_t is denoted by Φ.

An extension of the above persistence measure can be readily obtained for (6) using $\Delta y_t = (A(L), D(L))(\epsilon_t', \nu_t')'$ and noting that $\Sigma\Phi = 0$. Cross-sectoral persistence measure (4) can be extended to

$$P_{ij} = \frac{e_i' \left(A(1)\Sigma A(1)' + D(1)\Phi D(1)'\right) e_j}{(e_i' \left[D(0)\Phi D(0)' + \Sigma\right] e_i)^{1/2}(e_j' \left[D(0)\Phi D(0)' + \Sigma\right] e_j)^{1/2}}. \tag{7}$$

Furthermore sectoral persistence $P_i = \sqrt{P_{ii}}$ can be decomposed into two components. One component (P_{S_i}) is caused by the identified macro shocks ν_t, the other component (P_{O_i}) is due to "other" shocks ϵ_t:

$$P_i^2 = \lambda_i P_{S_i}^2 + (1-\lambda_i) P_{O_i}^2 \tag{8}$$

where

$$P_{S_i}^2 = \frac{e_i' D(1)\Phi D(1)' e_i}{e_i' D(0)\Phi D(0)' e_i}, \qquad P_{O_i}^2 = \frac{e_i' A(1)\Sigma A(1)' e_i}{e_i' \Sigma e_i}$$

and

$$\lambda_i = \frac{e_i' D(0)\Phi D(0)' e_i}{e_i' \left[D(0)\Phi D(0)' + A(0)\Sigma A(0)'\right] e_i}.$$

Note that λ_i is the relative size of macro shocks in sector i. Moreover the component due to macro shocks $P_{S_i}^2$ can be further decomposed into persistence due to particular shocks $\nu_{j,t}$:

$$P_{S_i}^2 = \sum_{j=1}^{p} \theta_j \left(P_{S_{ij}}^2 + P_{SX_{ij}}^2 \right) \tag{9}$$

where

$$P_{S_{ij}}^2 = \frac{e_i' d_j(1)\Phi_{jj} d_j(1)' e_i}{e_i' d_j(0)\Phi_{jj} d_j(0)' e_i} \qquad \text{and} \qquad \theta_j = \frac{e_i' d_j(0)\Phi_{jj} d_j(0)' e_i}{e_i' D(0)\Phi D(0)' e_i}.$$

$P_{S_{ij}}^2$ denotes the persistence caused by direct effects of shocks $\nu_{j,t}$. $d_j(z)$ is the jth row of $D(z)$. The interaction terms $P_{SX_{ij}}^2$ capture the effects due to the covariances between the particular shocks $\nu_{j,t}$ (see LPP).

The above decompositions (8) and (9), however, require $e_i' D(0)$ and $e_i' d_j(0)$, respectively, being different from zero, as can be easily verified. Consequently, estimates of $D(0)$ being not significantly different from zero may lead to unreliable estimates of $P_{S,i}^2$ and accordingly to high standard errors. In this case it may be preferable to use *weighted* persistence measures

$$P_{S_i}^{\lambda} = \sqrt{\lambda_i} P_{S_i} \qquad \text{and} \qquad P_{O_i}^{\lambda} = \sqrt{1-\lambda_i} P_{O_i} \tag{10}$$

for the relative size of the long-run-impacts of macro and other shocks and accordingly $\sqrt{\lambda_i \theta_j} P_{S_{ij}}$ for the relative size of long-run impact of particular macro shocks $\nu_{j,t}$.

The system is estimated by vector autoregression:

$$B(L)\Delta y_t = D^*(L)\nu_t + \epsilon_t \tag{11}$$

where $B(L) = I - \sum_{k=1}^{K} B_k L^k$ and $D^*(L) = \sum_{s=0}^{S} D_s^* L^s$. Moving average representation can be achieved by the transformation $\Delta y_t = B^{-1}(L) D^*(L)\nu_t + B^{-1}(L)\epsilon_t$.

3 Results

For analysis of Austrian output six sectors shown in Table 1 were considered. They account for about 80 % of total GNP. Sectors Government and Agriculture were omitted. Furthermore, in order to reduce the number of sectors, Mining was subsumed under sector Energy. The quarterly data range from 67:1 to 89:4. They are seasonally adjusted by CENSUS-X11.[3]

The remainder is organized as follows. First the results of unit root tests for the particular sectors are documented. As the unit root hypothesis is rejected for none of the sectors a cointegration analysis using Johansens's (1988) method is performed. The analysis yields only one cointegrating vector thus indicating the relative importance of sectorspecific shocks. Subsequently sectorspecific and aggregate persistence measures are reported for model (1). Finally the effects of several foreign shocks to Austrian economy are estimated. These shocks include innovations in GNP of European OECD-countries, oil prices, effective exchange rate and 3-month interest rate.

Unit Roots and Cointegration in Sectoral Outputs

The existence of unit roots in sectoral output levels is necessary for the ongoing analysis to be meaningful.

Additionally output growth rates must be stationary. Both conditions are well satisfied for all sectors. This can be seen from Table 1 where the results of augmented Dickey-Fuller-tests are reported. According to the procedure proposed by Perron (1989) also a break in time-trend in 1974:3 was considered. However, the results remain essentially the same.[4]

Cointegrating relationships

$$\Delta y_t = \mu + \sum_{k=1}^{K} B_k \Delta y_{t-k} + \Pi y_{t-K-1}$$

were estimated using Johansen's (1988) Maximum Likelihood method. The procedure identified one significant cointegrating vector at the 5 percent level according

[3]Data are taken from the database of the Austrian Institute for Economic Research.

[4]ADF-tests were computed with a simple time trend. The number of lagged differenced residuals was chosen so that the value of the Ljung-Box-statistics was insignificant at the 20 percent level.

TABLE 1. Augmented Dickey-Fuller-Tests and Cointegration analysis

	ADF - Tests		Cointegration analysis		
sector	output (level)	output (growth)	Eigen values	trace crit.	λ_{max} crit.
Total	-2.86	-3.95 *	.465	2.05	2.05
Manufacturing	-1.91	-3.97 **	.322	6.97	4.91
Energy	-2.05	-5.75 **	.281	28.35	21.37
Construction	-1.89	-6.34 **	.242	53.78	25.43
Trade	-2.40	-5.92 **	.061	83.72	29.94
Transportation	-2.61	-5.59 **	.026	131.92	48.19
Services	0.91	-6.25 **			

Notes :Estimation period is 68:2 to 89:1. Criticial values for ADF-tests are -3.41 and -3.96 ($\alpha = 0.01$), respectively. 5%-critical values for trace- and λ_{max}-criteria are 95 and 39 (see Johansen, 1998).

to both trace- and λ_{max}-criteria. Results are reported in Table 1 for 4 lags in accordance to subsequently used models. The number of cointegrating vectors, however, remains unchanged across different lag lengths from 2 to 4. Furthermore tests of pairwise cointegration in sectoral output confirm absence of a high number of cointegrating relationships. Cointegration tests between each possible pair of sectors were performed using the method by Engle & Granger (1987). A time trend was not included into the analysis. The null of no cointegration was rejected in 4 cases among 30 tests. These include Energy with Trade and Traffic and Trade with Traffic and Services.

As argued by PPL the small number of cointegrating vectors indicates that there are many independent sources of stochastic trends in aggregate output. This clearly is a strong rationale for multisectoral analysis.

Multisectoral Persistence Estimates

The above error-correction model for 6 sectors was estimated with 4 lags over the period 68:2 to 89:1. The unrestricted model contains 180 parameters, not counting the parameters of the covariance matrix, and thus clearly is overparametrized. Estimation of a restricted model seems more appropriate. However, estimation techniques for cointegration analysis within a restricted VAR-model remain to be developed. Thus in line with PPL and LPP the error-correction term will be ommitted in the ongoing analysis. While this does not affect interpretation of persistence matrix P, it represents a misspecification of the model. However, existence of only one significant cointegrating vector as well as the lack of pairwise cointegration suggest that the resulting estimation bias may not be too large.

PPL proposed a restriction where sector i is made dependent on its own past and the past sum of all other sectors. The system then also can be written in the form

$$\Delta y_t = \mu + \sum_{k=1}^{K} B_k \Delta y_{t-k} + \sum_{k=1}^{K} C_k \Delta y_{t-k}^{-} + \epsilon_t, \tag{12}$$

where $\Delta y_t^{-} = (\Delta y_t^{(-1)}, \ldots, \Delta y_t^{(-m)})$, the ith element $\Delta y_t^{(-i)} = \sum_{j \neq i} \Delta y_{i,t}$ being the sum of all sectors without sector i. B_k and C_k are diagonal m x m matrices. The number of parameters reduces to 54. The model can be efficiently estimated by seemingly unrelated regression.

TABLE 2. Likelihood and LR-Tests against model A

model	A	P	B	C
no. of				
parameters	150	54	77	51
ln L	2403.7	2298.3	2361.1	2357.8
LR-test		210.6	85.1	91.7
		(<.001)	(.27)	(.68)

Unfortunately, this kind of restriction is highly rejected for Austrian data. The value of the Likelihood Ratio test is 210.6, which is highly significant (see Table 2). Thus a relaxation of restriction (12) was considered by adding those sectors j whose influences on sector i deviate significantly from the influence of $\Delta y_t^{(-i)}$. These sectors may be regarded as responsible for rejection of (12). Search for those sectors was done by looking separately at each equation of model (12). Lagged values of sector j were added as explanatory variables to equation i. The equations

$$\Delta y_{i,t} = \mu_i + \sum_{k=1}^{K} B_k \Delta y_{i,t-k} + \sum_{k=1}^{K} C_k \Delta y_{t-k}^{(-i)} + + \sum_{k=1}^{K} c_k^{*} \Delta y_{j,t-k} + \epsilon_{i,t} \tag{13}$$

were estimated for each $j \neq i$. F-tests of

$$H_0 : c_1^{*} = c_2^{*} = \ldots = c_K^{*} = 0$$

were used to identify the sectors with deviate influence. In particular, Construction appeared to exert deviate influence on Manufacturing, Energy and Trade, furthermore Manufacturing on Trade, Traffic and Services and Services on Trade.

Adding significant lags to equations (12) lead to a restricted model B that was not rejected by the LR-Test. It contains 77 parameters. Finally, further restrictions were imposed by removing insignificant lags. In particular, parameters for lagged $\Delta y_t^{(-i)}$ were insignificant in most cases. The resulting model will be denoted by C and has 51 parameters.

Residuals of both models B and C pass Ljung-Box-tests for absence of serial correlation and tests for skewness and kurtosis. Table 2 reports the loglikelihoods for the unrestricted model A, PPL's restriction (P) and models B and C and corresponding Likelihood Ratio tests.

TABLE 3. Sectoral Persistence Estimates based on models A, B, C

Sector	A	B	C
Total	1.890	1.818	1.489
	(0.862)	(0.727)	(0.311)
Manufacturing	1.661	1.660	1.327
	(0.703)	(0.597)	(0.241)
Energy	0.693	0.744	0.699
	(0.277)	(0.255)	(0.133)
Construction	1.473	1.471	1.185
	(0.710)	(0.603)	(0.285)
Trade	1.075	0.871	0.752
	(0.453)	(0.328)	(0.147)
Transportation	1.809	1.658	1.367
	(0.793)	(0.643)	(0.303)
Services	1.710	1.638	1.664
	(0.761)	(0.610)	(0.407)

Notes: Simulated standard errors (1000 replications) in parentheses.

Table 3 presents persistence estimates for models A, B and C. Standard errors (in parentheses) were computed by Monte Carlo simulations with 1000 replications. Restriction B alters estimates only marginally. Exclusion of insignificant lags (model C), however, leads to considerably lower estimates. Furthermore it is evident that estimates of model C are much more accurate than those of A and B. Persistence measures vary considerably across sectors. Especially sectors Energy and Trade exhibit low, Services relatively high persistence.

It is of particular interest to compare the above aggregate persistence estimate with those obtained from univariate models. Univariate persistence estimates for the sum of sectoral output growths ΔY_t (Total) are shown in Table 4.[5]

TABLE 4. Univariate ARIMA-models for the Total

	(0,3)	(1,2)	(2,1)	(3,0)
LR-Test	4.751	5.269	4.616	5.230
AIC	-8.816	-8.224	-8.814	-8.821
SIC	-8.597	-8.603	-8.595	-8.603
LJ(12)	3.583	3.524	4.540	3.667
P_Y	1.292	1.411	1.460	1.440
	(0.221)	(0.389)	(0.446)	(0.375)

Decision for one model appeared to be extremely difficult. Based on AIC, SIC, autocorrelation of residuals and significance values of parameters several reasonable models were chosen. The most reasonable specification seems to be an ARMA(1,2) specification. None of the models, however, is rejected against a random walk with drift by the Likelihood Ratio test.

Univariate persistence estimates are quite close to that of model C. Their standard errors, however, are somewhat higher. They only exception is the ARMA(0,3)-model. It exhibits a lower persistence estimate with also a lower standard error. Apart from that model persistence estimates coincede well with model C.

Impact of Macroeconomic Shocks

A small open economy with a fixed exchange rate and high export shares like Austria is highly dependent on foreign events. Exports as well as interest rates and effective exchange rate can be regarded as exogenous. By the methodology described in

[5]Notes: AIC and SIC are presented in the formulation of Luetkepohl (1991, p. 129) as $AIC = \ln(\hat{\sigma}^2) + 2k/T$ and $SIC = \ln(\hat{\sigma}^2) + k \ln T/T$. Thus models with low values are to be selected. The critical value for the LR-test of the ARMA-models against $\Delta y_t = \mu + \epsilon_t$ is given by $\chi^2_{3;.05} = 7.81$ in all cases.

section 2 the following analysis pursues estimation of long-run impacts of foreign shocks to Austrian sectoral output. Unexpected changes in

- GNP of European OECD-countries (g_t)
- nominal effective exchange rate (e_t)
- Austrian 3-month interest rates (r_t)
- nominal oil prices in ATS (o_t)

will be taken into account.[6]

A look at the identifying condition $\mathbf{E}\epsilon_t\nu_t' = 0$ may be noteworthy. As the shocks ν_t are predetermined this condition can be regarded as an orthogonalization of ϵ_t on ν_t. The problems of interpreting impulse response curves in this context have been extensively discussed in the review of Sims (1980) error decompositions and the structural VAR approach (e.g. Blanchard & Watson, 1986; Bernanke, 1986). The arguments also hold for the ongoing persistence decomposition. For meaningful interpretation one must accept ϵ_t as conditional on given ν_t. By the use of the above mentioned innovations, however, it can be well argued that this assumption may not be too heavily violated. The reason for regarding interest rates and effective exchange rate as exogenous lies in the fact of Austria's highly committed exchange rate policy. Austrian currency (ATS) has been bound closely to the D-Mark over nearly the whole observation period with the exception of the period 1973-1974, where the ATS was devalued. Thus interest rates are largely determined by German ones. Furthermore, effective exchange rate is essentially determined by exchange rates to DM.

The particular shocks were constructed through autoregressive equations. ADF-tests did not reject the unit root hypothesis in any of the macro variables. Thus GNP of European OECD countries (Euro-GNP), effective exchange rate and nominal oil prices were taken in first differences of logs. Interest rate was differenced. Finally the particular shocks were estimated by the following equations. Estimation was done by OLS.

There appeared to be no significant influence of any lagged variable to the exchange rate. This may be due to little variation in the time series because of the constancy

[6] Interest rates, effective exchange rate and oil prices are from WIFO-database. OECD-GNP and GNP of European OECD-countries were taken from OECD 'Main Economic Indicators'.

$$
\begin{aligned}
g_t = {} & \underset{(3.01)}{.00394} + \underset{(1.01)}{.1006 w_{t-1}} + \underset{(1.69)}{.1674 w_{t-2}} \\
& + \underset{(1.43)}{.1376 w_{t-3}} - \underset{(-2.91)}{.0147 o_{t-4}} + \hat{\nu}_{g,t}, \quad R^2 = .234 \\
r_t = {} & \underset{(.02)}{.00003} + \underset{(2.00)}{.2133 r_{t-4}} - \underset{(-2.44)}{.2606 r_{t-5}} \\
& - \underset{(-2.58)}{.2764 r_{t-6}} + \hat{\nu}_{r,t}, \quad R^2 = .183 \\
e_t = {} & .00211 + \hat{\nu}_{e,t} \\
o_t = {} & \underset{(0.71)}{.01080} + \underset{(3.74)}{.3899 o_{t-1}} + \hat{\nu}_{o,t}, \quad R^2 = .147
\end{aligned}
$$

of the DM-rate which has highest weight in the effective rate. For Euro-GNP lagged values of total OECD-GNP (w_t) were a better predictor than lagged values of Euro-GNP itself. Additionally the fourth lag of oil prices (U.S. \$) was added. In the interest rate equation lags lower than 4 were highly insignificant and thus omitted. Table 5 shows the correlations between the shocks. These are rather independent with the exception of a highly negative correlation between exchange rate and oil price innovations.

All subsequent estimations are based on model C. Estimation of

$$B(L)\Delta y_t = D^*(L)\hat{\nu}_t + \epsilon_t \tag{14}$$

however, is complicated by the fact that the macro shocks $\hat{\nu}_t$ have been estimated in a previous step and thus contain an error. Least Squares estimates of $B(L)$ and $D^*(L)$ are still consistent (Pagan, 1984), but the estimate of the covariance matrix estimate is biased (Pesaran 1987, p. 174ff). The model is

$$
\begin{aligned}
B(L)\Delta y_t &= D(L)\nu_t + \epsilon_t \\
x_{j,t} &= z'_{j,t}\gamma_j + \nu_{j,t},
\end{aligned}
$$

but we estimate

$$
\begin{aligned}
B(L)\Delta y_t &= D(L)\hat{\nu}_t + u_t \\
&= D(L)\nu_t + D(L)(\hat{\nu}_t - \nu_t) + u_t \\
&= D(L)\nu_t + \epsilon_t
\end{aligned}
$$

Therefore the sample covariance matrix $\hat{\Sigma}_u$ is a biased estimate of Σ_ϵ and must be corrected accordingly. System (11) was estimated by SUR and $\hat{\Sigma}_u$ subsequently decomposed into $\hat{\Sigma}_\epsilon$ and $\hat{D}(1)\hat{\Sigma}_{\hat{\nu}_t - \nu_t}\hat{D}(1)'$. This was done using a multivariate extension of an approach proposed by Pesaran (1987, p.174ff). The derivation is shown in the appendix.

F-tests on the impact of particular macro shocks to sectoral output growth are presented in Table 6, two hypothesis being tested. Under null hypothesis

$$H_1 : d^*_{j,i} = 0, \; i = 0, \ldots, 4$$

shocks have no effect on output growth at all, while under

$$H_2 : \sum_{i=0}^{4} d^*_{j,i} = 0$$

macro shocks are allowed to have short run effects, but no long run impact on output growth. Clearly, H_1 implies H_2, but not vice versa.[7]

TABLE 5. Correlations between macro shocks

	Europ. GNP	3-month rate	oil price
exc. rate	-0.117	-0.021	-0.405
OECD-GNP		+0.253	-0.057
3-month rate			-0.009

Tests of H_1 indicate a significant impact on sectoral output growth rates only for Euro-GNP growth and interest rates. The impact of both shocks also is found to be significant by an overall Likelihood Ratio test of H_1. Shocks in Euro-GNP growth exert significant influence on Austrian GNP growth primarily through sectors Manufacturing, Trade and Transportation. Long-run impact proves to be significant also for sector Services. Significant effects of unexpected changes in interest rates are found in sectors Trade, Transportation and Energy. Long-run impact is not significant in any of the sectors. Significant effects of oil price shocks exist in sector Transportation. However, the only significant coefficient appearing at lag 0 surprisingly has positive sign indicating that price shocks may have been partially

[7]Sectoral output equations were estimated by OLS without correction of error variances, which nevertheless implies correct distribution of the test statistics under the null.

transferred to other sectors. The lack of an overall response to oil price shocks seems somewhat puzzling at first sight. However, it may be a reasonable explanation that at both large oil price shocks in 1973 and 1979 Austrian economy responded rather to the slowdown in OECD-countries than the oil price itself.[8]

Furthermore oil price shocks seem to have been partially offset by a countervally movement of effective exchange rate. Correlation between shocks in oil price and effective exchange rate appears to be −.40. Finally, the failure in finding statistical significant influences of effective exchange rate shocks again has to be seen in the light of high export shares to Germany and the fixed ATS-DM exchange rate.

TABLE 6. F-Tests for the impact of macro shocks (model C)

	H_1				H_2	
SECTOR	exchange rate	Euro GNP	3-month rate	oil price	Euro GNP	3-month rate
Manufacturing	1.01	4.82	0.82	1.06	9.10	0.70
	(.41)	(.00)	(.53)	(.38)	(.00)	(.40)
Energy	0.51	0.65	1.90	0.67	1.38	0.01
	(.77)	(.66)	(.10)	(.64)	(.24)	(.96)
Construction	0.87	0.92	1.24	1.05	0.16	2.01
	(.50)	(.47)	(.29)	(.39)	(.69)	(.16)
Trade	1.27	5.58	2.35	1.48	17.05	1.93
	(.28)	(.00)	(.05)	(.20)	(.00)	(.16)
Transportation	0.79	3.99	2.78	3.00	11.67	2.39
	(.55)	(.00)	(.02)	(.02)	(.00)	(.12)
Services	0.26	1.57	0.34	0.72	7.46	1.35
	(.93)	(.17)	(.88)	(.60)	(.01)	(.25)
LR-test	30.12	55.28	48.14	32.03	—	—
	(.45)	(.00)	(.02)	(.36)	—	—

Notes: Probability values in parentheses. Distributions of test statistics are $F_{5,70}$ (H_1) and $F_{1,70}$ (H_2). The LR-statistics is distributed with χ^2_{30}.

Before performing persistence decomposition, a view on possible cointegrating

[8]This is also reflected in the fact of Austria's output response to both oil price shocks occurring two quarters later than response of Euro-GNP. Accordingly, one can find significant influence of oil price shocks at higher lags. This influence largely disappears, however, when controlled for lagged Euro-GNP.

relationships between sectoral output and exogenous shocks may be useful. Since cointegrating relationships transform one by one into restrictions on the spectral density at frequency zero (see section 2), they impose restrictions on the decomposition of persistence measures. Particularly cointegration between Euro-GNP and sectoral output seems critical in this respect. The data support this possibility only weakly, however. Pairwise cointegration tests between sectoral output and integrated shocks on Euro-GNP $(1 - L)^{-1}\nu_{g,t} = \sum_{j=1}^{t} \nu_{g,j}$ were performed using the method by Engle & Granger (1987). Absence of cointegration between sectoral output and integrated shocks is accepted in all cases at the 10 percent level. The same holds for aggregate output. Cointegration tests with Euro-GNP itself indicate cointegration only for sectors Energy and Trade. Furthermore, absence of cointegration between aggregate output and Euro-GNP is accepted at the 10 percent level.

For persistence decomposition, innovations in Euro-GNP and 3-month interest rate were added to equations of model C. Again only significant lags were included, with a maximum of four lags. Contemporaneous innovations were always included, as long as a lag of the particular innovation appeared in the equation. In particular, innovations in Euro-GNP appear in all equations but Energy and Construction. Interest rate innovations were added to sectors Energy, Construction, Trade and Transportation. Persistence estimates are presented in Table 7. It shows the overall persistence estimate P_i and its decomposition into weighted persistence $P_{S_i}^{\lambda}$ and $P_{O_i}^{\lambda}$ of macro and other shocks, respectively.[9]

Persistence estimates for the particular shocks also are presented by their weighted values $\sqrt{\lambda_i \theta_j P_{S_{ij}}}$. Interaction effects are generally small and thus were omitted.

Inclusion of exogenous shocks leaves the total aggregate persistence estimate essentially unchanged. However, its standard error reduces slightly. Persistence estimates of macro shocks provide quite meaningful values. Weigthed persistence of Euro-GNP reaches values between 0.6 and 1.0 for sectors Services, Manufacturing, and Transportation. It is somewhat lower for Trade. With the exception of Services the results coincide with informal characterizations of these sectors as being highly dependent on exports. The remaining sectors face only little foreign competition also on home markets. In sum, a considerable part of total persistence is due to GNP of European OECD-countries. Innovations in Euro-GNP account for about thirty percent of variation in aggregate output growth in the long run. Unexpected changes in interest rates exhibit generally low long run impact. It is highest in sector Construction,

[9]Since coefficients of D_0^* were insignificant in many cases, decomposition (8) gives unreliable estimates (see section 2).

TABLE 7. Weighted persistence estimates of macro and sector specific shocks

SECTOR	total	macro	other	λ_i	Euro GNP	3-month rate
Total	1.422	0.774	1.192	.087	0.786	0.053
	(.245)	(.219)	(.186)	(.047)	(.219)	(.099)
Manufacturing	1.279	0.775	1.017	.209	0.775	0.000
	(.205)	(.201)	(.162)	(.086)	(.201)	(—)
Energy	0.645	0.179	0.619	.024	0.000	0.179
	(.106)	(.095)	(.088)	(.039)	(—)	(.075)
Construction	1.162	0.350	1.108	.005	0.000	0.350
	(.243)	(.183)	(.209)	(.021)	(—)	(.125)
Trade	0.739	0.395	0.624	.165	0.295	0.043
	(.129)	(.156)	(.090)	(.079)	(.165)	(.042)
Transportation	1.347	0.832	1.060	.175	0.601	0.078
	(.235)	(.234)	(.174)	(.082)	(.240)	(.052)
Services	1.639	1.007	1.307	.020	1.007	0.000
	(.349)	(.345)	(.254)	(.038)	(.345)	(—)

Notes: Simulated standard errors (1000 replications) in parentheses.

which is characterized by particularly long financing horizons.

Again it may be of interest to compare the results with univariate estimates for the sum of sectoral output growth ΔY_t. Estimates are based on an autoregressive model with three lags. Innovations again are added up to the fourth lag. F-tests on the impact of exogenous shocks indicate highly significant influence of Euro-GNP and also significant long run impact. Influence of interest rate is not significant, but long run impact approaches significance. [10]

Persistence estimates are presented for two models, model 1 unrestricted, model 2 containing lag 0 and the only significant lag 3 for both shocks. As can be seen from Table 8 the univariate approach leads to a considerably higher weight λ und thus to generally higher weighted persistence estimates of macro shocks. Standard errors are slightly higher than for extended model C.

[10] Values of test statistics for H_1 and H_2 are

European OECD-GNP	$H_1 : F_{5,70} = 7.89$	$(p < .001)$	$H_2 : F_{1,70} = 17.05$	$(p < .001)$
3-month rate	$H_1 : F_{5,70} = 1.27$	$(p = .29)$	$H_2 : F_{1,70} = 3.74$	$(p = .06)$

TABLE 8. Persistence estimates for univariate model of aggregate output

	total	macro	other	λ_i	Euro GNP	3-month rate
Model 1	1.627	1.313	0.961	0.291	1.350	0.474
(unrestricted)	(.340)	(.327)	(.245)	(.085)	(.341)	(.282)
Model 2	1.484	1.031	1.068	0.270	1.060	0.369
(restricted)	(.312)	(.255)	(.236)	(.081)	(.261)	(.196)

Notes: Simulated standard errors (1000 replications) in parentheses.

4 Conclusions

In their analysis of sectoral output growth for UK and US economies, PPL and LPP found considerably lower persistence estimates of the multisectoral measure compared to the univariate approach. Furthermore both studies report lower standard errors for the multivariate measure. Both has not been confirmed for Austrian data. Methodological differences lie primarily in the kind of restrictions imposed on the model. The comparison of models A, B and C proves the persistence measures to be quite sensitive with respect to overparametrization as well as in the univariate case. At least for Austrian data the difficulties in finding an appropriate model resulted in some arbitrariness of the final equations. Comparison of aggregate persistence measure in model C with univariate analysis indicates that the results of model C may be quite reasonable. In sum, however, there is little evidence for the multisectoral measure of aggregate persistence measure to be more accurate than the univariate one.

On the other hand relative size of sectoral persistent estimates remains roughly the same across different specifications. Regardless of the above mentioned shortcomings estimation of sectoral persistence and its decomposition due to particular sources proves to be an useful exercise in its own right. For Austrian data well-known informal facts have been confirmed, and quantified, in quite a reasonable manner. The analysis has given further insights into the transfer of foreign shocks to Austrian economy. Short-term interest rates seem to exert virtually no long-run impact on output growth. Innovations in GNP of European OECD-countries account for about thirty percent of variation in Austrian GNP growth in the long run, mainly through sectors Manufacturing, Transportation and Services. Oil price innovations seemingly have had primarily indirect impact on Austrian economy through European business cycle.

In the light of the very different reactions of particular sectors to exogenous shocks we may argue that an univariate approach cannot capture the impacts of shocks as well as multisectoral estimates. This can be illustrated for the case of interest rate shocks: in the disaggregated model there is very different sectoral response. Significant response in three sectors leads to overall rejection of H_1. However, for the univariate model, not only H_1 cannot be rejected, but contradictingly H_2 is rejected at a 6% level. Finally, estimation of long run response is based on only one significant parameter at lag 3.

Appendix: Correction of the Variance-Covariance-Matrix

According to the correction proposed in Pesaran (1987, p.174ff) the covariance matrix estimate $\hat{\Sigma}_\epsilon$ can be adjusted as follows. Consider model (6)

$$\begin{aligned} B(L)\Delta y_t &= D(L)\hat{\nu}_t + u_t \\ &= D(L)\nu_t + D(L)(\hat{\nu}_t - \nu_t) + u_t \\ &= D(L)\nu_t + \epsilon_t, \end{aligned}$$

where the regressors $\hat{\nu}_t = (\hat{\nu}_{1,t}, \ldots, \hat{\nu}_{p,t})$ have been estimated in a previous step by OLS of $x_{j,t} = z'_{j,t}\gamma_j + \nu_{j,t}$ These equations can be jointly written as

$$\begin{pmatrix} x_1 \\ x_2 \\ \vdots \\ x_p \end{pmatrix} = \begin{pmatrix} Z_1 & 0 & \ldots & 0 \\ 0 & Z_2 & \ldots & 0 \\ \vdots & \vdots & \vdots & \vdots \\ 0 & 0 & \ldots & Z_p \end{pmatrix} \begin{pmatrix} \gamma_1 \\ \gamma_2 \\ \vdots \\ \gamma_p \end{pmatrix} + \begin{pmatrix} \nu_1 \\ \nu_2 \\ \vdots \\ \nu_p \end{pmatrix}$$

or equivalently as

$$x = Z\gamma + \nu$$

As ν_t has been estimated by OLS we can use the relationship

$$\hat{\nu} - \nu = -Z(Z'Z)^{-1}Z'\nu = -P_z\nu.$$

It follows that $\hat{\nu}_{t-i} - \nu_{t-i} = -z'_{t-i}(Z'Z)^{-1}Z'\nu$. We are interested in an estimate of Σ_ϵ. From the above decomposition of $\hat{u}_t$

$$\begin{aligned} \epsilon_t &= \hat{u}_t - \textstyle\sum_{i=0}^{p} D_i(\hat{\nu}_{t-i} - \nu_{t-i}) \\ &= \hat{u}_t - \textstyle\sum_{i=0}^{p} D_i z'_{t-i}(Z'Z)^{-1}Z'\nu \end{aligned}$$

Using the restriction $\Sigma_{\epsilon\nu} = 0$ the covariance matrix Σ_ϵ can be decomposed into

$$\Sigma_\epsilon = \Sigma_u + \textstyle\sum_{i,j=0}^{p} D_i \mathbf{E}\left[(\hat{\nu}_{t-i} - \nu_{t-i})(\hat{\nu}_{t-j} - \nu_{t-j})'\right] D'_j$$

Thus with $\mathbf{E}\nu\nu' = \Omega \otimes I$ the covariance matrix $\mathbf{E}\left[(\hat{\nu}_{t-i} - \nu_{t-i})(\hat{\nu}_{t-j} - \nu_{t-j})'\right]$ can be consistently estimated by

$$\hat{\Sigma}_{ij} = (T - i - 1)^{-1} \sum_{t=i+1}^{T} z'_{t-i}(Z'Z)^{-1}Z'(\hat{\Omega} \otimes I)Z(Z'Z)^{-1}z'_{t-j}$$

and finally the estimate of Σ_ϵ is given by

$$\hat{\Sigma}_\epsilon = \Sigma_u + \sum_{i,j=0}^{p} \hat{D}_i \hat{\Sigma}_{ij} \hat{D}'_j$$

Consistency can be shown with the same arguments as in Pesaran (1987, p.174ff).

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