A VAR MODEL FOR MONETARY POLICY ANALYSIS IN A SMALL OPEN ECONOMY

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ABSTRACT: The interest in empirical studies of monetary policy has increased in the last decade. The deregulation of financial markets and the increased use of explicit policy rules and targets have made monetary policy more transparent and interesting for economic analysis. This paper demonstrates how a VAR model with long run restrictions justified by economic theory can be usefully applied in analyses of issues central to monetary policy: the effects of innovations in interest rates and other shocks; the short and long run relationships between prices and nominal and real exchange rates; the properties of an index of monetary conditions; dynamic forecasts of inflation; and the relation between inflation and the output gap.

KEYWORDS: Cointegration, common stochastic trends, monetary policy, vector autoregressions.

JEL CLASSIFICATION NUMBERS: C32, C52, C53, E31, E52.

1. INTRODUCTION

Vector autoregressive (VAR) models have been much used in empirical studies of macroeconomic issues since they were launched for such purposes by Sims (1980). They are now widely used in all kinds of empirical macroeconomic studies, from relatively atheoretical exercises such as data description and forecasting, to tests of fully specified economic models.[1] The purpose of this paper is to show how a VAR approach can be usefully applied in analyses of issues which are central to monetary policy in a small open economy under an inflation targeting regime.

The interest in empirical studies of monetary policy has increased in the last decade, possibly for the following two reasons. First, financial markets have been deregulated and monetary

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[1]In a study of U.S. monetary policy by Rotemberg and Woodford (1997), VAR models are used both for data description and as input in quantitative theoretical analyses. More generally, VAR models are often the starting point in estimations of so called error correction models, which have proven to have good properties as empirical models of macroeconomic time series; see e.g. Hendry and Mizon (1993) and Johansen and Juselius (1994).

policy therefore more oriented towards open market operations than regulatory measures. Second, monetary policy in many — especially small and relatively open — economies has been increasingly and more explicitly based on policy rules and monetary targets. Explicit inflation targets are for example used in Australia (since 1993), Canada (1991), Finland (1993), Israel (1991), New Zealand (1990), Spain (1995), Sweden (1993), and the U.K. (1992).[2] These developments have made monetary policy more transparent and more interesting for economic analyses.

A stylized picture of the monetary policy process in a country with an inflation target may look something like the following. Official central bank inflation forecasts are presented to the public rather infrequently (e.g. in quarterly "inflation reports" as in Sweden and in the U.K.). On these occasions attempts are made to measure and justify the overall stance of monetary policy, considering not only the development of inflation, but also other variables such as interest rates, the nominal exchange rate, indexes of "monetary conditions" (weighted averages of exchange and interest rates), and the "output gap" (the difference between actual and potential GDP). This work is partly based on econometric and statistical models, but the forecasters' judgments also play a central role. In the periods between the publications of the inflation forecasts, the development of various indicators — such as the output gap and the nominal exchange rate — are almost continuously used to update the forecasts and guide monetary policy.

In the process of conducting monetary policy analysis, central bank economists are faced with a number of empirical questions. Does the nominal exchange rate help to predict inflation? Does the nominal exchange rate adjust in response to the difference between domestic and foreign inflation, to restore some equilibrium level of the real exchange rate? How useful are various measures of the output gap and of monetary conditions? How fast do changes in monetary policy affect output and inflation? These questions concern complex relations between variables which are all endogenously and simultaneously determined in the economic system. It is hardly surprising that a common procedure is to develop partial models (or simple rules of thumb) which are intended to handle these questions one at a time, i.e. to analyze each issue under a set of *ceteris paribus* assumptions. The risk with such an approach is that the different partial models have properties that make them inconsistent with one another. If so, they provide a shaky ground for policy analysis.

[2]The experiences in these and other countries are discussed in Haldane (1995) and Leiderman and Svensson (1995).

We do certainly not expect that there is any single model that can provide the best possible answers to all relevant questions in the analysis of monetary policy, or that it yields exactly the same answers when estimated for different time periods. Nevertheless, we believe that it is necessary — not the least in order to make monetary policy transparent — to develop one, say one, model which offers a consistent framework for studying the above questions. Such an omnibus model can be used as a benchmark for partial models, i.e. models that are question specific in design, as well as for the informal analyses that have to be made frequently in the day to day operation of a central bank. In addition, it can provide useful information about aggregate relationships that are used in theoretical analyses of inflation targeting. We intend to show that a VAR model can serve these purposes.

While many previously used inflation forecasting models depend on exogenous variables, the VAR approach endogenously determines all the variables which make up the system. This means that it permits us to compute multi-step forecasts for each variable. Obviously, a model which can only be used to generate 1-step ahead forecasts is of limited use for monetary policy analysis.

Earlier VAR studies have in many cases been concerned with measuring monetary policy and its macroeconomic effects. See e.g. Gordon and Leeper (1994), Christiano, Eichenbaum, and Evans (1996), Leeper, Sims, and Zha (1996), and Bernanke and Mihov (1998) for studies of the U.S., and Cushman and Zha (1997) for a study of Canada and further references.[3] In contrast to these earlier VAR studies, the identification of a reaction function for monetary policy is not an issue in our paper. The reason is partly that we are interested in a broad set of questions, relevant for monetary policy, but not all directly related to the effects of changes in monetary policy. Another reason is that we expect it to be difficult to find a sufficiently long series of observations on a single policy instrument.

The rest of the paper is organized as follows. In Section 2 the VAR framework is presented. The empirical analysis, using Swedish data, begins in Section 3 with specification tests and tests of long run (cointegration) relationships. Throughout, asymptotic tests are augmented by parametric bootstrap analogues in order to make inference more reliable. In Section 4 we present identifying assumptions that make it possible to interpret the VAR model's residuals in terms of underlying structural shocks. In Section 5 we use our framework to shed light on a number of issues which are central in monetary policy analysis. Section 6 contains a summary and suggests some lessons.

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[3]Rudebusch (1996) criticizes VAR analyses of monetary policy, and Sims (1996) responds to this criticism. Pagan and Robertson (1995) review VAR analyses of monetary policy.

2. The VAR Model

2.1. *VAR Models with Cointegrated Variables*

A VAR model for x_t, a vector of n observable variables, may generally be written

$$\Pi(L)x_t = \delta_0 + \delta_1 D_t + \varepsilon_t, \tag{1}$$

where ε_t, a vector of residuals, is assumed to be i.i.d. Gaussian with zero mean and positive definite covariance matrix Σ. $\Pi(\lambda)$ is an $n \times n$ matrix polynomial of order p and defined by $\Pi(\lambda) = I_n - \sum_{j=1}^{p} \Pi_j \lambda^j$, where λ is a complex number and L is the lag operator such that $L^j x_t = x_{t-j}$. Furthermore, D_t is a vector of d observable deterministic variables. The VAR model may be viewed as the reduced form of an underlying structural model. What distinguishes the VAR from a structural model is that the latter is derived from economic theory, which usually implies a large number of restrictions on parameters; restrictions which are not imposed on the VAR model when it is estimated from macroeconomic data.

Macroeconomic time series are often characterized by a high degree of persistence. Frequently, the persistence is well described by a so called unit root process, e.g. a random walk. In such cases, at least some shocks have permanent effects on x_t (i.e. x_t is nonstationary) and standard asymptotic results may not be applicable. However, often one finds that changes in x_t, denoted by $\Delta x_t = (1-L)x_t$, are stationary (x_t is integrated of order one, $\mathsf{I}(1)$) and also that certain linear combinations of the variables in x_t are stationary (x_t is cointegrated, $\mathsf{CI}(1,1)$). If so, the VAR model may be rewritten as a so called vector error correction (VEC) model

$$\Gamma(L)\Delta x_t = \delta_0 + \delta_1 D_t - \alpha\left(\beta' x_{t-1}\right) + \varepsilon_t, \tag{2}$$

where $\Gamma(\lambda) = I_n - \sum_{i=1}^{p-1} \Gamma_i \lambda^i$, $\Gamma_i = -\sum_{j=i+1}^{p} \Pi_i$, and $\alpha\beta' = \Pi(1)$. The matrices α and β are $n \times r$ with rank equal to r. Let α_{ij} and β_{ij} denote the elements of α and β, respectively. The columns of β, denoted by $\beta_s, s = 1, \ldots, r$, are the so called cointegration vectors.

Since the VEC model is expressed in terms of stationary variables, Δx_t and the linear combinations $\beta' x_t$, standard theory can be used for inference about the parameters in Γ_i and α. Estimates of the cointegration rank r and β can be obtained through the maximum likelihood procedures suggested by Johansen (1988, 1991).

It is sometimes possible to give the cointegration vectors economic interpretations. In our case,

$$x_t = \begin{bmatrix} y_t & p_t & i_t & e_t & y_t^* & p_t^* & i_t^* \end{bmatrix}',$$

where y_t and y_t^* denote domestic (Swedish) and foreign GDP, respectively; p_t and p_t^* domestic and foreign consumer price indexes; i_t and i_t^* domestic and foreign nominal three month interest rates; and e_t the nominal exchange rate (price of foreign currency in terms of domestic currency).[4] If one cointegration vector takes the form

$$\beta_1 = \begin{bmatrix} 0 & -1 & 0 & 1 & 0 & 1 & 0 \end{bmatrix}',$$

the implication is that the real exchange rate, $e_t + p_t^* - p_t$, is stationary. That is, although the nominal exchange rate and the price levels are nonstationary, the relative price between domestic and foreign goods is not.

If the cointegration relations are interpreted as long run equilibrium conditions, then the elements of α have natural economic interpretations as adjustment coefficients. For instance, if the real exchange rate is "undervalued" in the sense that it is above its long run stationary level, we may expect the nominal exchange rate to appreciate ($\alpha_{41} > 0$), the domestic price level to rise ($\alpha_{21} < 0$), and/or the foreign price level to fall ($\alpha_{61} > 0$). Whether the real exchange rate is stationary or not, and whether it can be used to forecast the domestic price level and the nominal exchange rate, are clearly questions that are relevant for monetary policy. Such questions can be formulated as hypotheses about the parameters of the VEC model in (2).

A variable which is integrated of order one (contains a unit root) is also said to have a *stochastic trend* since (some) shocks to it have permanent effects. Variables which are cointegrated are analogously said to have *common stochastic trends*. If a vector of n variables has r cointegration relations, the variables are driven by $k = n - r$ stochastic trends (Stock and Watson, 1988). If the variables in the VAR model are found to be cointegrated, the model can thus be written as a so called common trends model:

$$x_t = A\tau_t + \Phi(L)\varphi_t + \mu_0 t + \mu_1 \sum_{i=1}^{t} D_i + \Theta(L)D_t + \kappa, \tag{3}$$

where τ_t is a k-dimensional vector of (unobservable) stochastic trends,

$$\tau_t = \tau_{t-1} + \psi_t. \tag{4}$$

The factor $\Phi(L)\varphi_t$ is mean zero stationary with $\varphi_t = F^{-1}\varepsilon_t$ and ψ_t being the first k elements of φ_t, whereas μ_i, $i = 0, 1$, and A are orthogonal to β'.[5]

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[4]All variables are expressed in terms of natural logarithms. The data are quarterly and from the period 1972:2–1996:4. More details are given in the Data Appendix.

[5]Accordingly, the cointegration relations can be expressed as $\beta' x_t = \beta'\Phi(L)\varphi_t + \beta'\Theta(L)D_t + \beta'\kappa$. This r dimensional stochastic process is stationary around the possibly time varying mean $\beta'(\kappa + \Theta(L)D_t)$.

2.2. *Hypotheses about Cointegration*

With the seven variables in our data set, how many common stochastic trends do we expect to find? In equilibrium business cycle models, production is often assumed to be driven by stochastic shocks to technology, while the price level is also affected by stochastic shocks to the money supply (see e.g. Cooley and Hansen, 1995). Empirically, technology (the "Solow residual") appears to be well described by a random walk with drift. Given the secular rise in the price levels in most countries during the last half century, it is also reasonable to assume the existence of a nominal stochastic trend, determined by the design of monetary policy. Accordingly, if the domestic and foreign real and nominal stochastic trends are independent, then we expect to find $k \geq 4$, and, hence, $r \leq 3$.[6]

Suppose there are four stochastic trends and hence three cointegration vectors. If there are two stochastic trends (e.g. technology and monetary policy) driving the three foreign variables, y_t^*, p_t^*, and i_t^*, these three variables must be cointegrated. But there is little reason for any of the domestic variables in a small country like Sweden to affect the long run stationary relationship between the foreign variables. Hence we expect to find one cointegration relation between the foreign variables only. One possibility is that i_t^* is stationary because the real interest rate and inflation expectations are stationary, and because of the Fisher relation (see e.g. the equilibrium business cycle model with stochastic trends in output and money supply in Söderlind and Vredin, 1996). If so, we expect i_t to be stationary for the same reasons.

If i_t^* and i_t are nonstationary, this may be because the real exchange rate or inflation expectations are nonstationary. The latter may not be a bad approximation in our sample, since inflation targeting is a relatively recent phenomenon. In any case, financial markets equilibrium implies that the difference between i_t and i_t^* is equal to the expected rate of change of the exchange rate plus, possibly, a risk premium. If the nominal exchange rate is not integrated of an order greater than one, and if the risk premium is zero (i.e. if domestic and foreign securities are perfect substitutes) or stationary, the interest rate differential is stationary. Alternatively, $i_t - i_t^*$ may be cointegrated with variables that affect the risk premium, possibly domestic and foreign output.

Goods market equilibrium may imply that the relative price between domestic and foreign goods $\left(e_t + p_t^* - p_t\right)$, the real exchange rate, is stationary, if domestic and foreign goods are so

[6]The stochastic trend in technology is usually assumed to be unobservable. If the money supply were exogenous, it should be possible to measure the stochastic trend in monetary policy directly from money stock data. We find it more reasonable, however, to treat monetary policy as partly endogenous. We also believe that price level and interest rate data are more informative about monetary policy (at least in Sweden) than some money stock series. See Bernanke and Mihov (1998) and Svensson (1998a) for discussions of related issues.

close substitutes that purchasing power parity holds on average. There are many studies which suggest that PPP does not hold (see e.g. Jacobson and Nessén, 1998), but we may still expect the real exchange rate to be cointegrated with variables that reflect the demand or supply of domestic and foreign goods, such as y_t and y_t^*.

We should thus look for three cointegration vectors which lend themselves to economic interpretations. Two vectors can be interpreted as equilibrium conditions for the goods and financial markets, and these should involve the relative price $\left(e_t + p_t^* - p_t\right)$ and the interest rate differential $i_t - i_t^*$, respectively. The third cointegration relation, interpreted as a world market equilibrium condition, should involve y_t^*, p_t^*, and i_t^* and reflect that these three foreign variables are not driven by three independent stochastic trends. That is, the cointegration vectors should have the following form:

$$\begin{aligned}
\beta_1 &= \begin{bmatrix}\beta_{11} & -1 & \beta_{13} & 1 & \beta_{15} & 1 & \beta_{17}\end{bmatrix}', \\
\beta_2 &= \begin{bmatrix}\beta_{21} & \beta_{22} & 1 & \beta_{24} & \beta_{25} & \beta_{26} & -1\end{bmatrix}', \\
\beta_3 &= \begin{bmatrix}0 & 0 & 0 & 0 & \beta_{35} & \beta_{36} & 1\end{bmatrix}'.
\end{aligned} \tag{5}$$

While the hypothesized vectors in (5) appear to be reasonable from an economic point of view, the second vector is generally not identified. We thus need to specify a set of vectors that satisfies our basic economic reasoning while — at the same time — can be uniquely determined from the data. One such set obtains when the real exchange rate and the two nominal interest rates are stationary, i.e.

$$\begin{aligned}
\beta_1 &= \begin{bmatrix}0 & -1 & 0 & 1 & 0 & 1 & 0\end{bmatrix}', \\
\beta_2 &= \begin{bmatrix}0 & 0 & 1 & 0 & 0 & 0 & 0\end{bmatrix}', \\
\beta_3 &= \begin{bmatrix}0 & 0 & 0 & 0 & 0 & 0 & 1\end{bmatrix}'.
\end{aligned} \tag{6}$$

This hypothesis holds in the theoretical model of inflation targeting by Svensson (1998b).[7]

Let us briefly discuss how a finding of $k \neq 4$ (and hence $r \neq 3$) may be interpreted. $k < 4$ may be consistent with a common trend in domestic and foreign technology and/or in domestic and foreign monetary policy. In the former case the additional cointegration vector could be associated with a stationary relationship between domestic and foreign GDP, and in the latter it might be the case that the nominal exchange rate is stationary. A finding that $k > 4$, on

[7] The reason why the hypothesis (6) is a special case of (5) is due to the fact that the space spanned by the vectors in (6) is equivalent to the space spanned by the vectors in (5) when all β_{ij} are equal to zero.

the other hand, implies that there are other shocks with permanent effects than innovations in technology and nominal trends. This is a feature of e.g. models with “bubbles” or multiple equilibria, where exogenous changes in expectations can become self fulfilling. The additional stochastic trend may thus reflect shocks to expectations. In that case it may not be possible to find a stationary financial markets equilibrium condition (or even a goods market equilibrium).

2.3. *Identifying Structural Shocks*

For forecasting purposes it is often sufficient to analyze the unrestricted VAR model in (1). VEC models like (2), however, often turn out to have superior empirical properties, possibly because they make use of economic theory for finding meaningful long run relationships, while the short run adjustment is left relatively unrestricted. For some purposes one may want to go further and identify the shocks to the common stochastic trends and their effects on the variables in the system.

The residuals, ε_t, in the VAR model can be seen as linear combinations of the structural shocks. We can identify at most n such shocks, which we denote by φ_t. Having k shocks ψ_t, with permanent effects on the variables in x_t (cf. (3)–(4)), there are $n-k=r$ other structural shocks with only transitory effects. Let these be contained in the vector ϕ_t, so that

$$\varepsilon_t = F\varphi_t = F\begin{bmatrix}\psi_t \\ \phi_t\end{bmatrix}. \tag{7}$$

Examples of transitory shocks are changes in aggregate demand that affect the cyclical evolution of the macroeconomy, but not its long run development.

In general, we cannot identify the structural shocks, even if we do know the cointegration vectors, unless we impose restrictions on the common trends model.[8] To separate the real (technology) trends from the nominal trends we may, for instance, assume that nominal shocks have no long run effect on real output. And, in order to separate the domestic trends from the foreign, we may assume that the former have no long run effect on foreign variables (the domestic economy being a small open economy). Similarly, we would presumably like to distinguish transitory shocks to domestic aggregate demand from shocks to foreign aggregate demand, e.g. by assuming that domestic demand shocks have no impact on foreign variables.

[8] Since the matrix F in (7) has n^2 parameters, it follows that n^2 restrictions have to be imposed on the parameters in (3). Whenever $k<n$, some of these restrictions are imposed via the assumption that A is $n\times k$ rather than $n\times n$. For details, the reader is referred to e.g. Warne (1993) and Englund, Vredin, and Warne (1994).

We will explain, implement, and evaluate identifying assumptions like these in Section 4 below. For the moment, let us just emphasize that the set of structural shocks and the associated identifying assumptions are closely related to assumptions which have been employed in previous analyses of monetary policy. For instance, the Svensson (1998b) model includes aggregate supply shocks, which in our model are labeled technology shocks, and these shocks can be distinguished from monetary policy shocks and from other shocks to aggregate demand.

3. An Empirical VAR Model with Cointegrated Variables

3.1. *Specification Analysis*

Prior to estimating lag length parameter in (1) and the cointegration rank of x_t, it is useful to consider and discuss the possibility that some of the nonstationary features in our system are due to deterministic breaks and regime shifts. If this is the case, a valid cointegration analysis requires conditioning on the nonstationary influence of such deterministic variables.

Figure 1 shows the time plots of the 7 endogenous variables being modeled. The domestic and foreign output series, y_t and y_t^*, are both characterized by strong seasonal patterns, whereas the remaining five series appear not to be affected by seasonality. Thus, rather than including a full set of seasonal dummies in the VAR model we have chosen to seasonally adjust the two output series only. This is done through regression analysis on seasonal dummy variables prior to the system estimation.

The domestic and foreign price variables, p_t and p_t^*, display the usual strong upward trends typical for price level variables. We note that there appear to be shifts in the trends of the price levels during the 1980s and 1990s. These may very well reflect changes arising from regime shifts in economic policy.

Economic policies directed towards low inflation were introduced in most Western European countries and in the U.S. already in the beginning of the 1980s, but it was not until the early 1990s that Sweden started to seriously pursue a less Keynesian (accommodating) stabilization policy. During the 1970s and 1980s, the Swedish krona was devalued on several occasions: 1973:1, 1976:4, 1977:2, 1977:3, 1981:3, and 1982:4. These devaluations are clearly born out in Figure 1, where the series for the nominal exchange rate, e_t, displays jumps at each of these dates. The two largest devaluations occurred in 1981:3 and 1982:4. However, the most dramatic change in the exchange rate series occurs when Sveriges Riksbank (the central bank of Sweden) abandoned the pegged exchange rate in favor of a floating exchange rate in 1992:4.

On the above grounds, we include the following set of deterministic variables:

$$D_{i,t} = \begin{cases} 1 & \text{if } t \in I_i,\ i = 1, 2, \ldots, 5, \\ 0 & \text{otherwise,} \end{cases}$$

where $I_1 = \{1981{:}1\}$, $I_2 = \{1981{:}3\}$, $I_3 = \{1982{:}4\}$, $I_4 = \{1979{:}2, \ldots, 1996{:}4\}$, and $I_5 = \{1992{:}4, \ldots, 1996{:}4\}$. In the terminology of Perron (1989), Dummy 1–3 represent "crashes", whereas Dummy 4 and 5 represent "changes in growth". We interpret the latter two as capturing the regime shifts in economic policy in foreign countries and Sweden, respectively. The timing is quite straightforward in the case of Sweden; Dummy 5 is associated with the floating exchange rate (inflation targeting) regime. The exact date of the foreign regime shift is harder to pinpoint, but events around the second oil crisis possibly acted as catalysts for shifts towards low inflation policies in many countries; hence the dating of Dummy 4.

It may be questioned, of course, whether a linear VAR model with constant parameters provides an adequate approximation of an economy that experiences different policy regimes. In practise, however, we find 5 dummy variables to be sufficient to overcome the specification problems we encounter. For example, non-normality problems in the exchange rate equation disappear when we include Dummy 2, 3, and 5.

Having settled on the conditioning set, we will now move on to the determination of the lag length of the VAR. This is usually done by checking the agreement between the properties of the estimated residuals and the assumption of independent and identically distributed normal zero mean errors in (1). Instead we propose to examine the residuals of the VEC model (2) for the following reasons:

(i) The specification tests employed are valid for, strictly speaking, stationary processes only.

(ii) The reference distributions for the specification tests are asymptotic and, hence, may constitute poor approximations to the small sample distributions required for reliable inference. In order to circumvent this problem we evaluate the residuals using parametrically bootstrapped versions of the asymptotic specification tests. But the bootstrap procedure also requires stationarity to be valid, and is thus better suited for the VEC rather than the VAR.

Panels A–C in Table 1 present misspecification analyses for $p \in \{2, 3, \ldots, 6\}$ and $r \in \{1, 2, \ldots, 5\}$ with respect to multivariate serial correlation, multivariate normality, and multivariate autoregressive conditional heteroscedasticity. Inference is given in terms of asymptotic as well as bootstrapped p-values.[9] Apart from providing robust inference, there is further interesting information to gain from the bootstrap tests; namely an evaluation of the size properties of the corresponding asymptotic tests. For instance, we see that the two sets of p-values for the multivariate ARCH test (Table 1 C) by and large coincide. On the other hand, the multivariate Portmanteau test (Table 1 A) yields extremely poor asymptotic inference. The overall impression is that all specifications involving $p = 4, 5, 6$ are acceptable, with some reservation for the normality tests when the rank r is set to 1 and 2 in the $p = 4$ case. The principle of parsimony suggests that 4 lags is the appropriate choice.

3.2. *Cointegration Tests*

The cointegration test results are reported in Table 2, and again we have estimated small sample reference distributions by bootstrapping. The first two panels show asymptotic critical values. In Panel A, where the effects on the critical values from the two regime shift dummies ($D_{4,t}$ and $D_{5,t}$) are ignored, we infer a cointegration rank of $r = 5$. In Panel B, where we have simulated asymptotic critical values with respect to the effects from the two regime shift dummies,[10] we find a rank of $r = 6$. Turning to Panel C, with small sample critical values, there is evidence of at least three cointegration vectors, but not more than four.

We have theoretical reasons to believe that there are at least 3 cointegration vectors and there are a number of empirical results which suggest that $r = 3$ is a reasonable choice. First, when evaluating the forecasting performance (Section 5.1 below) we find that a model with three vectors does fairly well. In particular, the three hypothesized cointegration vectors in (6) seem to work quite well. Second, the results from the univariate specification tests (Table 3) suggest that the model with $p = 4, r = 3$ has reasonable properties. Third, with four cointegration vectors, the fluctuations around the three stochastic trends do not look stationary.[11] These results, together with our theoretical reasons as laid out in Section 2.2, have led us to base the remaining analyses on a cointegration rank of $r = 3$.

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[9] The idea in bootstrap hypothesis testing is to estimate a reference distribution (critical values). This is done by first generating a large number of pseudo samples with the null hypothesis in question imposed. The next step is to evaluate the test function in each pseudo sample and then arrange the results in ascending order. We generate the pseudo samples by a parametric procedure where we substitute the parameters of the VEC with the estimates from the original sample and feed in pseudo random vectors, generated as $N(0, \widehat{\Sigma})$, in place of ε_t. All bootstrap results in this paper involve 100,000 generated pseudo samples.

[10] The remaining dummy variables do not affect the asymptotic distributions of the LR trace statistics.

[11] Graphs of the so called transitory components for $r = 3, 4$ are available on request. So are the results from parameter stability tests of the VEC model under $r = 3$.

3.3. *Cointegration Relations*

In (8) below we report the estimated cointegration vectors given a particular set of exactly identifying restrictions. The restrictions, implying a certain rotation of the ML estimator of β, have been selected in order to facilitate a comparison with the hypothesized vectors in Section 2.2. The first relation is interpreted as a goods market equilibrium condition and the second is related to a financial markets equilibrium condition. The third relation is intended to capture common trends and equilibrium conditions between the foreign variables, but the (apparent) nonzero coefficients on the domestic interest rate and the exchange rate make such an interpretation troublesome. The general impression is still that the estimated cointegration vectors in (8) are not widely at odds with the theoretical arguments in Section 2.2.

$$\begin{aligned}
\widehat{\beta}_1 &= \begin{bmatrix} -1.044 & -1 & -2.875 & 1 & 0.602 & 1 & 0.052 \end{bmatrix}' \\
\widehat{\beta}_2 &= \begin{bmatrix} 0.243 & 0.490 & 1 & 0.015 & 0 & -0.754 & -1 \end{bmatrix}' \\
\widehat{\beta}_3 &= \begin{bmatrix} 0 & 0 & -0.363 & -0.400 & -0.300 & .318 & 1 \end{bmatrix}'
\end{aligned} \tag{8}$$

It is natural to proceed by examining the overidentifying restrictions implied by the three theoretical vectors in (6). In earlier work (see e.g. Jacobson, Vredin, and Warne, 1998) we have found that the LR test for hypotheses about the cointegration vectors, for a given rank, is not well approximated by the limiting χ^2 distribution. In fact, it often tends to be seriously oversized. Therefore, we will also estimate the small sample distributions with parametric bootstrapping.[12]

Based on the limiting χ^2 distribution with 12 degrees of freedom, the LR test statistic of 110.3 firmly rejects (6). The following bootstrap percentiles were obtained:

	80%	90%	95%	97.5%	99%
$\chi^2(12)$	15.8	18.5	21.0	23.3	26.2
bootstrap	41.5	48.1	53.8	59.2	65.7

Hence, we may conclude that although the likelihood ratio test is indeed seriously oversized, inference does not change when using the bootstrap test.

In summary, we have found support for our hypothesis that the variables in x_t can be characterized by a VEC model like (2). This implies that they are driven by a reduced number

[12] Gredenhoff and Jacobson (1998) use response surface regressions to examine the size distortion problem as a function of the sample size and the complexity of the model, represented by the dimension of the system, the lag order, and the cointegration rank. The paper also evaluates the size properties of a bootstrap version of the test. It is found that for reasonable sample sizes, such as the one at hand, the bootstrapped test has an almost α exact size.

of stochastic trends and therefore are tied together in the long run. There are three long run, cointegration, relations, which can be given reasonable economic interpretations, although the specific hypothesis about these equilibrium relations in (6) is rejected.

4. A Common Trends Model

In discussions and analyses of monetary policy, the following type of questions are often raised: What is the effect on inflation from a depreciation of the nominal exchange rate? What happens to output and inflation if the domestic interest rate goes up? If all these variables are endogenously determined, such questions are, strictly speaking, not well defined. At the general level, it is useful to think of inflation and output as being influenced by many different shocks. The comovements between these variables on the one hand, and variables such as exchange and interest rates on the other, are determined by which particular shock the economy has been predominantly affected by. In order to identify the effects of various shocks we put restrictions on the parameters of a structural VEC model. These restrictions are consistent with assumptions commonly applied in theoretical analyses, such as the "small open economy" and the "monetary neutrality" hypotheses.

4.1. *Identification*

The cointegration analysis suggests that there are four common stochastic trends. As simply matters of notation, let us call the first stochastic trend a foreign real trend, the second a foreign nominal trend, the third a domestic real trend, and the fourth a domestic nominal trend.

As discussed in Section 2.2 above, the existence of real and nominal stochastic trends is consistent with equilibrium business cycle theory. Shocks to the real trends give rise to "Solow residuals". The nominal trends derive from monetary policy. It would probably be incorrect, however, to infer that the innovations to the nominal trends in our framework are entirely (or maybe even mostly) due to unexpected changes in monetary policy. Nominal prices and exchange rates may also change permanently because of other disturbances, such as wage shocks. It is the design of monetary policy, however, which permits such shocks to have permanent nominal effects (i.e. inflation is a "monetary phenomenon").[13]

It seems natural to associate innovations in interest rates with unexpected changes in monetary policy, although we are aware that interest rate shocks also have other sources (that may have different effects compared to monetary policy shocks). Bearing this in mind, we will label one of our three transitory shocks a domestic interest rate shock and another a foreign interest

[13]Crowder, Hoffman, and Rasche (1998) make a similar interpretation of the nominal trend in their study of U.S. data.

rate shock. The third transitory shock is just assumed to represent "other aggregate demand" shocks.

In order to estimate the effects of the structural shocks φ_t, i.e. to estimate the parameters of A and F, denoted a_{ij} and f_{ij}, in the common trends model (3), we can e.g. make the following identifying assumptions.[14]

(A1) The structural shocks are independent, and their variances are normalized to unity, i.e. $\mathrm{E}[\varphi_t \varphi_t'] = I_n$.

Intuitively it makes sense to assume that there are some innovations to nominal variables which are independent of, for instance, shocks to technology (and vice versa). This assumption is consistent with theoretical models (e.g. Cooley and Hansen, 1995, and Svensson, 1998b). The normalization to unit variances is simply for convenience and with no loss of generality.

In order to separate the effects due to the stochastic trends we have to impose (at least) $k(k-1)/2 = 6$ restrictions on A (see e.g. Englund et al., 1994). The following restrictions naturally come to mind:

(A2) Sweden is a small open economy; domestic technology and nominal trend shocks have no long run effects on the foreign variables, i.e. $a_{53} = a_{54} = a_{63} = a_{64} = a_{73} = a_{74} = 0$.

(A3) Long run monetary neutrality; nominal trend shocks do not affect GDP in the long run, i.e. $a_{12} = a_{14} = a_{52} = a_{54} = 0$.

To identify the transitory shocks, we have to impose (at least) $r(r-1)/2 = 3$ additional restrictions, e.g. on their contemporaneous effects, as is usually done in VAR analyses. Following many earlier VAR studies, we could assume that

(A4) Domestic interest rate shocks have no immediate effects on domestic and foreign output, and, in addition, no immediate effect on the foreign price level and interest rate, i.e. $f_{15} = f_{55} = f_{65} = f_{75} = 0$.

(A5) Foreign interest rate shocks have no immediate effects on domestic and foreign output, i.e. $f_{16} = f_{56} = 0$.

The assumption that interest rate shocks have no immediate real effects has — in VAR analyses of U.S. data — proved to be an identifying assumption which makes the effects of interest rate

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[14] For some examples of theoretical problems due to the assumption that the structural shocks are linear combinations of a (fairly low dimensional) vector of VAR innovations, see e.g. Lippi and Reichlin (1993) and Faust and Leeper (1997). To deal with the inference problem due to long run identifying restrictions, which is one of the problems that Faust and Leeper discuss, we rely on the assumption that 4 lags is sufficient to describe the serial correlation pattern in our data.

shocks look like ones prejudices about monetary policy shocks. (A4) is just an extension to the case of a small open economy.

Taken together, the restrictions (A1)–(A5) are overidentifying, which means that some of them are testable (given the others). We have thus chosen to impose, in addition to (A1), only the following exactly identifying restrictions:[15]

(A2′) $a_{53} = a_{54} = a_{63} = a_{64} = 0$.

(A3′) $a_{14} = a_{52} = a_{54} = 0$.

(A4′) $f_{15} = f_{55} = 0$.

(A5′) $f_{56} = 0$.

Since identifying assumptions are always crucial and controversial, some further discussion of our restrictions is warranted. First, we want to emphasize that the labels on the shocks are only meant to be indicative. We find it meaningful to separate permanent shocks from transitory, real shocks from nominal, and interest rate shocks from other transitory shocks, although it must be recognized that each of these categories contains a large number of unobserved underlying shocks to production possibilities, economic behavior, and policy. For instance, the domestic interest rate shock presumably comprises not only domestic monetary policy shocks, but also other sources of innovations in interest rates which are not contemporaneously correlated with innovations in output. Here we may find exogenous changes in expectations or risk premia, i.e. some kind of “credibility” shocks.[16] Second, it should be noted that given a cointegration rank of $r = 3$, the restrictions (A2′)–(A5′) are sufficient to distinguish all structural shocks. We do not have to restrict the three cointegration vectors. However, since the cointegration vectors (6) are theoretically interesting, the implications of imposing also these long run restrictions will be examined. Finally, it should be noted that if we were only interested in separating the permanent (real and nominal) shocks from the transitory (aggregate demand and interest rate) shocks, it would be sufficient to know the cointegration rank r, and we would not have to impose any further identifying restrictions.[17]

[15]Note that $a_{54} = 0$ (the domestic nominal trend does not affect foreign output in the long run) follows both from the hypothesis of monetary neutrality and the small open economy hypothesis. (A2′) and (A3′) thus involve only six restrictions.

[16]The Svensson (1998b) model contains exogenous “risk premium” shocks, while Leeper et al. (1996), and Cushman and Zha (1997) identify “information” shocks, and Smets (1996) “exchange rate” shocks.

[17]Note, however, that the maintained assumption that the structural shocks can be decomposed into k permanent and r transitory shocks, already implies kr identifying assumptions on the common trends model. With (A1) yielding $n(n+1)/2$ and (A2′)–(A5′) $k(k-1)/2 + r(r-1)/2$ restrictions, it follows that a total sum of n^2 identifying restrictions have been imposed on the parameters of the common trends model in (3), corresponding to the number of parameters in F. With $C = \beta_\perp \left(\alpha'_\perp \Gamma(1) \beta_\perp\right)^{-1} \alpha'_\perp$ and $\begin{bmatrix} A & 0 \end{bmatrix} = CF$, it follows that the parameters of A are directly related to the parameters of F.

4.2. *Common Trends Results*

The estimated A matrix of the common trends model (3)–(4), based on the restrictions (A1), (A2′)–(A5′), and the theoretical cointegration vectors in (6) is presented in Table 4 A. The following (significant) results are reasonable, given our labels and identifying assumptions.

(i) A positive foreign nominal trend shock raises the foreign price level in the long run. It also raises the domestic price level.
(ii) A positive domestic real shock raises domestic GDP in the long run.
(iii) A positive domestic nominal trend shock raises the domestic price level in the long run and depreciates the domestic currency.

All these results hold also for the model based on the empirical cointegration vectors in (8); see Table 4 B. But the A matrix in the latter model also involves the following (significant) long run multipliers:

(iv) A positive foreign real shock raises foreign GDP, which is a desirable property of the model.
(v) A positive foreign nominal shock depreciates the domestic currency, which is counter to our expectations. At the same time it raises the domestic price level more than the foreign price level.
(vi) The domestic interest rate increases if there is a positive shock to the domestic or foreign nominal trends, and decreases if there is a positive shock to the domestic real trend. These results are not unreasonable, given that the interest rates are not stationary according to the empirical cointegration vectors.
(vii) The foreign interest rate increases if there is a positive foreign real shock or if there is a positive domestic nominal shock. These are the least intuitive properties of the VEC model conditioned on the empirical cointegration vectors.

Comparing the monetary neutrality restrictions (A3) with their weaker version (A3′) we find that the overidentifying restriction $a_{12} = 0$ — a foreign monetary policy shock has no significant effect on domestic GDP — cannot be rejected for either of the A matrices in Table 4.

In the remainder of this subsection we will concentrate on the VEC model with the empirical cointegration vectors in (8). The reason is that the coefficients in this model are estimated with much greater precision, but many of the results that we will report are consistent with those obtained in the model with the theoretical vectors in (6).

In Table 5 we present decompositions of the forecast error variance of each variable with respect to different shocks. It deserves to be emphasized, again, that the decomposition with respect to permanent (real and nominal) shocks on the one hand and transitory (interest rate and aggregate demand) shocks on the other does *not* depend on the identifying assumptions on the common trends and the transitory shocks but only on the cointegration rank r. Only the distinctions between the different types of permanent and transitory shocks respectively, require identifying assumptions.

It can be seen that in the long run, domestic real shocks account for most of the forecast uncertainty in domestic output. In the short run, foreign interest rate shocks and other aggregate demand shocks are also important. At the two years horizon, which perhaps is the most interesting time perspective for monetary policy, real trend shocks dominate the forecast uncertainty in domestic output. This result is thus consistent with the "real business cycles" view that such cycles are driven by technology shocks.

Regarding fluctuations in the three month domestic interest rate, i_t, transitory shocks are somewhat more important than permanent shocks in the short run. Innovations to the domestic interest rate itself account for about 30 percent of the variance within the first quarter. In the longer run, shocks to the domestic real and foreign nominal trends become more important.

Permanent shocks explain virtually all of the forecast uncertainty for the domestic price level, not only in the long run, but also in the short run. We interpret this as a reflection of our sample period being dominated by an accommodative monetary policy regime, i.e. that a monetary policy aiming at price stability is a recent phenomenon. The negligible role of transitory shocks is nevertheless surprising, as is the fact that only the foreign permanent shocks seem to matter. It must be kept in mind, however, that all these results are derived from a VAR model where level shifts associated with some devaluations have been handled by the use of dummy variables. The results may therefore reflect that there is no strong domestic trend in the Swedish price level, over and above accommodative monetary policy interventions in the form of devaluations (recall that the cointegration tests suggested that there may be three trends rather than four).[18] Domestic real trend shocks explain half of the short run forecast error variance of the nominal exchange rate, as shown in Panel D of Table 5. Transitory interest rate shocks seem to be important for exchange rate fluctuations only in the very short run.

A common argument is that exchange rate fluctuations are largely driven by exogenous changes in expectations, or "credibility shocks", which are unrelated to changes in fundamental conditions such as technology, preferences, or economic policy. Our results show that about one quarter of the long run forecast error variance of the nominal exchange rate is accounted

[18] Another possibility is that there are measurement problems. For instance, it is well known that there have been persistent shifts in the price level due to changes in indirect taxes. This problem may be circumvented by looking at producer prices rather than consumer prices.

for by domestic nominal trend shocks, while the share of the foreign nominal trend is about three quarters. As noted above, the nominal trend shocks probably contain a large number of different shocks which — according to our identifying assumptions — have no long run effects on output, but which, through the accommodative behavior of monetary policy, have been allowed to have persistent nominal effects. In principle, exogenous changes in expectations may be included in the nominal trend shocks, but they may also show up as interest rate shocks. Neither of these categories of shocks are dominant in the short run, however, where domestic real and foreign interest rate shocks appear to be most important for exchange rate fluctuations. Hence, our results do not lend much support to the view that domestic nonfundamental shocks play a large role for exchange rate fluctuations. If nonfundamental shocks are important, they are likely to stem from international financial markets and affect the exchange rate through foreign interest rate changes.

Figures 2–4 show the impulse responses to a domestic nominal trend shock, a domestic interest rate shock, and an “other aggregate demand” shock, respectively. We have chosen to look at these impulse responses since they can shed light on the possible effects of an unexpected change in monetary policy. If monetary policy shocks lie behind nominal trend shocks, we have the so called “price puzzle” noted in many other VAR studies: as the domestic interest rate is raised, the price level also goes up. This puzzle also arises, although less clearly, if we assume that monetary policy shocks show up in the form of so called other aggregate demand shocks. The responses to the domestic interest rate shock, on the other hand, are more consistent with what seems to be many economists’ prejudices about the effects of a monetary policy shock: the increase in the domestic interest rate is associated with a fall in the domestic price level, although this effect is quickly reversed and furthermore very uncertain.

We have compared some of the impulse responses generated by our VAR model with the corresponding theoretical impulse responses generated by the inflation targeting model in Svensson (1998b) and get the following results.[19] First, the impact effects are often qualitatively the same in the VAR model(s) and the Svensson model. But, second, the impulse responses from the VAR model(s) display much more oscillations than Svensson’s theoretical impulse responses. This is a reflection of the fact that there are richer (less restricted) dynamic relations in the estimated VAR model(s) than in the inflation targeting model. Third, the peak effects of various shocks occur much faster in the VAR model(s) than in the theoretical model. These results

[19] We have examined the effects on the four domestic variables (inflation, output, the interest rate, and the exchange rate) from three shocks: a domestic interest rate shock, a domestic real trend shock, and a foreign nominal trend shock (a foreign inflation shock in Svensson’s terminology). The comparisons concern the so called flexible CPI targeting version of Svensson’s model.

are rather robust to the choice of VAR model, i.e. using theoretical or empirical cointegration vectors. Our findings thus suggest that it may be worthwhile to investigate the implications of theoretical inflation targeting models which allow for less restricted lag structures in aggregate demand and supply relations than have been studied so far. The same argument has been made by Sack (1998) in a study of U.S. monetary policy.

5. Monetary Policy Analysis

5.1. *Forecasting Inflation*

The conduct of monetary policy using an explicit inflation target hinges on the availability of a path of inflation forecasts. Thus, we have to establish that the VEC model is useful for forecasting in order to launch it as a serious contender as a general reference model in the analysis of monetary policy. One advantage with our approach is that all variables enter endogenously into the model. This means that it is able to produce genuine *ex ante* dynamic forecasts for each of the variables. Throughout, the forecasting horizon is 1994:1–1996:4 (12 quarters), and both dynamic (multi-step) and recursive 1-step ahead forecasts are evaluated. The dynamic forecasts approximately correspond to the information that a central bank needs to consider prior to a policy decision; the recursive 1-step ahead forecasts illustrate the situation that a central bank continuously has to envisage as new information becomes available.

In the analysis below we evaluate not only the theoretically motivated specification of the cointegration vectors in (6), but also an exactly identified set of vectors given by (8).

The first two graphs in Figure 5 give the results for the VEC model based on the empirical cointegration vectors in (8) without any restrictions on other parameters (UVEC). The corresponding graphs in Figure 6 give the results for the model based on the theoretical cointegration vectors in (6). Panel B of Table 6 summarises the results in terms of root mean square errors (RMSEs). The model with the theoretical vectors outperforms the one which uses the empirical vectors both in terms of dynamic and recursive 1-step ahead forecasts.[20]

Graphs III and IV in Figures 5 and 6 show the gains in forecasting performance from simplifying the models using the so called general to specific modeling strategy (see e.g. Hendry, 1995). The largest gains are obtained for the forecasts based on the empirical cointegration vectors, but also the forecasts conditioned on the theoretical vectors display improvements, especially for the dynamic forecasts; cf. Panels B and C in Table 6.

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[20]This is consistent with the results presented by Ingram and Whiteman (1994), who show that theoretical restrictions on a VAR can aid forecastability (although their framework is quite different from ours).

To sum up, although the testing evidence does not support the hypothesis of the three cointegration vectors being the real exchange rate and the two nominal interest rates, using these relationships does not result in a loss of forecasting accuracy. In fact, the model based on the empirical cointegration vectors only generates a lower RMSE in one out of four cases examined; and, in this case, the difference between the two models is small. Furthermore, the magnitudes of the RMSEs for the theoretical cointegration model are similar to those obtained in previous studies for other countries.[21] Having said this, it is important to remember that all models suffer from substantial overprediction problems at the end of the forecasting horizon.[22]

The uncertainty associated with the forecasts is very large: the width of the typical 95 percent (error variance based) confidence interval is around three to four percentage points.[23] This is important because some central banks that have chosen to base their monetary policy on inflation targeting announce not only a specific target but also a tolerance interval. Moreover, statistical model uncertainty has been used as one argument why such an interval is used. However, these intervals are typically quite narrow (for example, ± 1 percentage point on an annual basis as in Sweden), and their appropriateness for that purpose may thus be questionable. Finally, as shown in Table 6, the unrestricted VEC models generally perform worse than simple random walk models, while the opposite holds true for the restricted VEC models.[24]

What are then the most important characteristics of our inflation forecasting equations? The restricted (or parsimonious) inflation equation for the VEC model based on the theoretical cointegration vectors in (6) has the following appearance:

$$\begin{aligned}
\widehat{\Delta p_t} = & \underset{(.03)}{.06}\Delta e_{t-1} + \underset{(.21)}{.48}\Delta p^*_{t-1} - \underset{(.23)}{.50}\Delta p^*_{t-2} + \underset{(.22)}{1.19}\Delta p^*_{t-3} \\
& - \underset{(.05)}{.11}\Delta y_{t-1} - \underset{(.04)}{.07}\Delta y_{t-2} - \underset{(.05)}{.11}\Delta y_{t-3} - \underset{(.07)}{.05}\Delta y^*_{t-1} \\
& - \underset{(.10)}{.15}\Delta i_{t-2} - \underset{(.10)}{.07}\Delta i_{t-3} - \underset{(.11)}{.33}\Delta i^*_{t-1} - \underset{(.10)}{.09}\, i_{t-1} \\
& - \underset{(.06)}{.17}\, i^*_{t-1} + \big(\text{constant and dummy variables}\big),
\end{aligned} \tag{9}$$

[21] See for example Stevens and Debelle (1995), OECD (1993), and Artis (1997).

[22] Almost all official forecasts for Swedish CPI inflation in 1995 and 1996 were plagued by severe overprediction problems. For 1995, the 1-year ahead forecasts of the six most prominent official forecasters averaged around 3 percent, whereas the outcome was 2.4 percent. For 1996 the same six forecasters again averaged 3 percent, but now the outcome dropped to 0.1 percent.

[23] Parameter variance based confidence intervals imply even larger uncertainty; see Doornik and Hendry (1997, chapters 7 and 10).

[24] For purposes of comparisons, Table 6 also displays RMSEs for simple mean reverting models. The dynamic mean forecasts are computed using the mean of the inflation rate over the sample period 1973:2–1993:4. The recursive 1-step ahead mean forecasts are computed using a recursively updated mean over the sample periods 1973:2–t, where $t = 1993{:}4, 1994{:}1, \ldots, 1996{:}3$.

where standard errors are given within parentheses, and the estimation period is 1973:2–1993:4. The overall standard error for the equation is about 0.85 percent. It deserves to be emphasized that the presented parameter estimates are nonstructural and hence do not offer any direct economic interpretations. Nevertheless, the following observations are noteworthy.

First, equation (9) suggests that interest rate innovations, and hence monetary policy changes, affect inflation faster than is commonly assumed in theoretical inflation targeting models (where a one or two year lag is often imposed). This is confirmed by looking at the impulse response functions, which however, as noted in Section 4.2, also show that the effects are very uncertain.

Second, the inflation equation suggests that there is a significant pass through of exchange rate changes, i.e. a depreciation of the nominal exchange rate, *ceteris paribus*, indicates that inflation will rise. This explains why the development of the nominal exchange rate is given so much attention in monetary policy analysis in general and inflation forecasting in particular. But, again, the coefficient on the nominal exchange rate in (9) is not a measure of the "effect" of exchange rate changes on inflation since both variables are endogenous. Implicit elasticities may however be calculated, for different shocks and forecast horizons, by looking at the impulse responses, such as those in Figures 2–4, and at the long run multipliers, such as those in Table 4.

5.2. *Real and Nominal Exchange Rates*

The concept of an equilibrium exchange rate plays an important role in monetary policy analyses, in particular in small open economies. The reasoning goes along the following lines: suppose the equilibrium level could somehow be measured, then this information could be used to help predict the future path of exchange rates, inflation, and interest rates. *Nominal* exchange rates, however, appear to be almost unpredictable and show little tendencies to revert to any (even time varying) mean. An equilibrium level of the *real* exchange rate may be easier to define. It is not unreasonable to define the equilibrium relative price of domestic goods as the level of the real exchange rate which is consistent with an equilibrium in the domestic goods market.

Using the first cointegration relation in (5), the goods market equilibrium relation may be written as

$$y_t = -\frac{1}{\beta_{11}}\left(e_t + p_t^* - p_t\right) - \frac{\beta_{13}}{\beta_{11}} i_t - \frac{\beta_{15}}{\beta_{11}} y_t^* - \frac{\beta_{17}}{\beta_{11}} i_t^* + z_t^y, \tag{10}$$

where z_t^y is a measure of the equilibrium error in the goods market. Similarly, we may express a measure of the goods market equilibrium in terms of the real exchange rate:

$$e_t + p_t^* - p_t = -\beta_{11} y_t - \beta_{13} i_t - \beta_{15} y_t^* - \beta_{17} i_t^* + z_t^q. \tag{11}$$

The equilibrium level of the real exchange rate may thus be defined as $-\beta_{11} y_t - \beta_{13} i_t - \beta_{15} y_t^* - \beta_{17} i_t^*$ and the equilibrium error in the real exchange rate is (proportional to) the (long run) equilibrium error in the goods market.[25] In the case of long run PPP, the equilibrium real exchange rate is constant and fluctuations in the real exchange rate directly reflect deviations from the long run equilibrium.

As reported in Section 3.3, we can reject that the empirical VAR model with three cointegration vectors satisfy all the theoretical restrictions implied by equation (6). The weaker hypothesis of PPP, i.e. that one cointegration vector looks like β_1 in (6), is rejected using the asymptotic distribution (p-value = .00), but not according to the bootstrapped distribution (p-value = .07).

Irrespective of whether we use the theoretical cointegration vector that is consistent with long run PPP or the empirical cointegration vector β_1 in (8) to estimate the deviation of the real exchange rate from its equilibrium level, we have the problem that the equilibrium error may be nonstationary unless we condition on the influences from dummy variables (discussed in Section 3.1). There can thus be shifts in the equilibrium level of the real exchange rate in connection with the regime shifts in the early 1980s and early 1990s. This would be hard to explain if the regime shifts were only associated with changes in monetary policy. It is well known, however, that the 1980s and 1990s have been characterized by quite far reaching policy reforms, involving e.g. liberalization of capital markets and foreign trade, deregulation and privatization of many industries, etc.

In Graph I of Figure 7 we depict the real exchange rate, and in Graphs II–IV we show the equilibrium error, z_t^q, measured under two different assumptions about the cointegration vectors: the β_1 vectors in (8) and (6), respectively. The estimated z_t^q series have been adjusted for the influence of the dummy variables; this accounts for the difference between Graphs III and IV.[26] Notice that the trend in the real exchange rate is eliminated if the deterministic regime

[25] Since $z_t^y = (1/\beta_{11}) z_t^q$, the discussion hinges on the assumption that $\beta_{11} \neq 0$ and, thus, that the real exchange rate is nonstationary.

[26] The reduced form of the common trends model can be written as

$$x_t = C\xi_t + C^*(L)\varepsilon_t + \mu_0 t + \mu_1 \sum_{i=1}^{t} D_i + \kappa + \Theta(L) D_t,$$

where $\xi_t = \xi_{t-1} + \varepsilon_t$ and C is $n \times n$ with rank k such that $\beta' C = 0$. The vector x_t can now be decomposed into a permanent or trend component, $x_t^{\mathrm{p}} = C\xi_t + \mu_0 t + \mu_1 \sum_{i=1}^{t} D_i$, and a transitory component, $x_t^{\mathrm{s}} = C^*(L)\varepsilon_t + \kappa +$

shifts are taken into account. Our results thus suggest that the relative price of Swedish goods and services was kept at a higher level through some mechanisms that were prevalent before the regime shift (possibly market regulations).

It can be seen from these graphs that the real exchange rate seems to have been undervalued in 1994, irrespective of which definition of the equilibrium rate we choose to look at. This can thus explain the appreciation of the nominal exchange rate after 1994 (cf. Figure 1). Whether or not the real exchange rate was undervalued or overvalued at the end of the sample period (1996) is less clear.

The exchange rate equation of the restricted parsimonious VEC model based on the theoretical cointegration vectors (i.e., the exchange rate equation from the same system as the inflation equation (9)) looks as follows:

$$\begin{aligned}
\widehat{\Delta e_t} = &- \underset{(.29)}{.45}\,\Delta p_{t-3} - \underset{(.09)}{.12}\,\Delta e_{t-2} - \underset{(.09)}{.07}\,\Delta e_{t-3} + \underset{(.55)}{1.17}\Delta p^*_{t-3} \\
&- \underset{(.10)}{.08}\,\Delta y_{t-1} - \underset{(.10)}{.16}\,\Delta y_{t-3} - \underset{(.17)}{.17}\,\Delta y^*_{t-1} - \underset{(.16)}{.22}\,\Delta y^*_{t-2} \\
&- \underset{(.17)}{.12}\,\Delta y^*_{t-3} - \underset{(.25)}{.97}\,\Delta i_{t-1} - \underset{(.27)}{.32}\,\Delta i_{t-2} - \underset{(.24)}{.30}\,\Delta i_{t-3} \\
&+ \underset{(.27)}{.36}\,\Delta i^*_{t-1} + \underset{(.28)}{.55}\,\Delta i^*_{t-2} + \underset{(.24)}{.56}\,\Delta i^*_{t-3} \\
&- \underset{(.05)}{.14}\left(e_{t-1} + p^*_{t-1} - p_{t-1}\right) + \underset{(.18)}{.71}\,i_{t-1} - \underset{(.16)}{.58}\,i^*_{t-1} \\
&+ \left(\text{constant and dummy variables}\right).
\end{aligned} \tag{12}$$

It can be seen that the real exchange rate is useful for predicting the next period change in the nominal exchange rate.[27] As a partial relation, if the real exchange rate is above its equilibrium level (zero, conditioned on the dummies) by 1 percentage point, we expect the nominal exchange rate to appreciate by .14 percentage points. Hence, the nominal exchange rate does not follow a random walk, but does part of the job of closing the equilibrium error in the real exchange rate.

The VAR approach thus provides a framework for explaining why the concept of an equilibrium real exchange rate attracts interest in monetary policy analysis. A real exchange rate

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$\Theta(L)D_t$. The deviation from the goods market equilibrium, z_t^q, is measured by $\beta_1' C^*(L)\varepsilon_t$, i.e. by $\beta_1' x_t$ adjusted for the deterministic term $\beta_1'\left(\kappa + \Theta(L)D_t\right)$. Given that PPP is imposed on the parameters of the VEC model, the estimate of this deterministic term depends on which other restrictions are imposed on the parameters. In Graph III, there are no other restrictions, while in Graph IV the VEC model is estimated under the assumptions that the two nominal interest rates are stationary, i.e. the restrictions on β_2 and β_3 in (6).

[27] This is in contrast to the findings for some other countries by Norrbin, Reffett, and Ji (1997), who reported that the PPP deviation seemed to have little predictive power for nominal exchange rate changes, although it Granger caused U.S. inflation.

above its long run mean (a weak real exchange rate), *ceteris paribus*, can in certain models be shown to be associated with a lower (stronger) level of the nominal exchange rate and lower inflation in the future. More generally, however, changes in real and nominal exchange rates are driven by some underlying shocks, and the responses of exchange rates and inflation are quite different depending on the dominating shock. Impulse response analysis (as in Section 4.2) is one way to address such issues.

5.3. *Monetary Conditions Indexes*

Discussions of monetary policy are sometimes based on some index of "monetary conditions" (a so called MCI). The purpose of an MCI is to identify combinations of exchange and interest rates that leave the stance of monetary policy unaffected. In practice, MCIs are usually linear combinations of exchange and interest rates (real or nominal), where the weights are chosen so as to reflect these variables' effects on output; see e.g. Gerlach and Smets (1996) and Ericsson, Jansen, Kerbeshian, and Nymoen (1997).

As noted in the previous subsection, the goods market equilibrium condition associated with the first cointegration vector in (5) may be written as (10) and hence $y_t - z_t^y$ may be defined as the equilibrium level of y_t. If we define an MCI, denoted by m_t, as

$$m_t = \left(e_t + p_t^* - p_t\right) + \beta_{13} i_t,$$

we see that an unchanged level of the MCI will be associated with an unchanged level of actual output y_t, given y_t^*, i_t^*, and z_t^y; or, equivalently, an unchanged level of the MCI will be associated with an unchanged level of equilibrium output, given y_t^* and i_t^*.

Although this is certainly not the only possible definition or interpretation of a monetary conditions index, it suggests some limits to the usefulness of the MCI for monetary policy purposes. First, certain linear combinations of exchange and interest rates cannot in general be expected to be consistent with a stable level of output unless other determinants of output, e.g. shocks to world market conditions and other disturbances, are unchanged. Second, the MCI derived from the goods market equilibrium condition (10) is a long run relation. It is not informative about the relation between interest rates, exchange rates, and aggregate demand in the short run, i.e. it does not take account of the lags in the transmission mechanisms. Furthermore, if the real exchange rate is stationary, i.e. if the β_{ij}:s satisfy the restrictions given by β_1 in (6), there is no long run relation between the level of output and the real exchange rate (cf. footnote 25 on page 22). In that case, domestic and foreign goods are perfect substitutes, and in a small open economy domestic production is not affected by domestic aggregate demand.

Third, although there are certain circumstances under which the MCI concept makes theoretical sense (see Gerlach and Smets, 1996, and Svensson, 1998b, for discussions), there seems to be little reason to believe that the combinations of exchange and interest rates that can help predict inflation satisfy the restrictions implied by an MCI. But the usefulness of MCIs for inflation forecasting is an empirical issue of course.

To shed light on the properties of an MCI we perform two exercises. First we compare the effects, through impulse response analyses, of various shocks on m_t, output (y_t), and inflation (Δp_t). Second, we test whether the inflation equation in the VEC model satisfies the restrictions on $(e_t + p_t^* - p_t)$ and i_t implied by a certain definition of the MCI. In these exercises, the MCI is defined as

$$m_t = \left(e_t + p_t^* - p_t\right) - 2.875 i_t,$$

which is based on our estimate of β_{13} in (8). These exercises are only meant to be illustrative, but it should be noted that a weight of around -3 is rather typical for MCIs used by central banks and other institutions which analyze monetary policy.[28]

In Graphs I–VII of Figure 8 we depict the effects of the seven structural shocks on the MCI, domestic output, and domestic inflation in the common trends model with the estimated cointegration vectors. Graph II shows that after a shock to the foreign nominal trend, the three variables develop in a similar fashion. In general, however, they appear to react quite differently. It should be noted, though, that an optimal monetary policy does not imply that the MCI should change in a certain way, irrespective of what shock that hits the economy (see Gerlach and Smets, 1996). It is thus not unreasonable that monetary conditions are tightened (m decreases) when inflation rises after a transitory aggregate demand shock (see Figure 8 VII). What the graphs show is that there is no simple rule of thumb that can tell us how to infer the "inflationary pressure" from the development of the MCI.

The tests of the usefulness of the MCI for forecasting have been formulated as follows. The variable vector has been redefined by replacing e_t with m_t. We then test if all information about p_t^*, p_t, and i_t that is relevant for future inflation can be captured by the MCI variable. This is done for two versions of the VEC model, using the cointegration vectors (6) and (8), respectively. In the model based on the theoretical vectors (6), the hypothesis is that the coefficients on Δp_{t-s}, Δi_{t-s}, Δp_{t-s}^*, and the second cointegration relation (i_{t-1}) are equal to zero in the inflation equation. In the model based on the empirical cointegration vectors (8), the hypothesis

[28]The median of the weights reported by Ericsson et al. (1997), Table 1, is -3.5. Due to a few large outliers, the unweighted average is -4.6.

is that the coefficients on Δp_{t-s}, Δi_{t-s}, Δp^*_{t-s}, and the second and third cointegration relations are equal to zero. Both hypotheses are strongly rejected, which is hardly surprising (against the background of e.g. the good forecasting properties of the inflation equation (9)).

To sum up, using our cointegrated VAR approach we find that an MCI only to a very limited extent per se is useful in the analysis of monetary policy. Interest rates and exchange rates are, of course, central to the conduct of monetary policy, but some aggregated index which combines these variables to obtain a simple rule of thumb to infer the inflationary pressure may be misleading.

5.4. *Inflation and the Output Gap*

Theoretical models of inflation and monetary policy are often based on aggregate demand and supply relations which suggest that inflation is related not to the actual level of output, but to the difference between actual and some measure of potential (or natural, or long run) output, i.e. to a so called output gap. Potential output is however exogenous in those models. In empirical applications, the output gap is often measured as the deviation of actual output from a trend. A popular approach for determining the trend is to fit a linear deterministic trend to output, but more flexible trend models, e.g. with time varying growth rates, are also used.[29]

For the empirical questions that we have analyzed with our VAR framework so far, it has not been necessary to explicitly identify an output gap. For instance, if we look at the inflation equation (9) reported in Section 5.1, and the corresponding output equation

$$
\begin{aligned}
\widehat{\Delta y_t} = & \underset{(.20)}{.24}\Delta p_{t-1} - \underset{(.21)}{.47}\Delta p_{t-3} - \underset{(.06)}{.11}\Delta e_{t-1} - \underset{(.07)}{.08}\Delta e_{t-2} \\
& + \underset{(.07)}{.06}\Delta e_{t-3} - \underset{(.40)}{.82}\Delta p^*_{t-1} + \underset{(.38)}{1.19}\Delta p^*_{t-2} + \underset{(.42)}{.67}\Delta p^*_{t-3} \\
& - \underset{(.08)}{.86}\Delta y_{t-1} - \underset{(.07)}{.84}\Delta y_{t-2} - \underset{(.08)}{.74}\Delta y_{t-3} - \underset{(.13)}{.36}\Delta y^*_{t-2} \\
& - \underset{(.13)}{.23}\Delta y^*_{t-3} - \underset{(.20)}{.29}\Delta i_{t-2} + \underset{(.20)}{.40}\Delta i^*_{t-1} + \underset{(.21)}{.24}\Delta i^*_{t-2} \\
& + \underset{(.19)}{.28}\Delta i^*_{t-3} + \underset{(.04)}{.12}\left(e_{t-1} + p^*_{t-1} - p_{t-1}\right) - \underset{(.13)}{.26}\, i_{t-1} \\
& - \underset{(.12)}{.31}\, i^*_{t-1} + \left(\text{constant and dummy variables}\right),
\end{aligned}
\tag{13}
$$

we see that inflation and output growth are expressed as functions of lagged changes of y_t, but not the deviation of y_t from some trend. That does not mean that the VAR approach is inconsistent with common theoretical models of the relation between output and inflation. In fact, a model consistent way to measure the output gap is to define it as the demeaned transitory

[29] A review of various techniques for estimating the output gap is given in Apel and Jansson (1997).

component of output, i.e. the first element of $\Phi(L)\varphi_t = \Phi(L)F^{-1}\varepsilon_t = C^*(L)\varepsilon_t$ (cf. equation (3) and footnote 26 on page 22). Hence, this output gap measure does *not* depend on the identifying assumptions in the common trends model and can be viewed as an extension of the detrending technique discussed and applied in e.g. Vredin and Warne (1991).

The fact that the output gap plays no explicit role in the VAR model may imply that this model is inefficient, if the theoretical models of inflation and the output gap are correct. On the other hand, since most theoretical models of inflation and monetary policy do not suggest how potential output, and hence the output gap, should be defined, it appears to be an advantage of the VAR approach that many questions in monetary policy analysis can be studied empirically without any explicit identification of the output gap. Nevertheless, it is important to study how the measure of the transitory component of output compares with other commonly used measures of the output gap, and how much extra information about future inflation these other measures may contain.

In Graphs I–III of Figure 9 we depict the three measures of the output gap which are used in Sveriges Riksbank's "Inflation Reports". The first is output detrended by a Hodrick-Prescott (HP) filter (using a smoothness parameter of 6400). The second is a measure of the output gap derived from an unobserved components (UC) model by Apel and Jansson (1997), who use an Okun's law as well as a Phillips curve relationship to estimate the cyclical part of output from unemployment and inflation data. The third measure is based on a production function (PF) approach, which estimates potential output from capital stock and employment data (see Hansen, 1997). As can be seen from the graphs, the three measures are strongly correlated. However, the Hodrick-Prescott filtered series gives a much smaller negative output gap at the end of the sample.

Graphs IV and V show our measures of the demeaned transitory component of output derived from the VEC based on theoretical and empirical cointegration vectors, respectively.[30] It can be seen that the two measures follow each other closely until the beginning of the 90s. In 1993, the gap computed from the model with the theoretical cointegration vectors drops dramatically from around minus 1 percent to minus 3.5 percent, and it then decreases further to about minus 5.5 percent at the end of 1995. The gap measure for the empirical vectors, on the other hand, only decreases slightly during the 1993–95 period. Thus, while the former gap

[30]To make the comparison with the series in Graphs I–III meaningful, the series in Graphs IV–V have been computed as four-quarter moving averages of the demeaned transitory components. This dampens the influences from high frequency noise, which in the Riksbank's measures of the output gap is handled by the X11 seasonal filter.

measure displays a behavior quite similar to those using the UC and the PF approaches, the latter — during this episode — looks more like the HP filtered gap series.

A general comparison of the output gaps with respect to the interpretation of historical business cycles reveals an interesting pattern. While both VEC measures suggest that the early 1980s is almost neutral in terms of the business cycle, the UC, PF, and HP measures all indicate that this is a period where economic activity is strongly below trend. Hence, whereas the downturn in actual economic activity that characterized these years is a transitory phenomenon according to the UC, PF, and HP output gaps, the VEC measures, rather, suggest that the downturn is due to a drop in the permanent component of output.

In Panel A of Table 7 we report cross correlation statistics between the VEC measure based on the theoretical cointegration vectors and the HP, UC, and PF output gap measures. As can be seen, the correlations between the VEC gap and the UC and PF gaps are much stronger than the correlations between the VEC gap and the HP filtered output gap.[31]

In Panel B we show how these measures correlate with inflation. Somewhat surprisingly, the VEC gap displays the strongest correlation with inflation at both leads and lags. Concerning the other gap measures, inflation does not appear to be correlated with leads of the HP and PF measures, while lags of all three gaps are weakly correlated with inflation.

The finding that some output gap measures are correlated with inflation does, of course, not mean that an explicit measure of the output gap is needed for making good forecasts of future inflation. As shown in Section 5.1, our cointegrated VAR approach does relatively well in forecasting inflation, but does not require any such identification. As a simplistic way to check whether or not the three output gap measures add any predictive power for inflation in relation to the information already contained in the VAR system, we have simply augmented the inflation equation of the cointegrated VAR with these measures, one at a time. This has been done for four different versions of the VAR: using the empirical and theoretical cointegration vectors, and the unrestricted and the more parsimonious models, i.e. the same four models which were compared in the forecasting exercises in Section 5.1. The results are presented in Table 8.

[31] A feature of Table 7 which at first glance may appear a bit peculiar is that the cross correlations are somewhat stronger for $s < 0$ than for $s > 0$. To understand this result better, note that we have computed the VEC measure as a one sided (no leads) moving average of the original (demeaned) transitory component for output. This means that some of the covariation that is obtained for $s < 0$ actually relates to the contemporaneous covariation. While a two sided smoothing procedure (leads and lags) could be used to mitigate the issue, such a filter has the disadvantage of relying on future information for determining the current value. For practical purposes, a one sided (no leads) filter is preferable because it corresponds better to the information available to policy makers in "real time".

The output gap measures are not significant in the parsimoniously specified VAR models, which we know have good forecasting properties, but seem to add information about future inflation to the larger, relatively unrestricted VAR models.

It has to be emphasized that the comparisons undertaken above should not be interpreted as rigorous tests of whether or not explicit measures of the output gap are useful for the purpose of analyzing and conducting monetary policy. Nor should they be interpreted as formal tests of whether a particular measure is more useful than some other, since the measures are not computed using the same information set.[32]

What we think the comparison shows is that the VAR approach can do relatively well as an empirical model, and be used for quite different purposes, without imposing strong assumptions about how potential output is determined. This is encouraging, since there appears to be little consensus about how potential output should be defined, theoretically or empirically.

6. Conclusions

In this paper we have shown that a VAR model is a usable and flexible tool for analyses of many different issues that are relevant for monetary policy. The VAR framework can thus offer a benchmark with which other models, which are designed to handle specific issues, can be compared and consistently checked. We believe that it is necessary to develop such benchmark models in order to make monetary policy analysis as transparent as possible. The existence of a benchmark model does not, however, imply that in-depth analyses of various issues, using other models than VARs, are unnecessary. Quite the contrary, our results point to the need for more empirical studies of several problems.

We have shown that it is possible to formulate a VAR model that has reasonably good properties when it comes to forecasting inflation (compared to other commonly applied models). The VAR model allows for such complex dynamic relations that appear to be empirically important, yet allows us to impose (and test) long run restrictions that are suggested by economic theory. The VAR model can thus serve as a statistical tool, while at the same time be economically interpretable.

In analyses of monetary policy and in the process of inflation forecasting, it is common to examine different variables that are believed to be related to future inflation, such as nominal and real exchange rates, and various definitions of so called monetary conditions and output gaps. Our analyses show that such relations can be studied within one consistent framework, and we get the following results:

[32]For example, the UC model of Apel and Jansson (1997) is designed to explain the change rather than the level of inflation.

(i) It is possible to identify an equilibrium real exchange rate that can help predict future changes in the nominal exchange rate. The equilibrium rate seems to have been affected by changes in policy regimes, but once these are accounted for (via dummy variables) the real exchange rate appears to be stationary.

(ii) Nominal exchange rate fluctuations can help predict future inflation.

(iii) Monetary conditions indexes, i.e. certain linear combinations of exchange and interest rates, are however not likely to be very useful. It is informative to take account of interest rates, exchange rates, and prices, of course, but the restrictions one imposes when defining an MCI are not likely to be justified.

(iv) Inflation is significantly correlated with an output gap that can be calculated using a VAR model. That output gap, in turn, can be shown to be correlated with other measures of the output gap (which are not more strongly correlated with inflation). On the other hand, it is not necessary to identify any output gap in order to forecast inflation, which is encouraging since there is little consensus on how to measure potential output.

We have also obtained several other results that we believe are important for monetary policy in general and inflation targeting in particular. First, although the VAR approach can be shown to have as good forecasting properties as other commonly applied approaches, the uncertainty of inflation forecasts is generally very large. Second, the analysis suggests that a considerable share of the forecasting uncertainty of Swedish inflation stems from foreign shocks. Third, given the forecast uncertainty, the tolerance intervals that some central banks (including Sveriges Riksbank) define around their inflation targets are too narrow to be interpreted as confidence intervals in the usual sense. This means that they must be justified on some other grounds, e.g. that they can help to create accountability (as seems to be the case currently in the U.K.). Fourth, interest rate innovations affect inflation and output more rapidly than what is commonly assumed in theoretical models of inflation targeting. And finally, the empirical results suggest that the effects (of various shocks) are more oscillatory, i.e. that the dynamic interactions are more complex, than commonly assumed. The policy implication of the last observation should be investigated.

DATA APPENDIX

The quarterly data set runs from 1972:2 to 1996:4. Due to lags, the effective estimation period begins 1973:2. All series are seasonally unadjusted except for the two real GDP series which are adjusted through regressions on seasonal dummies prior to the system analysis. The exact definitions and sources of the variables are as follows. Real domestic output, y_t, is defined as $100 \ln Y_t$, where Y_t is Swedish real GDP in fixed 1991 prices (source: Statistics Sweden). The domestic price level, p_t, is given by $100 \ln P_t$, where P_t is the Swedish consumer price index in quarterly averages with 1991 as the base year (source: Statistics Sweden). The domestic nominal interest rate, i_t, is defined as $100 \ln\left(1 + I_t/100\right)$, where I_t is the Swedish three month treasury bills rate in percent, ultimo (source: Sveriges Riksbank). The nominal exchange rate, e_t, is defined as $100 \ln S_t$, where S_t is the geometric sum (using IMF's TCW, Total Competitiveness Weights) of the nominal Krona exchange rate of Sweden's 20 most important trading partners (source: Sveriges Riksbank and IMF). Foreign real output, y_t^*, is given by $100 \ln Y_t^*$, where Y_t^* is German real GDP in fixed 1991 prices (source: Bundesbank). The foreign price level, p_t^*, is equal to $100 \ln P_t^*$, where P_t^* is the geometric sum (IMF's TCW) of the CPIs (quarterly averages, 1991 is the base year) of Sweden's 20 most important trading partners (source: Sveriges Riksbank and IMF). Finally, the foreign nominal interest rate, i_t^*, is calculated as $100 \ln\left(1 + I_t^*/100\right)$, where I_t^* is the German three month treasury bills rate in percent, ultimo (source: Sveriges Riksbank).

TABLE 1: Asymptotic and bootstrapped p-values for multivariate specification tests in percent.

(A) Serial correlation tests

model	Portmanteau asymp	Portmanteau boot	$LM(1)$ asymp	$LM(1)$ boot	$LM(4)$ asymp	$LM(4)$ boot
$p=2, r=1$	.00	.00	.31	.39	.00	.00
$r=2$	.00	.00	.82	1.12	.00	.00
$r=3$	.00	.00	.04	.07	.00	.00
$r=4$	.00	.00	.35	.67	.00	.00
$r=5$	.00	.00	.05	.15	.00	.00
$p=3, r=1$	14.55	68.17	.00	.00	.10	.15
$r=2$	3.99	46.85	.00	.00	.06	.10
$r=3$	6.22	64.49	.00	.00	.56	.83
$r=4$	4.64	69.98	.00	.01	.22	.38
$r=5$	9.35	50.51	.00	.00	.28	.55
$p=4, r=1$	.01	24.93	2.88	3.55	4.69	7.15
$r=2$	.01	52.96	10.69	13.56	12.11	18.58
$r=3$	.00	44.14	24.55	32.03	18.57	26.45
$r=4$	.00	46.15	27.90	37.88	12.39	20.61
$r=5$	.00	44.63	57.53	70.70	9.00	18.16
$p=5, r=1$	.00	31.92	17.12	18.43	61.00	71.11
$r=2$	.00	50.36	4.49	5.85	33.64	45.72
$r=3$	.00	45.88	4.51	7.09	12.10	20.51
$r=4$	.00	42.45	4.48	8.43	25.09	38.86
$r=5$	.00	38.20	8.23	15.90	29.77	47.17
$p=6, r=1$	.00	15.79	43.79	45.38	9.20	14.73
$r=2$	.00	50.72	78.96	83.12	24.18	34.89
$r=3$	.00	51.80	57.09	67.82	25.24	38.27
$r=4$	.00	58.97	41.05	57.90	39.76	58.00
$r=5$	.00	64.80	36.60	58.04	18.16	35.28

NOTES: The Portmanteau statistic is asymptotically χ^2 with $n^2([T/4]-p+1)-nr$ degrees of freedom. $LM(q)$ is a Lagrange Multiplier test with respect to the q:th lag. It is asymptotically χ^2 with n^2 degrees of freedom.

TABLE 1: (Continued) Asymptotic and bootstrapped p-values for multivariate specification tests in percent.

(B) Normality tests

	Omnibus		Skewness		Kurtosis	
model	asymp	boot	asymp	boot	asymp	boot
$p = 2, r = 1$	.00	.00	1.78	4.25	.00	.00
$r = 2$	.00	.00	9.50	14.40	.00	.01
$r = 3$	.00	.00	25.99	31.68	.00	.01
$r = 4$	.00	.00	.00	.00	.00	.00
$r = 5$	.00	.00	.00	.00	.00	.00
$p = 3, r = 1$	.00	.00	.00	.00	.00	.00
$r = 2$	.00	.00	.04	.37	.00	.00
$r = 3$	.00	.00	1.09	3.12	.00	.01
$r = 4$	.00	.00	.00	.05	.00	.00
$r = 5$	.00	.00	.00	.02	.00	.00
$p = 4, r = 1$	.06	.58	.02	.46	.00	.14
$r = 2$	1.60	4.76	.70	3.73	.00	1.42
$r = 3$	14.29	21.62	2.14	7.66	.00	7.65
$r = 4$	6.59	12.90	9.65	21.51	.00	30.94
$r = 5$	2.25	6.63	7.93	19.61	.00	31.97
$p = 5, r = 1$	6.36	12.60	6.36	16.66	.00	8.34
$r = 2$	15.43	25.66	7.25	21.06	.00	13.22
$r = 3$	22.14	33.40	11.93	29.42	.00	32.21
$r = 4$	26.73	39.03	12.93	32.07	.00	39.11
$r = 5$	15.86	28.76	8.65	26.74	.00	35.57
$p = 6, r = 1$	2.36	9.00	1.53	10.63	.00	6.38
$r = 2$	10.52	23.63	18.46	43.66	.00	8.76
$r = 3$	3.74	13.33	14.49	39.79	.00	4.83
$r = 4$	5.88	18.01	6.96	28.13	.00	11.30
$r = 5$	13.98	31.30	24.30	54.57	.00	38.59

NOTES: Omnibus refers to the multivariate test for normality, suggested by Doornik and Hansen (1994), which is asymptotically χ^2 with $2n$ degrees of freedom. The skewness and kurtosis statistics are the multivariate tests for excess skewness and kurtosis in Mardia (1970). They are both asymptotically χ^2 with $n(n+1)(n+2)/6$ and 1 degree(s) of freedom, respectively.

TABLE 1: (Continued) Asymptotic and bootstrapped p-values for multivariate specification tests in percent.

(C) ARCH test

	M-ARCH	
model	asymp	boot
$p = 2, r = 1$	66.04	69.16
$r = 2$	51.92	59.28
$r = 3$	68.23	70.74
$r = 4$	97.81	95.44
$r = 5$	97.58	95.08
$p = 3, r = 1$	78.83	76.03
$r = 2$	70.40	70.47
$r = 3$	67.07	67.75
$r = 4$	91.07	87.41
$r = 5$	92.21	88.73
$p = 4, r = 1$	.02	3.11
$r = 2$	.01	3.26
$r = 3$	.01	3.20
$r = 4$	.02	4.55
$r = 5$	1.09	18.78
$p = 5, r = 1$	.40	9.46
$r = 2$	.01	3.55
$r = 3$	.45	12.86
$r = 4$	1.34	20.58
$r = 5$	1.95	24.26
$p = 6, r = 1$	.96	20.05
$r = 2$	.02	6.52
$r = 3$	.00	2.57
$r = 4$	.00	11.30
$r = 5$	.01	7.14

NOTES: M-ARCH is a multivariate version of the univariate Lagrange Multiplier test against ARCH suggested by Granger and Teräsvirta (1993) (cf. McLeod and Li, 1983). Since the limiting distribution of this statistic remains to be derived, we have used critical values from a χ^2 with $n^2(T/4)$ degrees of freedom for asymptotic inference.

TABLE 2: Asymptotic and bootstrapped quantiles for the LR cointegration test.

(A) Asymptotic distribution with unrestricted constant

r	$7-r$	80%	90%	95%	97.5%	99%	LR_{tr}
0	7	111.79	117.73	123.04	127.59	133.04	323.24
1	6	84.10	89.37	93.92	97.97	102.95	213.73
2	5	60.23	64.74	68.68	72.21	76.37	120.39
3	4	40.08	43.84	47.21	50.19	53.91	74.32
4	3	23.72	26.70	29.38	31.76	34.87	35.46
5	2	11.06	13.31	15.34	17.24	19.69	10.70
6	1	1.64	2.71	3.84	5.02	6.64	.68

(B) Asymptotic distribution with unrestricted constant
and 2 regime dummies

r	$7-r$	80%	90%	95%	97.5%	99%	LR_{tr}
0	7	111.30	117.57	123.04	127.64	133.15	323.24
1	6	81.06	86.58	91.16	95.34	100.34	213.73
2	5	54.53	59.15	63.33	66.76	70.56	120.39
3	4	31.60	35.41	38.65	41.67	45.11	74.32
4	3	12.27	14.77	16.98	19.05	21.91	35.46
5	2	5.99	7.78	9.49	11.10	13.30	10.70
6	1	1.64	2.71	3.84	5.02	6.64	.68

(C) Bootstrapped distribution for the empirical VAR model

r	$7-r$	80%	90%	95%	97.5%	99%	LR_{tr}
0	7	200.29	211.47	220.83	229.27	239.46	323.24
1	6	134.96	144.26	152.28	159.31	168.01	213.73
2	5	95.50	104.70	111.69	117.96	125.22	120.39
3	4	60.99	67.66	73.47	78.51	84.68	74.32
4	3	36.61	42.07	46.68	50.91	56.01	35.46
5	2	15.17	18.63	21.71	24.63	28.05	10.70
6	1	3.65	5.58	7.46	9.16	11.42	.68

NOTES: The critical values in Panel A are taken from Johansen (1995, Table 15.3). Those in Panel B have been obtained by simulation using Bent Nielsen's program "Disco". The empirical distributions in Panel C have been estimated using a parametric bootstrap procedure.

TABLE 3: Univariate specification tests for the empirical VEC model with $p = 4$ and $r = 3$.

equation	AR(5)	p-value	normality	p-value	ARCH(4)	p-value
Δy_t	1.45	21.91	5.47	6.49	.84	50.23
Δp_t	.56	73.28	5.10	7.82	1.00	45.10
Δi_t	.83	53.05	1.58	45.34	.43	78.76
Δe_t	3.22	1.21	2.31	31.53	1.33	26.86
Δy_t^*	1.50	20.40	.62	73.24	.58	67.73
Δp_t^*	.31	90.72	4.48	10.67	.19	94.03
Δi_t^*	.41	84.27	2.03	36.17	.54	71.09

NOTES: The tests have been calculated using PcFiml 9.0. For technical details, see Doornik and Hendry (1997, Chapter 10). AR(5) is an F-test against the hypothesis of 5:th order serial correlation, with p-values computed from the $F(5,60)$ distribution. The normality test is the omnibus statistic (cf. Doornik and Hansen, 1994), which is asymptotically χ^2 with 2 degrees of freedom. Finally, ARCH(4) is an F-test against the hypothesis of 4:th order ARCH, with p-values calculated from the $F(4,57)$ distribution. All p-values are given in percent.

TABLE 4: Estimated elements of the A matrix in the common trends model.

(A) Theoretical cointegration relations

Equation	Foreign real trend	Foreign nominal trend	Domestic real trend	Domestic nominal trend
y_t	.044	−.262	.730	.000
	(.663)	(.438)	(.156)	
p_t	−2.146	1.986	−.056	.712
	(3.204)	(1.168)	(.221)	(.162)
i_t	.000	.000	.000	.000
e_t	.370	.445	−.056	.712
	(.696)	(.498)	(.221)	(.162)
y_t^*	1.136	.000	.000	.000
	(.902)			
p_t^*	−2.517	1.541	.000	.000
	(2.988)	(.796)		
i_t^*	.000	.000	.000	.000

(B) Estimated cointegration relations

Equation	Foreign real trend	Foreign nominal trend	Domestic real trend	Domestic nominal trend
y_t	−.052	−.099	.577	.000
	(.193)	(.273)	(.100)	
p_t	.354	1.690	−.002	.339
	(.897)	(.766)	(.077)	(.050)
i_t	.014	.270	−.197	.092
	(.156)	(.155)	(.038)	(.013)
e_t	.107	.975	.036	.589
	(.555)	(.522)	(.134)	(.086)
y_t^*	1.000	.000	.000	.000
	(.278)			
p_t^*	−.394	1.387	.000	.000
	(.728)	(.606)		
i_t^*	.473	.044	−.057	.267
	(.167)	(.128)	(.061)	(.039)

NOTES: Estimated asymptotic standard errors within parentheses; see Warne (1993) for details.

TABLE 5: Forecast error variance decompositions for the common trends model with estimated cointegration vectors in percent.

(A) Forecast error variance of y_t

Quarters	ψ_{fr}	ψ_{fn}	ψ_{dr}	ψ_{dn}	ϕ_{di}	ϕ_{fi}	ϕ_{ad}
1	.71	4.68	22.80	3.64	.00	25.31	42.86
8	6.75	4.90	50.29	3.56	.76	11.97	21.77
16	5.42	4.46	63.62	2.53	.58	7.91	15.47
∞	.80	2.83	96.37	.00	.00	.00	.00

(B) Forecast error variance of p_t

Quarters	ψ_{fr}	ψ_{fn}	ψ_{dr}	ψ_{dn}	ϕ_{di}	ϕ_{fi}	ϕ_{ad}
1	28.02	47.94	3.56	12.80	2.16	1.66	3.86
8	16.18	73.92	1.01	7.62	.30	.22	.74
16	8.95	85.24	.40	4.88	.17	.09	.28
∞	4.03	92.25	.00	3.72	.00	.00	.00

(C) Forecast error variance of i_t

Quarters	ψ_{fr}	ψ_{fn}	ψ_{dr}	ψ_{dn}	ϕ_{di}	ϕ_{fi}	ϕ_{ad}
1	4.48	4.69	14.98	17.13	29.48	19.06	10.18
8	4.33	20.07	38.39	7.64	14.76	5.71	9.09
16	4.09	29.27	34.01	9.25	12.05	4.25	7.08
∞	.17	60.63	32.19	7.01	.00	.00	.00

(D) Forecast error variance of e_t

Quarters	ψ_{fr}	ψ_{fn}	ψ_{dr}	ψ_{dn}	ϕ_{di}	ϕ_{fi}	ϕ_{ad}
1	.93	6.19	50.18	1.40	13.01	28.18	.11
8	.84	27.78	37.31	18.89	3.77	8.79	2.60
16	1.25	35.54	27.11	24.88	2.84	6.46	1.92
∞	.87	72.54	.10	26.49	.00	.00	.00

NOTES: The trend innovations are denoted by ψ while the transitory innovations are denoted by ϕ. The subscript f denotes foreign, d stands for domestic, r for real, n for nominal, i for interest rate, while ad denotes aggregate demand.

TABLE 6: RMSE analysis of inflation forecasts.

(A) Naive models

Forecasting rule	Dynamic	Recursive
random walk	.645	.578
mean	1.619	1.540

(B) Unrestricted VEC models

Cointegration model	Dynamic	Recursive
Estimated (eq. 8)	.833	.864
Theoretical (eq. 6)	.706	.499

(C) Restricted VEC models

Cointegration model	Dynamic	Recursive
Estimated (eq. 8)	.509	.717
Theoretical (eq. 6)	.535	.486

NOTES: The dynamic forecasts are formed using models estimated over the sample period 1973:2–1993:4. The recursive 1-step ahead forecasts are based on an updating procedure where the models are reestimated each period on the maximum sample length prior to forecasting. The dynamic random walk rule means a no change forecast of inflation conditional on the information available in 1993:4. The recursive 1-step ahead random walk rule means a no change forecast of inflation conditional on the information available in the quarter immediately prior to forecasting. The dynamic mean rule is a forecast of inflation which is equal to the average value over the sample period 1973:2–1993:4. Finally, the recursive 1-step ahead mean rule is a forecast based on a recursively updated average value which is obtained using the maximum sample length available prior to forecasting.

TABLE 7: Cross correlations between 4 output gap measures and inflation for the sample period 1977:1–1995:4.

(A) Output gap correlations

Output gap	$\rho(\text{VEC}_t, \text{OG}_{t+s})$ -4	-3	-2	-1	0	1	2	3	4
UC	.69	.65	.60	.55	.47	.40	.35	.28	.21
HP	.58	.51	.43	.34	.23	.12	.06	$-.01$	$-.09$
PF	.58	.52	.46	.40	.31	.24	.22	.19	.14

(B) Inflation-output gap correlations

Output gap	$\rho(\Delta p_t, \text{OG}_{t+s})$ -4	-3	-2	-1	0	1	2	3	4
VEC	.43	.41	.41	.50	.60	.54	.49	.54	.61
UC	.28	.28	.28	.30	.28	.22	.22	.26	.18
HP	.26	.23	.22	.24	.18	.06	.07	.14	.04
PF	.23	.21	.22	.25	.16	.06	.10	.18	.10

NOTES: VEC is a smoothed 4-quarter moving average (no leads) estimate of the estimated demeaned transitory component in output when the theoretical cointegration vectors in (6) are used. UC is an unobserved components model estimate of the deviations from trend (cf. Apel and Jansson, 1997), HP is the deviation from trend using the Hodrick-Prescott filter, while PF is the deviation from trend using a production function approach to measuring the output gap (cf. Hansen, 1997). As a rule of thumb for determining significance, we may use the formula $1/\sqrt{T}$ as a proxy for the standard errors of the correlations. Here we obtain $1/\sqrt{76} \approx .12$, so that a value outside the interval $[-.24, .24]$ can be viewed as different from zero at the 5 percent level of marginal significance.

TABLE 8: Wald tests of the hypothesis that 3 lags of output gap measures are not useful for predicting inflation for the sample period 1976:4–1996:4.

(A) Theoretical cointegration relations

Output gap	Restricted model	Unrestricted model
HP	3.93	11.16
	[.27]	[.01]
UC	7.19	15.96
	[.07]	[.00]
PF	7.07	14.73
	[.07]	[.00]

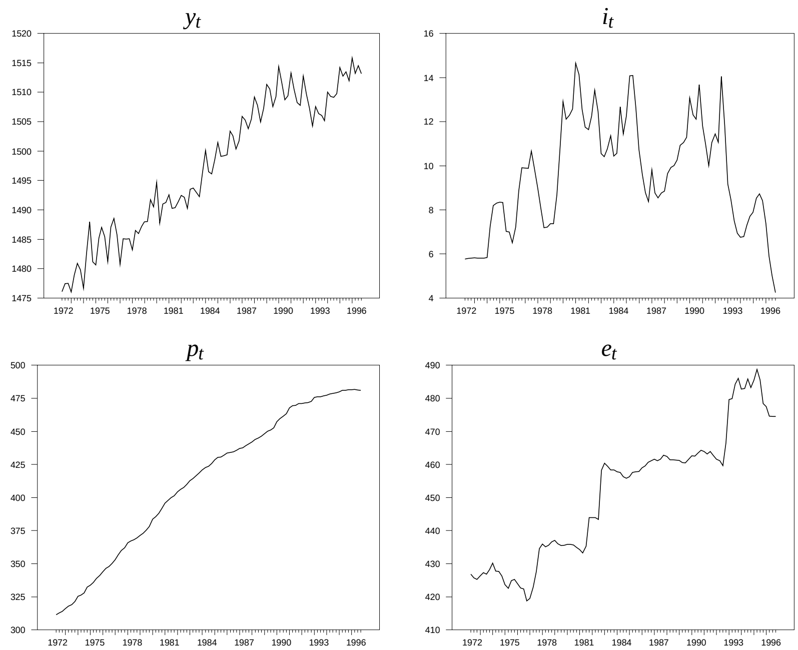
(B) Estimated cointegration relations

Output gap	Restricted model	Unrestricted model
HP	1.31	2.88
	[.73]	[.41]
UC	5.99	11.71
	[.11]	[.01]
PF	7.43	9.04
	[.06]	[.03]

NOTES: The reference distribution for the Wald statistic is χ^2 with 3 degrees of freedom (3 lags), and p-values are reported within brackets.

FIGURE 1: Time series observations of the 7 domestic and foreign variables in the vector x_t for the period 1972:2–1996:4.

(I) The domestic variables



(II) The foreign variables

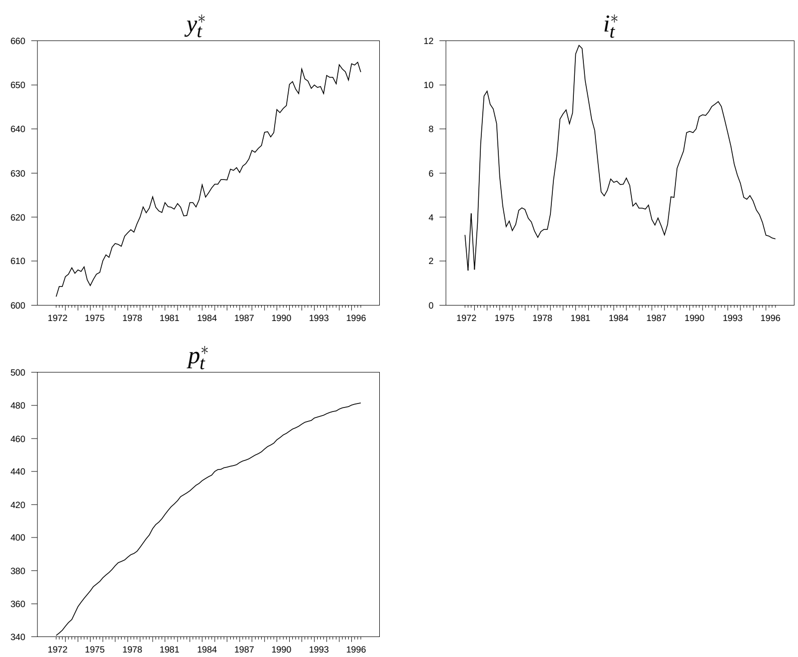


FIGURE 2: Effects on the domestic variables from a shock to the domestic nominal trend, with 95 percent confidence intervals using the estimated asymptotic distribution.

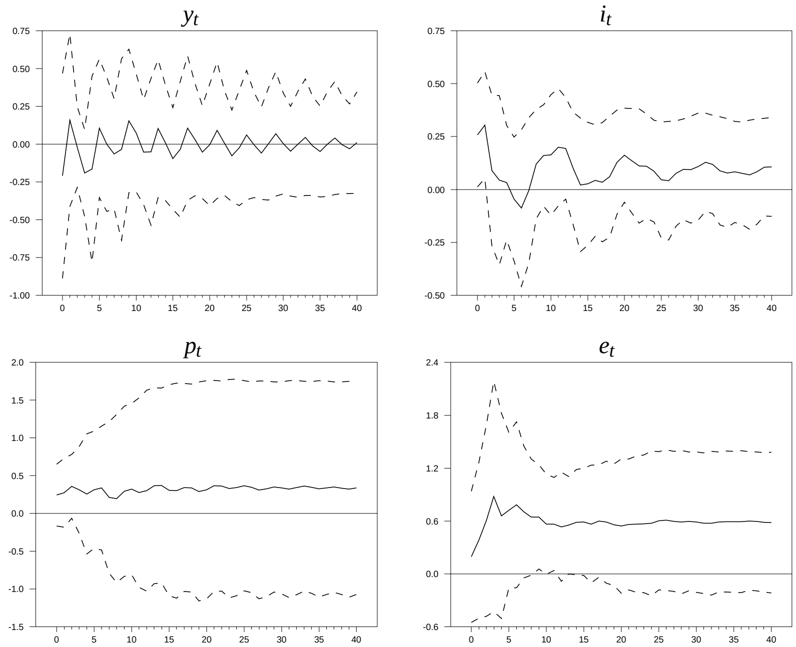


FIGURE 3: Effects on the domestic variables from a shock to the domestic interest rate, with 95 percent confidence intervals using the estimated asymptotic distribution.

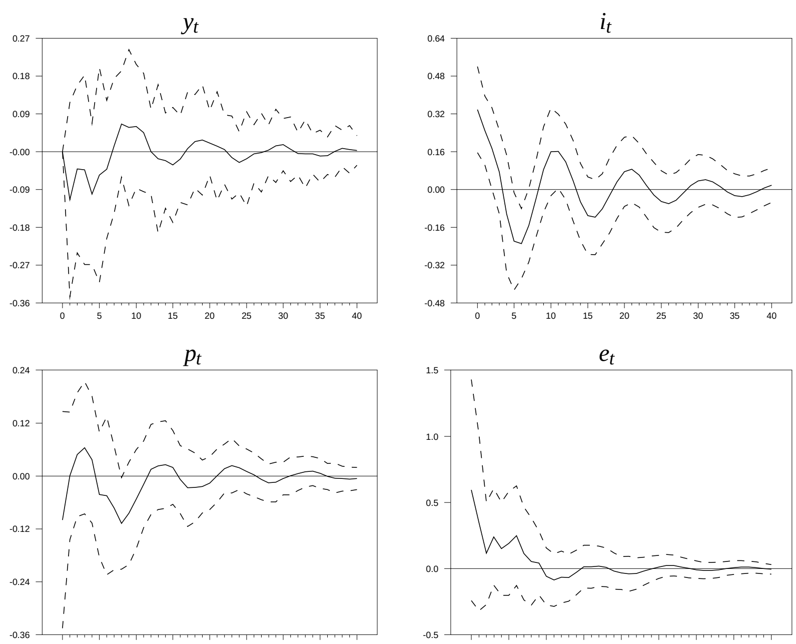


FIGURE 4: Effects on the domestic variables from a shock to aggregate demand, with 95 percent confidence intervals using the estimated asymptotic distribution.

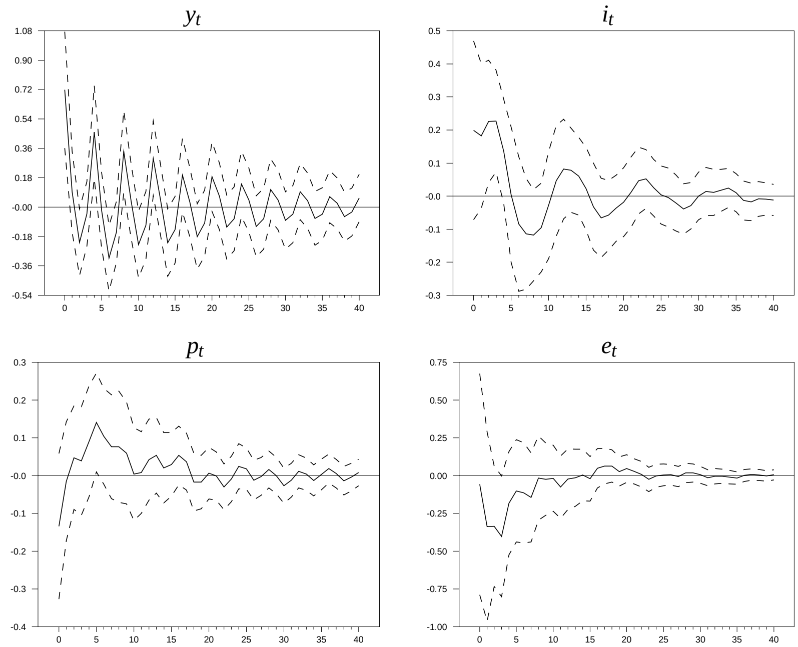


FIGURE 5: Inflation (solid line) and inflation forecasts 1994:1–1996:4 (dashed line) with estimated 95 percent confidence intervals using the empirical cointegration vectors in equation (8).

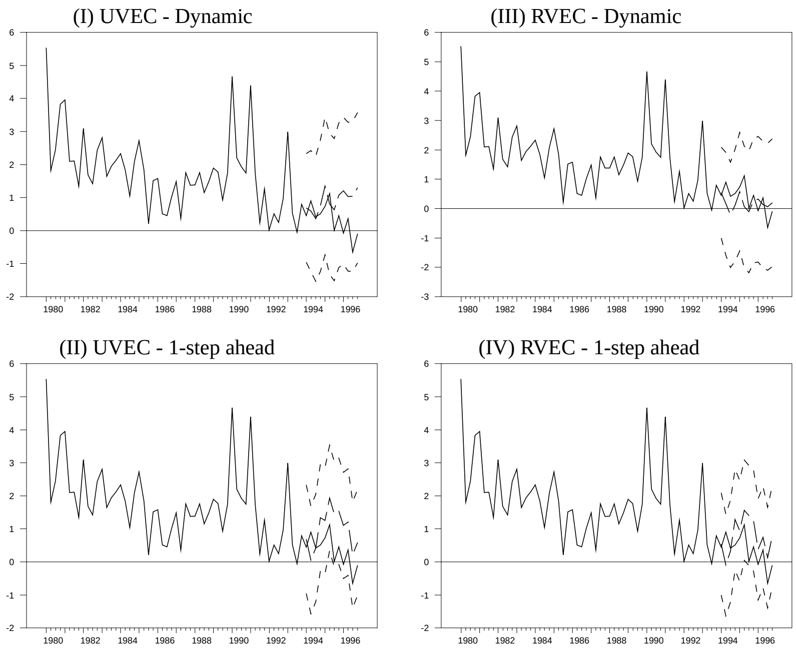


FIGURE 6: Inflation (solid line) and inflation forecasts 1994:1–1996:4 (dashed line) with estimated 95 percent confidence intervals using the theoretical cointegration vectors in equation (6).

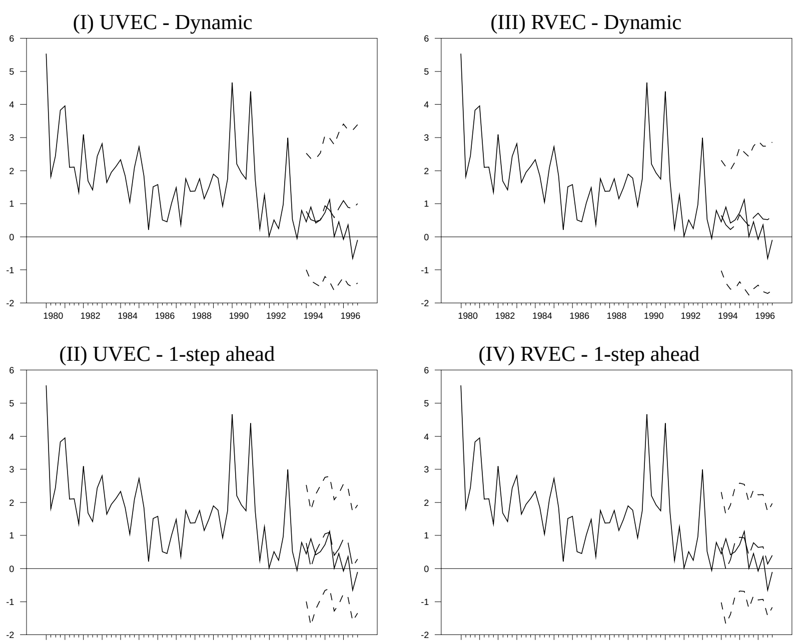


FIGURE 7: The real exchange rate and three estimates of the deviations from the long run goods market equilibrium.

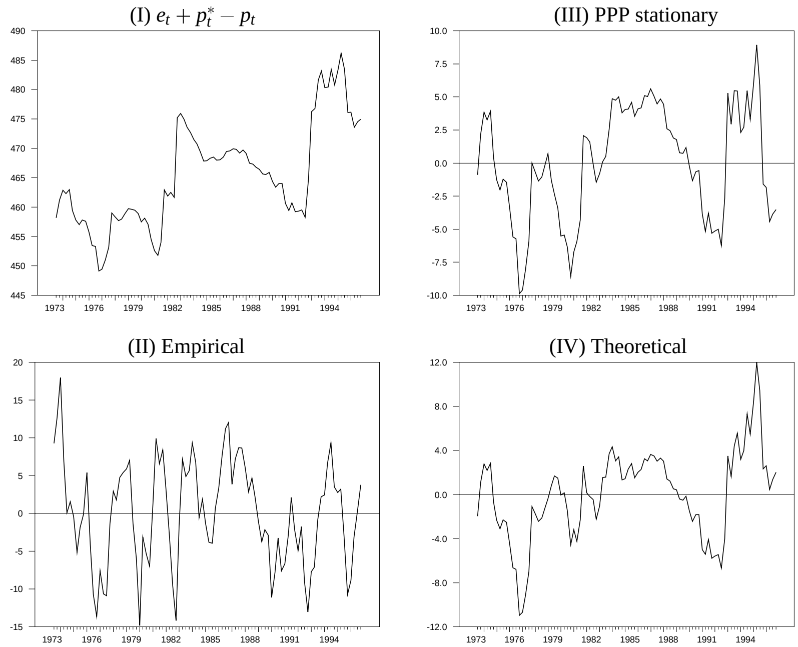
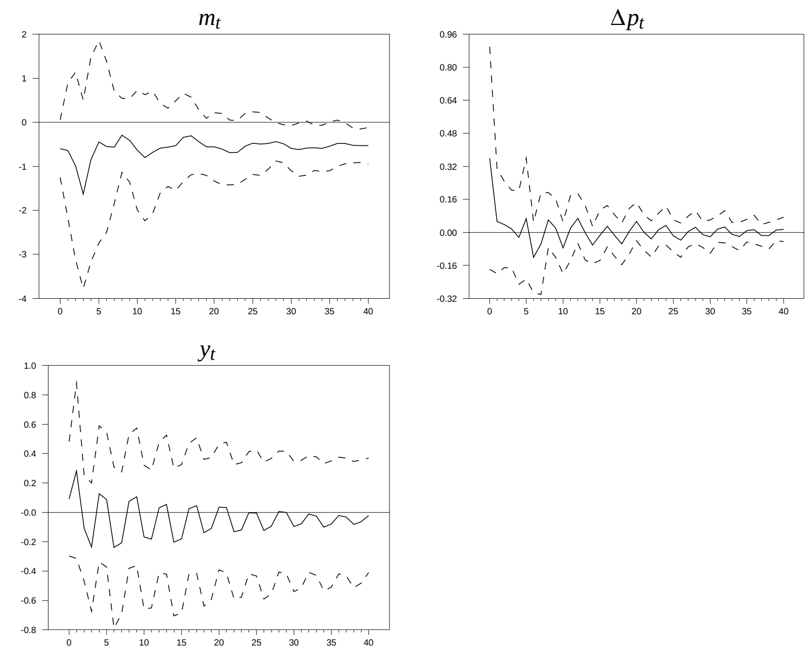


FIGURE 8: Effects on the MCI measure $m_t = \left(e_t + p_t^* - p_t - 2.875 i_t\right)$, output, and inflation from the structural shocks, with 95 percent confidence intervals using the estimated asymptotic distribution.

(I) The foreign real trend shock



(II) The foreign nominal shock

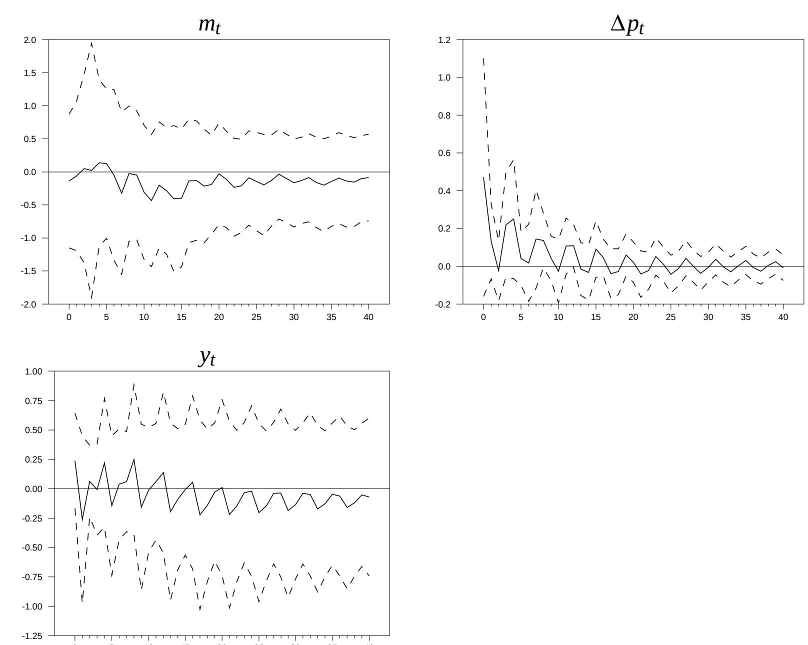
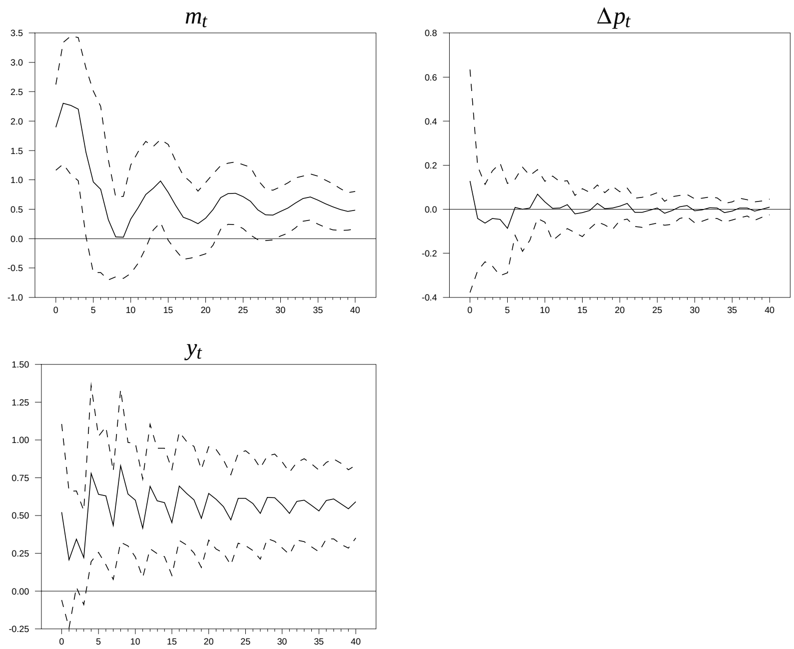


FIGURE 8: (Continued) Effects on the MCI measure $m_t = \left(e_t + p_t^* - p_t - 2.875 i_t\right)$, output, and inflation from the structural shocks, with 95 percent confidence intervals using the estimated asymptotic distribution.

(III) The domestic real trend shock



(IV) The domestic nominal shock

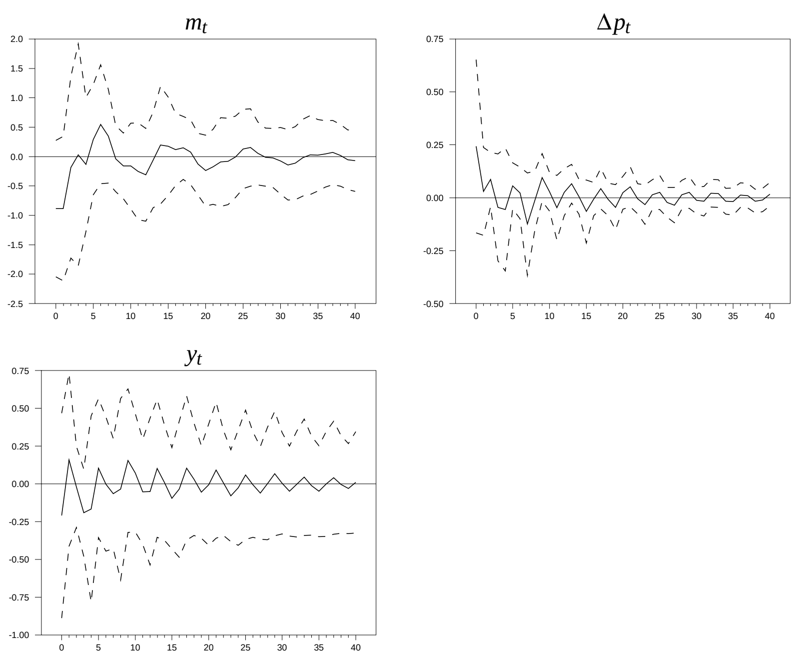
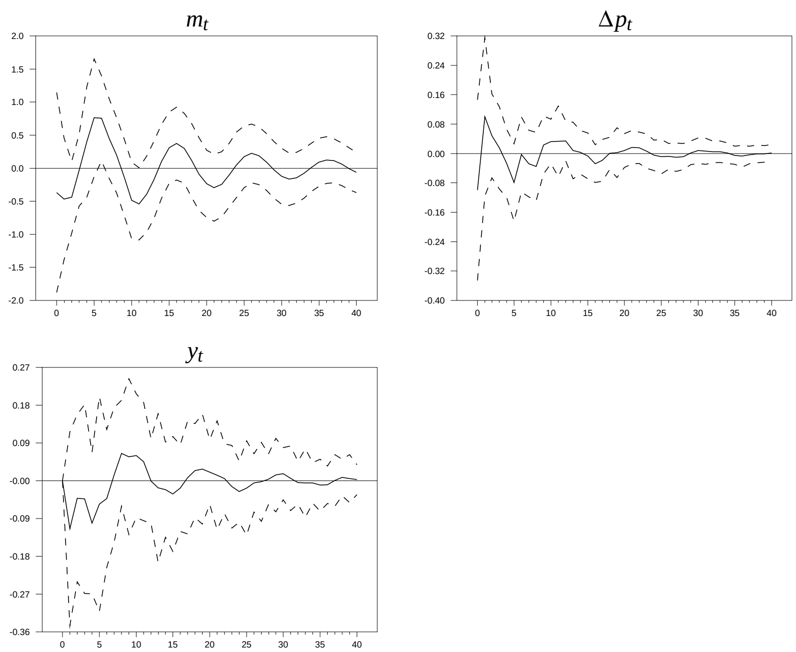


FIGURE 8: (Continued) Effects on the MCI measure $m_t = \left(e_t + p_t^* - p_t - 2.875 i_t\right)$, output, and inflation from the structural shocks, with 95 percent confidence intervals using the estimated asymptotic distribution.

(V) The domestic interest rate shock



(VI) The foreign interest shock

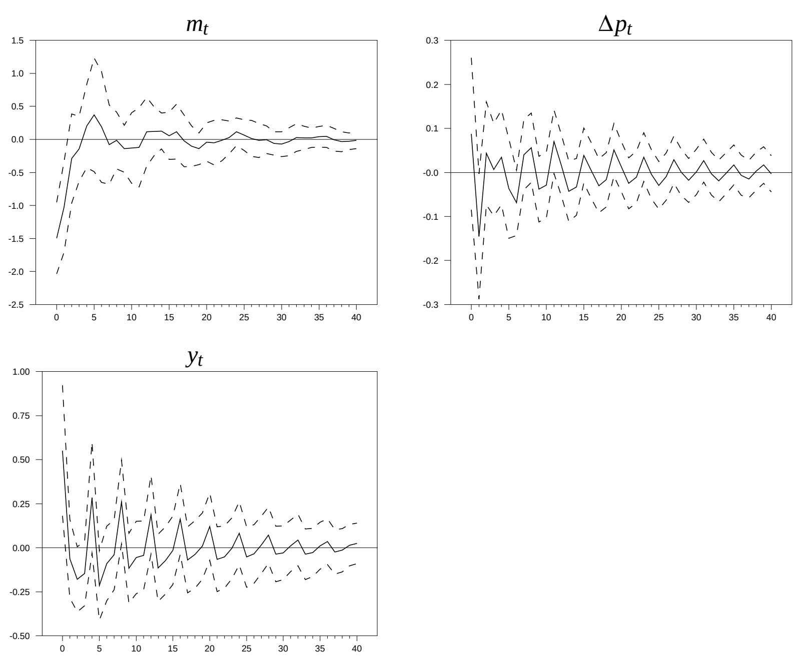


FIGURE 8: (Continued) Effects on the MCI measure $m_t = \left(e_t + p_t^* - p_t - 2.875 i_t\right)$, output, and inflation from the structural shocks, with 95 percent confidence intervals using the estimated asymptotic distribution.

(VII) The aggregate demand shock

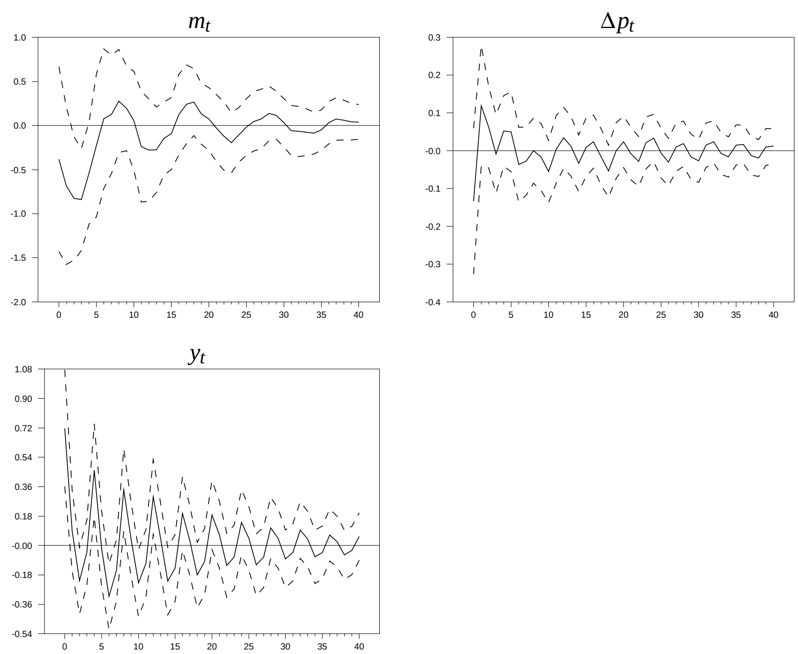
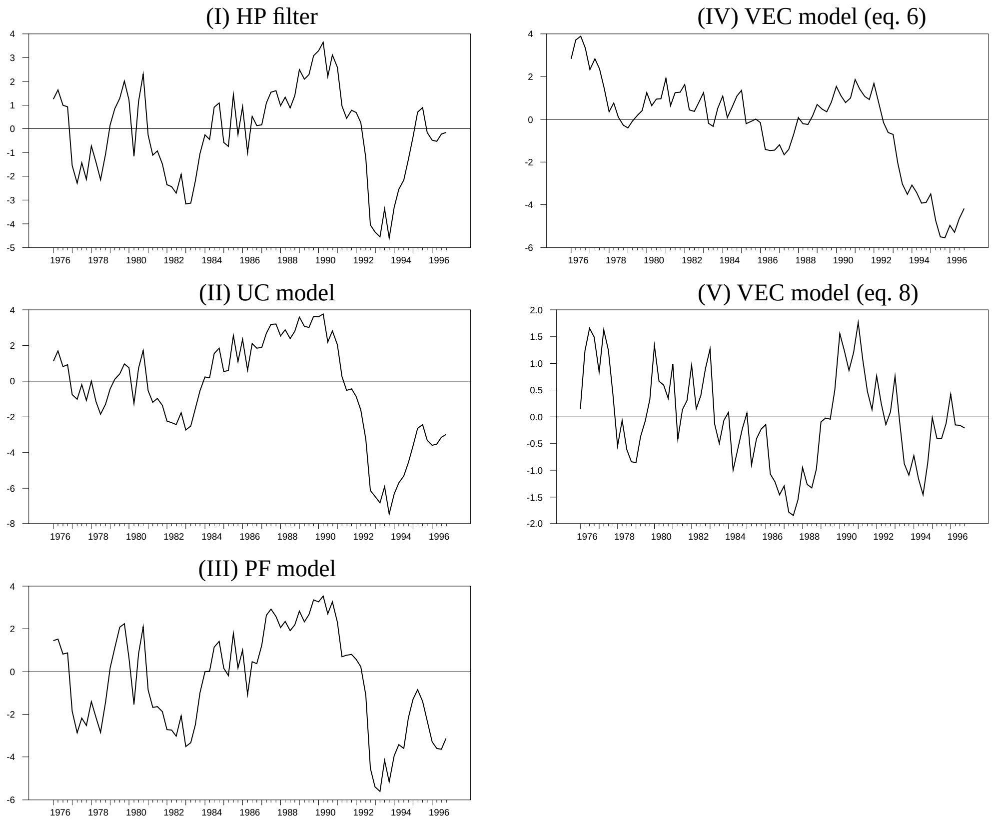


FIGURE 9: Estimates of the output gaps for the VEC models based on the theoretical and the estimated cointegration vectors and three alternative measures used by Sveriges Riksbank for the period 1976:1–1996:4.



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