Modeling Monetary Transmission in Switzerland with a Structural Cointegrated VAR Model

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1. Introduction

Central banks are interested in learning about the effect of a change in their policy rate – typically a short-term interest rate – on their target variables, inflation and output growth. Consequently, a rich empirical literature on the monetary transmission mechanism exists, mainly covering the United States and the euro area. By contrast, only few studies focus on Switzerland.[1]

This paper investigates the transmission of monetary policy shocks in a structural cointegrated vector-autoregressive (VAR) model for Switzerland that allows us to impose and test a long-run structure as well as a short-run structure on the data. We include exogenous variables in the model, using the methodology of Pesaran, Shin and Smith (2000) and Garratt, Lee, Pesaran and Shin (2003). The five cointegrating relations among the variables in the model are interpreted as capturing money demand, the real interest rate, a term spread, uncovered interest parity and an output demand schedule. To investigate whether economic relations have remained stable after the Swiss National Bank (SNB) adopted a new monetary policy framework in 2000, we perform stability tests and recursive analyses. A monetary policy shock is identified by interpreting the

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1 Examples are Kugler and Rich (2002), Jordan, Kugler, Lenz and Savioz (2002), Natal (2002, 2004), Kugler and Jordan (2004) and Kugler, Jordan, Lenz and Savioz (2005).

contemporaneous relations between the variables as interactions between money demand and a monetary policy reaction function. Impulse responses show that a contractionary monetary policy shock leads to a decline in inflation and output. Though the model contains the price of oil as a commodity price, an exchange rate puzzle, meaning that the exchange rate depreciates on impact after a contractionary monetary policy shock, remains present.

The paper is structured as follows. The next section provides a brief overview of the existing empirical studies on monetary transmission in Switzerland. Most of the literature for Switzerland focusses on structural vector autoregressive (SVAR) models that include only a small number of variables and neglect the openness of the Swiss economy. Section 3 discusses the data and their time-series properties. Since we find that all macroeconomic variables can be considered as non-stationary, we estimate a cointegrated VAR model that is discussed in Section 4. Section 4 also presents the results from the empirical analysis of the cointegrating relations. In Section 5 we impose an economic structure on the covariance matrix of the residuals and identify a monetary policy shock from the interaction of money demand and a monetary policy reaction function. Finally, Section 6 concludes.

2. Related Literature

Structural VAR models have been widely used to investigate the transmission of monetary policy shocks to macroeconomic variables. These models impose restrictions on the covariance matrix of the residuals of a reduced-form VAR to provide an economic interpretation of the shocks. By contrast, a structural cointegrated VAR model also specifies the number of cointegrating relations and possibly imposes restrictions on the cointegrating space.[2] As in SVAR models, a structural identification of the covariance matrix can be added and the usual impulse-response, variance-decomposition and historical decomposition analyses can be performed.

Cointegrated VAR studies of monetary policy effects have been conducted for several countries, including the euro area (Vlaar and Schuberth, 1998; Coenen and Vega, 2001; Vlaar, 2004), Germany (Hubrich and Vlaar, 2004;

2 In particular, a cointegrating VAR model implies that the dimension of the cointegrating space is larger than zero and smaller than the number of variables in the system since otherwise the model would correspond either to a VAR in first differences or a VAR in levels.

Brüggemann, 2003), Spain (Juselius and Toro, 2006) and the United Kingdom (Dhar, Pain and Thomas, 2000; Garratt, Lee, Pesaran and Shin, 2003). Baltensperger, Jordan and Savioz (2001) estimate a cointegrated VAR model for Switzerland comprising nominal M3, real gross domestic product (GDP), the GDP deflator and the government bond yield. Imposing a long-run money demand function and using a Choleski decomposition to identify monetary policy shocks, they find that the residuals from the long-run money demand equation help forecast inflation. Assenmacher-Wesche and Pesaran (2007) compute generalized impulse-response functions to an interest rate shock in a cointegrated VAR model for Switzerland but they do not identify structural shocks.

The studies by Jordan, Kugler, Lenz and Savioz (2002), Kugler and Rich (2002) and Kugler and Jordan (2004) are based on structural VAR models estimated in first differences, comprising M1, GDP, the consumer price index (CPI) and the three-month interest rate. They use long and short-run restrictions and find plausible reactions to a monetary policy shock. The paper by Kugler, Jordan, Lenz and Savioz (2005) uses long-run identifying restrictions only, but is able to generate similar impulse responses to a monetary policy shock. Natal (2002; 2004) extends the analysis by considering foreign variables. He also includes money and interprets the contemporaneous interactions as reflecting money-supply and money-demand shocks. In contrast to this paper, however, he does not identify cointegrating relations between the variables. Since VAR models contain many parameters, it seems desirable to impose restrictions on the long-run behavior of the variables. In addition, we include foreign variables that appear important, providing a more complete analysis of the transmission mechanism in Switzerland than the previous studies.

3. Data

After the breakdown of the Bretton-Woods System in 1973 the SNB started to target M_1 in 1975, which we choose as the starting date for the empirical analysis.[3] Since the model will include two lags, the effective sample runs from the third quarter of 1974 to the last quarter of 2006.

The first step in the construction of the model is the choice of variables.[4] Since our goal is to analyze monetary transmission, the system needs to include a short-

3 Peytrignet (1999) surveys the experience with monetary targeting in Switzerland.

4 A detailed description of the data can be found in the appendix.

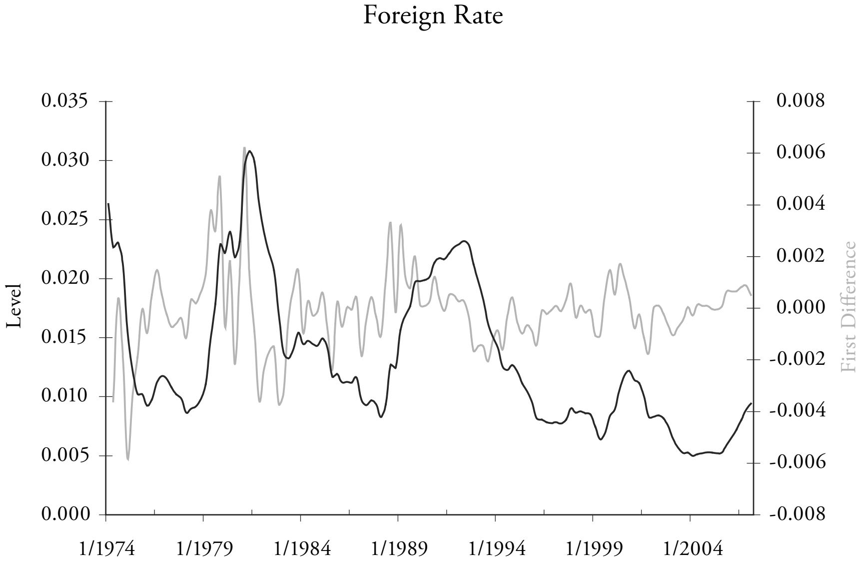
term interest rate, r^s, as the policy instrument. Here, we use the three-month London interbank offered rate (LIBOR). In addition, we consider output, y, and the rate of inflation, π, because these are the variables we are most interested in when assessing the effects of monetary policy. Inflation is measured as the quarterly change in the consumer price index (CPI) because the SNB focusses on the CPI in its formulation of monetary policy.

The SNB pursued for long a strategy of monetary targeting that was abandoned only in 2000. Nevertheless, money has remained important in the new policy framework (see JORDAN, PEYTRIGNET and RICH, 2001). Since available empirical evidence indicates that the demand for M_2 has been stable in the past (FISCHER and PEYTRIGNET, 1991; PEYTRIGNET and STAHEL, 1998), we include real M_2, deflated with the CPI. The nominal effective exchange rate of the Swiss franc (defined as foreign currency per Swiss franc), e, is included because it is related to both the course of monetary policy and exerts an influence on output and inflation. Finally, the long-term interest rate, r^l, which can be expected to play a role in monetary transmission from policy rates to output and inflation, is included as well.

Switzerland is a small open economy and heavily influenced by developments abroad. To capture the links with the international economy we add the price of oil (in US dollar), p^{oil}, and the foreign interest rate, r^*, to the system. Oil prices show large fluctuations that generally transmit quickly to domestic prices. Interest rates in Switzerland are strongly influenced by foreign interest rates, with the euro area being the most important neighboring country. We therefore proxy the foreign interest rate by the three-month euro rate. Following ARTIS and BEYER (2004) and BRÜGGEMANN and LÜTKEPOHL (2006) we link the euro-area rate to the German three-month rate rather than to an average of the euro-area member countries before 1999, since arguably the European Central Bank resembles most closely the Bundesbank in its monetary policy.

All variables are in logarithms. Interest rates are expressed as $0.25\ln(1 + R/100)$ where R is the interest rate in percent per annum to make units comparable to the quarterly inflation rate. Figure 1 presents the variables in levels and first differences. It is apparent that interest rates and inflation show similar behavior over time so that we may expect to find some cointegrating relations between them. Moreover, first differences seem to be stationary, though formal evidence from unit-root tests will be presented next.

Figure 1: Data in Levels (Black) and First Differences (Grey)



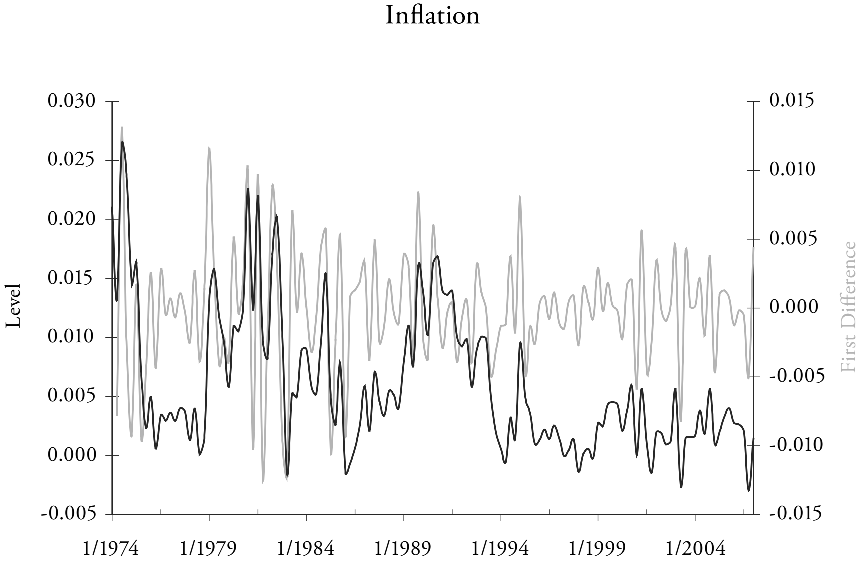
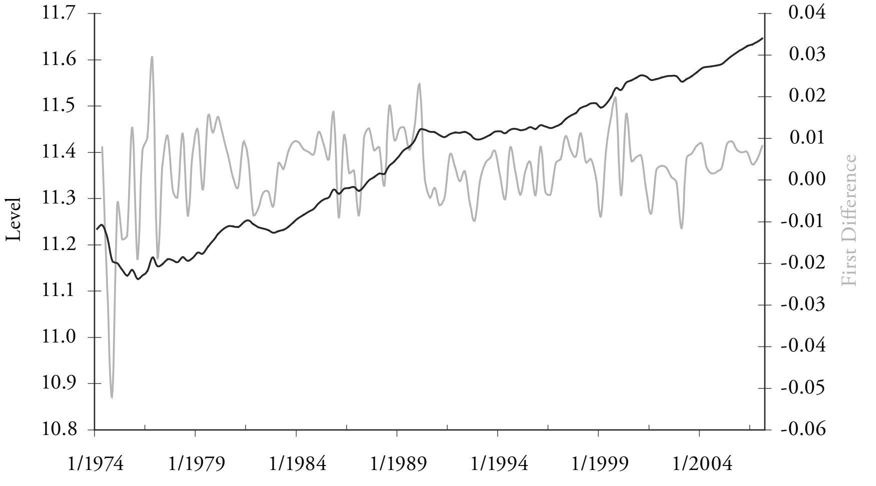


Figure 1 (continued)

Output



Real M2

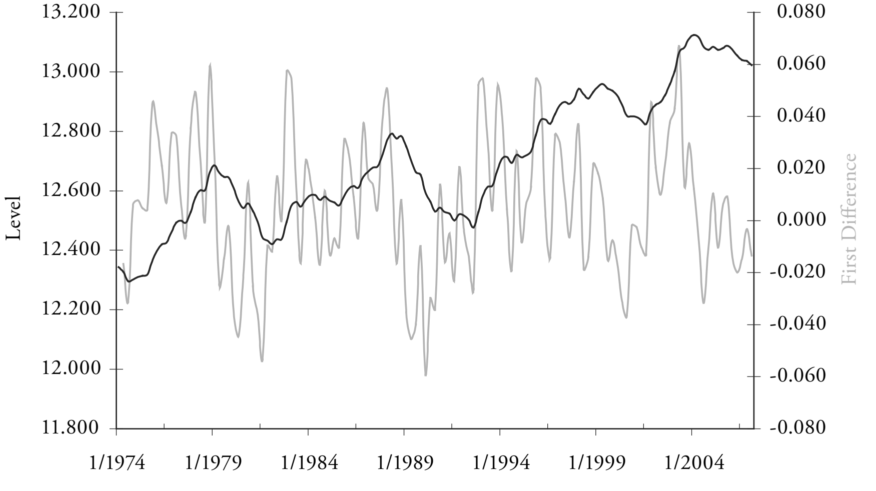
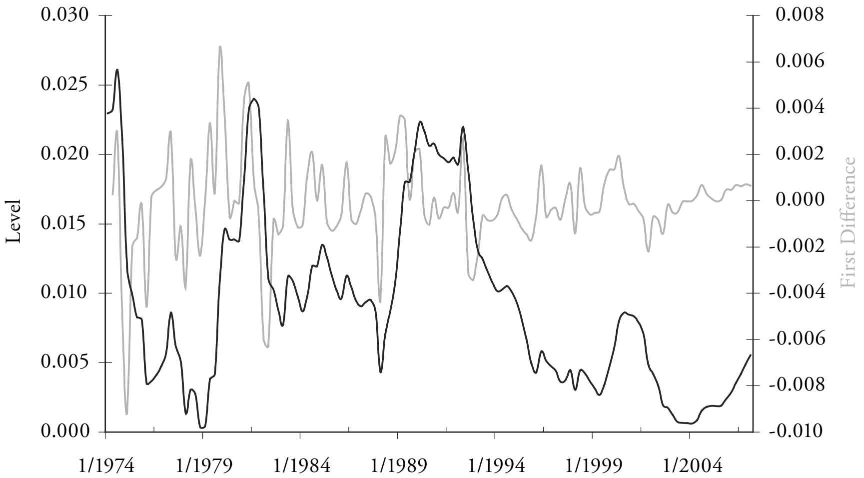


Figure 1 (continued)

Short Rate



Long Rate

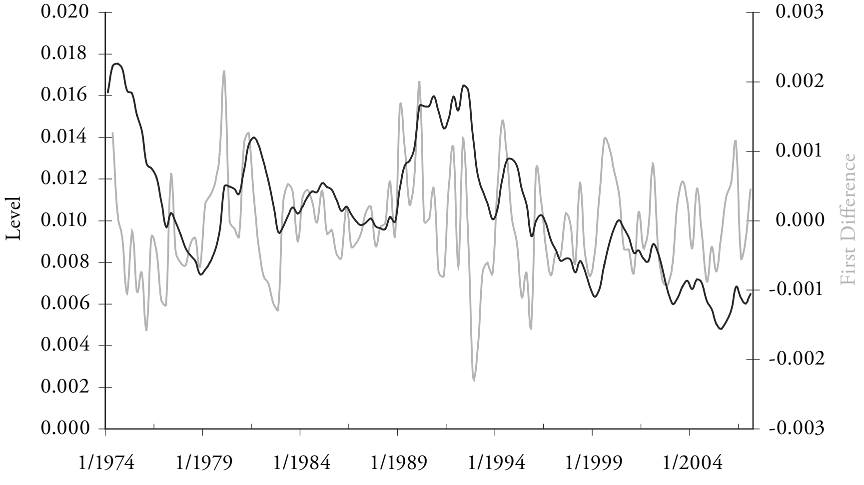
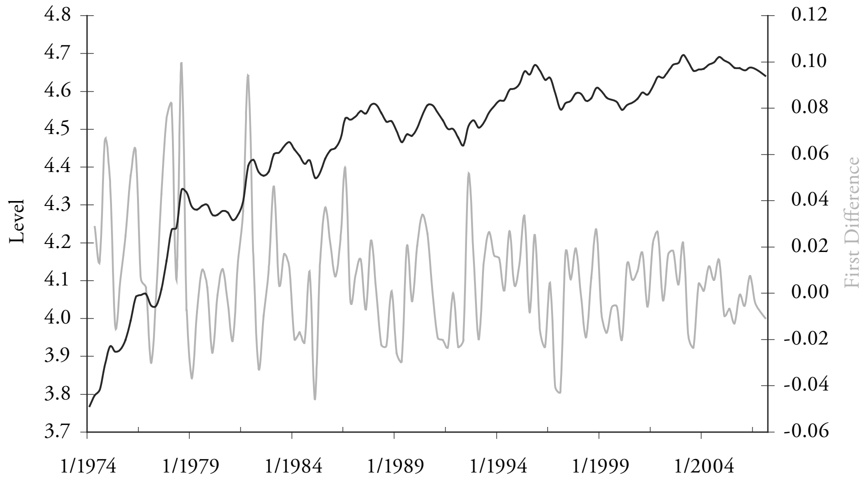
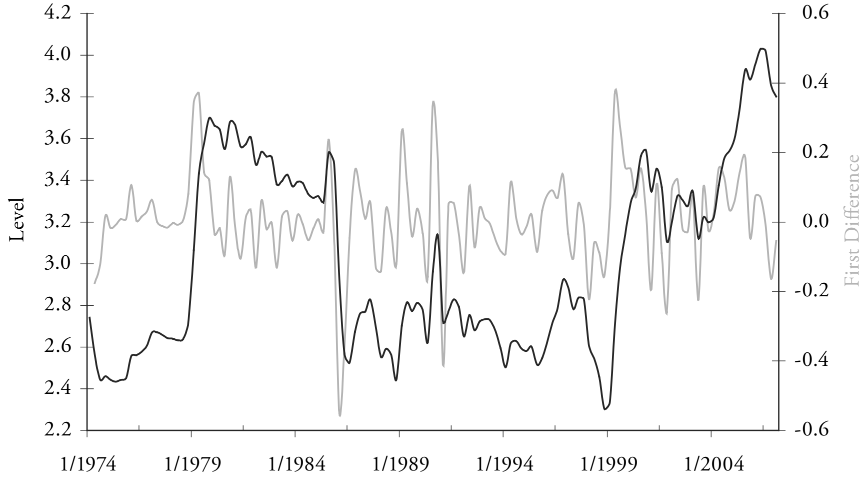


Figure 1 (continued)

Exchange Rate



Oil Price



It is important to first investigate the time-series characteristics of the data since they have implications for the econometric methodology used and the long-run cointegrating relations one would expect to find between the variables. We therefore perform Augmented Dickey Fuller (ADF) tests, allowing for up to six lags. From Figure 1 it is apparent that e_t, $(m-p)$, y_t, and possibly p_t^{oil} trend over time whereas inflation and interest rates do not appear to do so. The regressions in levels therefore include a trend and an intercept for the former group of variables and an intercept only for the latter group. All ADF regressions applied to the first differences include an intercept. The results are shown in Table 1. Entries in italics show the lag length that was selected by the Akaike criterion (AIC). Since the results of the ADF tests may depend on the number of lags included in the regressions we show results including up to six lagged differences of the variable to be tested. After accounting for a maximum of six lags, the sample period for the unit-root tests runs from 1976Q1 to 2006Q4 for all tests.

In establishing the unit-root properties of the variables we first check whether their first differences are in fact stationary. The ADF test results for the first differences, which are reported in the upper panel of Table 1, reject the presence of unit roots in all the series, except for real M_2 when three lags are included in the regression. Since the AIC favors the inclusion of four lags and this gives a clear indication of stationarity, we proceed with the assumption that all the first differences are stationary.

Turning to the level of the variables, the ADF-test results in the bottom panel of Table 1 suggest that the unit-root hypothesis cannot be rejected in all the variables, when the order of augmentation indicated by the AIC is used. In general, this result continues to hold if other augmentation orders are used. The only exceptions are real money when four lags are included and the domestic interest rate when less than two lags are included. Summing up, we interpret the unit-root tests as indicating that all series under consideration can be regarded as nonstationary.

4. The Structural Cointegrated VAR Model

The starting point for the empirical analysis is the reduced form of a cointegrated VAR model with p lags, written in error-correction form,

$$\Delta \mathbf{z}_t = -\mathbf{\Pi} \mathbf{z}_{t-1} + \sum_{i=1}^{p-1} \mathbf{\Gamma}_i \Delta \mathbf{z}_{t-i} + \mathbf{b}_0 + \mathbf{b}_1 t + \mathbf{u}_t, \qquad (1)$$

Table 1: Unit Root Tests

				First Differences				
Lags	Δr^*	$\Delta\pi$	$\Delta(m-p)$	Δy	Δr^s	Δr^l	Δe	Δp^{oil}
0	*–6.72*	–15.51	–6.63	*–10.27*	*–7.68*	*–7.17*	*–8.30*	–8.56
1	–5.35	–11.69	–5.87	–6.66	–6.32	–6.54	–6.45	–8.02
2	–5.01	–8.77	–4.30	–5.26	–5.69	–5.43	–5.57	–5.40
3	–4.54	*–8.36*	–2.80	–4.51	–5.38	–5.20	–5.75	–5.68
4	–4.02	–6.93	*–4.75*	–4.45	–4.97	–5.77	–5.69	–5.71
5	–4.08	–6.43	–4.28	–4.18	–4.24	–4.45	–5.28	–4.35
6	–4.07	–4.86	–4.05	–4.48	–4.40	–3.73	–4.56	*–4.66*
				Levels				
Lags	r^*	π	$(m-p)$	y	r^s	r^l	e	p^{oil}
0	–1.27	–4.75	–1.55	–1.91	–1.65	–1.62	–2.11	–1.55
1	*–2.16*	–3.51	–2.59	–2.09	*–2.47*	*–2.46*	*–2.76*	–2.14
2	–2.38	*–2.82*	–2.56	–2.41	–2.52	–2.32	–2.90	–1.73
3	–2.33	–2.89	–3.16	*–2.70*	–2.47	–2.42	–2.98	–2.23
4	–2.45	–2.35	–4.83	–2.97	–2.42	–2.31	–2.66	–1.84
5	–2.68	–2.40	*–2.88*	–2.95	–2.44	–1.95	–2.47	–1.57
6	–2.52	–2.26	–3.00	–3.14	–2.71	–2.22	–2.42	*–1.99*

Note: The first column shows the number of lags included in the test. All regressions in first differences include an intercept, whereas the regressions in levels include a trend and an intercept for e, $m-p$, y and p^{oil}, and an intercept only for r^*, π, r^s, and r^l. After accounting for the necessary lags, the sample period for all tests runs from 1976Q1 to 2006Q4. Entries in italics denote the lag length selected by the AIC criterion. The 5 percent critical values are –3.45 for the tests including an intercept and a trend and –2.89 for tests including a trend only.

where $\mathbf{z}_t = \{\pi_t, y_t, (m-p)_t, r_t^s, r_t^l, e_t, r_t^*, p_t^{oil}\}'$ is a $m \times 1$ vector including inflation, real output, real M_2, the short-term interest rate, the long-term interest rate, the exchange rate, the foreign interest rate and the oil price. The matrices $\{\boldsymbol{\Gamma}_i\}_{i=1}^{p-1}$ contain the short-run responses, $\mathbf{b}_0$ denotes a vector of constants and $\mathbf{b}_1$ a vector of trend coefficients that are restricted to lie in the cointegration space.[5] The residuals, $\mathbf{u}_t$, are distributed normally with mean zero and covariance matrix

5 See Garratt, Lee, Pesaran and Shin (2006, Section 6.3) for details.

Σ, $\mathbf{u}_t \sim \mathrm{N}(0,\Sigma)$. The matrix Π is a $m \times m$ matrix of long-run multipliers. If there exist r cointegrating relations, $0 < r < m$, between the variables, Π has reduced rank and can be written as

$$\Pi = \alpha\beta', \tag{2}$$

where α is a $m \times r$ matrix of loading coefficients and β is a $m \times r$ matrix of long-run coefficients.[6]

The reduced-form model in equation (1) can be transformed into its structural form by pre-multiplying a nonsingular matrix $\mathbf{A}_0$,

$$\mathbf{A}_0 \Delta \mathbf{z}_t = -\mathbf{a}\beta' \mathbf{z}_{t-1} + \sum_{i=1}^{p-1} \mathbf{A}_i \Delta \mathbf{z}_{t-i} + \mathbf{a}_0 + \mathbf{a}_1 t + \varepsilon_t, \tag{3}$$

where $\varepsilon_t \sim \mathrm{N}(0,\Omega)$ and the short-run parameters of the structural form are related to the short-run parameters of the reduced form by the following relations:

$$\mathbf{A}_i = \mathbf{A}_0 \boldsymbol{\Gamma}_i, \ \mathbf{a} = \mathbf{A}_0 \alpha, \ \varepsilon_t = \mathbf{A}_0 \mathbf{u}_t, \ \mathbf{a}_0 = \mathbf{A}_0 \mathbf{b}_0, \ \mathbf{a}_1 = \mathbf{A}_0 \mathbf{b}_1, \tag{4}$$

and $\Omega = \mathbf{A}_0 \Sigma \mathbf{A}'_0$. While the reduced-form parameters can be estimated from the data one has to impose additional constraints to obtain $\mathbf{A}_0$.

Because β is the same in both the structural and the reduced form, estimation of the cointegrating relations can be done in either form. Identification of a long-run structure and a short-run structure thus can be treated as two separate statistical problems (see Juselius, 2006). We therefore first deal with the identification of the long-run structure before turning to the identification of the $\mathbf{A}_0$ matrix.

4.1 Modelling Choices

Before estimating the model in equation (1), the number of lags, p, to be included in the estimation has to be selected.[7] To this end, different lag-selection criteria are computed for an unrestricted VAR model, considering a maximum lag

6 We speak of loading coefficients instead of error-correction coefficients because these pertain to the reduced form and cannot be given an economic interpretation. By contrast, **a** shows the structural error correction, see equation (3).

7 Estimation of the cointegrating relations was done with the program by Anders Warne for Matlab which is available under http://texlips.hypermart.net/svar/index.html.

length of four.[8] The Akaike information criterion (AIC) and the forecast prediction error (FPE) select a specification with two lags, whereas the Hannan-Quinn (HQ) criterion, the Schwarz criterion and the fractional marginal likelihood (FML) favor one lag. Since first-order serial correlation was present when choosing $p = 1$, we select $p = 1$.

Table 2: Cointegration Test Statistics

Rank	Eigenvalue	Trace statistic	90% Asymptotic critical value
1	0.492	288.26	181.08
2	0.422	201.67	144.81
3	0.305	131.50	112.54
4	0.188	84.99	84.27
5	0.152	58.36	60.00
6	0.139	37.25	39.73
7	0.100	18.13	23.32
8	0.036	4.68	10.68

Note: The system includes r^*, π, y, $m-p$, r^s, r^l, e, p^{oil}, a constant and a restricted linear trend and is estimated with two lags. The sample period is 1975Q1 to 2006Q4. The asymptotic critical values are from Juselius (2006).

The cointegration test statistics are shown in Table 2. The null hypothesis of no cointegration is investigated by testing the rank of Π in equation (1). We find four eigenvectors that are significantly different from zero at the 90 percent level. Since the system with eight variables is rather large, we next test for weak exogeneity. This amounts to testing whether the error-correction terms enter significantly the equation for the respective variable. We find that the oil price can be considered as weakly exogenous with a p-value of 0.20. Though it is unlikely that Swiss disequilibria have an influence on the euro-area interest rate, the rejection of the foreign interest rate being weakly exogenous is likely to be caused by temporal aggregation or expectation effects. If the SNB is able to predict future policy moves of the European Central Bank and takes them into account when

8 Asymptotic results for the lag order selection criteria continue to hold if cointegrated $I(1)$ processes are investigated instead of $I(0)$ processes, see Paulsen (1984).

setting the domestic short-term rate, we would expect to find an influence of the error-correction terms on the foreign interest rate. We here include r^* among the endogenous variables but results for the cointegration space remain unaffected if the euro-area rate is treated as being weakly exogenous.[9]

4.2 The Conditional Model

In estimating the model we follow the approach developed by Pesaran, Shin and Smith (2000) and Garratt, Lee, Pesaran and Shin (2003; 2006). Weak exogeneity of the oil price implies that this variable has a direct, contemporaneous influence on the endogenous variables but is itself not affected by the error-correction terms, i.e., the disequilibria in the Swiss economy.

Relying on these results we next investigate a partial system that is conditioned on the oil price. To this end, we partition the vector $\mathbf{z}_t$ into $\mathbf{z}_t = \{\mathbf{x}_t', \mathbf{x}_t^{*\prime}\}'$ where $x_t = \{r_t^*, \pi_t, y_t, (m-p)_t, r_t^s, r_t^l, e_t\}'$ is a $m_x \times 1$ vector of endogenous variables and $\mathbf{x}_t^* = \{p_t^{oil}\}'$ a $m_{x^*} \times 1$ vector with the exogenous variable. By partitioning the error term $\mathbf{u}_t$ as $\mathbf{u}_t = (\mathbf{u}'_{xt}, \mathbf{u}'_{x^*t})'$ and its variance matrix as

$$\Sigma = \begin{pmatrix} \Sigma_{xx} & \Sigma_{xx^*} \\ \Sigma_{x^*x} & \Sigma_{x^*x^*} \end{pmatrix}, \tag{5}$$

$\mathbf{u}_{xt}$ can be expressed conditionally in terms of $\mathbf{u}_{x^*t}$ as $\mathbf{u}_{xt} = \Sigma_{xx^*}\Sigma_{x^*x^*}^{-1}\mathbf{u}_{x^*t} + \upsilon_t$. The cointegrated VAR model in equation (1) then can be split into a conditional model, describing the evolution of the endogenous variables,

$$\Delta\mathbf{x}_t = -\alpha_x \beta \mathbf{z}_{t-1} + \Lambda\Delta\mathbf{x}_t^* + \sum_{i=1}^{p-1}\Psi_i \Delta\mathbf{z}_{t-i} + \tilde{\mathbf{b}}_{x0} + \mathbf{b}_{x1}t + \upsilon_t, \tag{6}$$

and a marginal model for the exogenous variables,

$$\Delta\mathbf{x}_t^* = \sum_{i=1}^{p-1}\Gamma_{x^*i}\Delta\mathbf{z}_{t-i} + \mathbf{b}_{x^*0} + \mathbf{u}_{x^*t}, \tag{7}$$

9 Distinguishing between endogenous and weakly exogenous variables does not change the number of cointegrating relations nor the identification of a long-run structure. By contrast, it will have an effect on the identification of a short-run structure.

where $v_t \sim \mathrm{N}(0, \Sigma_{xx} - \Sigma_{xx^*}\Sigma_{x^*x^*}^{-1}\Sigma_{x^*x})$ is uncorrelated with $\mathbf{u}_{x^*t}$ by construction. The coefficients in equations (6) and (7) are related to the coefficients in equations (1) and (2) by the following relations:

$$\begin{array}{lll} \alpha = (\alpha'_x, 0)', & \Lambda = \Sigma_{xx^*}\Sigma_{x^*x^*}^{-1}, & \Gamma_i = (\Gamma_{x,i}, \Gamma_{x^*,i})', \\ \Psi_i = \Gamma_{x,i} - \Lambda\Gamma_{x^*,i}, & i = 1, \ldots, p-1, & \mathbf{b}_1 = (\mathbf{b}_{x1}, 0)', \\ \mathbf{b}_0 = (\mathbf{b}_{x0}, \mathbf{b}_{x^*0})' \text{ and} & \tilde{\mathbf{b}}_{x0} = \mathbf{b}_{x0} - \Lambda\mathbf{b}_{x^*0}. & \end{array}$$

Estimation of the cointegrating relations based on the conditional model in equation (6) is as efficient as maximum likelihood estimation of the whole system because the information available from the model for $\Delta\mathbf{x}_t^*$ is redundant for estimation of the parameters in the model for $\Delta\mathbf{x}_t$. By contrast, investigation of the dynamic properties of the system, e.g., through impulse-response analysis, requires the inclusion of the marginal model in equation (7) because the dynamic properties depend on the processes driving the exogenous variables.

4.3 Identification of the Long-Run Structure

Since we have already tested for cointegration in the unconditional system, there is no need to recalculate the cointegration test statistics for the conditional model. We therefore continue by testing identification restrictions on β.[10] Among the variables in the model, we would expect to find a money-demand relation, linking money to output and the opportunity cost of holding real balances. Since M_2 includes interest bearing components, the opportunity cost will be the difference between the yield on alternative assets and the own rate of return. Nevertheless, since the inclusion of a yield spread causes the rate of inflation to drop out, inflation should enter the cointegrating relation as additional variable. In addition, a Fisher-parity relationship might be present, implying that the real interest rate is stationary in the long run. The expectations theory of the term structure suggests that the term spread is stationary. Moreover, if the change in the exchange rate is stationary, the domestic interest rate should be cointegrated with the foreign interest rate. Finally, there could exist an aggregate-demand relation between the deviations of output from trend and real interest and exchange rates.

10 The cointegrated VAR approach aims at describing the data in a statistical way. Inside this model, certain hypothesis coming from economic theory then can be tested. For a more formal exposition of similar long-run relations as we use here, see Garratt, Lee, Pesaran and Shin (2006).

Table 3: Cointegration Properties

	r^*	π	y	$(m-p)$	r^s	r^l	e	p^{oil}	t	χ^2 (df)	p
	Money demand relations										
H_1	0	–21.38	–1.04	1	48.00	–26.62	0	0	0	2.99 (2)	0.22
H_2	0	18.61	–1	1	31.39	–32.78	0	0	0	3.78 (3)	0.34
H_3	0	16.06	–1	1	0	0	0	0	0	13.22 (4)	0.01
	Output relations										
H_4	0	8.52	1	0	0	–8.52	0.07	0	–0.0033	0.38 (2)	0.83
H_5	0	25.48	1	0	0	0	0	0	–0.0017	6.99 (3)	0.07
	Interest rate relations										
H_6	0	–1	0	0	1	0	0	0	0	15.12 (5)	0.01
H_7	0	–1	0	0	0	1	0	0	0	13.49 (5)	0.02
H_8	0	0	0	0	1	–1	0	0	0	19.70 (5)	0.00
H_9	–1	0	0	0	1	0	0	0	0	17.14 (5)	0.00
H_{10}	0	–1.06	0	0	1	0	0	0	0	15.09 (4)	0.00
H_{11}	0	–0.64	0	0	0	1	0	0	0	11.48 (4)	0.02
H_{12}	0	0	0	0	1	–1.81	0	0	0	4.31 (4)	0.37
H_{13}	–1.11	0	0	0	1	0	0	0	0	17.02 (4)	0.00

Note: The hypotheses H_1 to H_{13} are described in the text. χ^2 is the likelihood ratio test statistic of the overidentifying restrictions, *df* denotes degrees of freedom and *p* is the associated *p*-value.

Since imposing all economically motivated restrictions at once makes it difficult to find out which relations are not well supported by the data, we first test these relations one by one, leaving the other cointegrating vectors unrestricted (see Juselius and Toro, 2006). Results are shown in Table 3. The hypothesis H_1 describes agents' demand for money as depending on real output, the difference between the alternative yield, r^l, and the own yield on money, proxied by the three-month interest rate, r^s, and the difference between the rate of inflation and r^s,

$$(m-p)_t = \beta_1^{H_1} y_t + \beta_2^{H_1}(r_t^l - r_t^s) + \beta_3^{H_1}(\pi_t - r_t^s) + b_0^{H_1} + u_t^{H_1}. \tag{8}$$

The last column in Table 3 shows that the implied restrictions are not rejected with a *p*-value of 0.22. Since the estimated coefficient on output is close to unity, H_2 sets $\beta_1^{H_2} = 1$, which increases the p-value to 0.34. We would expect $\beta_2^{H_1} < 0$ and $\beta_3^{H_1} < 0$, but the estimates in Table 3 imply positive coefficients on both opportunity cost variables. One reason may be that the short-term interest rate actually is a better measure of the opportunity cost than the long-term interest rate or inflation because the components of M_2 are of short maturity and typically close substitutes to short-term assets. Another explanation is that we capture effects of the expectation hypothesis or the Fisher parity in this single cointegration vector that can be better disentangled when imposing restrictions on all cointegrating vectors together. We will come back to this issue later.

H_3 tests whether the behavior of the SNB can be characterized by targeting the liquidity ratio at a level consistent with a constant inflation target, which can be formulated as

$$(m-p)_t - y_t = \beta_1^{H_3}\pi_t + b_0^{H_3} + u_t^{H_3} \tag{9}$$

and is rejected.

The second panel in Table 3 tests different specifications of an output relation. H_4 implies that the deviation of output from trend can be explained by the long-term real interest rate and the exchange rate,

$$y_t = \beta_1^{H_4}\pi_t + \beta_2^{H_4} r_t^l + \beta_3^{H_4} e_t + \beta_4^{H_4} t + b_0^{H_4} + u_t^{H_4}, \tag{10}$$

with $\beta_2^{H_4} > 0$, $\beta_1^{H4} = -\beta_2^{H4}$, $\beta_3^{H_4} < 0$ and $\beta_4^{H_4} > 0$. [11]

This hypothesis cannot be rejected, but whereas the coefficients on the trend and the exchange rate show the expected sign, the coefficients on the interest rate and inflation have the wrong sign. An interpretation of the output relation as a Phillips-curve relation with inflation adjusting to output deviations from trend is also not rejected at the 5 percent significance level, though with a lower p-value. We would expect a positive relation between output and inflation in H_5, which again does not come out in the results; see Table 3.

11 Ideally, the real exchange rate should enter this relation. We included the nominal exchange rate in the system because of its relevance for monetary policy and inflation. A possible solution would be to include the ratio between domestic and foreign prices as additional variable but since the system is already large with eight variables, we did not pursue this approach.

The last panel tests different hypothesis on the interest-rate relations. H_6 and H_7 test the Fisher parity between inflation and the short-term or the long-term interest rate. Both hypotheses are rejected. H_8 tests the expectations hypothesis and H_9 a long-run relation between the domestic and the foreign interest rate. Again, both hypotheses are rejected. We next relax the assumption of a unit coefficient in H_6 to H_8, which results in hypotheses H_{10} to H_{13}. Only H_{12}, which represents a relaxed version of the expectations hypothesis, is not rejected with a point estimate of –1.81 on the long-term rate, which is much smaller than the theoretically expected value of –1.

To identify the whole β matrix we next impose the restrictions that were not rejected together on the cointegrating vectors. Because we have only three interpretable cointegrating relations that are not rejected, we leave the fourth cointegration vector unrestricted. When considering several hypotheses together, one needs to keep in mind that linear combinations of the cointegration vectors also lie in the cointegration space. If, e.g., a term-structure relation is imposed, the coefficients on the long-term and the short-term interest rate in the money-demand relation are no longer identified.[12] To achieve identification, the coefficient on one interest rate in H_2 thus has to be restricted to zero. Here, we report the interest elasticity of money demand with respect to the long-term rate but results would remain unchanged if we included the short rate in the money-demand relation instead. Table 4 shows that H_2, H_5 and H_{12} together are rejected with a p-value of 0.02.

We obtain somewhat more appealing results when five cointegrating vectors are considered. Though the fifth eigenvector is not significantly different from zero at the 90 percent level, the trace statistic is close to the critical value. Despite only one interest relation being not rejected by the data, we would expect to have three theoretically motivated relations between the three interest rates and inflation in the model. It thus seems reasonable to investigate also a system with five cointegrating relations, in particular since the results for four cointegration vectors do not match with our theoretic priors and thus are difficult to interpret.

Imposing H_2, H_5 and H_{12} together with the Fisher parity on a system with five cointegration vectors gives indeed more plausible results. The implied restrictions are not rejected with a likelihood ratio test statistic of 16.35 and a p-value of 0.13. Nevertheless, some caveats are in place. First, the coefficient on the long rate in the term-structure relation (H_{12}) is still about half the size theory would predict. Second, the exchange rate enters the Fisher parity. Third, though the last

12 See Hoffman and Rasche (1996, p. 194) for a more detailed discussion.

Table 4: Cointegration Properties

	r^*	π	y	$(m-p)$	r^s	r^l	e	p^{oil}	t
					Three overidentified vectors, one exactly identified vector				
						$\chi^2(9) = 19.42$		$p = 0.02$	
H_1	0	–23.17	–1	1	52.20	–29.03	0	0	0
		(2.12)			(2.30)	2.42			
H_4	0	17.61	1	0	0	–17.61	–0.26	0	–0.0036
		(0.45)				(0.45)	(0.02)		(0.0001)
H_{12}	0	0	0	0	1	–1.96	0	0	0
						(0.04)			
	1.04	–5.53	0	0	1	0	0.10	0.008	–0.0007
	(0.15)	(0.18)					(0.01)	(0.001)	(0.0001)
					Four overidentified vectors, one exactly identified vector				
						$\chi^2(11) = 16.35$		$p = 0.13$	
H_1	0	0	–1	1	0	19.30	0	0	0
						(2.56)			
H_4	0	34.95	1	0	0	–34.95	–0.26	0	–0.0010
		(1.70)				(1.70)	(0.06)		(0.0003)
H_{12}	0	0	0	0	1	–2.08	0	0	0
						(0.12)			
H_6	0	–1	0	0	1	0	0.01	0	0
							(0.001)		
	–1	0	–0.33	0	1	0	0	–0.01	0.0013
			(0.02)					(0.001)	(0.0001)
					Five overidentified vectors				
						$\chi^2(11) = 48.94$		$p = 0.00$	
H_1	0	0	–1	1	0	12.64	0	0	0
						(2.98)			
H_4	0	0	1	0	0	0	0.07	0.026	–0.0044
							(0.02)	(0.004)	(0.0001)
H_{12}	0	0	0	0	1	–1	0	0	0
H_6	0	–1	0	0	1	0	0	0	0
H_9	–1	0	0	0	1	0	0	0	0

Note: Standard errors in parentheses. The reported p-values are from the χ^2 distribution. The bootstrapped p-value for the restrictions in the last panel is 0.09.

cointegrating vector implies a proportional relation between the domestic and the foreign interest rate, output, the oil price and a trend cannot be excluded.

Thus, the estimated system is still difficult to interpret in terms of economic relations. We therefore also impose all theoretically motivated restrictions together. The resulting coefficients are reported in the last panel of Table 4. Not surprisingly, the restrictions are rejected with a test statistic of 48.94. Nevertheless, the asymptotic critical values of the χ^2 distribution may not be valid in small samples (Gredenhoff and Jacobson, 2001). Using bootstrapped critical values leads to a *p*-value of 0.09 and thus a non-rejection of the restrictions on the cointegrating vectors at the 5 percent significance level. Since these vectors are more easily interpretable in economic terms, we proceed with them. Moreover, when checking the stability of the coefficients, which we will discuss in the next section, the coefficients turn out to be more stable in the more restricted system.

We next discuss the estimates for this system. The first cointegrating relation represents money demand. The semi-elasticity of money demand with respect to the interest rate is negative and implies that real money demand decreases by about three percent if the interest rate increases by one percentage point.

We interpret the second cointegrating relation as an output relation, linking deviations of output from trend to the exchange rate and the oil price. The coefficient on the oil price is significant and implies a reduction in output when oil prices increase.[13] The positive coefficient on the exchange rate is consistent with a decline in output when the exchange rate appreciates. The estimate of the trend implies an average annual growth rate of GDP of about 1.8 percent.

Note, however, that with this interpretation two problems arise. First, the real exchange rate and second, the relative price of oil in Swiss francs should appear in this relation. Ideally, the cointegrating relation should include the exchange rate minus the price level and the oil price minus the nominal exchange rate and the price level. Since prices turned out to be $I(2)$, they are not included in the system so that this specification is unfeasible in the present setup. One possible interpretation of our results is that the estimated vector reflects the price of oil converted into Swiss francs. For this interpretation, however, the coefficient on the exchange rate shows the wrong sign. In addition, the oil price is denoted in US dollar, whereas the exchange rate index contains mainly European countries as Switzerland's main trading partners. Alternatively, one could subtract the rate of inflation from both variables. The LR-test for this restrictions is 49.02 with a bootstrapped *p*-value of 0.09. Since the results are almost identical and both

13 Deviating from the previous analysis, we allow for a non-zero coefficient on oil prices in the output relation. This improved the fit considerably.

specifications do not match exactly with the theoretical priors, we proceed with the specification in Table 4 that involves fewer estimated parameters.

The three last cointegrating vectors involve no estimated parameters and represent a stationary real rate, the spread between the short-term and the long-term interest rate and long-run uncovered interest-rate parity, which reflects the close integration of the Swiss economy with the financial markets in the euro area.

Finally, Table 5 shows summary statistics for the estimated error-correction equations in the system. Serial correlation is absent except for the money equation. The other tests do also not point to serious misspecification of the model. The R^2 ranges from 0.26 for the output equation to 0.64 for the inflation equation, indicating that the model is able to explain a large part of the quarterly changes in the variables.

Table 5: Specification Tests for Error-Correction Equations

Variable	*LM*	*RESET*	*JB*	*ARCH*	R^2
r^*	0.07	0.01	0.00	0.02	0.49
π	0.07	0.45	0.01	0.92	0.64
y	0.56	0.88	0.00	0.05	0.26
$m-p$	0.00	0.09	0.56	0.07	0.58
r^s	0.11	0.02	0.05	0.00	0.49
r^l	0.57	0.13	0.37	0.00	0.43
e	0.35	0.18	0.00	0.05	0.31

Note: Entries show p-values. The endogenous variables are r^*, π, y, $m-p$, r^s,r^l, and e; the exogenous variable is p^{oil}. The sample period is 1975Q1 to 2006Q4. The system is estimated with two lags of the endogenous and the exogenous variables, a constant and a restricted linear trend. *LM* is a Lagrange-multiplier test for 4th order serial correlation, *RESET* is Ramsey's reset test of functional form, *JB* is the Jarque-Bera test for normality, and *ARCH* is a Lagrange-multiplier test for autoregressive conditional heteroscedasticity.

4.4 Stability of the Model

We next test stability by performing recursive analyses for the last 40 quarters, starting in 1997Q1. After having pursued a strategy of monetary targeting – first with targets for M1, later for M0 – continuing stability problems with M0 due to financial innovation lead the SNB to announce a new monetary policy framework in 2000. We thus investigate recursive stability of the model around this date. Stability of a cointegrating VAR model can be assessed with respect to the long-run relations as well as with respect to the short-run parameters. Since the long-run relations are based on economic theory we would expect them to be relatively more stable after a change in the monetary policy framework than the short-run coefficients that are entirely driven by the data. We will investigate both issues below.

Figure 2 shows recursive estimates for the long-run coefficients. While most coefficients seem reasonably stable, there is evidence that the interest elasticity of money demand has decreased until 2002, whereas it remained stable thereafter. One possible reason is that interest rates declined sharply to exceptionally low values whereas money growth increased to almost 20 percent at the time. Nevertheless, including the logarithm of the interest rate instead of the interest rate itself, which allows a stronger reaction of money demand in times of low interest rates, did not improve results. Applying a Nyblom test (see Hansen and Johansen 1999) for the constancy of β leads to a bootstrapped p-value of 0.23 for the supremum test (0.19 for the mean test) and thus a non-rejection of constancy.

Figure 3 shows the recursive likelihood-ratio test of the overidentifying restrictions as imposed in the bottom panel of Table 4. Except for the period 2003/2004 the test statistic remains roughly constant, indicating that the restrictions fit the data equally well during the recursive estimation period.

The next two figures consider the stability of the short-run coefficients. Figure 4 shows the results from recursive Chow breakpoint tests, which tests whether the short-run coefficients in the error-correction equations are constant for date $t+1$ relative to date t. There are a few rejections of parameter constancy in the equations for $m-p$ and y especially around 2003/04 when – in response to the recession after the stock-market bubble in 2000 – the SNB lowered the interest rate to 0.25 percent. Without including stock prices the model apparently has difficulties to reconcile high money growth and low interest rates with the low output growth at the time. In general, however, the figures indicate that the model is reasonably stable.

Figure 5 shows the Ploberger-Krämer-Kontrus (1989) test that checks whether the short-run coefficients are constant over time. Also here no evidence of instability is visible. We thus conclude that the model is reasonably stable also after the introduction of the SNB's new monetary policy framework.

Finally, we investigate the stability of the cointegrating relations over the full sample period. Figure 6 shows the cointegration relations, corrected for the short-run dynamics. In general, the relations seem to be reasonably stable and consistent with error-correction towards equilibrium, though in two cases somewhat larger errors had built up. In the first cointegrating vector, which we interpret as representing money demand, quite persistent errors are present in the early 1990s while the uncovered interest parity shows large deviations from equilibrium around the early 1980s. Thereafter, however, the cointegrating relations continue to fluctuate around their previous levels. Nevertheless, there appears to be a level shift in the term-structure relation from the mid-1990s on, which might have led to rejecting the unit coefficient in H_8. Since Switzerland was not among the countries that lifted capital controls in the early 1990s, the cause for this shift is not obvious. In total, however, the relations look reasonably stable and display little persistence. We thus feel confident that the specification with the five overidentified vectors (as reported in the bottom panel of Table 4) provides an appropriate model for the Swiss economy.

Figure 2: Recursive Estimates of β with 95 Percent Confidence Bands

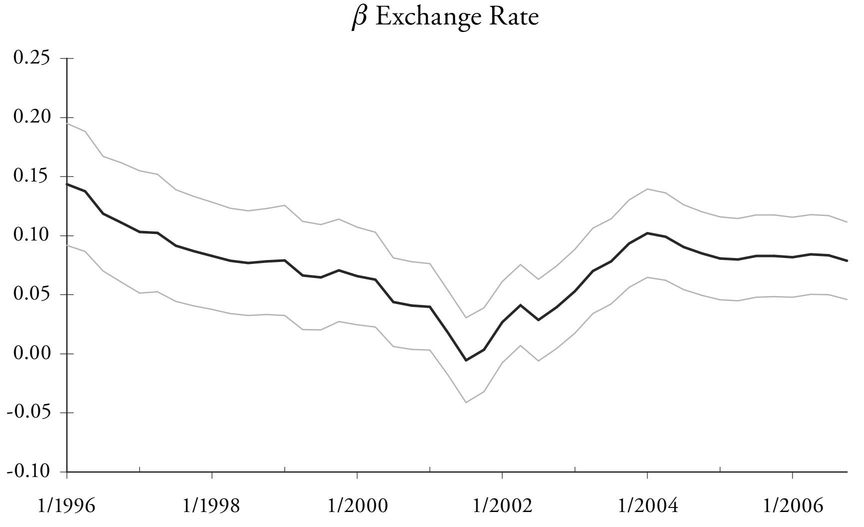
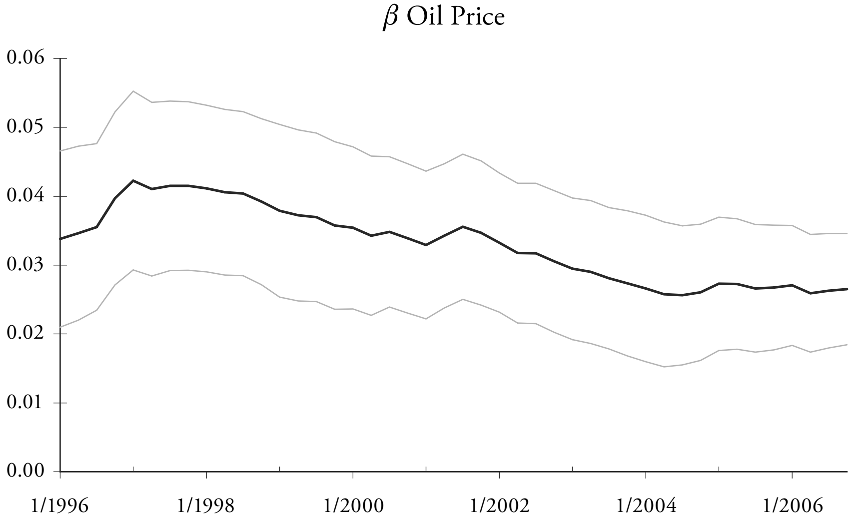


Figure 2 (continued)



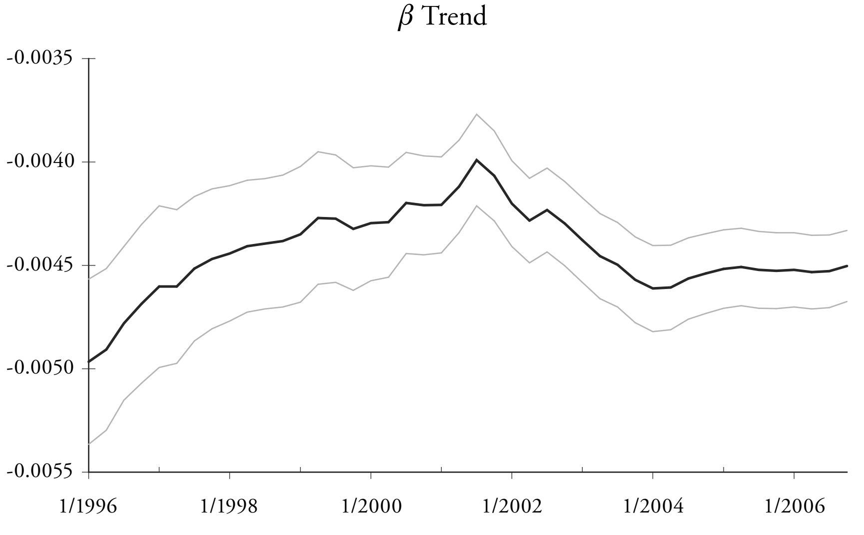


Figure 3: Recursive *LR* Tests of β Restrictions

Note: The asymptotic critical value for the 5 percent significance level is 26.3.

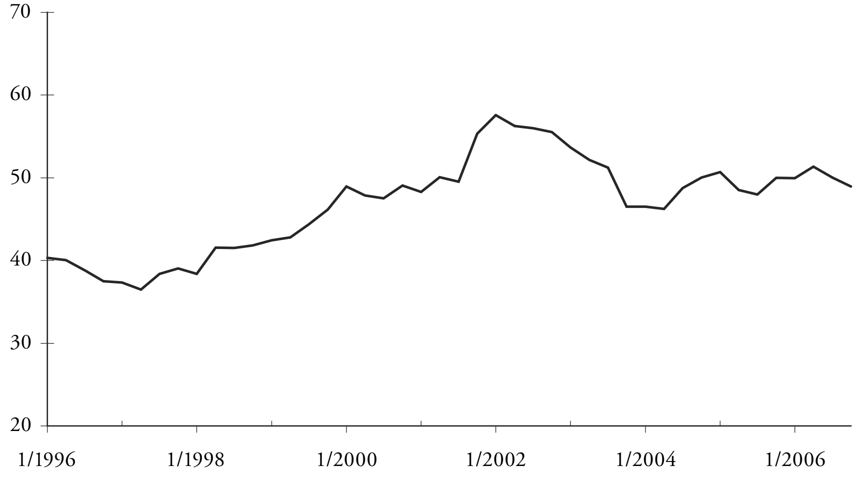


Figure 4: Recursive Chow Tests

Note: The tests are normalized so that the 5 percent critical value is unity.

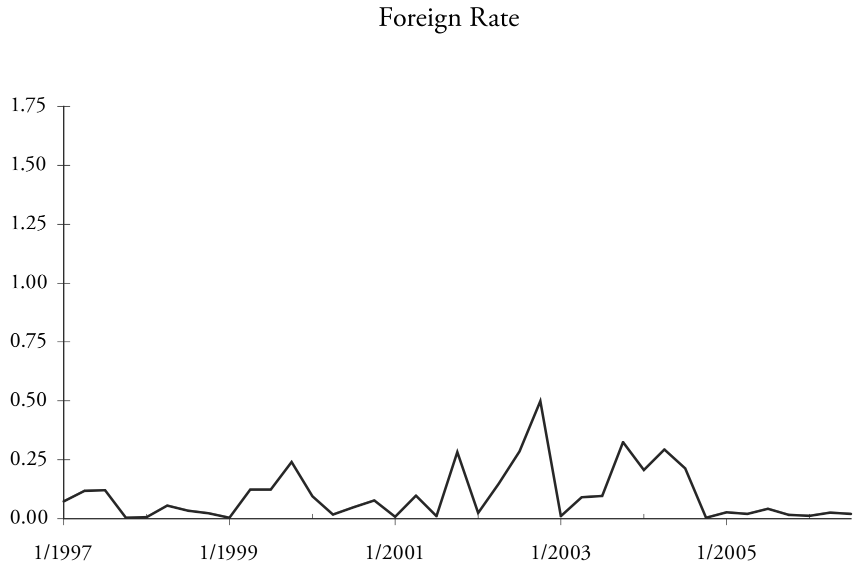
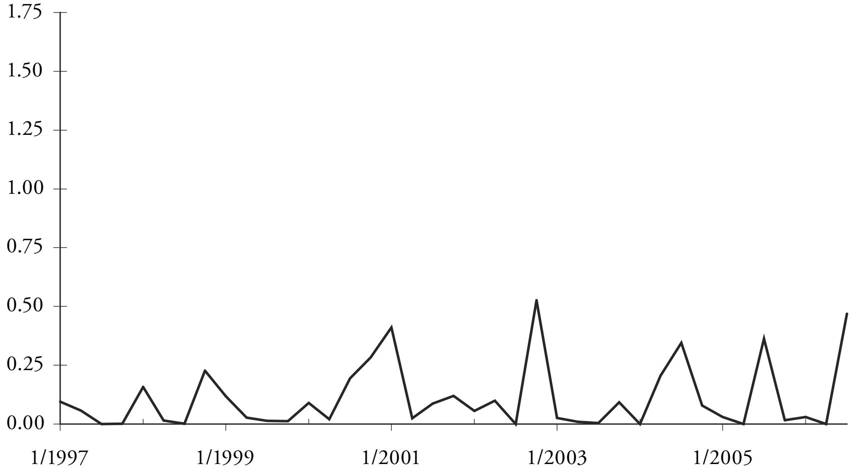


Figure 4 (continued)

Inflation



Output

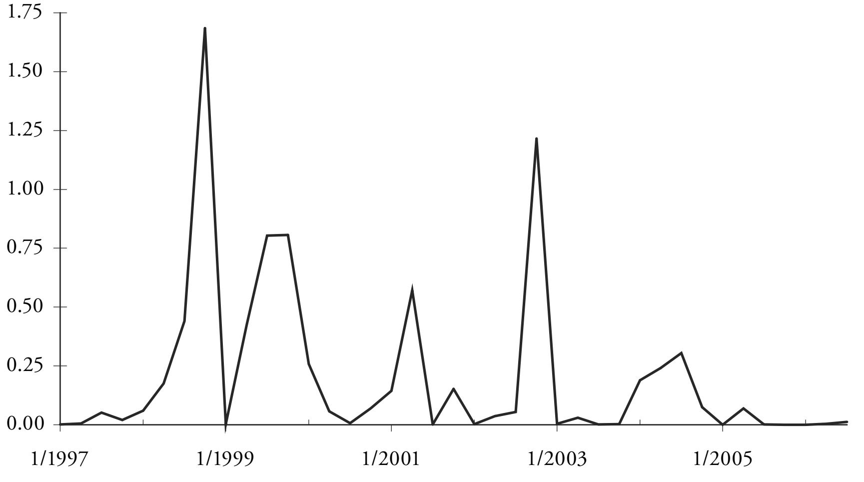
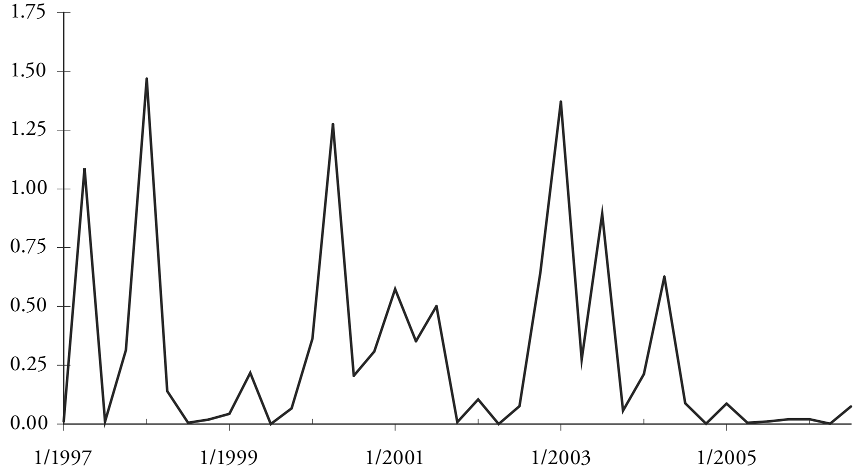


Figure 4 (continued)

Real M2



Short Rate

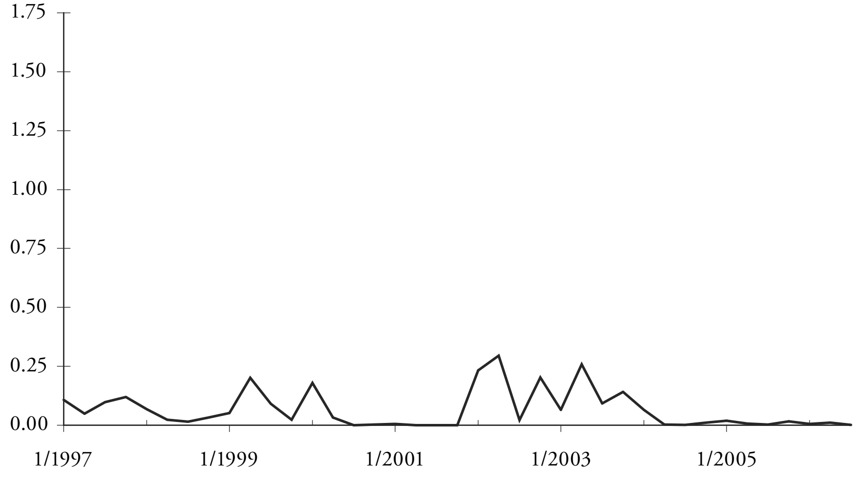
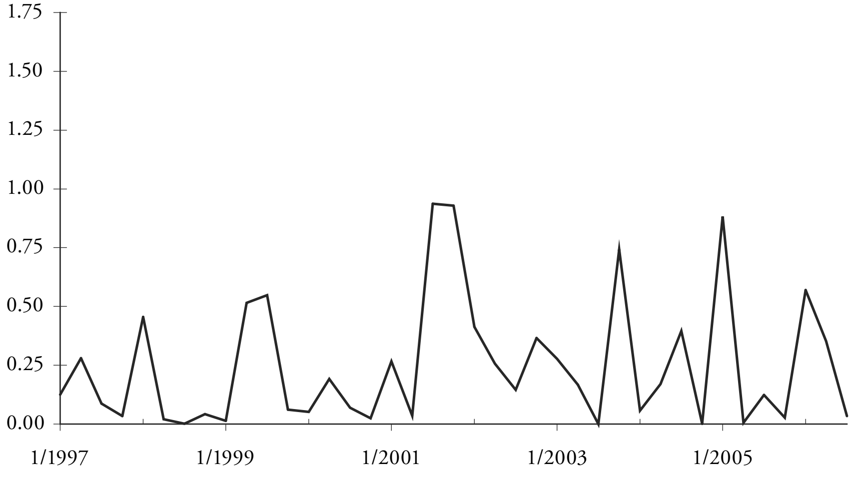


Figure 4 (continued)

Long Rate



Exchange Rate

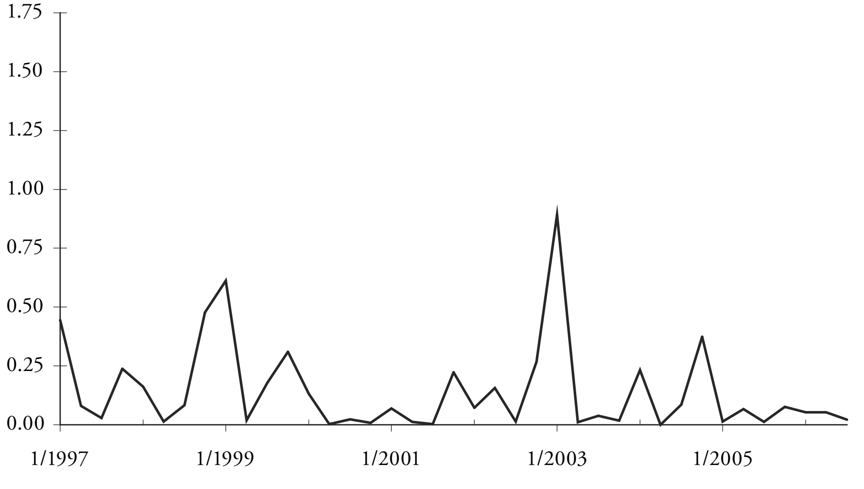
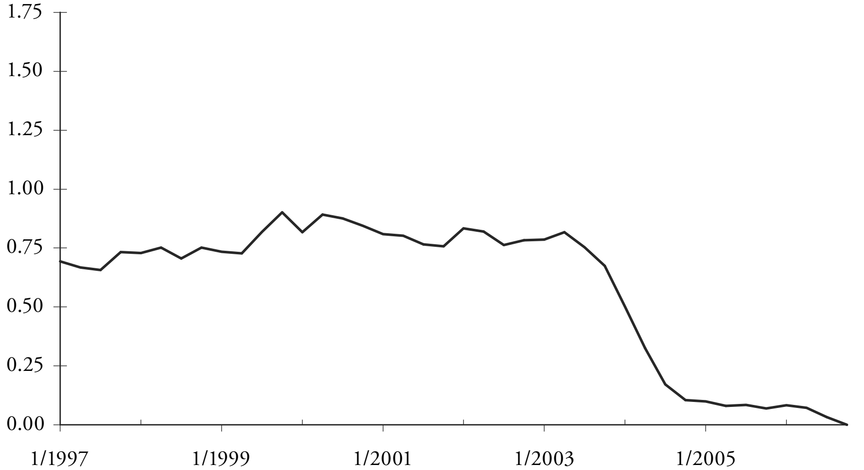


Figure 5: Recursive Ploberger-Krämer-Kontrus (1989) Tests

Note: The test statistic is distributed as Student's *t* with 15 degrees of freedom. The 5 percent critical value is 1.75.

Foreign Rate



Inflation

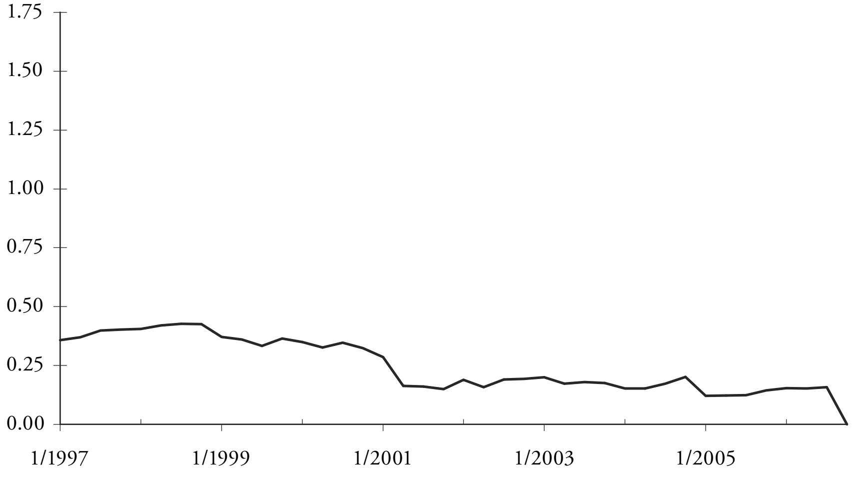
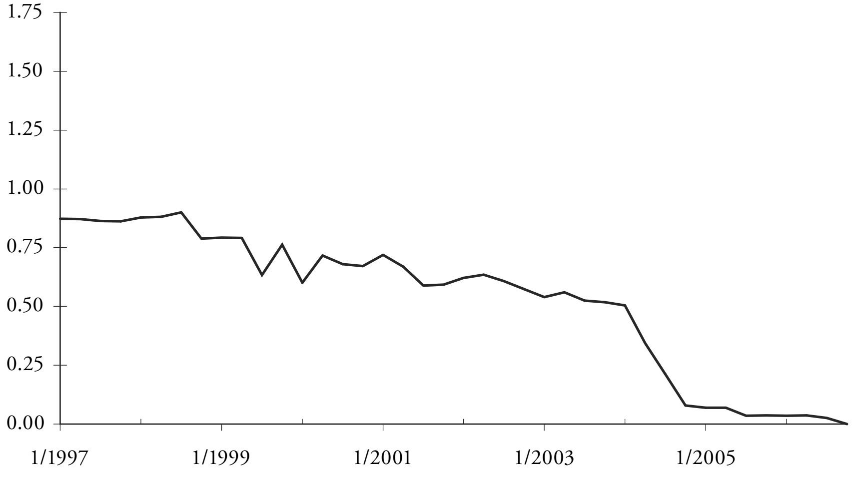


Figure 5 (continued)

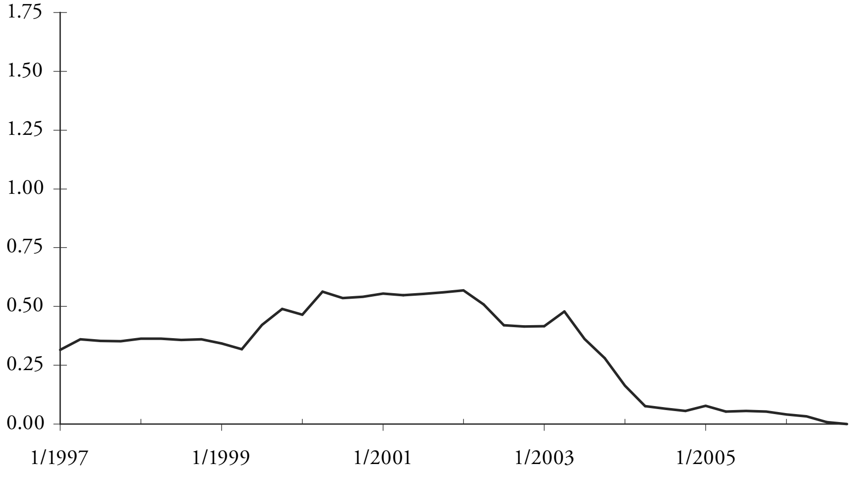
Output



Real M2

Figure 5 (continued)

Short Rate



Long Rate

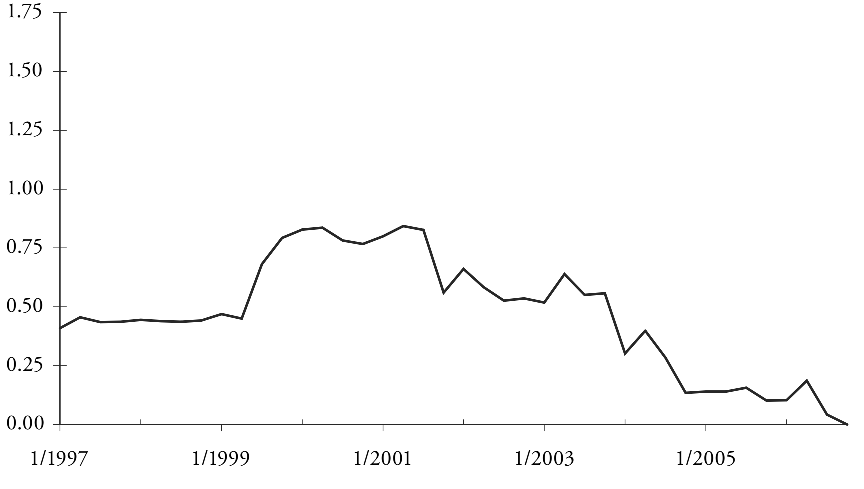


Figure 5 (continued)

Exchange Rate

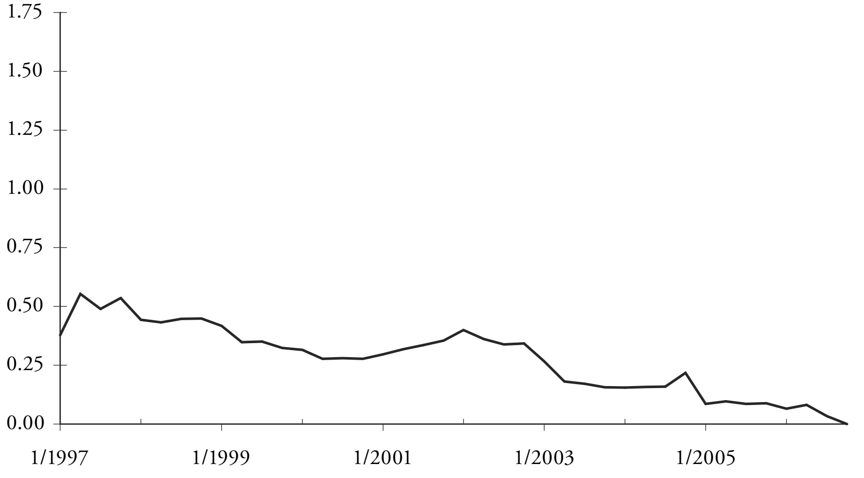


Figure 6: Cointegrating Relations

Money Demand

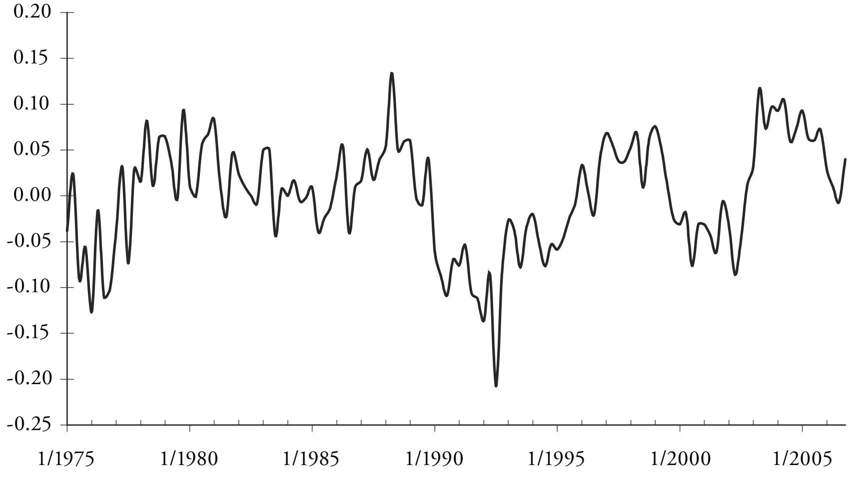
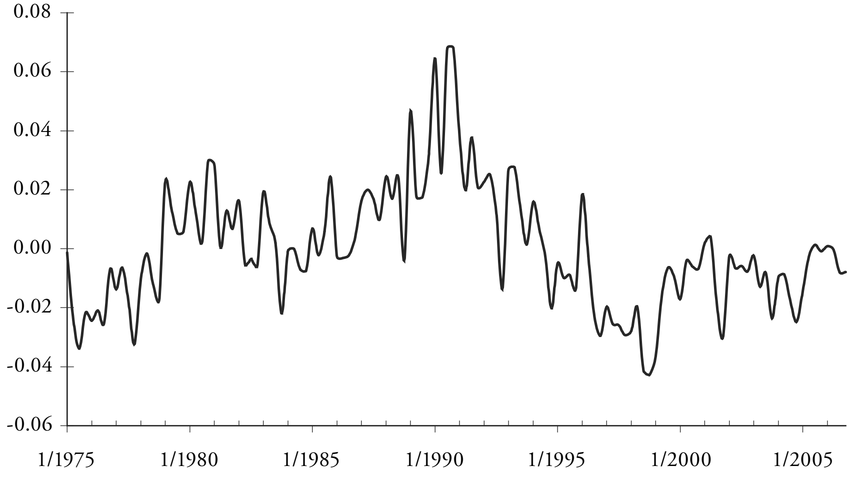


Figure 6 (continued)

Output Relation



Term Spread

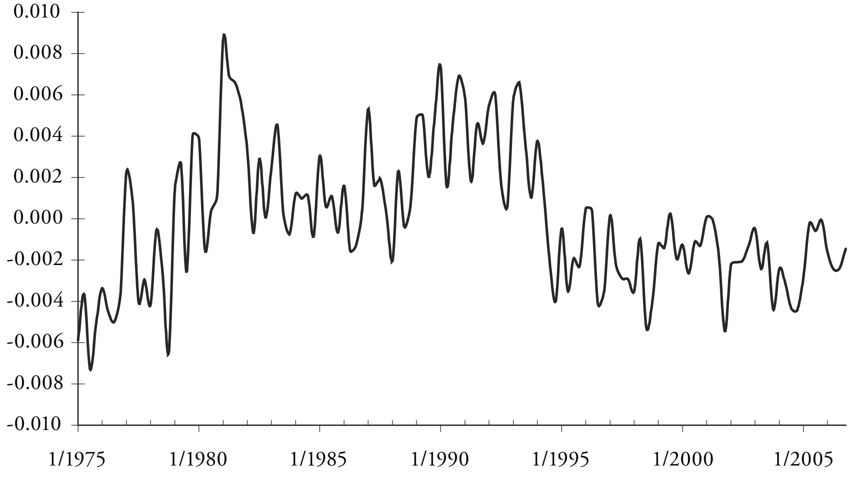
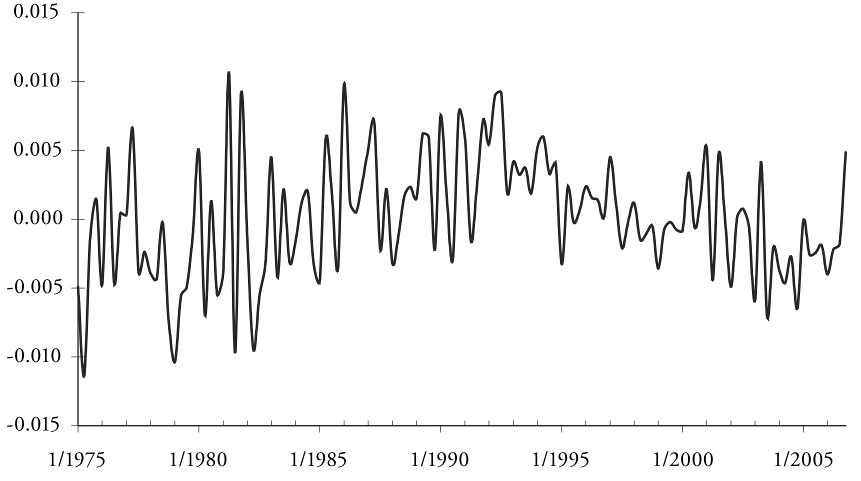
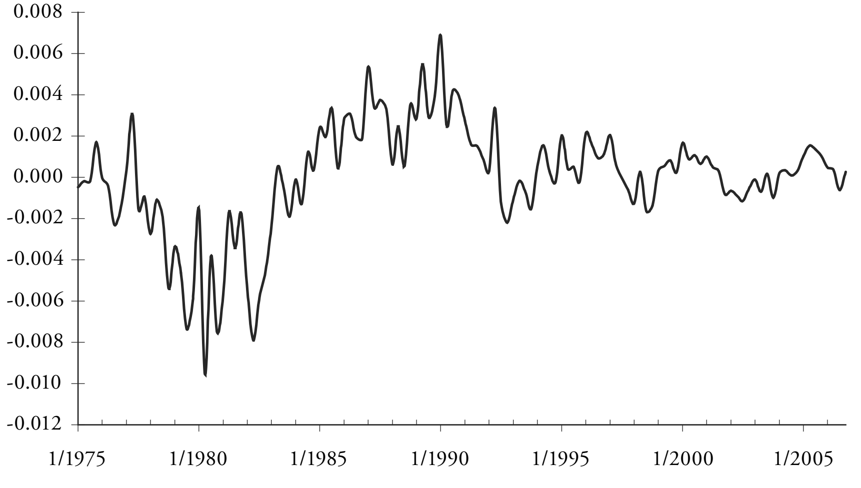


Figure 6 (continued)

Fisher Parity



Uncovered Interest Parity



5. A Structural Model for the Short Run

After having identified and checked the long-run structure of the model, we now proceed to short-run identification. To compute impulse responses to a monetary policy shock in the cointegrated VAR model with exogenous $I(1)$ variables, the conditional model for $\Delta\mathbf{x}_t$ in equation (6) together with the marginal model for $\Delta\mathbf{x}_t^*$ in equation (7) is required.[14]

In a cointegrated VAR model two possible approaches to the identification of structural shocks exist. First, we can identify a short-run structure by placing restrictions on the contemporaneous relations in the VAR. Second, the VAR can be identified based on the common trends driving the system by distinguishing between permanent and transitory shocks. In both cases, the residuals of the structural form are assumed to be orthogonal. We will discuss both approaches in turn.

The impulse responses for the cointegrated VAR model are derived from the moving-average representation of equation (3), which can be written as

$$\mathbf{z}_t = \mathbf{C}\mathbf{A}_0^{-1}\sum_{j=1}^{t}\varepsilon_j + \mathbf{C}^*(L)\mathbf{A}_0^{-1}(\varepsilon_t - \varepsilon_0) + \mathbf{c}_0 t + \mathbf{z}_0, \tag{11}$$

where $\mathbf{C} = \beta_\perp(\alpha_\perp'(\mathbf{I}_m - \sum_{i=1}^{p-1}\Gamma_i)\beta_\perp)^{-1}\alpha_\perp')$ describes how the common trends contribute to the variables in $\mathbf{z}_t$, $\beta_\perp$ and $\alpha_\perp$ denote the orthogonal complements of β and α, $\mathbf{c}_0 = \mathbf{C}\mathbf{a}_0 + \mathbf{C}^*\mathbf{a}_1$, and $\mathbf{C}^*(L) = \sum_{j=0}^{\infty}\mathbf{C}_j^* L^j$ is an infinite-order polynomial in the lag operator with coefficient matrices $\mathbf{C}_j^*$ that go to zero as $j \rightarrow \infty$.[15] Given a set of initial values for the random-walk components, $\mathbf{z}_0$, equation (11) decomposes the process $\mathbf{z}_t$ into $m - r$ random-walk components given by $\mathbf{C}\mathbf{A}_0^{-1}\sum_{j=1}^{t}\varepsilon_j$, m stationary components given by $\mathbf{C}^*(L)\mathbf{A}_0^{-1}(\varepsilon_t - \varepsilon_0)$ and m different linear trends, $\mathbf{c}_0 t$.[16] While $\mathbf{C}$ represents the long-run effects of the shocks to the system, the $\mathbf{C}_j^*$'s contain the transitory effects.

14 Though in principle the lag length in the conditional and the marginal model can differ, we choose a lag length of two for the marginal model as well.

15 See Johansen (1995) or Pesaran, Shin and Smith (2000) for a derivation of the moving-average representation for a cointegrated VAR model.

16 There are no quadratic trends in the level moving-average representation since the trend has been restricted to lie in the cointegration space.

Using this structure it is straightforward to split the model into shocks having permanent and shocks having only transitory effects.[17] The permanent shocks then are identified using restrictions on the long-run multipliers of the system, whereas the transitory shocks are identified by interpreting the reduced-form residuals $\mathbf{u}_t$ as functions of the structural shocks, i.e., by imposing restrictions on the matrix $\mathbf{A}_0^{-1}$. Identifying the whole system with long-run restrictions only is not feasible as the existence of r cointegrating relations implies that only $m - r$ shocks can have a permanent impact. Since our model contains seven endogenous variables and five cointegrating relations, only two independent common trends remain and five shocks would have to be identified by restrictions on $\mathbf{A}_0^{-1}$.

Alternatively, one can proceed as in the SVAR case and obtain impulse-response functions for the cointegrated VAR model by restricting $\mathbf{A}_0$ directly, while taking into account that the coefficients in the moving-average representation of the model in equation (11) incorporate the restrictions implied by the (overidentified) cointegrating relations. In this approach, however, there is no necessary distinction between permanent and transitory shocks and it is likely that all shocks will have permanent components (LEVTCHENKOVA, PAGAN and ROBERTSON, 1998, p. 521).

We use the relationship $\mathbf{A}_0\boldsymbol{\Sigma}\mathbf{A}_0' = \boldsymbol{\Omega}$ between the covariance matrix of the reduced form and the structural-form errors. Since this relation has $m \times m$ unknowns and only m equations, we need additional restrictions to be able to determine the matrix A_0. First, splitting up the model into endogenous and exogenous variables helps with short-run identification, as the $\mathbf{A}_0$ matrix can be partitioned into

$$\mathbf{A}_0 = \begin{pmatrix} \mathbf{A}_{x0} & 0 \\ -\boldsymbol{\Lambda} & \mathbf{A}_{x^*0} \end{pmatrix}. \tag{12}$$

In our case, $\mathbf{A}_{x^*0}$ is 1×1 and thus no further restrictions are necessary.[18] Assuming that the structural shocks are uncorrelated provides $m_x(m_x - 1)/2$ restrictions on $\mathbf{A}_{x0}$, which, together with a normalization of the variance of the shocks, leaves another $m_x(m_x - 1)/2$ restrictions to be determined by economic theory.

17 See, e.g., KING, PLOSSER, STOCK and WATSON (1991), VLAAR and SCHUBERTH (1999), BRÜGGEMANN (2003), BREITUNG, BRÜGGEMANN and LÜTKEPOHL (2004) and VLAAR (2004). The decomposition of the stochastic part of $\mathbf{z}_t$, however, is not unique and raises a number of identification problems discussed in LEVTCHENKOVA, PAGAN and ROBERTSON (1998).

18 If additional exogenous variables were present, a recursive identification scheme could be used since the monetary policy shock is not affected by the ordering of the variables in $\mathbf{A}_{x^*0}$.

5.1 Results for the Short-Run Identification

Most of the SVAR literature identifies a monetary policy shock by specifying the central bank's reaction function, which relates the residual from the short-term interest rate equation to the variables that the central bank is assumed to observe during the current period.[19] We follow this literature and assume that the monetary policy reaction function takes the form of a Taylor rule with the central bank responding to current values of output, inflation and the foreign interest rate.[20] In addition, we interpret the residual from the equation for real money as representing a short-run money demand function, relating real money balances to the long-term interest rate and output.[21]

While these two equations receive a structural interpretation, we assign no structural interpretation to the other shocks in the system and use a triangular identification scheme for them. This implies that the other equations affect the reactions of the variables in the system to a monetary policy shock only in terms of their ordering relative to the monetary policy equations. We order r^*, y, and π first, assuming that they do not react to a monetary policy shock during the period, which is common in the literature. The long-term interest rate and the exchange rate are included as information variables that are affected by all other variables during the current period and thus ordered after the monetary policy equations.

Table 6 shows the estimated $\mathbf{A}_0$ matrix. Since many coefficients in the $\mathbf{A}_{x0}$ matrix turned out to be insignificant, they were restricted to zero.[22] The overidentifying restrictions cannot be rejected by a likelihood-ratio test with a p-value of 0.13. The resulting monetary policy equations (with standard errors reported in parentheses) are

$$\underset{}{u^{r^s}} \underset{(0.12)}{-0.54u^{r^*}} \underset{(0.05)}{+0.07u^{\pi}} \underset{(0.02)}{-0.04u^{y}} = \varepsilon^{MP} \tag{13}$$

for the monetary policy shock ε^{MP}, and

$$u^{m-p} \underset{(0.28)}{-0.69u^{y}} \underset{(5.04)}{+19.84u^{r^l}} = \varepsilon^{MD} \tag{14}$$

19 See Christiano, Eichenbaum and Evans (1999).

20 Gerdesmeier, Roffia and Eleftheriou (2006) find that for most countries in the euro area monetary policy is well described by a reaction function of that form.

21 See Leeper and Roush (2003), Kim and Roubini (2000) or Cushman and Zha (1997).

22 For identification of the monetary policy shock only the i^{th} column of $\mathbf{A}_0$ is relevant. The zero restrictions imposed thus affect the estimated monetary policy shock only insofar as they alter the estimate of the coefficients in the respective column.

Table 6: Estimated A_0 Matrix

	r	π	y	$m-p$	r^s	r^l	e
ε^1	1	0	0	0	0	0	0
ε^2	**–0.90** (0.20)	1	0	0	0	0	0
ε^3	**–2.12** (0.54)	0	1	0	0	0	0
ε^{MD}	0	0	**–0.69** (0.28)	1	0	**19.84** (5.04)	0
ε^{MS}	**–0.54** (0.12)	0.07 (0.05)	**–0.04** (0.02)	0	1	0	0
ε^6	0	0	0	**–0.012** (0.005)	**–0.28** (0.03)	1	0
ε^7	0	**1.88** (0.69)	0	0	**2.42** (1.26)	0	1

Note: Asymptotic standard errors in parentheses. A likelihood-ratio test of ten overidentifying restrictions does not reject the restrictions with a test statistic of 15.11 and a p-value of 0.13. Coefficients significant at the 5 percent level are indicated in boldface.

for the money demand shock ε^{MD}. Except for the coefficient on inflation in the monetary policy reaction function, all coefficients are significant and show the expected sign. The monetary policy shock shows a strong reaction with respect to movements in the foreign interest rate. It also reacts positively to short-run increases in real GDP. By contrast, the reaction to inflation shows the wrong sign and the coefficient is insignificant. Short-run money demand reacts negatively to increases in the long-term interest rate and positively to short-run movements in real GDP with a coefficient of less than unity.

5.2 Impulse Responses

Having identified a short-run structure for the model we now can perform impulse-response analysis. Figure 7 plots the impulse responses to a typical monetary policy shock, together with their bootstrapped 90 percent confidence intervals.[23] After a typical monetary policy shock the short-term rate rises by 50 basis

23 Impulse responses are robust to different identification schemes as long as a reaction of monetary policy to contemporaneous inflation is allowed.

points on impact.[24] The interest rate reaction is larger than the 13 basis points Natal (2004) estimates but more in line with the results of Kugler and Rich (2002) and Kugler and Jordan (2004), who obtain an interest-rate reaction of 40 basis points. The long-term interest rate shows a similar pattern as the short rate, but moves by only about 10 basis points.

Inflation falls by about 30 basis points after one quarter and returns to practically zero in the third and fourth quarter. This pattern is generally found in studies for Switzerland and is ascribed to the indexing of housing rents to the short-term interest rate.[25] Thereafter, inflation decreases again and bottoms out after about ten quarters. While inflation returns to zero eventually, the price level is permanently lower after a contractionary monetary policy shock.

Output increases during the first quarters, though not significantly, and starts to fall thereafter. Again, this pattern is also found by Kugler and Rich (2002) and Kugler, Jordan, Lenz and Savioz (2005) in small SVAR systems with different identification restrictions. Like inflation output reaches its minimum around ten quarters after the shock. Real M_2 falls on impact but rises to its previous level after about twelve quarters.

While the reactions of inflation, output and real balances agree with theory, the exchange rate depreciates on impact, so that an exchange-rate puzzle is present despite the inclusion of money in the model. In an overshooting model we would expect the exchange rate to appreciate on impact after a contractionary monetary policy shock. In the long run prices should rise, the interest differential between the domestic and the foreign assets should disappear and the exchange rate should depreciate. We find that a contractionary monetary policy shock leads to a persistent appreciation of the exchange rate in the long run, which is consistent with the price reaction. Nevertheless, uncovered interest parity is violated because the domestic rate increases above the foreign rate (see Figure 7). Together with the appreciation this implies excess returns for the domestic currency.[26] Of the studies for Switzerland only Natal (2003; 2004) includes an exchange rate. He does not obtain an exchange rate puzzle which may be due to the fact that in addition to M_2 he includes credit variables.

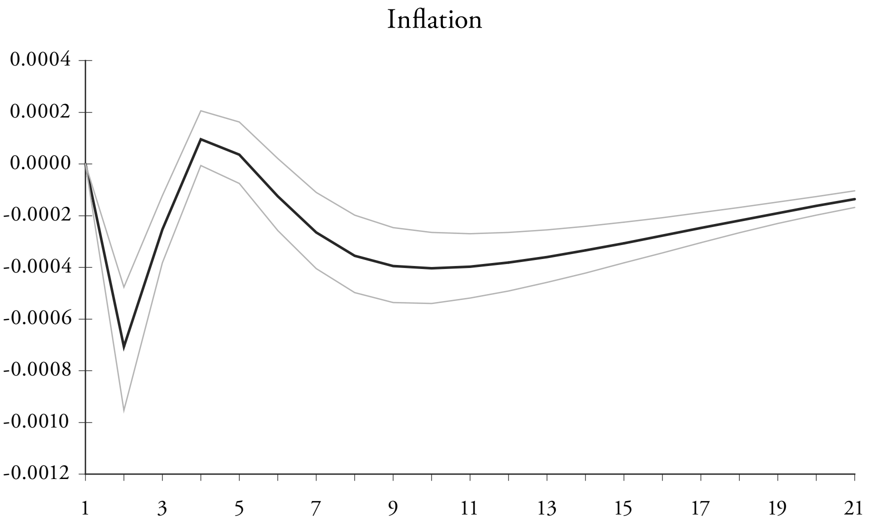
24 For the interpretation of the impulse responses recall that interest rates are expressed as $0.25 \ln(1 + R/100)$ where R is the interest rate in percent per annum.

25 Swiss legislation allows house owners to pass changes in the mortgage rate to tenants. Stalder (2002) estimates that rents increase by 4.5 percent in reaction to a one percentage point increase in the mortgage rate. Since rents have a share of 20 percent in the CPI, prices tend to rise a few quarters after a rise in the interest rate.

26 The same results are found for the US by Eichenbaum and Evans (1995).

Figure 7: Impulse Responses (Black Line) to a Monetary Policy Shock with Boot-strapped 90% Confidence Bounds (Grey Lines)

Confidence bounds are based on 1000 non-parametric bootstrap replications with β and the marginal-model parameters assumed fixed.



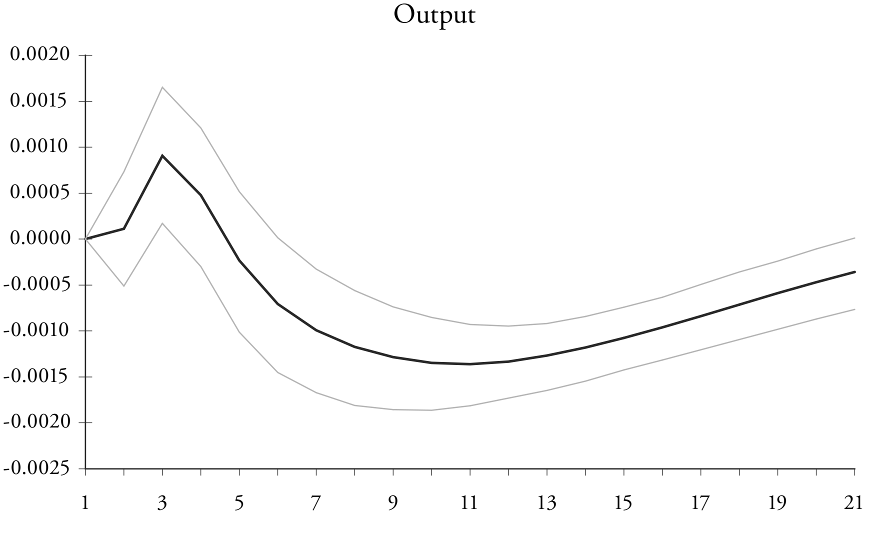
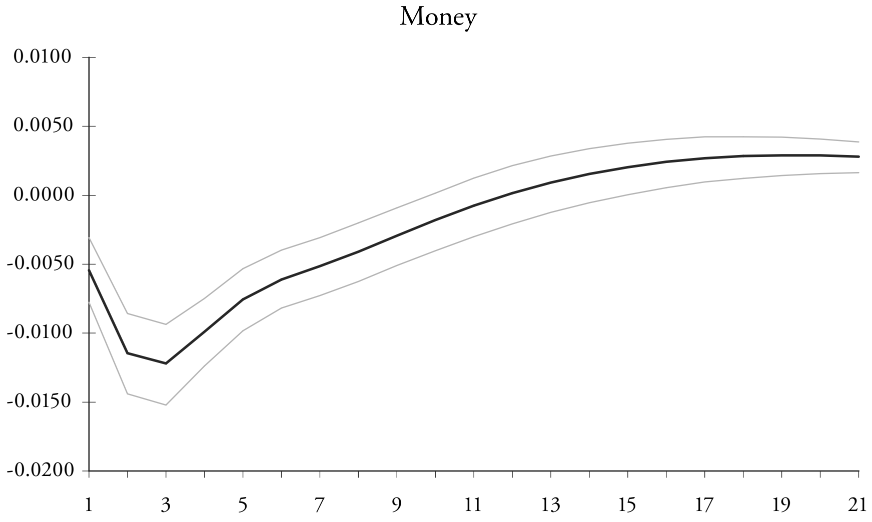


Figure 7 (continued)



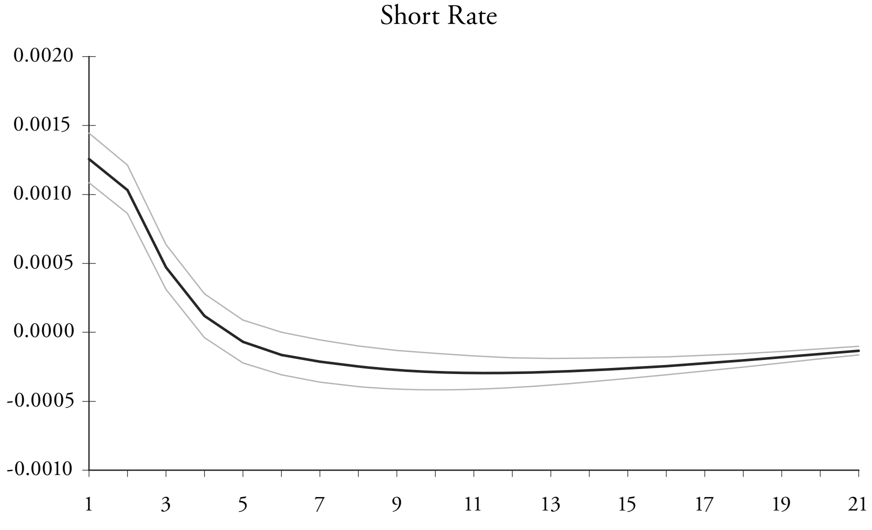
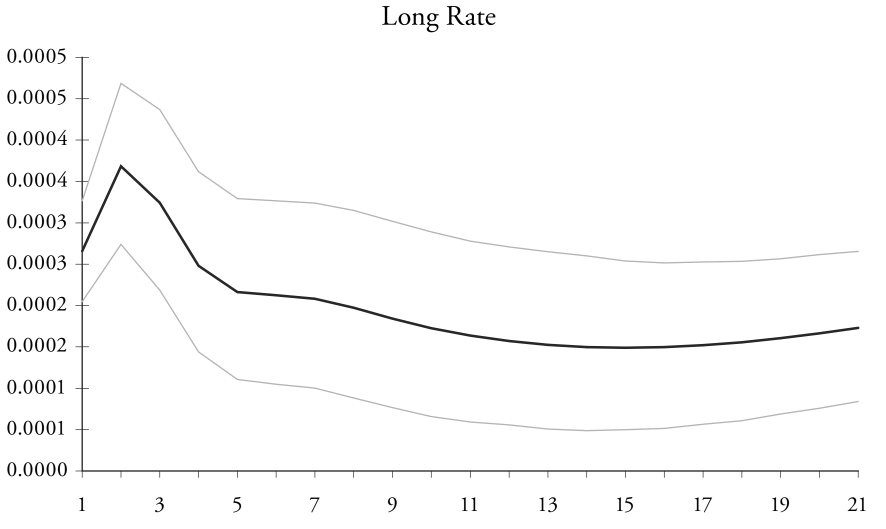


Figure 7 (continued)



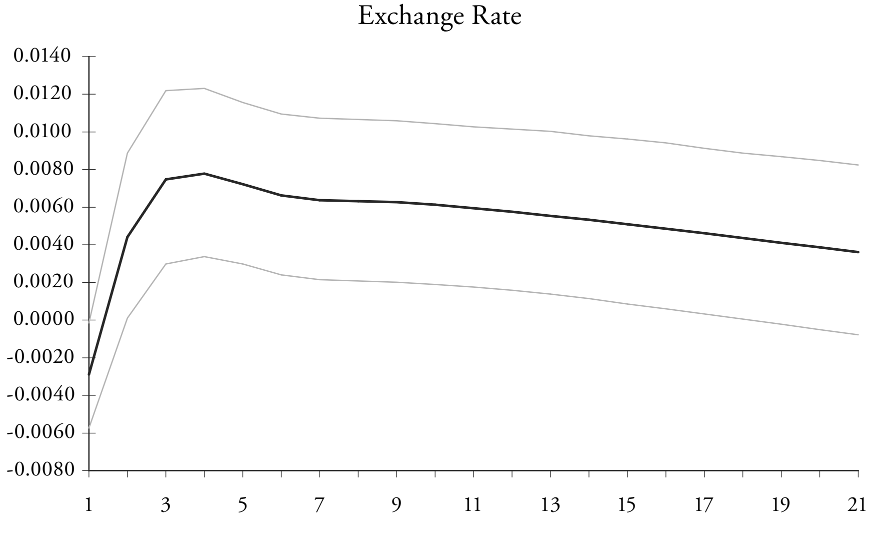
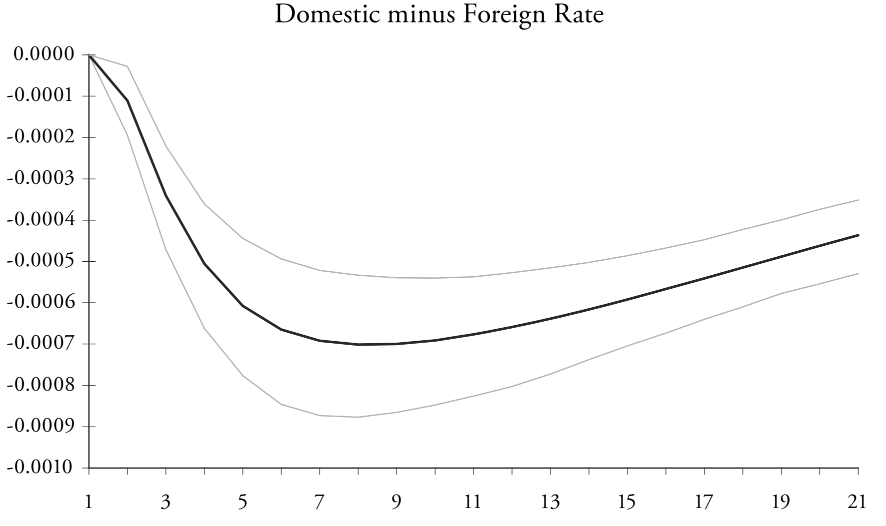
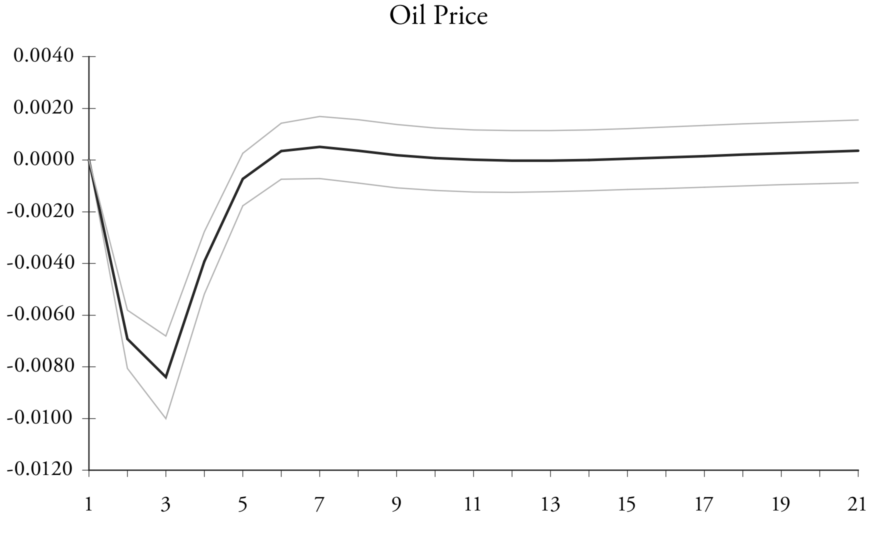


Figure 7 (continued)





6. Conclusions

In this paper we study a cointegrating VAR model of the monetary transmission mechanism in Switzerland, which incorporates inflation, real GDP, real M_2, a short and a long-term interest rate, the exchange rate and the foreign interest rate as endogenous variables and the oil price as exogenous variable. We identify a long-run structure with five cointegrating vectors that are interpreted as a money-demand function, a stationary real interest rate, a stationary term spread, uncovered interest parity and an aggregate-demand schedule. For M_2 a unitary income elasticity of money demand is not rejected. Stability of the model after the introduction of the new monetary policy framework by the SNB is tested and cannot be rejected.

We identify a monetary policy shock by interpreting the contemporaneous interactions between the variables in the system as a money demand relation and a monetary policy reaction function. Impulse responses of inflation, money and output to a monetary policy shock are consistent with theoretical priors. By contrast, we obtain an impact depreciation of the exchange rate after a contractionary monetary policy shock – the so-called exchange rate puzzle – that is often found in the SVAR literature. Overall, the model appears to provide a plausible description of the monetary transmission mechanism in Switzerland while taking account of the openness of the Swiss economy.

A. Appendix: Data

The Swiss data are from the data base of the SNB. The price level is the consumer price index (CPI) with the base of December 2005 = 100. Inflation is measured as the quarterly change in the CPI. For the CPI an adjustment was made to overcome breaks due to new data collection procedures at the Federal Statistical Office. From 2000 on the CPI includes end-of-season sales. This introduces marked seasonality into the sub-index for clothing and footwear, as can be seen in Figure A.1.

In addition, the data collection shifted from the end of the month to the beginning of the month in January 2002, which introduced another break into the series. We adjust for these changes by shifting the series by one month backward between January 2000 and January 2002, the period indicated by the vertical lines in Figure A.1. The resulting missing value is filled by inserting the December 2001 value of the sub-index. The series is smoothed by computing a twelve-month backward moving average. The smoothed sub-index is added to the CPI without clothing and footwear, using the weight of this subindex in the CPI.

Figure A.2 shows inflation computed from the original and the adjusted inflation series. Though the weight of the clothing-and-footwear subindex is less than 5 percent since 2000, it is clearly visible that the adjustment reduces the seasonal variability of the inflation rate since 2002 considerably.

M_2 is in the definition of 1995 (excluding Liechtenstein) and includes cash, sight deposits and savings deposits. Real money is calculated by deflating M_2 with the adjusted CPI. Output is the seasonally adjusted quarterly real gross domestic product (GDP) computed by the SECO (Secrétariat d'état à l'économie) from 1981 on. Quarterly output estimates before 1981 were interpolated from the official annual data by the SNB. The short-term interest rate is the nominal end-of-month three-month London Interbank Offered Rate (LIBOR) for Swiss francs, which is the operating target rate for the SNB. The long-term interest rate is the nominal yield on ten-year government bonds. The foreign interest rate is the nominal three-month money market rate for Germany until 1998 and the nominal three-month EURIBOR rate thereafter. All interest rates are expressed as $0.25\ln(1 + R/100)$, where R is the interest rate measured in percent per annum, so that they are in the same unit of measurement as the quarterly inflation rate.

The effective nominal exchange rate is a export-weighted index of the Swiss franc, calculated and published by the SNB, that includes Switzerland's 24 major trading partners.[27] The oil price is the price of crude oil (Brent) in US dollar. Monthly data for real money, the CPI, interest rates, the exchange rate and the oil price are aggregated into quarterly averages of monthly figures. Inflation is the quarterly difference of the logarithm of CPI.

27 See Fluri and Müller (2001).

Figure A.1: Price Index for Clothing and Footwear

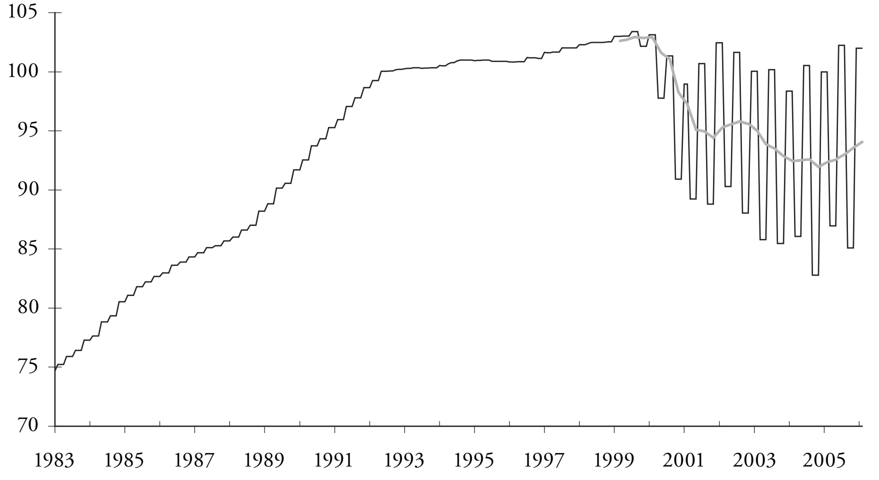
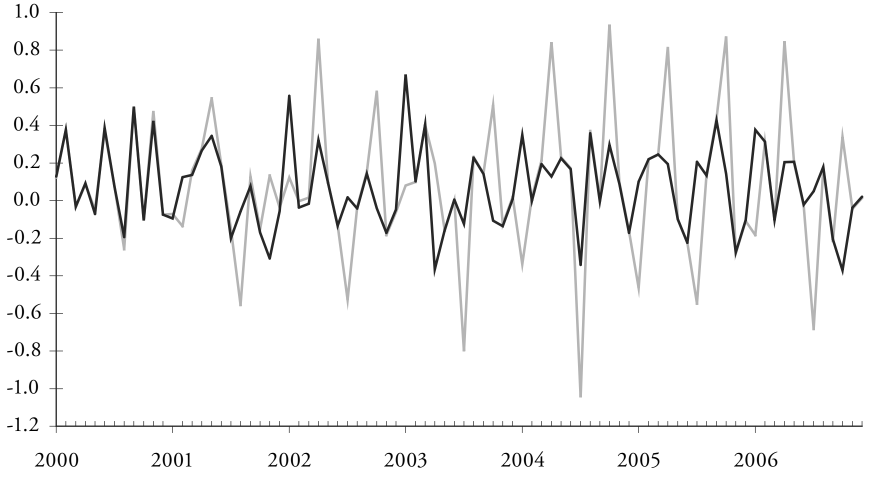


Figure A.2: Monthly Inflation Rate without (Black Line) and with Adjustment (Grey Line) of the CPI



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SUMMARY

This paper examines the transmission of monetary policy in Switzerland using a structural cointegrated VAR model that includes real money, real output, a long and short-term interest rate, inflation, the exchange rate and a foreign interest rate as endogenous variables and oil prices as exogenous variables. The model takes account of five cointegrating relations that are interpreted as money demand, the real interest rate, the term spread, uncovered interest parity and an aggregate demand schedule. Recursive analysis confirms that the model remains stable after the adoption of a new monetary policy framework of the Swiss National Bank in 2000. After identifying a monetary policy shock, the model is used for impulse-response analysis. We obtain plausible responses of inflation and output to a monetary policy shock but despite the inclusion of money and oil prices an exchange rate puzzle remains present.