The Short and Long-run Interdependencies Between the Eurozone and the U.S.A.

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Abstract: We estimate a quarterly structural cointegrating vector autoregression for the Eurozone and the U.S.A. based on long-run restrictions derived from a dynamic open economy model. Three long-run relations between the Eurozone and the U.S.A. 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, although growth rates will react over the medium-term.

Keywords: Structural vector error correction model, steady state, business cycle.

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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 analysis presented in this volume casts doubts on such a view. We address the interaction between long- and short-run fluctuations in an empirical model that explicitly relates business cycle variations to deviations from long-run equilibrium relations. In this model shocks to an economy push it away from the steady state, subsequently an adjustment process starts, which drives the system towards a new steady state trajectory.

We implement such an adjustment process in a Structural Vector Error Correction Model (SVECM) as suggested by Garratt et al. (2006). The autoregressive part models the short-term adjustment process. The cointegration part determines how far and for how long 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 assumption of orthogonality between cointegrating vectors, as in Johansen (1988, 1991). Instead, we derive long-run steady state restrictions from economic theory.

We define two SVECMs using linearized versions of the steady state equations of a dynamic open economy model: one for the Eurozone and another one for the U.S.A. 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 cycle. In each of the two models we combine the contemporary interaction between the Eurozone and the U.S.A. with an intertemporal, dynamic interaction between long-run cointegrating relations and short-run variations.

Most of the previous work on the relation between the U.S.A. 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. Dalsgaard et al. (2001) for example study a change in U.S. fiscal expenditures by one percent of the GDP using the OECD Interlink model. The resulting interaction between the two areas is asymmetric, in the sense that U.S. fiscal shocks have a higher impact on Eurozone output as compared to the effect of a reciprocal Eurozone-fiscal shock on U.S. output. Nevertheless, most of the work on the U.S.A. and the Eurozone treats both areas as large closed economies and therefore ignores international spill-overs; cf. Christiano et al. (1999) for the U.S.A. or Vlaar (2004) for the Eurozone.

In the next section we state the steady-state conditions of an open economy model. The model provides a set of endogenous variables for a SVECM 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 Vector Autoregression (VAR) in levels and test for the number of cointegrating relations among our endogenous variables. We then discuss the results of the estimated SVECM, describe the three long-run equilibrium relations that we identify, test for overidentifying restrictions implied by the open economy model and present generalized impulse response functions that show the dynamic response of the model to unexpected variations in output, the interest rate, and the exchange rate.

Steady-State Conditions

In this section we derive steady-state relations from a dynamic open economy model. The starting point for the analysis is a representative household's optimization problem. An infinitely lived representative household seeks to maximize the expected utility function,

$$E_0 \sum_{t=0}^{\infty} \beta^t U\left(C_t, \frac{M_t}{P_t}, L_t \right), \tag{1}$$

subject to the budget constraint (or current account balance),

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

by choosing infinite sequences of optimal consumption, $\{C_t\}_{t=0}^{\infty}$, nominal money holdings in home currency, $\{M_t\}_{t=0}^{\infty}$, labour 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, $\{B_t^*\}_{t=0}^{\infty}$. 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, δ, and the discount factor, β, 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 equation (2) represents total period spending (in consumption units), whereas the right hand side of equation (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:

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

where $x_t \equiv (C_t, M_t, L_t, K_{t+1}, B_t, B_t^*)$ 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 parameterisation for the period utility function:

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

That is, households have additively separable CRRA preferences for consumption, real balances, and leisure. The first order conditions corresponding to the maximization problem on the right hand side of equation (3) together with the functional assumption in equation (4) can be used to derive the uncovered interest parity, the Fisher inflation parity, and a money demand equation. The log-linearized versions of these equilibrium relations (as local approximations around their steady state values $(\bar{c}, \bar{i}, \bar{i}^*, \overline{m}_r, \bar{p}, \bar{r}, \overline{w}_r, \bar{\pi},)$ are:

$$(\hat{m}_t - \hat{p}_t) = \frac{\sigma_1}{\sigma_2} \hat{c}_t - \frac{\beta}{\sigma_2} \hat{i}_t \tag{5}$$

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

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

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 (MME), (6) represents the Fisher inflation parity (FIP) and (7) represents the uncovered interest rate parity.

Directly from the first order conditions for the optimal holdings of home and foreign bonds it follows that:

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

Since the domestic and foreign country are completely symmetric with respect to their tastes and expectation formation we will have that in equilibrium $C_t = C^*_t$ holds at any point in time and hence (8) can only be true if

$$P_t = e_t P^*_t, \tag{9}$$

i. e. the purchasing power parity (PPP) condition must hold.

We focused on the consumer problem so far. Firms are assumed to be perfectly competitive as in Garratt et al. (2006) and produce real aggregate output, Y_t, according to the constant returns to scale production function (10) using labor, L_t, and capital, K_t, as inputs:

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

Where we assume the $F(.)$ 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). Incomplete adjustment can be introduced 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, $y_t = \ln(Y_t/L_t)$. Under these assumptions, the steady state output gap

(OG) 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 regulation may result in non-identical production technologies. Garratt et al. (2006) provide a more generalized version of equation (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 an open economy converges in the long-run. To show how these theoretical steady state conditions can be used to derive restrictions for the cointegrating vectors of a SVECM 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. $\boldsymbol{\beta}'\mathbf{y}_{t-1}=\boldsymbol{\xi}_t$, and equilibrium errors, $\xi_{i,t}$, i=1,2,....5 have zero means. Under the assumption of stationary expectation errors and real interest rates the terms involving expectation operators in equations (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:

$$\begin{aligned} m_t - \beta_{22} y_t + \beta_{23} i_t &= b_{10} + \xi_{1t+1} && (13)\\ i_t - \Delta p_t &= b_{20} + \xi_{2t+1} && (14)\\ i_t - i_t^* &= b_{30} + \xi_{3t+1} \quad . && (15)\\ p_t - p_t^* - e_t &= b_{40} + \xi_{4,t+1} && (16)\\ y_t - y_t^* &= b_{50} + \xi_{5t+1} && (17) \end{aligned}$$

$$(18)$$

As can be seen from equations (13) through (18), most of the variables have either zero or one restrictions on their coefficients. Only the first equation features two free coefficients that must be estimated. These coefficients are combinations of parameters in the period utility function and the discount factor. The constant b_{20} has the interesting interpretation as an estimate of the natural interest rate in an open economy. Consequently ξ_{3t} is a measure of the stance of monetary policy or disequilibrium due to international capital movements.

Structural Vector Error Correction Model

The SVECM extends a conventional reduced form vector autoregressive models by adding error correction terms and imposing long-run restrictions on the cointegrating vectors. This makes the estimation of short-term dynamics entirely driven by data, whereas the steady-state conditions derived from economic theory provide the background for restricted estimates of the dynamic response to past deviations from the steady state. Stationary combinations of I(1)-variables can be interpreted as deviations from the long-run equilibrium. Johansen (1988, 1991) developed maximum likelihood estimators for the following general reduced form VEC:

$$\Delta\mathbf{y}_t = \mathbf{a}_0 - \mathbf{\Pi}\mathbf{y}_{t-1} + \sum_{i=1}^{p-1}\mathbf{\Gamma}_i\Delta\mathbf{y}_{t-1} + \mathbf{u}_t\,, \tag{19}$$

where $\Delta\mathbf{y}_t$ is the $m\times1$ vector of endogenous variables in first differences and $\mathbf{a}_0$ is an $m\times1$ vector of constants. The $m\times m$ coefficient matrices $\mathbf{\Gamma}_i$ 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 $m\times1$ vector of i.i.d. $(\mathbf{0}, \mathbf{\Sigma})$ errors. The matrix $\mathbf{\Pi}$ 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($\mathbf{\Pi}$)=$r<m$. In this case $\mathbf{\Pi}=\boldsymbol{\alpha}\boldsymbol{\beta}'$ where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are $m\times r$ matrices of full column rank. The linear cointegrating relations $\boldsymbol{\xi}_t=\boldsymbol{\beta}'\mathbf{y}_{t-1}$ 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. $\boldsymbol{\xi}_t\neq\mathbf{0}$.

Since there are many observationally equivalent factorizations of $\mathbf{\Pi}$ 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 normalisation conditions, another r^2-r restrictions will be needed for a unique factorization. Johansen (1988, 1991) uses the statistically motivated assumption that the columns of $\boldsymbol{\beta}$ form an orthogonal set but this identification strategy lacks economic meaning and renders the interpretation of cointegrating vectors difficult if r>1. Garratt et al. (1999, 2003, 2006) and Pesaran – Shin (2002) suggest to rely on economic theory and possibly also on other a priori information to obtain the necessary restrictions. Since cointegrating relations represent fluctuations around long-run equilibriums, the steady-state solutions from theoretical dynamic models are a promising starting point. We impose the steady state equilibrium conditions (13) to (18) 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.

The Data

The steady-state conditions of Section 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 SVECM of an open economy. We set up models for the Eurozone and the U.S.A., each containing eight endogenous variables: the domestic per capita money stock relative to per capita output, m_t, domestic per capita income, y_t, the domestic short term interest rate, i_t, domestic inflation, Δp_t, the foreign short term interest rate, i_t^*, the price differential, $\left(p_t - p_t^*\right)$, the Euro per U.S.-Dollar exchange rate, e_t, and foreign per capita income, y_t^*. The vector of endogenous variables is then given by $\mathbf{y}_t = \left(m_t, y_t, i_t, \Delta p_t, i_t^*, \left(p_t - p_t^*\right), e_t, y_t^*\right)$. 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 (for details see data appendix). We take logarithms of all variables except interest rates, $i_t=\ln(1+i_t/100)$. We divide the real money stock by real output. This allows equation (13) to be interpreted as a description for the velocity of money, if $\beta_{22}=1$.

SVECMs 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 a shock has a permanent effect on their level (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 almost uniform evidence in favor of a unit root in the levels[1]. Interestingly, the more sensible models for the interest rate, including only a constant, do not reject the null of a unit root. The

[1]) Estimation was carried out using the software package EViews 5.1. The workfile can be obtained from the corresponding author on request.

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 the more plausible set up 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 $\mathbf{y}_t$.

Results

The open economy model presented in Section 2 suggests five long-run steady state conditions which can be used to decompose the matrix $\Pi=\alpha\beta'$. 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:

$$\Delta \mathbf{y}_t = \boldsymbol{\alpha\beta}'\mathbf{a}_0 - \boldsymbol{\alpha\beta}' y_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta y_{t-i} + \sum_{i=0}^{p-1} \boldsymbol{\psi}_i \Delta poil_{t-i} + \mathbf{u}_t \,. \qquad (20)$$

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 (20) by premultiplying the constant $\mathbf{a}_0$ with $\alpha\beta'$. 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 in the appendix indicates that 1 to 2 lags should be included for the Eurozone and the U.S.A. The corresponding VECMs do not pass misspecification tests due to serial correlation in residuals thus we increase the lag order for cointegration tests to p=3. 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)-(18) 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 (OG), the international interest rate parity (IIP), and the purchasing power parity (PPP) fulfill these criteria. The resulting matrix of cointegrating vectors $\boldsymbol{\beta}'$ is

$$\boldsymbol{\beta}' = \begin{pmatrix} 0 & 1 & 0 & 0 & 0 & 0 & 0 & -1 \\ 0 & 0 & 1 & 0 & -1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & -1 & 0 \end{pmatrix}. \tag{21}$$

We impose 24 restrictions on the matrix $\boldsymbol{\beta}$, whereas exact identification requires only 9, and so 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 small sample size properties 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 3000 replications of the VEC model subject to the over-identifying restrictions and compare it to the ones obtained from an appropriately chosen exactly identified specification[2]. The resulting set of log-likelihood statistics was used to compute the upper 5%-critical value of 52.8. The likelihood ratio test for the Eurozone delivers a test statistic of 47.5, so we cannot reject the 15 over-identifying restrictions in the one sided test at the conventional 5% level. For the U.S. the likelihood ratio test statistic is 32.1 which also lies below the 5% critical value (50.8).

Figure 1 presents the resulting three long-run equilibrium errors as percentage deviations from steady state after subtracting constants. The purchasing power parity error shows the biggest fluctuations. A positive deviation from equilibrium can be interpreted as an overvalued Euro, whereas a negative deviation indicates an overvalued U.S.-Dollar. The most pronounced periods of US-Dollar overvaluation occurred around 1970, 1985, and 2001. The trough of 1985 clearly reflects the sharp turn around 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.

The other two equilibrium errors show smaller deviations from the steady state in the range of a few percentage points. In the 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 U.S.A. The output gap equilibrium error describes the deviation from the average ratio of Eurozone to U.S. output; it is positive if the Eurozone per-capita output exceeds this long-run average. The mean adjusted output gap in Figure 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-

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

capita output gained some edge over the U.S. 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 U.S. growth until the second half of 2006 is also interesting. This period corresponds surprisingly well to the U.S. productivity resurgence commencing around the mid of the 1990s and lasting until recently (cf. Oliner – Sichel, 2002; Gordon, 2003; Jorgenson – Stiroh, 2000).

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 (20) for the Eurozone and the U.S.A. but rely on bootstrapping techniques to determine confidence intervals for the following impulse response functions.

Generalized Impulse Response Functions

The SVECMs 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). In contrast to conventional impulse response functions, GIRFs do not need identifying assumptions for the transformation of estimation errors, $\mathbf{u}_t$, to a vector of orthogonal innovations. Instead, the system is directly shocked by one standard deviation of the estimation error from the equation of interest, while recognizing the contemporaneous correlation among errors in the computation of the system's response. This procedure does not allow us to give u_{it} a direct economic interpretation, but we are able to show the response of the model to, for example, a unit change in the interest rate, whatever the reason for this change may be.

Christiano et al. (1999) provide a standard reference for impulse responses showing the transmission of monetary shocks in the U.S. economy. For the Eurozone, Vlaar (2004) shows the dynamic response in a SVECM 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 directly. In the short-run, our U.S. 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. For output shocks, our results 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. Additionally to the point estimates for the GIRFs we provide lower and upper bounds for the 95 percent confidence interval using a non-parametric bootstrap method based on 2000 replications. A forecast horizon for the impulse responses of 40 quarters is sufficiently long to illustrate the working of the long-run error correction mechanism.

The interaction between the Eurozone and the U.S.A. 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 one percentage point shock to U.S. output in the Eurozone model. After a one percent level shift in the first quarter, U.S. output reaches a peak in the third quarter. Afterwards output converges slowly towards the new steady-state level, which is 0.8 percent above the original level. The 95 percent confidence interval clearly indicates a significant non-zero response. The Eurozone, on the other hand, starts out slowly and takes more than four 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 percentage points. Although this value is comparatively small it remains significant.

We now turn to the U.S. model's response to output shocks in the Eurozone. The bottom panel of Figure 2b shows that the Eurozone's output response to its own shock has a similar hump-shaped pattern as the U.S. economy, although, the persistence in the Eurozone is much weaker. Interestingly, U.S. output responds more quickly to a shock in EU-output but decreases considerably after the fourth quarter, converging to the new steady state from below. This pattern provides further evidence for the asymmetry of output shocks in both areas. Shocks originating in the U.S.A. tend to have a larger overall impact on output as compared to shocks originating in the Eurozone. The U.S. inflation rate goes up significantly by 0.1 percentage points.

We next shock the foreign interest rate in each model by one percentage point and plot the response of outputs and inflation rates in Figures 3a (Eurozone model) and 3b (U.S. model). In the Eurozone model, the response of U.S. output to a U.S. interest rate shock is muted, negative and converges slowly to the original path. This is mainly due to a sharp correction of the U.S. 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 U.S. output to interest rate changes in the Eurozone in Figure 3b is surprisingly large and reaches its new steady state level already after four quarters. The Eurozone itself shows a more sluggish response but keeps on a downward path over the full forecast horizon. Compared to the U.S.A. an interest rate shock in the Eurozone is more persistent. We attribute the higher persistence in the Eurozone to the fact that the Eurozone interest rate before 1999 is a weighted average of individual country values that includes spreads over the German short term interest rate. Giavazzi – Giovannini (1991) show how sensitive these spreads reacted to questions of credibility of the exchange rate target. The U.S. inflation rate shows a minor positive response owing to the depreciation of the U.S.-Dollar created by the interest rate differential of more than half a percentage point during the first three years after the shock.

Finally, we present the effect of a one percent increase in the Euro-U.S.-Dollar exchange rate. Figure 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 U.S. output to the appreciation is permanently positive. This counterintuitive result is due to a sharp reduction in short-term U.S. interest rates associated with the exchange rate shock.

Conclusions

We estimate structural vector error correction models (SVECMs) for the Eurozone and the U.S.A. by imposing long-run stead-state conditions consistent with a dynamic open economy model. We impose no short-term restrictions on the model. The open economy model describes the interaction between large open economies with free trade, no restrictions to capital transactions, and at least partial international diffusion technology. The estimated SVECMs include eight endogenous variables: home and foreign interest rates, the money stock, inflation rates, the price differential, the Euro-Dollar exchange rate, and home 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 interest rate parity or a stable money demand function. The development of the resulting equilibrium errors can be matched with well-known economic policy episodes.

Despite having only lag-order two, the preferred SVECM 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 previous papers based on a closed economy setup. The long-run response of our SVECM instead, is clearly driven by adjustment towards long-run international equilibrium relations and thus departs from stylized facts.

Models of endogenous growth that link business cycle variations with the long-term development of an economy typically assume closed economies and derive their feedback mechanism from national stock-flow relations (Comin, 2008). The empirical implication of such models would be a non-stationary rate of change in real output. We clearly reject unit roots in the differenced output series of both areas and thus find no evidence for prolonged shifts in output growth. On the other hand, the three cointegrating relations that we find provide a framework for international spillovers from output, interest rate, and even exchange rate shocks in large open economies. Although these shocks affect output levels over a considerable period of time they cause only medium-term shifts in the growth rate. Our computations show that the Eurozone output needs a period of 16 to 20 quarters to adjust to an output shock originating in the U.S.A., augmenting the medium-term annual growth rate in the Eurozone by 0.2 percent.

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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 U.S.-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 U.S. M1 real money stock per capita 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 U.S.-Dollar.

pus natural logarithm of the U.S. 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 U.S.A.

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 U.S.A. (base: first quarter of 2000 = 1).

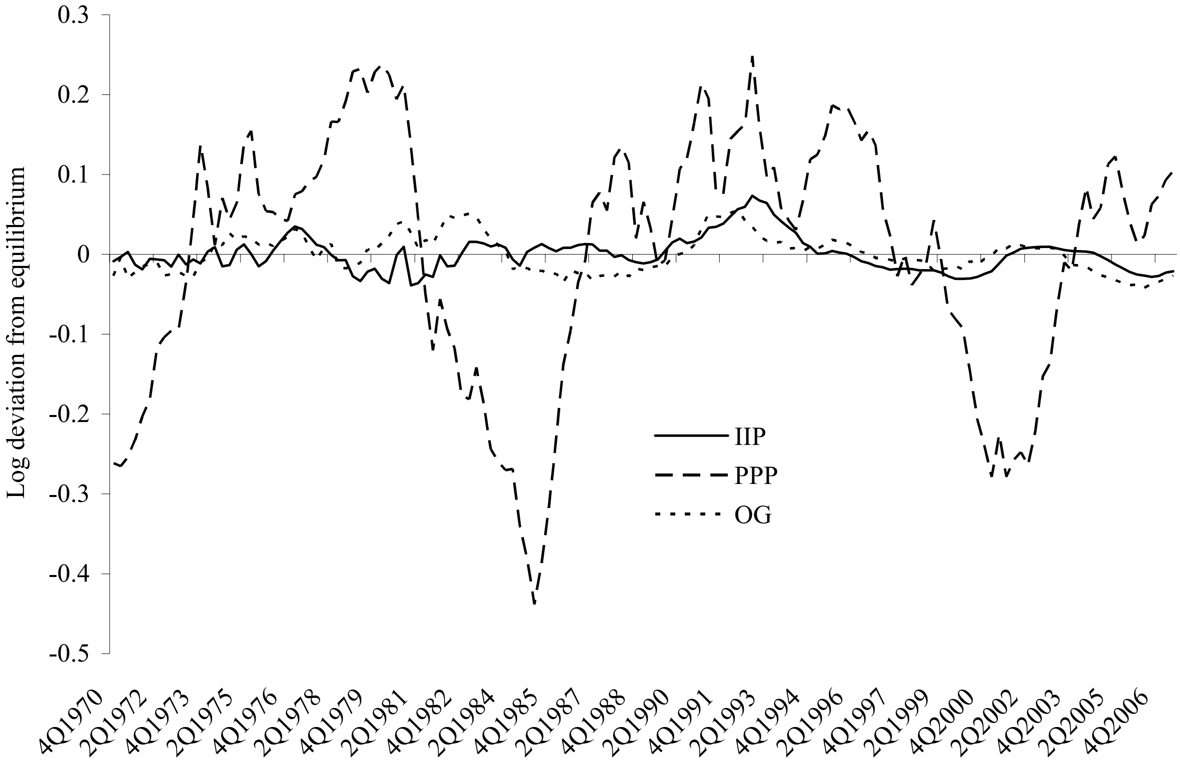
We define the Eurozone as a twelve countries' aggregate with Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal and Spain. Slovenia joined the monetary union in 2007 and is left out of our Eurozone aggregate. In case that no aggregate series for the 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.

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

Number of cointegrating relations		Eurozone trace statistic	Eurozone p-value		U.S.A. trace statistic	U.S.A. p-value	
r=0	r=1	265.2	0.00	***	259.4	0.00	***
r≤1	r=2	160.8	0.00	***	161.9	0.00	***
r≤2	r=3	99.3	0.03	**	104.3	0.01	**
r≤3	r=4	63.3	0.15		60.7	0.22	
r≤4	r=5	41.6	0.17		39.6	0.24	
r≤5	r=6	22.3	0.28		22.3	0.28	

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 U.S.A.. The model allows for a restricted constant and includes oil prices as exogenous variables. ***) indicates significance at the 1-percent level, **) indicates significance at the 5-percent level, * indicates significance at the 10-percent level.

Figure 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 U.S.A.

Figure 2a: Response to a positive 1-percent shock to U.S. output

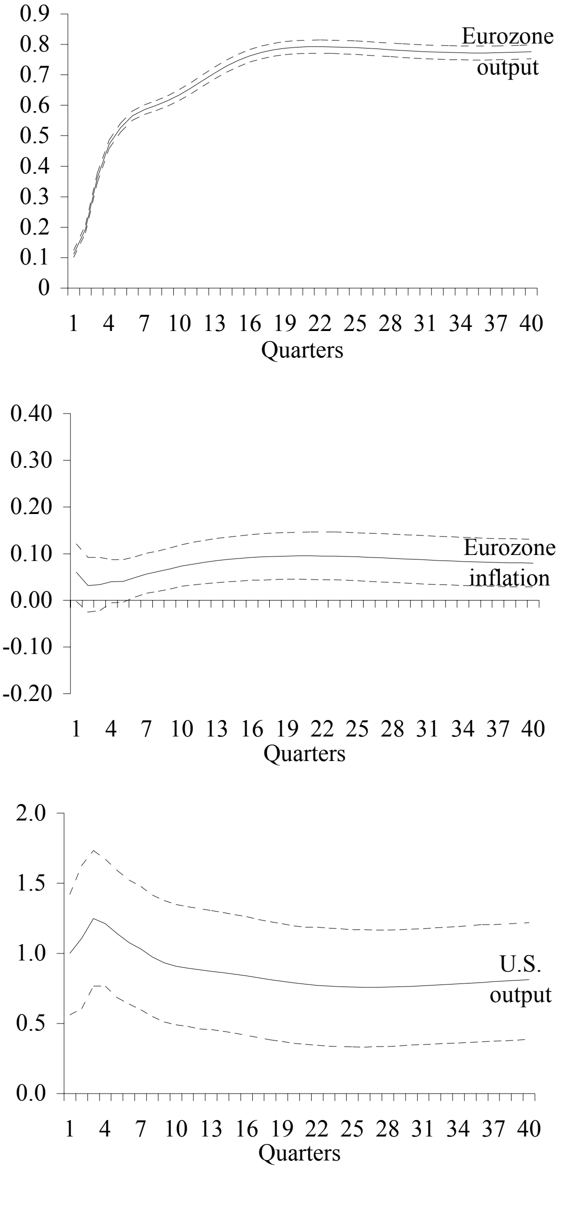
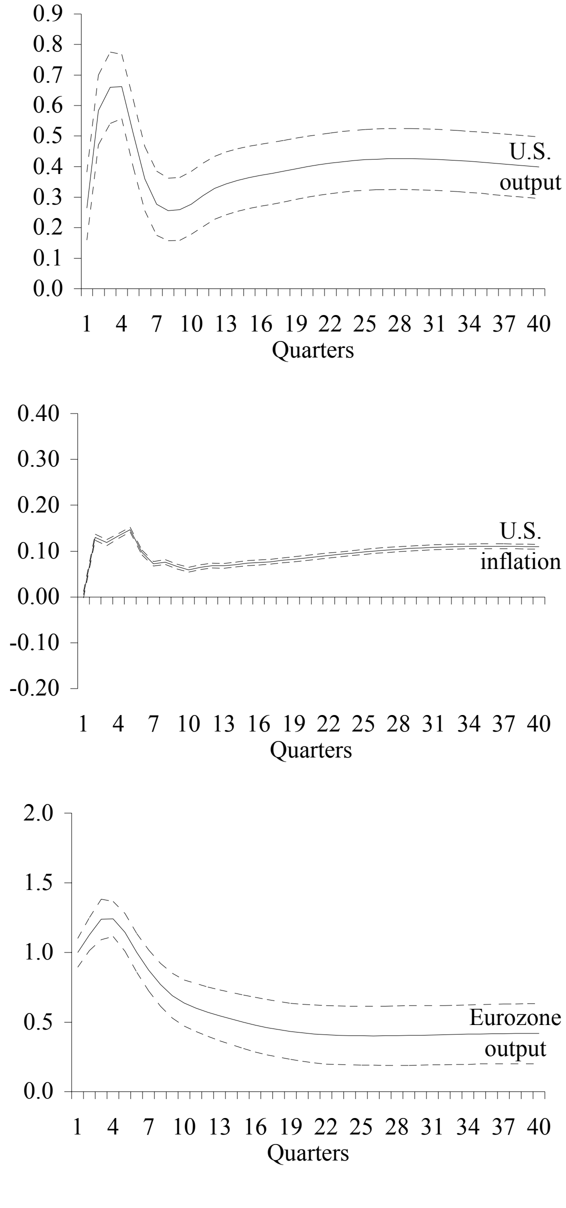


Figure 2b: Response to a positive 1-percent shock to Eurozone output



Note: Generalized impulse response functions according to Koop et al. (1996). Figure 2a shows the response of the Eurozone model and Figure 2b shows the response of the U.S. model. Dotted lines plot upper and lower bounds of 95-percent confidence intervalls, which are based on a non-parametric bootstrap using 2000 replications.

Figure 3a: Response to a positive 1-percentage point shock to U.S. interest rate

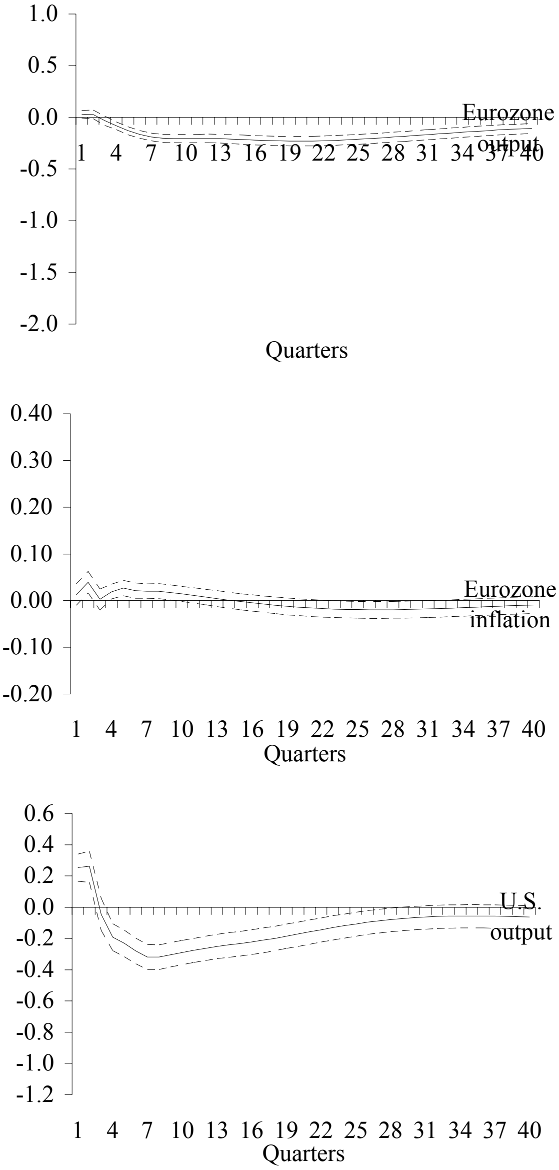
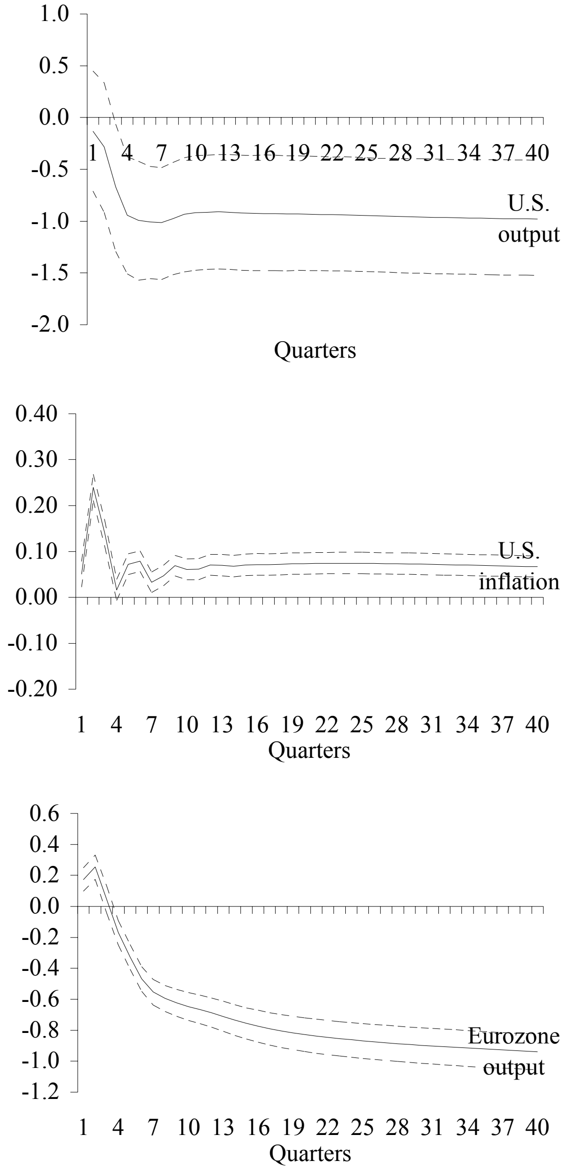


Figure 3b: Response to a positive 1-perc. point shock to Eurozone interest rate



Note: Generalized impulse response functions according to Koop et al. (1996). Figure 3a shows the response of the Eurozone model and Figure 3b shows the response of the U.S. model. Dotted lines plot upper and lower bounds of 95-percent confidence intervalls, which are based on a non-parametric bootstrap using 2000 replications.

Figure 4a: Response to a positive 1-percent shock to the Euro-Dollar exchange rate

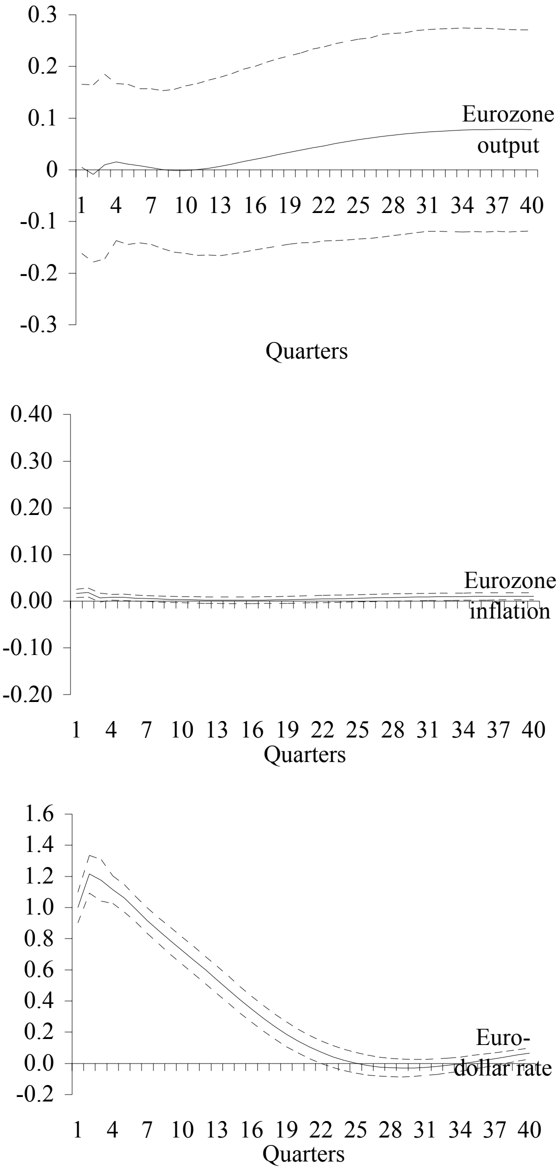
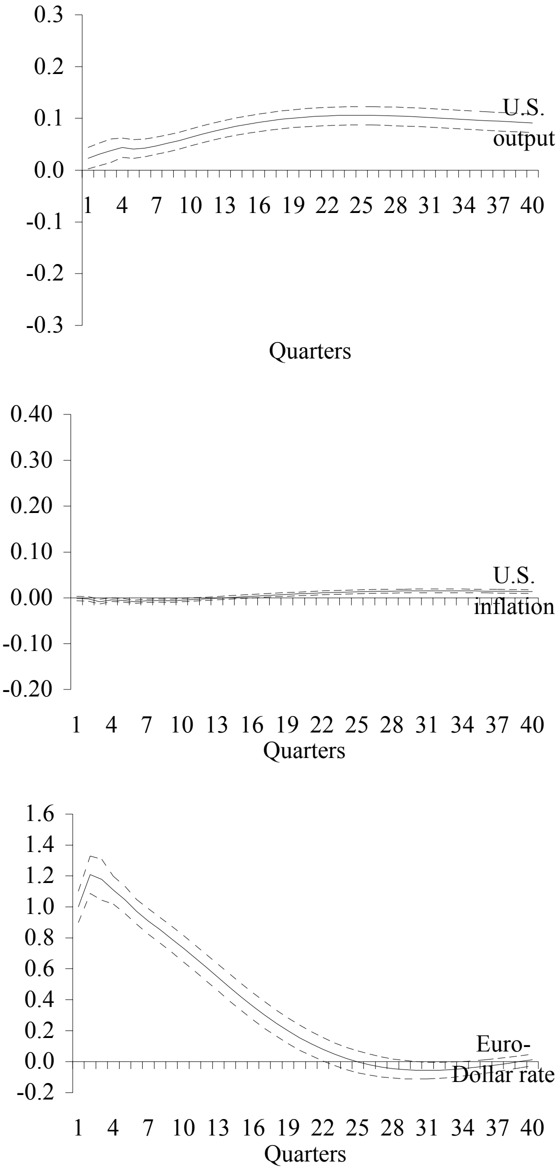


Figure 4b: Response to a positive 1-percent shock to the Euro-Dollar exchange rate



Note: Generalized impulse response functions according to Koop et al. (1996). Figure 4a shows the response of the Eurozone model and Figure 4b shows the response of the U.S. model. Dotted lines plot upper and lower bounds of 95-percent confidence intervalls, which are based on a non-parametric bootstrap using 2000 replications.

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

	ADF		PP		KPSS	
	Test statistics based on model including					
	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 *
$(p_{ez}-p_{us})$	0.00 ***	0.01 ***	0.18	0.45	0.13	0.13 *
p_{ez}	0.05 **	0.58	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					
Δe	0.00 ***	0.00 ***	0.00 ***	0.00 ***	0.07	0.06
Δm_{ez}	0.00 ***	0.00 ***	0.00 ***	0.00 ***	0.08	0.08
Δm_{us}	0.00 ***	0.00 ***	0.00 ***	0.00 ***	0.08	0.08
$\Delta(p_{ez}-p_{us})$	0.01 ***	0.03 **	0.00 ***	0.00 ***	0.11	0.04
Δp_{ez}	0.75	0.37	0.00 ***	0.00 ***	1.07 ***	0.10
Δp_{oil}	0.00 ***	0.00 ***	0.00 ***	0.00 ***	0.19	0.13 *
Δp_{us}	0.13	0.05 *	0.00 ***	0.00 ***	0.82 ***	0.09
Δi_{ez}	0.00 ***	0.00 ***	0.00 ***	0.00 ***	0.08	0.05
Δi_{us}	0.00 ***	0.00 ***	0.00 ***	0.00 ***	0.05	0.05
Δy_{ez}	0.00 ***	0.00 ***	0.00 ***	0.00 ***	0.25	0.06
Δy_{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 ***	0.04	0.04
$\Delta\Delta p_{oil}$	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
$\Delta\Delta y_{us}$	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		U.S.A.	
p	AIC	BIC	AIC	BIC
1	-56.34	**-54.39**	-56.34	**-54.39**
2	**-57.03**	-53.78	-57.07	-53.82
3	-57.02	-52.44	**-57.37**	-52.79
4	-57.01	-51.10	-57.21	-51.30

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.