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# The short and long-run interdependencies between the Eurozone and the USA

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**Abstract** We estimate quarterly cointegrating vector autoregressive models for the Eurozone and the USA based on long-run restrictions derived from a dynamic open economy model. Three long-run relations between the Eurozone and the USA emerge: relative purchasing power parity, international interest parity and a stationary output gap between the two economies. Generalized impulse response functions show differences in the dynamic adjustment of the two economies. Due to the I(1)-characteristic of both output series and the stability conditions imposed by the long-run equilibrium relationships, shocks to the model produce level effects only, while growth rates converge to their long-run averages.

**Keywords** Cointegration · VAR · Steady state · Business cycle

**JEL Classification** F41 · E32 · C32

## 1 Introduction

The discussion on the link between short- and long-run variations in output focuses on the question whether it is reasonable to decompose output fluctuations into a trend and a cyclical component, with a growth theory explaining the trend and a business cycle theory explaining deviations from the trend. The analyses presented in this volume cast doubts on such a view. We address the interaction between long- and short-run fluctuations in an empirical model that explicitly relates business

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cycle variations to deviations from long-run equilibrium relations. In this model shocks to an economy push it away from the steady state, but subsequently an adjustment process starts, which drives the system towards a new steady state.

We implement such an adjustment process in a cointegrated vector autoregression (VAR). The autoregressive part models the short-term adjustment process, while the cointegration part determines how far the economy deviates from the long-run equilibrium. The model is structural in the sense that the identification of the long-run relationships in the error correction part does not rely on the restrictions implied by the solution of the eigenvalue problem (Johansen 1988, 1991), instead, we derive a priori steady state restrictions from economic theory, as suggested by Garratt et al. (2006).

We estimate two cointegrated VAR models using linearized restrictions based on the steady state equations of a dynamic open economy model: one for the Eurozone and another one for the USA. Direct and indirect trade links, knowledge flows, and strongly integrated financial markets create international transmission channels that are likely to affect both the trend and the cyclical component. In each of the two models we combine the contemporary interaction between the Eurozone and the USA with an intertemporal, dynamic interaction between long-run cointegrating relations and short-run variations.

Most of the previous work on the relation between the USA and the Eurozone involves simulations of large-scale macroeconomic models. In such models, the transmission of a shock to an exogenous variable or a policy measure is shown by comparing the baseline with alternative scenarios. For this purpose, policy reaction functions and exchange rate regimes have to be assumed. For example, Dalsgaard et al. (2001) study a change in US fiscal expenditures by 1% of GDP using the OECD Interlink model, and find an asymmetry in the interaction between the two areas. Fiscal shocks in the USA have a higher impact on Eurozone output than corresponding Eurozone shocks on output in the USA. Nevertheless, most of the work on the USA and the Eurozone treats both areas as large closed economies and therefore ignores international feedbacks; cf. Christiano et al. (1999) for the USA, Vlaar (2004) for the Eurozone. The closest in spirit to our model is the global VAR by Dees et al. (2007), which also captures interaction between the USA and the Eurozone. The cointegrated VAR approach allows taking account of international spillovers in a small and largely data-driven dynamic model that highlights the interaction between short-run and long-run effects.

In the next section we state the steady-state conditions of an open economy model. The model provides a set of endogenous variables for the cointegrated VAR and defines restrictions for the steady-state equilibrium. We then show the relation between theoretical steady-state conditions and the error correction vector of a cointegrated system. After describing the data and testing their time series properties, we define a VAR in levels and test for the number of cointegrating relations among our endogenous variables. We then discuss the results of the cointegrated VAR, describe the three long-run equilibrium relations that we identify, test for over-identifying restrictions implied by the open economy model and finally present generalized impulse response functions that show the dynamic adjustment of the model to unexpected variations in output, the interest rate, and the exchange rate. The last section offers concluding remarks.

## 2 Steady-state conditions

In this section we derive steady-state relations from a dynamic open economy model. We start with the optimization problem of a representative infinitely lived household. The household seeks to maximize the expected utility function,

$$E_{t=0}\sum_{t=0}^{\infty}\rho^t U\left(C_t, \frac{M_t}{P_t}, L_t\right), \tag{1}$$

where $E_{t=0}$ denotes the expectations operator conditional on information at $t = 0$, subject to the budget constraint (or current account balance),

$$C_t + K_{t+1} + \frac{M_t}{P_t} + \frac{B_t}{P_t} + \frac{e_t B_t^*}{P_t} = \frac{w_t L_t}{P_t} + (1-\delta+r_t)K_t + \frac{M_{t-1}}{P_t} + \frac{(1+i_{t-1})B_{t-1}}{P_t} + \frac{\left(1+i_{t-1}^*\right)e_t B_{t-1}^*}{P_t} + T_t, \tag{2}$$

by choosing infinite sequences of optimal consumption, $\{C_t\}_{t=0}^{\infty}$, nominal money holdings in home currency, $\{M_t\}_{t=0}^{\infty}$, labor in hours, $\{L_t\}_{t=0}^{\infty}$, investment, which in turn implies a sequence of capital stocks, $\{K_t\}_{t=0}^{\infty}$, home nominal bonds, $\{B_t\}_{t=0}^{\infty}$, and foreign nominal bonds, $\left\{B_t^*\right\}_{t=0}^{\infty}$. Foreign bonds are traded in foreign currency purchased at a nominal exchange rate, $e_t$. Further, $T_t$ denotes government transfers and $i_t$ and $i_t^*$ represent nominal interest rates on home and foreign bonds, respectively (here and below the asterisk denotes a foreign variable). Moreover, households earn a nominal wage, $w_t$, and firms pay a real rental rate on capital, $r_t$, each period. The rate of capital depreciation, $\delta$, and the discount factor, $\rho$, are constant. All units are deflated by the domestic price index, $P_t$.

From the perspective of the representative household the left hand side of Eq. 2 represents total period spending (in consumption units), whereas the right hand side of Eq. 2 reflects total real income in period $t$. Next, we formulate the dynamic optimization problem as a dynamic programming problem using the following Bellman equation:

$$\begin{aligned} V(s_t) = \max_{x_t} \Bigg[ & U\left(C_t, \frac{M_t}{P_t}, L_t\right) + \rho E_t\{V(s_{t+1})\} \\ & -\lambda_t\left(C_t + K_{t+1} + \frac{M_t}{P_t} + \frac{B_t}{P_t} + \frac{e_t B_t^*}{P_t} - \frac{w_t L_t}{P_t} - (1-\delta-r_t)K_t - \frac{M_{t-1}}{P_t}\right. \\ & \left. - \frac{(1+i_{t-1})B_{t-1}}{P_t} - \frac{\left(1+i_{t-1}^*\right)e_t B_{t-1}^*}{P_t} - T_t\right)\Bigg], \end{aligned} \tag{3}$$

where $x_t \equiv \left(C_t, M_t, L_t, K_{t+1}, B_t, B_t^*\right)$ is the control vector and $s_t \equiv (K_t, r_t, w_t, i_t, i_t^*, B_t, B_t^*, P_t, e_t, T_t, M_{t-1})$ is the state vector. We use the following parameterization for the period utility function:

$$U\left(C_t, \frac{M_t}{P_t}, L_t\right) = \frac{C_t^{1-\sigma_1}}{1-\sigma_1} + \frac{(M_t/p_T)^{1-\sigma_2}}{1-\sigma_2} - \frac{\sigma_3}{1+\sigma_3}\frac{L_t^{1-\sigma_3}}{1-\sigma_3}. \quad (4)$$

That is, households have additively separable constant relative risk aversion preferences for consumption, real balances, and leisure. The first order conditions corresponding to the maximization problem together with the functional assumption in Eq. 4 can be used to derive the uncovered interest parity, the Fisher parity, and a money demand equation. The log-linearized versions of these equilibrium relations (as local approximations around their steady state values $(\bar{c}, \bar{i}, \bar{i}^*, \bar{m}_r, \bar{p}, \bar{r}, \bar{w}_r, \bar{\pi}, ))$ are:

$$(\hat{m}_t - \hat{p}_t) = \frac{\sigma_1}{\sigma_2}\hat{c}_t - \frac{\rho}{\sigma_2}\hat{i}_t, \quad (5)$$

$$\hat{i}_t = \frac{1}{1-\rho}\frac{\bar{\pi}}{1-\bar{\pi}}E_t\{\hat{\pi}_{t+1}\} + (1-\bar{\pi})E_t\{\hat{r}_{t+1}\}, \quad (6)$$

$$\hat{i}_t - i_t^* = \frac{1}{1-\rho}(E_t\{\hat{e}_{t+1}\} - \hat{e}_t), \quad (7)$$

where $\pi_t \equiv (P_{t+1} - P_t)/P_t$ denotes the rate of inflation and a hat denotes percentage deviations from the steady state. Equation 5 describes the money market equilibrium, (6) represents the Fisher parity, and (7) is the uncovered interest rate parity.

From the first order conditions for the optimal holdings of home and foreign bonds follows:

$$E_t\left\{\left(\frac{U_C(C_{t+1},\cdot)}{U_C(C_t,\cdot)}\right)\frac{e_{t+1}}{e_t}\frac{P_t}{P_{t+1}}\right\} = E_t\left\{\left(\frac{U_C(C_{t+1}^*,\cdot)}{U_C(C_t^*,\cdot)}\right)\frac{P_{t+1}^*}{P_{t+1}^*}\right\}. \quad (8)$$

Since the two economies are symmetric with respect to their tastes and expectation formation, $C_t = C_t^*$ holds in equilibrium, hence (8) can only be true if the purchasing power parity condition holds

$$P_t = e_t P_t^*. \quad (9)$$

We focused on the consumer's problem so far. Firms, on the other hand, are assumed to be perfectly competitive, as in Garratt et al. (2006), and produce real aggregate output, $Y_t$, according to a constant returns to scale production function using labor, $L_t$, and capital, $K_t$, as inputs:

$$Y_t = F(K_t, L_t, A_t) = A_t L_t F\left(\frac{K_t}{A_t L_t}, 1\right), \quad (10)$$

where $F(\cdot)$ satisfies the Inada conditions. Labor augmenting technical progress, $A_t$, is represented as an index. Assuming free international technological diffusion ensures that in the long-run domestic technical progress is linked to technical progress in the rest of the world, $A_t^*$. Yet differences in levels may persist if the process of diffusion is incomplete (cf. Parente–Prescott 1994). This possibility is described by a factor $0 < \gamma \leq 1$ in the technology diffusion equation:

$$A_t = \gamma A_t^*. \tag{11}$$

Suppose the two countries are identical with respect to technology, $F(\cdot)$, labor and capital. Further, we assume a time-invariant employment rate, transform output into per-capita terms, and take logarithms, such that $y_t = \ln(Y_t/L_t)$. Under these assumptions, the steady state output gap between home and foreign is completely determined by the extent of impediments to full technological diffusion:

$$y_t - y_t^* = \ln(\gamma). \tag{12}$$

Different social security and taxation regimes may create a deviation of employment rates and thus capital intensities will not be identical. Equally important, country specific labor or goods market regulations may result in non-identical production technologies. Garratt et al. (2006) provide a more general version of Eq. 12, but the conclusion that the output gap will be constant in the long-run still holds.

Equations 5 through 7, 9, and 12 provide five long-run equilibrium relations to which the two economies converge in the long-run. To show how these theoretical steady state conditions can be used to derive restrictions for the cointegrating vectors of a VAR, it is useful to collect all endogenous variables from the steady state conditions into a vector $\mathbf{y}_t = \left(m_t, y_t, i_t, \Delta p_t, i_t^*, \left(p_t - p_t^*\right), e_t, y_t^*\right)$. This vector includes the real money stock, real output levels of home and foreign, nominal interest rates, the inflation rate, the price differential between home and foreign, and the exchange rate. Since the model generates a constant relation between consumption and output in the steady state, we can substitute output for consumption in the vector of endogenous variables. The cointegrating vectors of this model are defined as linear combinations of elements in $\mathbf{y}_t$ which are stationary, i.e. $\boldsymbol{\beta}'\mathbf{y}_{t-1} = \boldsymbol{\xi}_t$ with equilibrium errors, $\xi_{it}$ $i = 1,2,\ldots,5$, having mean zero. Under the assumption of stationary expectation errors and real interest rates, the terms involving expectation operators in Eqs. 6 and 7 can be expressed in terms of observables (Garratt et al. 2006). In this case the expectation errors are subsumed into the long-run equilibrium errors $\boldsymbol{\xi}_t$. The steady state equilibrium conditions then suggest the following set of restrictions on the coefficients of the matrix $\boldsymbol{\beta}$ containing the cointegrating vectors:

$$m_t - \beta_{22} y_t + \beta_{23} i_t = b_{10} + \xi_{1t+1} \tag{13}$$

$$i_t - \Delta p_t = b_{20} + \xi_{2t+1} \tag{14}$$

$$i_t - i_t^* = b_{30} + \xi_{3t+1} \tag{15}$$

$$p_t - p_t^* - e_t = b_{40} + \xi_{4t+1} \tag{16}$$

$$y_t - y_t^* = b_{50} + \xi_{5t+1} \tag{17}$$

Equations 13 through 17 feature either 0 or 1 restrictions, except Eq. 13 that contains two free coefficients, which are combinations of parameters in the period utility function and the discount factor. Our steady-state conditions thus provide more restrictions than necessary to exactly identify $\boldsymbol{\beta}$. The constant $b_{20}$ has the interesting interpretation as an estimate of the natural interest rate in an open economy.

## 3 Cointegrated vector autoregressive model

The cointegrated VAR adds error correction terms and imposes long-run restrictions on an unrestricted VAR in differences. This leaves the estimation of short-term dynamics entirely data-driven, while we use steady-state conditions derived from economic theory as identifying restrictions for the estimation of the cointegrating vectors. Stationary combinations of I(1)-variables can be interpreted as deviations from the long-run equilibrium. Johansen (1988, 1991) developed a maximum likelihood estimator for the following general vector error correction model:

$$\Delta \mathbf{y}_t = \mathbf{a}_0 - \boldsymbol{\Pi}\mathbf{y}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta \mathbf{y}_{t-1} + \mathbf{u}_t, \tag{18}$$

where $\Delta \mathbf{y}_t$ is the $m \times 1$ vector of endogenous variables in first differences and $\mathbf{a}_0$ is an $m \times 1$ vector of constants. The $m \times m$ coefficient matrices $\Gamma_i$ describe the short-term response to past variations in lagged endogenous variables, $p$ is the order of the vector autoregressive process in levels, and $\mathbf{u}_t$ is an $m \times 1$ vector of i.i.d. $(\mathbf{0}, \boldsymbol{\Sigma})$ errors. The matrix $\boldsymbol{\Pi}$ relates $\Delta \mathbf{y}_t$ to past values of $\mathbf{y}_t$ and works as an error correction mechanism if the elements of $\mathbf{y}_t$ are integrated of order one and rank $(\boldsymbol{\Pi}) = r < m$. In this case $\boldsymbol{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$ where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are $m \times r$ matrices of full column rank. The linear cointegrating relations $\xi_t = \boldsymbol{\beta}'\mathbf{y}_{t-1}$ are I(0) and the elements of $\boldsymbol{\alpha}$ define the rate at which the system corrects deviations from the long term equilibrium, i. e. $\xi_t \neq \mathbf{0}$.

Since there are many observationally equivalent factorizations of $\boldsymbol{\Pi}$ into $\boldsymbol{\alpha}$, and $\boldsymbol{\beta}$, we need to impose at least $r$ identifying restrictions on each of the $r$ cointegrating relations to uniquely identify cointegrating vectors. Since $r$ restrictions already result from the normalization conditions, another $r^2 - r$ restrictions will be needed for a unique factorization. Johansen (1988, 1991) uses the statistically motivated restrictions resulting from a solution to the eigenvalue problem to obtain a maximum likelihood estimate of $\boldsymbol{\beta}$. This identification strategy ignores a priori economic information on coefficients and possible parameter restrictions, and renders the interpretation of cointegrating vectors difficult if $r > 1$. This problem has been addressed in Johansen–Juselius (1992). They obtained restrictions from a set of economic relationships such as the purchasing power parity, and developed related likelihood ratio tests. This approach has found its most extensive development in Garratt et al. (1999, 2003, 2006) and Pesaran–Shin (2002). Since cointegrating relations represent fluctuations around long-run equilibriums, the steady-state solutions from theoretical dynamic models provide appropriate restrictions. We impose the steady state equilibrium conditions (13) to (17) on the cointegrating vectors. This approach contrasts with structural VARs based on restricting contemporaneous short-run effects of structural disturbances (e.g., Bernanke 1986; Blanchard–Quah 1989; Gali 1992; Christiano et al. 1999). The main advantage of switching attention from restrictions on short-run effects to long-run cointegrating vectors is the usually broad consensus among economists about the validity of steady-state conditions.

## 4 The data

The steady-state conditions of Sect. 2 do not only suggest a number of restrictions for the decomposition of the reduced rank matrix $\mathbf{\Pi}$, but also define the set of variables for the cointegrated VAR of an open economy. We estimate separate models for the Eurozone and the USA, each containing eight endogenous variables: the domestic per capita money stock relative to real per capita GDP, $m_t$, domestic per capita income, $y_t$, the domestic short term interest rate, $i_t$, domestic inflation, $\Delta p_t$, the foreign short term interest rate, $i_t^*$, the price differential, $\left(p_t - p_t^*\right)$, the Euro per US-Dollar exchange rate, $e_t$, and foreign per capita income, $y_t^*$. Therefore the vector of endogenous variables reads $\mathbf{y}_t = \left(m_t, y_t, i_t, \Delta p_t, i_t^*, \left(p_t - p_t^*\right), e_t, y_t^*\right)$. Additionally, we use the oil price, $poil_t$, as a strictly exogenous variable.

We use the Main Economic Indicators and the Economic Outlook data bases from the OECD and the IMF International Financial Statistics. Most of the variables are transferred into indices with base year 2000 (cf. Appendix). We take logarithms of all variables except interest rates, to which we apply the following transformation: $i_t = \ln(1 + r_t/100)$. Since we already divided the real money stock by real output, Eq. 13 can be interpreted as a description for the inverse of the velocity of money, if $\beta_{22} = 0$.

Cointegrated VARs are based on the assumption that the endogenous variables are integrated of order one, I(1), i.e. the variables are non-stationary in the sense that shocks have permanent effects on their levels (Nelson–Plosser 1982). The unit root property can be removed by taking first differences such that the resulting series is stationary or integrated of order zero, I(0). In the case of cointegration there exist long-run stationary relations between the endogenous I(1) variables. We therefore test in a first step for the unit root properties of our time series by applying three procedures (Pfaff 2006): the augmented Dickey–Fuller Test (ADF), the Phillips–Perron Test (PP) and the Kwiatkowski–Phillips–Schmidt–Shin Test (KPSS). We then test for cointegration between non-stationary variables. Table A1 in the appendix provides strong evidence in favor of a unit root in the levels. The results confirm the unit root hypothesis in the levels, while those for first differences clearly point towards stationary time series, with mixed evidence only for inflation. For the inflation rates the ADF-test does not reject the null of a unit root, whereas the PP-test does. On the other hand, the KPSS-test rejects the null of no unit root in a more plausible setup without a trend. We take this as evidence of a unit root in the inflation series, and include inflation rates rather than price levels into $\mathbf{y}_t$.

## 5 Results

The open economy model presented in Sect. 2 suggests five long-run steady state conditions which can be used to decompose the matrix $\mathbf{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$. Before imposing these conditions on the data we test for the number of cointegrating vectors, i.e. the number long-run equilibriums in our data. We assume the following model structure:

$$\Delta \mathbf{y}_t = \boldsymbol{\alpha}\boldsymbol{\beta}'\mathbf{a}_0 - \boldsymbol{\alpha}\boldsymbol{\beta}' y_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta y_{t-i} + \sum_{i=0}^{p-1} \boldsymbol{\psi}_i \Delta poil_{t-i} + \mathbf{u}_t. \tag{19}$$

The 8 × 1 coefficient vectors $\boldsymbol{\Psi}_i$ show the dynamic response of the system to current and previous oil price shocks. We allow for a restricted constant in the cointegrating equations of model (19) by premultiplying the constant $\mathbf{a}_0$ with $\boldsymbol{\alpha}\boldsymbol{\beta}'$. This assumption is justified by the results of the unit root tests and conforms to stochastic trends in our series.

We obtain the optimal number of lags for the cointegration rank test by comparing Akaike (AIC) and Bayes information criteria (BIC) of unrestricted VAR($p$) models for $p$ = 1,…, 4. Table A2 indicates that up to two lags should be included in both models. The corresponding vector error correction models (VECMs) do not pass misspecification tests due to serial correlation in residuals. Thus, we increase the lag order for cointegration tests. Table 1 shows trace statistics and associated $p$-values. The results indicate the presence of three cointegrating vectors for both economies.

Since the model suggests five steady state conditions but the empirical results only provide evidence for three, we estimate VECMs with all possible combinations of the available potential long run restrictions 13–17 imposed on the 3-dimensional cointegration space. We choose the model for which the set of long run restrictions minimizes the AIC criterion, provided the long-run equilibrium errors are stationary. In both areas, the output gap, the international interest rate parity, and the purchasing power parity fulfill these criteria. The resulting matrix of cointegrating vectors $\boldsymbol{\beta}'$ is

$$\boldsymbol{\beta}' = \begin{pmatrix} 0 & 1 & 0 & 0 & 0 & 0 & 0 & -1 \\ 0 & 0 & 1 & 0 & -1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & -1 & 0 \end{pmatrix}. \tag{20}$$

**Table 1** Results of the trace test on the number of cointegrating relations

| Number of cointegrating relations | | Eurozone | | USA | |
|---|---|---|---|---|---|
| | | Trace statistic | $p$-Value | Trace statistic | $p$-Value |
| $r = 0$ | $r = 1$ | 239.4 | 0.00*** | 223.8 | 0.00*** |
| $r \leq 1$ | $r = 2$ | 165.4 | 0.04** | 162.3 | 0.00*** |
| $r \leq 2$ | $r = 3$ | 103.5 | 0.18 | 105.8 | 0.01*** |
| $r \leq 3$ | $r = 4$ | 65.6 | 0.22 | 62.4 | 0.17 |
| $r \leq 4$ | $r = 5$ | 42.5 | 0.38 | 39.5 | 0.24 |
| $r \leq 5$ | $r = 6$ | 23.7 | 0.35 | 22.4 | 0.27 |

*Note*. Results from cointegration tests according to Johansen (1995). $p$-Values for trace statistics are based on MacKinnon et al. (1999). The order of the underlying VAR($p$) model is 2 for the Eurozone and 3 for the USA. The model allows for a restricted constant and includes oil prices as exogenous variables. ** Indicates significance at the 5% level, *** Indicates significance at the 1% level

This means that we impose 24 restrictions on the matrix **β**. Since an exact identification requires only 9, the remaining 15 over-identifying restrictions can be tested. Garratt et al. (2006) point out that the related likelihood ratio test does not have good properties in small samples, as asymptotic critical values are substantially biased. We therefore apply the non-parametric bootstrap with resampling suggested by Garratt et al. (2006). Specifically, we compute 3,000 replications of the cointegrated VAR model subject to the over-identifying restrictions and compare it to the ones obtained from an appropriately chosen exactly identified specification.[1] The resulting set of log-likelihood statistics was used to compute the upper 5%-critical value of 51.2. The likelihood ratio test for the Eurozone delivers a test statistic of 17.1, so we cannot reject the 15 over-identifying restrictions in the one sided test at the 5% level. For the US model, the likelihood ratio test statistic is 46.4, which also lies below the 5% critical value (51.2).

Figure 1 presents the resulting three long-run equilibrium errors as percentage deviations from steady state after subtracting constants. The purchasing power parity errors show the largest fluctuations. A positive deviation from equilibrium can be interpreted as an overvalued Euro, whereas a negative deviation indicates an overvalued US-Dollar. The most pronounced periods of US-Dollar overvaluation occurred around 1970, 1985, and 2001. The trough of 1985 clearly reflects the sharp turnaround after signing the Plaza accord. The succeeding steep correction period ended in the mid of 1987 coinciding with the Louvre accord. Interestingly, substantial overvaluations of the Euro were less marked and often interrupted. As in Johansen–Juselius (1992), the purchasing power parity relationship passes a multivariate stationarity test, which is often rejected in univariate setups.

The other two equilibrium errors show smaller deviations from the steady state in the range of a few percentage points. In case of the international interest rate parity, a positive error indicates higher rates in the Eurozone. There is only one remarkable period of positive errors coinciding with the period after the German unification. Around the beginning of 1993 interest rates in the Eurozone were markedly higher as compared to the USA. The output gap equilibrium error describes the deviation from the average ratio of Eurozone to US output; it is positive if the Eurozone per-capita output exceeds this long-run average. The mean adjusted output gap in Fig. 1 was fairly close to zero in the beginning of the sample. In the wake of the Volcker disinflation policy starting in late 1979 (Romer–Romer 1989), Eurozone per-capita output gained some edge over the US counterpart peaking at the end of the disinflation policy in 1983 (Goodfriend–King 2005). The peak at the beginning of 1992 which was followed by a more or less steady lead in US growth until the second half of 2006 is also interesting. This period corresponds surprisingly well to the US productivity resurgence commencing around the mid of the 1990s and lasting until recently (cf. Oliner–Sichel 2002; Gordon 2003; Jorgenson–Stiroh 2000).

[1] See Johnston–DiNardo (1997) for a description of the bootstrapping procedure. For the exactly identified model we lift the zero restrictions on $\beta_{11}$, $\beta_{13}$, $\beta_{15}$, $\beta_{16}$, $\beta_{18}$, $\beta_{21}$, $\beta_{22}$, $\beta_{23}$, $\beta_{25}$, $\beta_{26}$, $\beta_{31}$, $\beta_{32}$, $\beta_{33}$, $\beta_{34}$, and $\beta_{36}$ in Eq. 20.

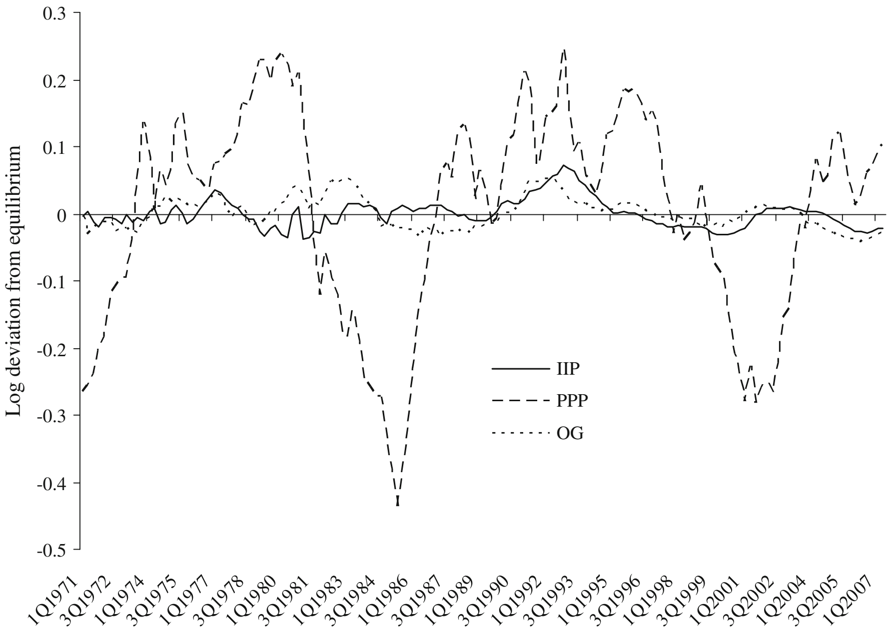


**Fig. 1** Long-run equilibrium errors from structural vector error correction model of the Eurozone, 1970Q4 through 2007Q2. *Note*. IIP represents international interest parity, PPP purchasing power parity, and OG is the output gap between the Eurozone and the USA

Before presenting the dynamic features of the two models, we report the fit for the reduced form systems. We provide the adjusted $R^2$ as a measure for the goodness of fit, the Jarque-Bera test on normality of residuals, the White test on heteroscedasticity, and the Portmanteau test on serial correlation in the residuals at lag 4 in Table 2. There is some evidence of heteroscedasticity in the interest rate equation which is due to higher volatility during the 1970s and the early 1980s as compared to the end of the sample. Given these confirmative results we continue with a specification using two lags in the differenced model (19) for the Eurozone and the USA but rely on bootstrapping techniques to determine confidence intervals for the following impulse response functions.

### 5.1 Generalized impulse response functions

The cointegrated VARs can be used to study how shocks in one economic area affect the overall economic performance in the other area. We carry out this analysis by computing generalized impulse response functions (GIRFs) as developed by Koop et al. (1996) and refined in Pesaran–Shin (1998). The system is directly shocked by one standard deviation of the estimation error from the equation of interest, taking account of contemporaneous correlation among errors in the computation of the system's response. This procedure does not endow $u_{it}$ with a direct economic interpretation. It shows the response of the model to a unit change in an endogenous variable, whatever the reason for this change may be.

**Table 2** Equation fit and misspecification tests

| | $DDp_{ez}$ | $De$ | $Dm_{ez}$ | $D(p_{ez} - p_{us})$ | $Di_{ez}$ | $Di_{us}$ | $Dy_{ez}$ | $Dy_{us}$ |
|---|---|---|---|---|---|---|---|---|
| Eurozone | | | | | | | | |
| Adjusted $R^2$ | 0.38 | 0.14 | 0.21 | 0.49 | 0.46 | 0.33 | 0.27 | 0.25 |
| Jarque–Bera | 0.41 | **0.01** | 0.05 | 0.45 | 0.21 | **0.00** | 0.09 | 0.07 |
| White test | 0.69 | 0.27 | 0.85 | 0.53 | **0.00** | **0.00** | 0.68 | 0.54 |
| Portmanteau test | 0.27 | 0.99 | 0.06 | 0.63 | 0.48 | 0.12 | 0.76 | 0.41 |
| | $DDp_{us}$ | $De$ | $Dm_{us}$ | $D(p_{ez} - p_{us})$ | $Di_{ez}$ | $Di_{us}$ | $Dy_{ez}$ | $Dy_{us}$ |
| USA | | | | | | | | |
| Adjusted $R^2$ | 0.55 | 0.15 | 0.48 | 0.51 | 0.48 | 0.34 | 0.27 | 0.23 |
| Jarque–Bera | 0.14 | **0.01** | **0.05** | 0.32 | 0.65 | **0.01** | 0.10 | 0.13 |
| White test | 0.76 | 0.72 | 0.75 | 0.83 | **0.04** | **0.00** | 0.88 | 0.58 |
| Portmanteau test | 0.51 | 1.00 | 0.65 | 0.88 | 0.44 | 0.20 | 0.76 | 0.31 |

*Note*. Structural vector error correction model for each area with restricted constant and oil prices as exogenous variables. Numbers in rows of test statistics are $p$-values, bold values indicate significant rejection of the null hypothesis. Null hypothesis of Jarque–Bera test is normally distributed of residuals, null hypothesis of White Test is homoscedastic residuals, null hypothesis of Portmanteau test at lag 12 is no serial correlation

Christiano et al. (1999) provide a standard reference for impulse responses showing the transmission of monetary shocks in the US economy. For the Eurozone, Vlaar (2004) shows the dynamic response in a cointegrated VAR based on a closed economy setting. Given that our definition of variables differs slightly from both papers, we can only compare interest rate and output shocks. In the short-run, our US model replicates the shape and size of comparable home interest rate shocks quite closely. The shape and size of GIRFs in the Eurozone model match quite well for interest rate shocks. Our results for output shocks are similar to the effect of demand shocks in Vlaar et al. (2004). The results are very different in the long-run, when the error correction mechanisms with respect to international equilibrium conditions unfold. In the following we concentrate our presentation on three shocks; each related to one of the three equilibrium conditions. In addition to the point estimates for the GIRFs we provide lower and upper bounds for the 95% confidence interval using a non-parametric bootstrap method based on 2,000 replications. We correct for possible bias in the estimate of the impulse response function by using Hall's (1992) percentile method for the computation of confidence intervals as proposed by Benkwitz et al. (2001). A simulation horizon of 40 quarters is sufficiently long to illustrate the working of the long-run error correction mechanism.

The interaction between the Eurozone and the USA can best be analyzed by comparing the response to shocks originating in the other area. First we introduce a shock to the foreign output equation in each area and plot GIRFs for the inflation rate and both per-capita output levels. Figure 2a plots the GIRF for a positive 1% point shock to US output in the Eurozone model. Following a 1% level shift in the first quarter, US output reaches a peak in the third quarter. Afterwards output converges slowly towards the new steady-state level, which is 0.8% above the

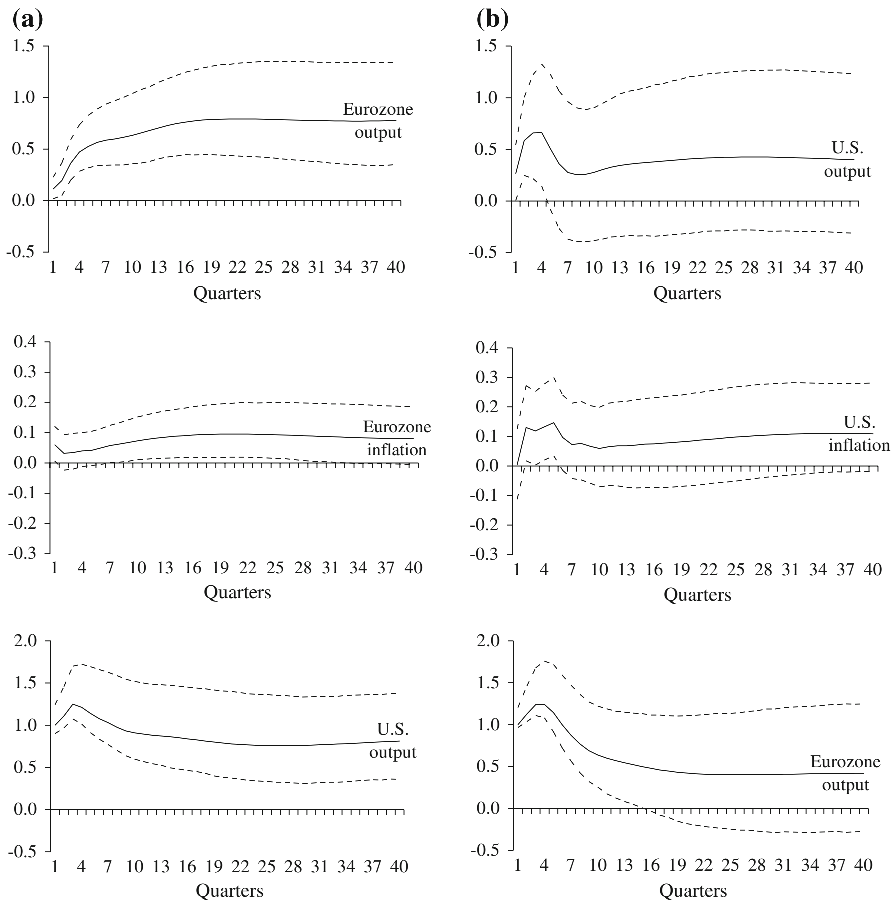


**Fig. 2** (**a**) Response to a positive 1% shock to US output. (**b**) Response to a positive 1% shock to Eurozone output. *Note*. Generalized impulse response functions according to Koop et al. (1996). (**a**) Shows the response of the Eurozone model and (**b**) shows the response of the US model. Dotted lines plot upper and lower bounds of 95% confidence intervals, which are based on a non-parametric bootstrap using 2,000 replications

original level. The 95% confidence interval clearly indicates a significant non-zero response. The Eurozone, on the other hand, starts out slowly and takes more than 4 years to approach the new steady-state level. The output gap restriction on the cointegration vector enforces the same long-term response for both areas. After a short phasing-in period the Eurozone inflation rate increases permanently by about 0.05% points. Although this value is comparatively small, it is significant.

We now turn to the US model's response to output shocks in the Eurozone. The bottom panel of Fig. 2b shows that the Eurozone's output response to its own shock has a similar hump-shaped pattern, although the persistence is lower. Interestingly, US output responds quicker to a shock in EU-output but decreases considerably after the fourth quarter, converging to the new steady state from below. This pattern points to an asymmetry of output shocks in both areas. Shocks originating in the

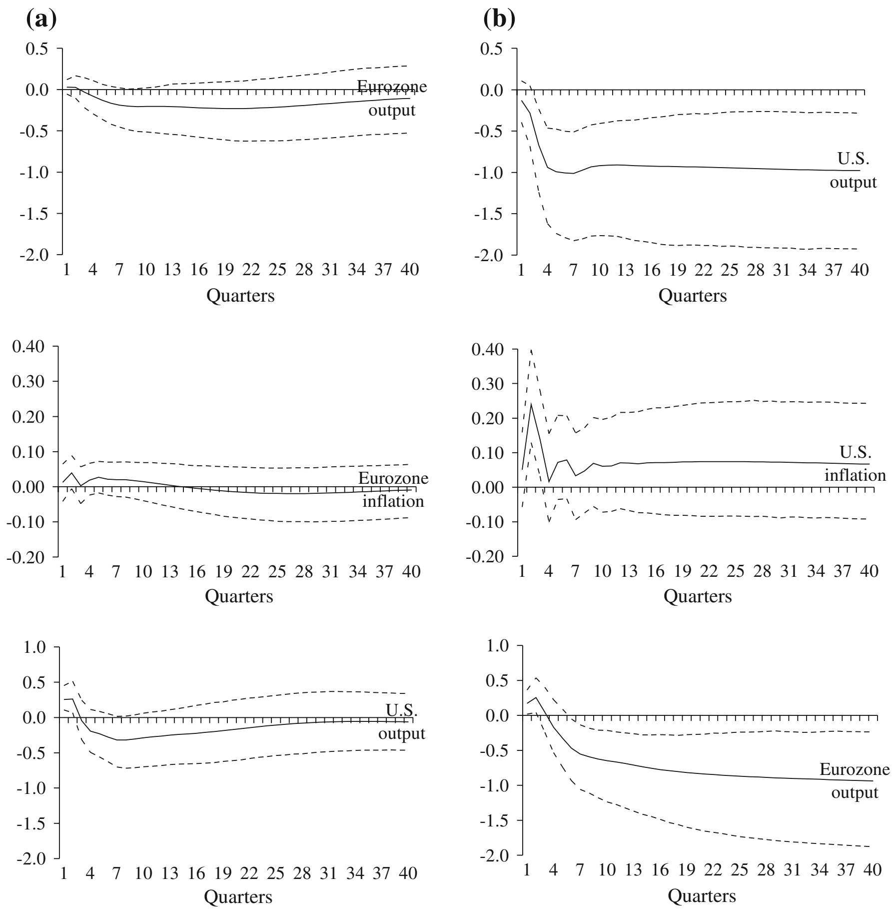


**Fig. 3** (**a**) Response to a positive 1% point shock to US interest rate. (**b**) Response to a positive 1% point shock to Eurozone interest rate. *Note*. Generalized impulse response functions according to Koop et al. (1996). (**a**) Shows the response of the Eurozone model and (**b**) shows the response of the US model. Dotted lines plot upper and lower bounds of 95% confidence intervals, which are based on a non-parametric bootstrap using 2,000 replications

USA tend to have a larger impact on output in both economies than shocks originating in the Eurozone. The US inflation rate rises significantly by 0.1% points.

We next shock the foreign interest rate in each model by 1% point and plot the response of outputs and inflation rates in Fig. 3a (Eurozone model) and 3b (US model). In the Eurozone model, the response of US output to a US interest rate shock is muted, negative and converges slowly to the original path. Such a pattern is similar to impulse responses from the Global VAR model by Dees et al. (2007). This is mainly due to a quick correction of the US interest rate hike in the second quarter caused by a large negative AR(2) coefficient. The Eurozone output decreases only slightly, while the inflation rate remains almost constant. The response of US output to interest rate changes in the Eurozone in Fig. 3b is surprisingly large and reaches its new steady state level after four quarters. The Eurozone itself shows a sluggish response but keeps on a downward path over the full forecast horizon. An interest rate shock in the Eurozone is therefore more persistent. We attribute the higher persistence in the Eurozone to the fact that the pre-1999 Eurozone interest rate is a weighted average of individual country rates that includes spreads over the German short term interest rate. Giavazzi–Giovannini (1991) show the high sensitivity of these spreads to concerns about the credibility of exchange rate targets, and the sluggish adjustment to their fundamental values. The US inflation rate shows a minor positive response owing to the depreciation of the US-Dollar created by the interest rate differential of more than half a percentage point in the first 3 years after the shock.

Finally, Fig. 4a suggests that neither output nor inflation in the Eurozone responds significantly to the depreciation of the Euro. The exchange rate shock is transient because the purchasing power parity condition works as an error correction mechanism and inflation rates in both areas hardly react. Contrary to our expectations, the response of US output to the appreciation is permanently positive. This counterintuitive result is due to a sharp reduction in short-term US interest rates associated with the exchange rate shock.

## 6 Conclusions

We estimate cointegrated vector autoregressive models for the Eurozone and the USA by imposing long-run stead-state conditions consistent with a dynamic open economy model, but we do not impose any short-term identifying restrictions. The theoretical model describes the interaction between large open economies with free trade, no restrictions to capital transactions, and at least partial international diffusion of technology. The estimated models include eight endogenous variables: domestic and foreign interest rates, the money stock, inflation rates, the price differential, the Euro–Dollar exchange rate, and domestic and foreign real output. A test for overidentifying restrictions on the cointegrating vectors cannot reject the presence of three long-run equilibrium relationships. These are the relative purchasing power parity, the international interest parity, and an international output gap relation. The data support international steady-state conditions rather than domestic conditions like the Fisher parity or a stable money demand function. The time pattern of the resulting equilibrium errors matches to well-known economic policy episodes.

The preferred specification produces non-trivial dynamic medium-term adjustment patterns in response to unexpected variations in one of the endogenous variables. For comparable shocks and in the short-run, both models closely resemble the dynamic response in existing studies based on a closed economy setup. The long-run response of our cointegrated VAR, in contrast, is driven by adjustment towards long-run international equilibrium relations, thus departing from traditional findings.

Models of endogenous growth that link business cycle variations with the long-term development of an economy typically assume closed economies and derive

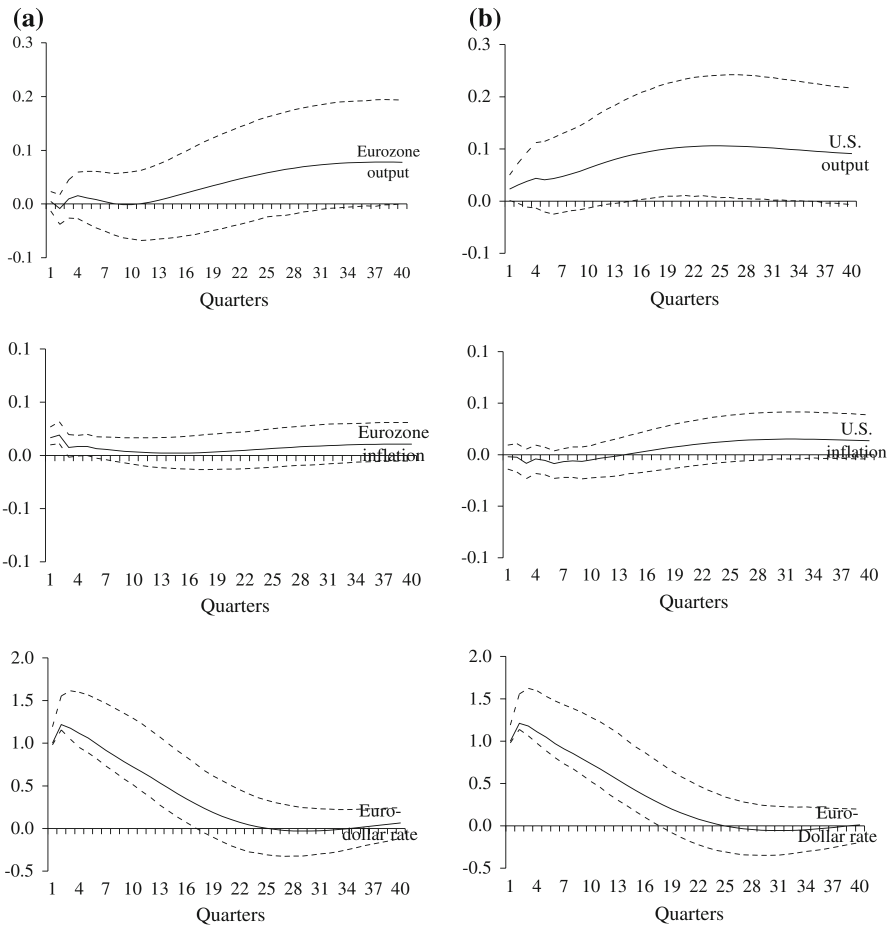


**Fig. 4** (**a**) Response to a positive 1% shock to the Euro–Dollar exchange rate. (**b**) Response to a positive 1% shock to the Euro–Dollar exchange rate. *Note*. Generalized impulse response functions according to Koop et al. (1996). (**a**) Shows the response of the Eurozone model and (**b**) shows the response of the US model. Dotted lines plot upper and lower bounds of 95 % confidence intervals, which are based on a non-parametric bootstrap using 2,000 replications

their feedback mechanism from national stock-flow relations, cf. Comin (2008) for a model with feedback between business cycles and R&D activity. The models by Comin–Gertler (2006) and Comin (2008) introduce time varying parameters for the number of firms and the markup over marginal costs in the final goods sector into the reduced-form equation for net value added. These models imply that the autocovariance structure for aggregate total factor productivity—and consequently also net value added—depends not only on the length of time separating a pair of observations but also their date. In this case, the rate of change in real output (net value added) would be non-stationary and the roots of the lag-polynomial of an AR(p) process fitted to these data would lie on or inside the unit circle. We clearly reject unit roots in the differenced output series of both areas and thus find no evidence for prolonged time-dependent shifts in output growth. In our model, high persistence arises due to three cointegrating relations that capture international spillovers from output, interest rate, and exchange rate fluctuations. These fluctuations affect output levels in the long-run, but cause medium-term shifts in the growth rate only.

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## Appendix: the data

The following series were obtained from the Main Economic Indicators and the Economic Outlook data bases of the OECD or the IMF International Financial Statistics

- *e* natural logarithm of the normalized nominal Euro per US-Dollar exchange rate (base: first quarter 2000 = 1).
- *hez* natural logarithm of the normalized Eurozone M1 real per capita money stock in relation to real per capita GDP (base: first quarter 2000 = 1).
- *hus* natural logarithm of the normalized US M1 real per capita money stock in relation to real per capita GDP (base: first quarter 2000 = 1).
- *pd* price differential measured as *pez-pus* (see below).
- *pez* natural logarithm of the Eurozone consumer price index (base: first quarter 2000 = 1).
- *poil* natural logarithm of import price for crude oil in US-Dollar.
- *pus* natural logarithm of the US consumer price index (base: first quarter of 2000 = 1).
- *rez* natural logarithm of $(1 + r_{ez}/100)$, where $r_{ez}$ is the annualized average 3 month interest rate in the Eurozone.
- *rus* natural logarithm of $(1 + r_{us}/100)$, where $r_{us}$ is the annualized average 3 month interest rate in the USA.
- *yez* natural logarithm of the normalized real per capita GDP in the Eurozone (base: first quarter of 2000 = 1).
- *yus* natural logarithm of the normalized real per capita GDP in the USA (base: first quarter of 2000 = 1).

We define the Eurozone as a twelve countries' aggregate with Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal and Spain. Slovenia joined the monetary union in 2007 and is left out of our Eurozone aggregate. In case that no aggregate series for the Eurozone was available, a weighted series was obtained out of individual data for the 12 member countries. We use the share of individual countries in the Eurozone aggregate GDP during the respective quarters as weights. Annual population data have been interpolated with Ecotrim using the Boot et al. (1967) method. Results for tests on unit root characteristics of the time series and on the VAR lag length are given in Tables A1 and A2, respectively.

**Table A1** Augmented Dickey–Fuller (ADF), Phillips–Perron (PP), and Kwiatkowski–Phillips–Schmid–Shin (KPSS) unit root tests, 1970Q1 through 2007Q2

Test statistics based on model including

| | ADF | | PP | | KPSS | |
|---|---|---|---|---|---|---|
| | Const | Const + Trend | Const | Const + Trend | Const | Const + Trend |
| For levels | | | | | | |
| $e$ | 0.12 | 0.33 | 0.12 | 0.34 | 0.07 | 0.06 |
| $m_{ez}$ | 0.49 | 1.00 | 0.55 | 1.00 | 0.82*** | 0.33*** |
| $m_{us}$ | 0.87 | 0.25 | 0.85 | 0.65 | 1.26*** | 0.14* |
| $(p_{ez} - p_{us})$ | 0.01*** | 0.03** | 0.19 | 0.45 | 0.13 | 0.13* |
| $p_{ez}$ | 0.02** | 0.71 | 0.00*** | 0.99 | 1.39*** | 0.36*** |
| $p_{oil}$ | 0.12 | 0.25 | 0.17 | 0.39 | 0.76*** | 0.21** |
| $p_{us}$ | 0.06* | 0.72 | 0.01*** | 0.96 | 1.39*** | 0.35*** |
| $i_{ez}$ | 0.49 | 0.33 | 0.55 | 0.46 | 0.82*** | 0.28*** |
| $i_{us}$ | 0.39 | 0.02** | 0.21 | 0.23 | 0.75*** | 0.14* |
| $y_{ez}$ | 0.53 | 0.31 | 0.32 | 0.12 | 1.47*** | 0.15* |
| $y_{us}$ | 0.87 | 0.02** | 0.88 | 0.10* | 1.46*** | 0.04 |
| For first differences | | | | | | |
| $\Delta e$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.07 | 0.06 |
| $\Delta m_{ez}$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.08 | 0.08 |
| $\Delta m_{us}$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.08 | 0.08 |
| $\Delta(p_{ez} - p_{us})$ | 0.02** | 0.06* | 0.00*** | 0.00*** | 0.12 | 0.04 |
| $\Delta p_{ez}$ | 0.62 | 0.15 | 0.15 | 0.00*** | 1.07*** | 0.10 |
| $\Delta p_{oil}$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.19 | 0.13* |
| $\Delta p_{us}$ | 0.13 | 0.05* | 0.00*** | 0.00*** | 0.82*** | 0.09 |
| $\Delta i_{ez}$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.08 | 0.05 |
| $\Delta i_{us}$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.05 | 0.05 |
| $\Delta y_{ez}$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.25 | 0.06 |
| $\Delta y_{us}$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.03 | 0.03 |
| For second differences | | | | | | |
| $\Delta\Delta e$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.07 | 0.06 |
| $\Delta\Delta m_{ez}$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.05 | 0.03 |
| $\Delta\Delta m_{us}$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.03 | 0.03 |
| $\Delta\Delta(p_{ez} - p_{us})$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.03 | 0.03 |
| $\Delta\Delta p_{ez}$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.05 | 0.05 |
| $\Delta\Delta p_{oil}$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.03 | 0.03 |
| $\Delta\Delta p_{us}$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.04 | 0.03 |
| $\Delta\Delta i_{ez}$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.05 | 0.05 |
| $\Delta\Delta i_{us}$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.20 | 0.20** |
| $\Delta\Delta y_{ez}$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.10 | 0.05 |

**Table A1** continued

Test statistics based on model including

| | ADF | | PP | | KPSS | |
|---|---|---|---|---|---|---|
| | Const | Const + Trend | Const | Const + Trend | Const | Const + Trend |
| $\Delta\Delta y_{us}$ | 0.00*** | 0.00*** | 0.00*** | 0.00*** | 0.07 | 0.06 |

*Note*. Values for the augmented Dickey–Fuller and the Phillips–Perron tests are $p$-values for the null hypothesis of a unit root in the series. Values for the KPSS test are the test statistics for the null hypothesis of no unit root in the series. The associated critical values for the model including a constant are 0.739 (1%), 0.463 (5%), and 0.347 (10%). The critical values for the model including a constant and a trend are 0.216 (1%), 0.146 (5%), and 0.119 (10%). The number of lags in the augmented Dickey–Fuller tests is chosen according to the Schwarz information criterion. The window length for Phillips–Perron test is based on Newey–West (1994) using a Bartlett kernel. * Indicates significance at the 10%-level, ** Indicates significance at the 5%-level, and *** Indicates significance at the 1%-level

**Table A2** Tests on lag length in VAR for levels

| Number of lags | Eurozone | | USA | |
|---|---|---|---|---|
| $p$ | AIC | BIC | AIC | BIC |
| 1 | −56.88 | **−55.10** | −56.87 | **−55.09** |
| 2 | **−57.44** | −54.18 | −57.65 | −54.39 |
| 3 | −57.42 | −52.68 | **−57.89** | −53.15 |
| 4 | −57.19 | −50.95 | −57.69 | −51.45 |

*Note*. Akaike information criterion (AIC) and Bayes information criterion (BIC) for unrestricted VAR($p$)-models estimated over the period 1970Q1 through 2007Q2. Bold numbers indicate minimum values. The VAR($p$)-model has an unrestricted constant

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