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Identifying the Effects of Monetary Policy Shocks in an Open Economy

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Abstract: This paper presents estimates of the effects of monetary policy shocks on the Swedish economy. A theoretical model of an open economy is used to identify a structural VAR model. The empirical results from the identified VAR model are compared with two less structural approaches for identification of monetary policy shocks. The first assumes that shocks can be measured as deviations from a forward looking interest rate rule, estimated using Sveriges Riksbank's (Swedish central bank) own forecasts. The second approach focuses on the effects of "narrative" monetary policy shocks as given by devaluations of the Swedish currency. We find that plausible theoretical restrictions often result in price puzzles. Although conventional results obtain with certain theoretical restrictions imposed on the VAR, another way to achieve this is by using external information about large policy shocks. Thus, we find that the effects of some devaluations are consistent with the conventional wisdom about the effects of monetary policy shock.

Keywords: Common trends, devaluations, identification, inflation, monetary policy shocks, open economy, structural vector autoregression.

JEL Classification Numbers: C32, E31, E52.

1. Introduction

The effects of monetary policy on inflation and the real economy have been the subject of much research. While many issues remain debatable, there seems to be broad agreement on what monetary policy, at a general level, can and cannot do. Most economists would, for example, probably agree that persistent movements in inflation can be controlled by the monetary authority. Similarly, it is generally accepted that persistent movements in real variables such as GDP and unemployment are due to real shocks that are unrelated to monetary policy.

Meaningful policy recommendations require — explicit or implicit — estimates of how a change in monetary policy influences the rest of the economy. This is true irrespective of whether the policy question concerns, e.g., the appropriate objectives for monetary policy, or, the central bank's optimal response to certain disturbances to the economy.

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Remarks: Preliminary and incomplete versions of this paper have been circulated under the title "Monetary Policy, Inflation and Unemployment: A Common Trends Model". We have benefitted from comments by seminar participants at the Conference on Bridging Economics and Econometrics (European University Institute), the European Central Bank, Eesti Pank, the EEA Congress and the ESEM in Lausanne, and the Workshop on VAR Models for Monetary Policy (Sveriges Riksbank). Discussions with Lawrence Christiano, Carlo Favero, Katarina Juselius, Eric Leeper, Jesper Lindé, and Frank Smets have been particularly useful. David Cesarini and Daniel Hammarsten provided excellent reasearch assistance. The views expressed in this paper are solely the responsibility of the authors and should not be interpreted as reflecting the views of the Executive Board of Sveriges Riksbank. In addition, the views do not necessarily reflect those of the European Central Bank.

From empirical analyses, in particular vector autoregressive (VAR) models, and from theoretical work, e.g., dynamic general equilibrium models, there has emerged a relatively clear cut picture of the effects on the U. S. economy from changes in the Fed's monetary policy. It is often taken as a "stylized fact" that after an unexpected easening of monetary policy (manifested in a drop in the short term interest rate) output, inflation, and productivity go up. The responses are "hump shaped" and the maximum effects are recorded after about 1 1/2 to 2 years. Quantitatively, an initial decline in the short term interest rate by 1/2 percent (which gradually disappears during 8 quarters) has a maximum expected effect on real GDP of around 1/2 percentage point and on inflation of around 0.2 percentage points.[1] Although these estimates are very uncertain, there seems to be more agreement about the effects of monetary policy shocks for the relatively closed U. S. economy compared with more open economies.

The theoretical open economy models presented by Svensson (2000) and McCallum (2001) imply effects of monetary policy shocks that are in agreement with those reported for the U. S., although they suggest that the effects occur somewhat faster in an open economy. Figure 1 reproduces the impulse response in the domestic interest rate from Svensson (2000, Fig. 3). Moreover, the impulse responses for domestic prices and the consumer price level have been computed by accumulating the effects on CPI inflation and domestic price inflation. If the short term interest rate goes up by 1 percentage point for 4 periods, and slowly declines to its original level, it is found that the consumer price level immediately falls (because of an appreciation of the nominal exchange rate that is fully passed through to import prices), while the domestic goods price is unaffected for 2 periods and thereafter declines. We think that the shapes of these impulse responses are broadly in line with monetary policy makers' prejudices, as reflected, e.g., in inflation reports from central banks in New Zealand, Sweden, and the U. K.

However, Svensson's results are not in line with most results from VAR models of open economies. Some results that we believe are representative are reproduced in Figure 2.[2] The top left graph shows the effects of a domestic monetary policy shock on the U. S. price level, according to Dedola and Lippi (2001). The results are perfectly consistent with many earlier findings about the U. S. economy, and can be compared with the effects of the same experiment in VAR models of the U. K., Germany, France, and Italy, also taken from Dedola and Lippi (2001). We also have results about Germany from Hubrich and Vlaar (2000) and Bagliano, Favero, and Franco (1999), and about Canada, from Cushman and Zha (1997).

The results that Dedola and Lippi report for the U. K. are similar to their findings for the U. S., although the "price puzzle", i.e., the initial increase in prices after a contractionary monetary policy shock, lasts longer. The U. K. price level does not decrease, persistently, until after 10 months, and after two years it is still as high as initially. The price puzzle effect is even stronger in the cases of France and Italy, and for Germany the price level stays above its baseline value for more than 4

[1] These results are from Christiano, Eichenbaum, and Evans (2001). Alternative identification schemes, sometimes with other results, have been presented by Leeper, Sims, and Zha (1996), Faust (1997), and Uhlig (2001).

[2] We gratefully acknowledge the help from Luca Dedola, Carlo Favero, Marianne Nessén, Lars Svensson, Peter Vlaar, and Tao Zha that made the reproduction of the graphs in Figures 1 and 2 possible.

years. Bagliano et al. (1999) also report that the price level in Germany is expected to rise in the long run after a temporary increase in the short term interest rate, although there is no price puzzle for the first 16 months. This is in contrast with the findings by Hubrich and Vlaar (2000), who report that inflation falls in Germany only in the first quarter after a contractionary monetary policy shock. Finally, looking at the results for Canada given by Cushman and Zha (1997), the development of the price level after a contractionary monetary policy shock looks similar to the stylized facts about the U. S.; but note that the experiment here is somewhat different, since the interest rate goes up only for a few months.

Judging from these results, it is difficult to argue that the effects of a monetary policy shock in an open economy can be approximated by the conventional wisdom, as partly given by theoretical models like those of Svensson (2000) and McCallum (2001), and partly by theoretical and empirical findings for the U. S. economy. In fact, it seems hard to get empirical support even for the very reasonable idea that a contractionary monetary policy shock leads to lower inflation.

The purpose of this paper is to examine the gap between the theoretical analyses of monetary policy in open economies and the empirical results from VAR models. We develop a structural VAR model and apply it to data from Sweden. The VAR model is structural in the sense that it is subject to (a relatively small set of) restrictions that will enable identification of various shocks and their individual effects on the Swedish economy. However, there are good reasons to believe that restrictions that have been imposed on VAR models of closed economies are less suitable for open economy analyses. In particular, the standard assumptions that a monetary policy shock has an immediate effect on the domestic interest rate, but neither on inflation, nor on output, are presumably not sufficient. One reason is that the domestic interest rate is also affected, immediately, by shocks to foreign interest rates and other disturbances on international financial markets.

The theoretical model that will be used to interpret various restrictions on the VAR model is presented in Section 2. Certain features of this model should be stressed right from the outset. First, its primary purpose is to entail a better understanding of alternative identifying restrictions on the VAR model. Hence, we present a solution of the theoretical model on a so called common trends form, a particular representation of structural VAR models pioneered by Shapiro and Watson (1988) and King, Plosser, Stock, and Watson (1991). The solution has the property that all domestic shocks influence some of the variables in the long run, thus making the usual dichotomy between permanent and transitory shocks invalid. We emphasize the theoretical model's long run properties. Dynamic relations — which may be empirically important but not essential for understanding identifying restrictions or long run relations — are largely overlooked in the theoretical analyses. Second, the properties of the model that we make use of are consistent with a broad class of open economy models. Formally, it combines features of earlier models of European labor markets applied by, e.g., Bean, Layard, and Nickell (1986), Blanchard (2000), and Bean (1994), with models of aggregate demand, price setting and monetary policy presented by, e.g., Svensson (2000), McCallum (2001), and Galí and Monacelli (1999). In contrast to the latter models, however, our model allows for incomplete exchange rate pass through (as in Adolfson, 2001, and Obstfeld and Rogoff, 2000) and also allows for analyses of the effects of monetary policy on unemployment.

The first empirical analyses of the VAR model are presented in Section 3. The discussion focuses on the problem of attaining a well specified empirical model with steady state (cointegration) relations that are interpretable using the theoretical model.

In Section 4 we investigate the effects of various shocks to the economy. Our attention is on the effects of monetary policy shocks and on identifying restrictions that both resemble those that have been applied in earlier VAR analyses, and that can be interpreted using our theoretical model. This involves combining long run restrictions of the type advocated by Shapiro and Watson (1988), Blanchard and Quah (1989), and King et al. (1991) with restrictions on contemporaneous relations that have been more common in structural VAR models of the U. S. economy. In contrast to these studies, our long run restrictions allow all shocks to have permanent effects on some variable. Thus, the identification procedure advocated by, e.g., King et al. (1991) is not applicable in our case and instead we rely on an extension for identifying and estimating VAR models with contemporaneous and long run restrictions that has been developed by Vlaar (1998).

In Section 5 we examine two alternative analyses of monetary policy shocks. The idea that actual policy shocks can be identified in VAR models has been questioned by Rudebusch (1998), among others. We therefore compare the VAR results from Section 4 with alternative measures of monetary policy shocks. In Section 5.1 we identify monetary policy shocks by estimating a forward looking interest rate rule on Swedish data. We then calculate impulse response functions from such shocks when they are fed into the VAR model from Section 3. Similar approaches have been applied by Blanchard and Watson (1986), to identify the effects of fiscal policy shocks, and by Bagliano et al. (1999), to derive the effects reported in Figure 2 above. Finally, in Section 5.3, we take a look at the effects of policy interventions using a "narrative" approach similar to Ramey and Shapiro's (1998) and Eichenbaum, Fisher, and Edelberg's (1998) analyses of effects on U. S. government purchases from exogenous shocks as given by events of large military buildups. The main conclusions are discussed in Section 6.

2. A Theoretical Framework for Monetary Policy in a Small Open Economy

In this section we first present and then solve numerically a simple macroeconomic model for a small open economy. The main purpose is to derive (i) a set of hypotheses about long run (cointegration) relations between the variables of interest, and (ii) a set of restrictions which may be used to identify monetary policy shocks from the residuals of a VAR model. The former will be examined in our empirical analysis of Swedish quarterly data in Section 3. Section 4 makes use of both the cointegration relations and the identifying assumptions.

2.1. *The Model*

Our model is basically a model of the labor market augmented by a model of aggregate demand, price setting, and monetary policy. The labor market model consists of equations describing labor demand, labor supply, and wage setting. This approach builds on earlier work on unemployment in

European countries by, e.g., Bean et al. (1986); see Bean (1994) for a survey.[3] The model of aggregate demand, price setting, and monetary policy builds on the recent literature on inflation targeting in open economies (see McCallum and Nelson, 1999, McCallum, 2001, Svensson, 2000, Galí and Monacelli, 1999, Walsh, 1999, and Adolfson, 2001).

Production possibilities are described by the function

$$y_t = \tau_t^T + \kappa e_t, \qquad t = 1, 2, \ldots. \tag{1}$$

Here, y_t is output, e_t is employment, the parameter κ measures returns to scale, while τ_t^T is a stochastic technology variable. All variables are expressed in natural logarithms. It is assumed that technology is exogenous and evolves according to

$$\tau_t^T = \mu^T + \rho_T \tau_{t-1}^T + \varepsilon_t^T, \qquad t = 1, 2, \ldots,$$

where ε_t^T is a pure technology innovation. In this section, ε_t^i will be used to denote i.i.d. shocks with zero mean, and covariance σ_{ij}, where the covariance is positive if $i = j$ and zero otherwhise.

We assume that the labor force (labor supply), l_t, is related to the real wage according to:

$$l_t = \eta(w_t - \nu_{y,t} - p_t) + \tau_t^L, \qquad t = 1, 2, \ldots, \tag{2}$$

where $w_t - \nu_{y,t}$ is the nominal after-tax wage and p_t the consumer price index.[4] Here, τ_t^L is a stochastic component of labor supply which follows the process:

$$\tau_t^L = \mu^L + \rho_L \tau_{t-1}^L + \varepsilon_t^L, \qquad t = 1, 2, \ldots.$$

Empirically, it is commonly observed that labor supply appears to have no trend, although productivity and the real wage do, which is often interpreted as reflecting that income and substitution effects on labor supply balance each other. In our model, this could arise if the labor supply elasticity $\eta = 0$ and $|\rho_L| < 1$, or if the stochastic trend in the real wage is proportional to the stochastic trend in labor supply, $\rho_L = 1$.

A wage setting relation is used to describe the fact that the labor market is not perfectly competitive. The nominal wage is supposed to be determined by the price level, productivity, unemployment, and taxes according to:

$$\begin{aligned} w_t = {} & p_{t_W} + \theta_1(y_{t_W} - e_{t_W}) - \theta_2(l_{t_W} - e_{t_W}) \\ & + \theta_3(\nu_{y,t_W} + \nu_{p,t_W} + \nu_{i,t_W}) + \tau_t^W, \qquad t = 1, 2, \ldots, \end{aligned} \tag{3}$$

where $t_W \in \{t, t-1\}$, $\nu_{p,t}$ is the pay roll tax rate, and $\nu_{i,t}$ is the indirect tax rate (VAT less subsidies, etc.).[5] This equation should primarily be interpreted as capturing the long run (average) relation between the real wage, productivity, unemployment, and the tax wedge (the ratio between the real wage that is relevant for the producer and the consumer, respectively). It may be viewed as the

[3] For earlier applications of common trends models to labor market issues and further references, see Jacobson, Vredin, and Warne (1997, 1998).

[4] Let $T_{y,t}$ denote the income tax rate. Then $\nu_{y,t} = -\ln(1 - T_{y,t})$.

[5] Let $W_t(1 + T_{p,t})$ denote the nominal wage including pay roll taxes and $P_t^d(1 + T_{i,t})$ denote the price of domestic goods including indirect taxes. Then $\nu_{p,t} = \ln(1 + T_{p,t})$ and $\nu_{i,t} = \ln(1 + T_{i,t})$.

outcome of a wage bargaining process. In the short run, nominal and real wages will presumably be influenced by expectations, and expectational errors, e.g., because of staggered contracts. In equation (3) we allow for "wage stickiness" when $t_W = t - 1$. In our empirical model, the estimated wage equations will have rich dynamics allowing for more complex mechanisms, but the simple wage relation (3) suffices for the discussions here.

The exogenous stochastic shock process to wage setting is given by:

$$\tau_t^W = \mu^W + \rho_W \tau_{t-1}^W + \varepsilon_t^W, \qquad t = 1, 2, \ldots.$$

Shocks to ε_t^W, may be interpreted as reflecting changes in the equilibrium level of unemployment.[6] Since unemployment may be influenced in the long run by other shocks as well, we shall instead refer to this shock as a wage shock. In any event, the parameter ρ_W may be interpreted as an overall measure of rigidities in the labor market.

Domestic producers set prices as a mark-up on labor costs. Letting p_t^d denote the price of domestic goods, we assume the following price setting relation

$$p_t^d = \psi(w_{t_p} + v_{p,t_p} + e_{t_p} - y_{t_p}) + \varepsilon_t^p, \qquad t = 1, 2, \ldots, \tag{4}$$

where $t_p \in \{t, t-1\}$ and the price setting (or "cost push") shock is assumed to have zero mean. As in the wage setting equation (3), we here allow for sticky domestic prices when the parameter $t_p = t - 1$. Again, the empirical model will allow for richer dynamics, along the lines suggested by, e.g., Galí and Gertler (1999), but the simple dynamics allowed for in relation (4) suffices for a discussion of our model's long run properties as well as for the identification of shocks. As noted by Galí and Gertler, a mark-up pricing relation like (4) is more directly linked to microeconomic theories of firms' behavior than the commonly used aggregate supply relations between inflation and some measure of the deviation between the actual and steady state levels of output (the output gap). It has also been noted by Walsh (1999), that an open economy aggregate supply relation needs to take exchange rate fluctuations into account, since labor supply (and, hence, wage costs) will be affected by the exchange rate (through the price of imports). This is true also in our model.

Prices of imports, p_t^m, are determined by the price level in the rest of the world, p_t^f, and the nominal exchange rate (price of foreign currency in terms of domestic currency), s_t, but also by the relation between activity (output) in the domestic economy in relation to the rest of the world:

$$p_t^m = \xi_1(p_{t_m}^f + s_{t_m}) + \xi_2 y_{t_m} - \xi_3 y_{t_m}^f + \tau_t^m, \qquad t = 1, 2, \ldots, \tag{5}$$

where $t_m \in \{t, t-1\}$ and

$$\tau_t^m = \mu^m + \rho_m \tau_{t-1}^m + \varepsilon_t^m, \qquad t = 1, 2, \ldots.$$

[6] Defining this level as $\tau_t^* = \tau_t^W / \theta_2$, equation (3) can be rewritten as:

$$\begin{aligned} w_t - p_t = p_{t_W} - p_t + \theta_1(y_{t_W} - e_{t_W}) - \theta_2(l_{t_W} - e_{t_W} - \tau_{t_W}^*) \\ + \theta_3(v_{y,t_W} + v_{p,t_W} + v_{i,t_W}) + \theta_2(\tau_t^* - \tau_{t_W}^*). \end{aligned}$$

This suggests that what matters for the real wage is not the level of actual unemployment per se, but its level in relation to some perceived equilibrium level of unemployment. As Barro and Gordon (1983) and Ireland (1999), among many others, we treat this possible equilibrium level of unemployment as an exogenous variable. It is time varying, however, with shocks denoted by ε_t^W.

This relation is intended to allow for pricing-to-market behavior, which may be relevant, not only in the short run (outside steady state); see, e.g., the survey by Goldberg and Knetter (1997).[7] Also, it allows for import price stickiness when $t_m = t-1$. In the empirical application, we will of course allow both the law of one price as one special case ($\xi_1 = 1, \xi_2 = \xi_3 = \rho_m = 0$) of the long run solution, and for more complex dynamics.[8]

The aggregate demand function for domestic goods depends on income (output) in the rest of the world, the real interest rate, and the relative price between domestic and foreign goods. Because this relative price may differ between the domestic market and the world market (due to potential pricing-to-market of imports), two relative price relations may influence aggregate demand:

$$\begin{aligned} y_t = {} & \gamma_1 y^f_{t_y} + \gamma_2 (p^m_{t_y} - p^d_{t_y}) + \gamma_3 (p^f_{t_y} + s_{t_y} - p^d_{t_y}) \\ & - \gamma_4 (i_{t_y} - E_{t_y}[p_{t_y+1} - p_{t_y}]) + \varepsilon^y_t, \qquad t = 1, 2, \ldots, \end{aligned} \tag{6}$$

where $t_y \in \{t, t-1\}$.

The aggregate consumer price index is simply an average of the prices of domestic goods and imports (including indirect taxes):

$$p_t = \delta p^d_t + (1-\delta) p^m_t + \nu_{i,t}, \qquad t = 1, 2, \ldots. \tag{7}$$

The financial markets equilibrium condition is a modified version of uncovered interest parity,

$$i_t = i^f_t + E_t[s_{t+1} - s_t] + \zeta d_{\mathrm{fm},t} + \tau^I_t, \qquad t = 1, 2, \ldots, \tag{8}$$

where i^f_t denotes the foreign nominal interest rate, and

$$\tau^I_t = \mu^I + \rho_I \tau^I_{t-1} + \varepsilon^I_t, \qquad t = 1, 2, \ldots.$$

The intercept dummy $d_{\mathrm{fm},t}$ and the stochastic component τ^I_t are intended to capture deviations from UIP that are related to credit market imperfections (such as capital controls, which were abandoned in Sweden during the 1980s) and risk premia (which are believed to be relevant, but also hard to model).

Finally, monetary policy is described by a Taylor-type rule:

$$\begin{aligned} i_t = {} & \phi_1 (p_t - p_{t-1}) + \phi_2 [w_t + \nu_{p,t} + e_t - y_t - p^d_t] \\ & + \phi_3 i_{t-1} + \phi_4 d_{\mathrm{mp},t} + \tau^M_t, \qquad t = 1, 2, \ldots, \end{aligned} \tag{9}$$

where

$$\tau^M_t = \mu^M + \rho_M \tau^M_{t-1} + \varepsilon^M_t, \qquad t = 1, 2, \ldots.$$

This formulation differs from common Taylor-type rules in a couple of ways. Monetary policy is assumed to respond to the wage share rather than the output gap. Although this model makes it possible to distinguish between transitory and permanent movements in output, it would have been technically rather complicated to include the transitory component of output in the central

[7] See Obstfeld and Rogoff (2000) for a critique of the pricing-to-market assumption.

[8] Note that domestic exporters, unlike foreign exporters, are not allowed to be pricing-to-market in the long run.

bank's reaction function. Therefore, we have chosen to exploit the idea that the wage share, rather than the output gap, enters the reaction function of monetary policy (see, e.g., Galí and Gertler, 1999). In addition, regime shifts in monetary policy are assumed to be captured by the intercept dummy $d_{\mathrm{mp},t}$ and the stochastic component τ_t^M, while ϕ_3 reflects possible interest rate smoothing objectives.

This model has eight endogenous variables – y, e, l, w, p (or p^d), s, i, and p^m – which are driven by eight shock processes and eight exogenous variables – ν_y , ν_p , ν_i, p^f, y^f, i^f, d_{fm}, and d_{mp}. There are six potential "domestic" common stochastic trends that may give rise to unit roots: there may be a stochastic trend in technology ($\rho_T = 1$), labor supply ($\rho_L = 1$), equilibrium unemployment ($\rho_W = 1$), import prices ($\rho_m = 1$), the risk premium ($\rho_I = 1$), and/or monetary policy ($\rho_M = 1$). We have assumed that the price setting relation (4) and the aggregate demand relation (6) are stationary, but further cointegration relations are possible, depending on how many common stochastic trends we find. Hence, the model suggests that there may be from two and up to eight cointegration relations.[9]

2.2. *Solving the Model*

We have solved the model numerically using Sims's (2000) `gensys` MatLab function.[10] The solution of the model is conditioned on assumptions about its parameters. In particular, we have to express the model in terms of assumed stationary variables only, which involves making assumptions about the ρ_i parameters. When choosing numerical values for the parameters, the following circumstances have been important:

(i) for this particular theoretical model, certain parameter values are inconsistent with the existence of a unique solution;

(ii) to identify the empirical counterparts to the shocks in the theoretical model, certain parameter restrictions will be needed; and

(iii) the empirical results reported in Section 3 below provide suggestions about the stationary (cointegration) relations.

We have thus calculated a benchmark solution based on certain numerical values for the parameters. These are given in Table 1, and the variable transformations used to solve the model are discussed in the Appendix. The resulting long run and contemporaneous responses of the endogenous variables are presented in Table 2, Panel A and Panel B, respectively. Some assumptions are absolutely critical for the model's qualitative properties, while others only affect the quantitative results.

Real Stochastic Trends

We have assumed that there are 3 stochastic trends arising from technology, labor supply, and import prices, i.e. $\rho_T = \rho_L = \rho_m = 1$. This means that there can be at most 5 cointegration relations. These assumptions are partly based on the empirical results presented in Section 3 and

[9] The number of "domestic" common stochastic trends is equal to the number of endogenous variables less the number of cointegration relations; see Stock and Watson (1988).

[10] Sims' software is available for download from: `http://www.princeton.edu/~sims/`. See also Klein (1999) and references therein for alternative ways to solve linear rational expectations models.

partly on what we think are the most relevant candidates. For simplicity we have also assumed that the other ρ parameters are zero, but this is inconsequential.

Stationary Wage Share

We have also assumed that $\psi = 1$, such that the wage share (from the firms' perspective) is stationary. This implies (from the reaction function (9)) that the domestic interest rate is stationary, given that $|\phi_3| < 1$, and inflation is stationary.

Stationary Foreign Interest Rate

In order to treat the integration properties of the domestic and foreign interest rates symmetrically, we assume that the foreign interest rate is stationary. Together with the previous assumption, this implies, through equation (8), that the first difference of the nominal exchange rate is stationary.

Nominal Long Run Neutrality

Next, we assume that $\xi_1 = 1$. The importance of this assumption stems from the fact that except for the domestic price and import price equations, all other equations include only relative prices. Imposing $\psi = \xi_1 = 1$ implies that real variables are only linked to relative prices and *not* to the level of any individual price variable.

Price and Wage Stickiness

In the absence of wage or price rigidities there is not really an interesting role for monetary policy to play. For that reason, as well as for the reason of allowing for some "realistic" short run behavior and for identification purposes, we assume that $t_W = t_p = t_m = t - 1$. In addition, we suppose that $t_y = t - 1$, which implies that output is only allowed to respond to monetary policy shocks with a lag. These assumptions have only minor effects on the long run impulse responses; they affect the long run responses in nominal variables, but not in real variables.

The Nominal Stochastic Trend

Note that our assumptions imply that the long run development of output, employment, the labor force (and, hence, unemployment), and all relative prices are driven by the 3 stochastic trends to technology, labor supply, and import prices. It is natural to think of the stochastic trend in import prices as reflecting terms-of-trade shocks. In contrast, shocks to monetary policy, aggregate demand, domestic price setting, the modified UIP relation, and wage setting all have permanent effects on nominal variables, but not on real variables. There is thus a fourth common trend in our model, which is a weighted sum of the 5 shocks that only influence nominal variables in the long run. We shall refer to this trend as the nominal trend, although shocks to, e.g., aggregate demand may not be pure nominal shocks.

Discussion

The contemporaneous effects of unit domestic shocks are given in Panel B of Table 2. For example, a unit shock to monetary policy raises the interest rate by 1 unit since domestic and import prices are sticky (inflation does not change) and the wage share only reacts directly to wage, price, and technology shocks, i.e., the output gap measure is unaffected. As a consequence of the financial market relation (8), the expected change in the exchange rate increases by 1 unit. The exchange rate itself is temporarily strengthened and overshoots the long run effect.[11] The overshooting effect is also reflected in inflation expectations, where prices in the following period are expected to drop more than in the long run.

Due to lags in aggregate demand, output only responds contemporaneously to demand shocks. As a consequence of the production function in equation (1), employment therefore falls by 1 unit from a unit shock to technology. Because of price and wage rigidities, it follows that unemployment increases by an equal amount from such shocks, i.e., labor supply does not respond. The immediate effect on the domestic interest rate from technology shocks is therefore a drop given by the value of ϕ_2 (the weight on the wage share in the reaction function). The result that the domestic interest rate does not respond to the price shock contemporaneously is only due to our choice of parameter values.[12]

Sensitivity Analysis

We have also examined how the benchmark results in Table 2, Panels A and B, change when we vary the parameters in Table 1 one by one. As noted above, $\psi \neq 1$ or $\xi_1 \neq 1$ implies that nominal long run neutrality does not hold. Different assumptions about the sources of the common trends (i.e., setting other ρ_i's to unity) would of course change the model fundamentally. The numerical values of most other parameters are much less important, and changes in the assumptions affect the model's properties in quite intuitive ways.

For instance, the parameters γ_4 and the ϕ_i's do not matter for the long run responses of the real variables. Hence, neither the systematic not the unsystematic part of monetary policy matters for the trending behavior of the real variables. Changing the assumptions about the central bank's reaction function do, of course, affect nominal variables both in the short and in the long run. For instance, increasing ϕ_1 has the desired effect of dampening the effects of all shocks on the price level. A higher degree of interest rate smoothing changes the relative importance of shocks comprising the nominal trend.

Alternative parameterizations of the wage setting relation also affect the results in very intuitive ways. For instance, if wage setting responds less to productivity (lower θ_1), technology shocks have a smaller positive effect on output, but larger effects on unemployment. Increasing the response to

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[11] The overshooting effect is a consequence of the effects on the expected change being 1 and minus the long run response being less than one half. For the benchmark model, minus the contemporaneous effect on the nominal exchange rate from a monetary policy, demand, or a financial market shock is equal to the contemporaneous effect on the expected change in the nominal exchange rate plus the long run effect on the nominal exchange rate. Once we change the ϕ_i parameters or δ, the relation between the contemporaneous response in the nominal exchange rate and the expected change and the long run response in the nominal exchange rate from these shocks becomes more complex.

[12] The immediate effect on the domestic interest rate from a price shock is equal to $\phi_1\delta - \phi_2$. Hence, if we increase the value of ϕ_1 or δ, then price shocks will have a positive effect, while raising the value of ϕ_2 means that such shocks will have a negative effect on the interest rate.

unemployment (higher θ_2) has similar effects. This is worth noting, because our empirical results suggest that θ_1 may be lower, and θ_2 higher, than assumed in Table 1.[13]

In the next section we shall the theoretical model's steady state (cointegration) relations and number of common trends on Swedish quarterly data. The model's potential for identification of monetary policy shocks will be examined in Section 4.

3. An Empirical VAR Model with Cointegrated Variables

3.1. *Stationary Relations: Preliminary Considerations*

The solution of the theoretical model presented above was derived under the assumption that there are $r = 4$ cointegration relations. Typically, these relations are complicated functions of the parameters of the model. But, in some cases, the stationary relations can be directly observed in the data. In this subsection we will undertake a simple graphical analysis of some relations to see how well they fit the assumptions made in the theoretical model. Here and in what follows, the empirical analysis is undertaken using Swedish quarterly data running from the first quarter 1970 to the last quarter 1999.

First, consider the theoretical price setting relation (4). It implies that the wage share from the firms' point of view is stationary provided that $\psi = 1$. In Figure 3 we find graphs of the development of the wage share both from the firms' and workers' point of view.[14] The difference between these two measures of the wage share depends on terms of trade $p_t^d - p_t^m$, the income tax $v_{y,t}$, the pay roll tax $v_{p,t}$, and the indirect tax $v_{i,t}$. Graphs of these taxes, as well as of the terms of trade ($p_t^x - p_t^m$, where p_t^x is the log of the export price index), are shown in Figure 3. As can be seen, the wage share from the workers' point of view does not appear to have a constant mean; rather, there seems to be a downward trend. On the other hand, the wage share from the firms' point of view does not display any strong trend. Looking at the relevant graphs we can see that the reason is that the terms of trade and, in particular, taxes are characterized by changes that tend to decrease $w_t - v_{y,t} + e_t - y_t - p_t$ relative to $w_t + v_{p,t} + e_t - y_t - p_t^d$. Consequently, as long as we focus on price setting of domestic goods, a stationary wage share does not seem to violate the observed data.

Second, according to our theoretical reaction function for monetary policy, a stationary wage share (from the firms' perspective) together with a stationary inflation rate (and the parameter restrictions $|\rho_M|, |\phi_3| < 1$) implies that the domestic interest rate is stationary. The cointegration space should thus contain a trivial cointegration vector that equals the domestic interest rate. The time series graph of the domestic interest rate i_t is shown in Figure 4. As can be seen, this variable is likely a borderline case. As concerns inflation, there is a tendency of a shift in the mean around the introduction of the inflation target in 1993. This can be seen in the graph of the price level (shown in Figure 4), where the growth rate of the trend declines around that date. Although these variables

[13] In our theoretical model, given the assumptions in Table 1, it turns out that there does not exist any solution of $\theta_2 \geq 0.5$. For the same reason, not all assumptions about δ and the ϕ_i's are feasible. Since such properties of the theoretical model hinge on its unrealistic shortage of dynamic adjustment mechamisms, we do not dwell on the matter here. As noted above, the main purposes of the model is to provide guidance for our choices of cointegration and indentifying restrictions.

[14] The domestic price, p_t^d, has been calculated using equation (7) with $\delta = 1/2$.

may not be stationary off hand, there is thus a possibility that they may be stationary conditional on certain deterministic breaks and regime shifts.[15] The particular deterministic variables that we have chosen to condition our empirical analysis on are discussed in the next subsection.

3.2. *Cointegrated VAR Models: Specification Issues*

A VAR model for x_t, a vector of n endogenous variables, with $q = q_1 + q_0$ exogenous random variables, $z_t = (z_{1,t}, z_{0,t})$, can be written as

$$\Pi(L)x_t = \delta_0 + \delta_1 D_t + \Phi_1(L)z_{1,t} + \Phi_0(L)z_{0,t} + u_t, \tag{10}$$

where u_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 $L^j x_t = x_{t-j}$. Similarly, $\Phi_k(\lambda)$ is an $n \times q_k$ matrix polynomial of order p for $k = 0, 1$, given by $\Phi_k(\lambda) = \sum_{i=0}^{p} \Phi_{k,i}\lambda^i$. Furthermore, D_t is a vector of d observable deterministic variables.

The vector x_t may be nonstationary, reflecting the fact that some of the endogenous variables can e.g. be I(1). If this is the case, then standard asymptotic inference may not be applicable when testing certain hypotheses. However, if the changes in x_t, denoted by $\Delta x_t = (1-L)x_t$, and, moreover, certain linear combinations of the levels of x_t are found to be stationary, then x_t is said to be cointegrated.

With stochastic exogenous variables in the model, we also need to consider their orders of integration and their relation to the endogenous variables. Letting $z_{k,t}$ be integrated of order k, we thus assume that $z_{0,t}$ is stationary, while $z_{1,t}$ ($\Delta z_{1,t}$) is nonstationary (stationary). Moreover, there does *not* exist a linear combination between any of the variables in $z_{1,t}$ that is stationary, but instead there may exist linear combinations between the variables in x_t and $z_{1,t}$ that are stationary. If so, the VAR model in (10) can be reparameterized into a vector error correction (VEC) model

$$\Gamma(L)\Delta x_t = \delta_0 + \delta_1 D_t + \alpha\beta' X_{t-1} + \Psi(L)\Delta z_{1,t} + \Phi_0(L)z_{0,t} + u_t, \tag{11}$$

where $X_t = (x_t, z_{1,t})$, $\Gamma(\lambda) = I_n - \sum_{i=1}^{p-1}\Gamma_i\lambda^i$, $\Gamma_i = -\sum_{j=i+1}^{p}\Pi_j$, $\Psi(\lambda) = \sum_{i=0}^{p-1}\Psi_i\lambda^i$, $\Psi_0 = \Phi_{1,0}$, $\Psi_i = -\sum_{j=i+1}^{p}\Phi_{1,j}$, and $\alpha\beta' = -[\Pi(1) \quad \Phi_1(1)]$. The matrix α is $n \times r$ while β is $n + q_1 \times r$ with rank equal to r and the columns of β, denoted by β_j, $j = 1, \ldots, r$, are the so called cointegration vectors, i.e. $\beta' X_t$ is I(0). Inference in the VEC model (11) is feasible to the extent that asymptotic distributions do exist for tests of economically relevant hypotheses.

An empirical cointegration model corresponding to the theoretical common trends model presented in Section 2 involves the endogenous vector x_t with $n = 8$ variables and the exogenous vector $z_{1,t}$ with $q_1 = 5$ variables. These variable vectors comprise the cointegrating space and are defined as

$$\begin{bmatrix} x_t & \vdots & z_{1,t} \end{bmatrix} = \begin{bmatrix} y_t & i_t & p_t & s_t & e_t & p_t^m & w_t & u_t & \vdots & y_t^f & p_t^f & v_{y,t} & v_{p,t} & v_{i,t} \end{bmatrix}$$

In addition, $z_{0,t} = i_t^f$ so that $q_0 = 1$. That is, the foreign interest rate is both exogenous and stationary.

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[15] Apel and Jansson (1999), who investigate this hypothesis in detail, reach the same conclusion.

The domestic output series, y_t, and employment series, e_t, are both characterized by strong seasonal patterns. Moreover, the seasonal pattern in the output series shifts around 1980, requiring special attention. Rather than including a full set of seasonal dummies in the VAR model we have chosen to seasonally adjust these two series only. This is done through regression analysis on seasonal dummy variables prior to the system estimation. In the case of y_t two separate regressions are run, one for the subsample 1970:1–1979:4 and one for 1980:1–1999:4. The time plots of the adjusted output and employment series and the remaining six endogenous variables are shown in Figure 4.

It is possible that some of the nonstationary features in our system are due to deterministic breaks and regime shifts. To allow for this, the VEC model (11) has been augmented by a set of dummy variables. The theoretical model in Section 2 contains two (types of) dummy variables assumed to capture changes in financial market regulations and regime shifts in monetary policy, respectively.

During the sample period at hand, the most dramatic shift in Swedish monetary policy occurred when Sveriges Riksbank (the central bank of Sweden) abandoned the fixed exchange rate in 1992:4 and introduced an explicit inflation target (in 1993:1). From the early 1970s until the beginning of the 1990s, both the mean and variance of inflation were high and recurrent cost crises were accommodated by several devaluations of the fixed exchange rate. Devaluations occurred in 1973:1, 1976:4, 1977:2, 1977:3, 1981:3, and 1982:4. These are clearly born out in Figure 4, where the series for the nominal exchange rate, s_t, displays distinct jumps at each of these dates. The two largest devaluations occurred in 1981:3 and 1982:4.

Against this background and denoting the empirical counterpart to d_{mp} by D_{mp}, we can write:

$$D_{\mathrm{mp},i,t} = \begin{cases} 1 & \text{if } t \in \mathbb{I}_i,\ i = 1, 2, \dots, 8, \\ 0 & \text{otherwise,} \end{cases}$$

where $\mathbb{I}_1 = \{1973{:}1\}$, $\mathbb{I}_2 = \{1976{:}4\}$, $\mathbb{I}_3 = \{1977{:}2\}$, $\mathbb{I}_4 = \{1977{:}3\}$, $\mathbb{I}_5 = \{1981{:}3\}$, $\mathbb{I}_6 = \{1982{:}4\}$, $\mathbb{I}_7 = \{1992{:}4\}$, and $\mathbb{I}_8 = \{1993{:}1, \dots, 1999{:}4\}$. In the terminology of Perron (1989), dummies 1–7 represent "crashes" while dummy 8 captures "changes in growth".

It is harder to pinpoint the exact date of the policy shift towards deregulation of financial markets, but the process seems to have started with the introduction of Certificates of Deposits and the loosening of liquidity requirements in the early 1980's. Hence, the empirical counterpart to the dummy d_{fm} is:

$$D_{\mathrm{fm},t} = \begin{cases} 1 & \text{if } t \in \{1970{:}1, \dots, 1979{:}4\} \\ 0 & \text{otherwise.} \end{cases}$$

Having discussed the endogenous and exogenous variables, we now turn our attention to the determination of the lag length. This can be, and usually is done by evaluating the estimated residuals from a fitted VAR model by means of univariate and multivariate asymptotic specification tests.

The strategy is to include as few lags as possible, given associated residuals that are largely consistent with the assumption of independent and identically distributed normal zero mean errors. However, we will depart from this approach in two respects. First, by considering the residuals of the VEC in (11) rather than of the VAR in (10), since the specification tests are valid for, strictly speaking, stationary processes only. Second, by using bootstrapped analogues to the asymptotic specification tests in order to circumvent the size problems associated with the asymptotic reference distributions. In a strict sense, the bootstrap procedure also requires stationarity and is thus better suited for the VEC than the unrestricted VAR.[16]

To this end we will perform misspecification analyses for the VEC models defined by lag orders $p \in \{2, 3, \ldots, 4\}$ and cointegration ranks $r \in \{1, 2, \ldots, 8\}$ with respect to multivariate serial correlation, multivariate normality, and multivariate autoregressive conditional heteroskedasticity. Table 3 presents the evidence for these tests by means of both asymptotic and bootstrapped p-values. From the table it can be seen that the asymptotic reference distributions can be very misleading when applying these tests (see in particular the tests against serial correlation and normality). Based on the outcome of the bootstrap tests, our conclusion is that all models with 3 lags and 3 or more cointegration vectors give error terms with acceptable statistical properties.

3.3. *Cointegration Tests*

Given the specification of our model with two regime dummies and six exogenous (possibly nonstationary) variables, testing for the cointegrating rank using asymptotic critical values is not straightforward since these have to be simulated. But, again, given the size problem for asymptotic tests a safer route is bootstrap estimation of small sample critical values. Table 4 presents such estimated values for lag orders 2–4. It is interesting to note that the small sample distributions shift substantially to the right as the lag order increases, confirming the appropriateness of a bootstrap test procedure. For all three lag orders the test outcomes point in the direction of 3 or 4 cointegration vectors. Focusing on the case of 3 lags (favored by the specification tests) we reject, at conventional test levels, $r = 4$ in favor of $r = 3$. However, given the fact that the hypothesis of 4 cointegrating vectors is not resoundingly rejected and that we know that these tests have low power, choosing $r = 4$ does not appear unreasonable in light of the theoretical model.

The next task is to explore the cointegrating space. This is done imposing linear restrictions on the empirical vectors as implied by the theoretical model. In Table 5 we thus present four cointegration and three noncointegration relations that are consistent with the theoretical model in Section 2. These theoretical vectors will guide estimation of the restricted cointegration space and result in an aggregate demand, a price setting, a financial markets equilibrium, and a wage setting relation. Overall, when we impose the additional restrictions $\gamma_2 = \gamma_3 = 1/20$ and $\delta = 1/5$,

[16] 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 $N(0, \hat{\Sigma})$, in place of u_t. All bootstrap results in this paper involve 100,000 generated pseudo samples.

the estimated free parameters, given in Table 6, are not unreasonable in terms of magnitude and sign.[17]

The LR test of the linear restrictions results in a test statistic of 241.02 and using an asymptotic $\chi^2(32)$ critical value, we reject the null hypothesis that these four relations constitute a valid rotation of the empirical vectors. While this test, again, is known to be badly approximated by the limiting χ^2 distribution (see Gredenhoff and Jacobson, 2001), the magnitude of the test statistic is such that we need not bootstrap the critical values to conclude that the theoretical vectors are indeed rejected. For this reason we have chosen to undertake the rest of the empirical analyses for two VEC models, using both the four unrestricted cointegration vectors and the four theoretically restricted vectors. The advantage of using the latter is of course that our empirical VEC in that case has an economically well-defined equilibrium, which may make results easier to interpret.

4. The Effects of Monetary Policy

4.1. *Identification*

The solution to the theoretical model in Section 2.2 allows us to identify all the domestic shocks in the model. In this subsection we examine how to use the theoretically suggested identification scheme in the empirical analysis of a structural VAR model with cointegration relations. Since all theoretical shocks have long run effects on some variables, we cannot use the approach suggested by King et al. (1991) for identifying shocks to the common trends and extended to the identification of transitory shocks by Warne (1993). The reason is that in these studies it is assumed that the number of shocks with nonzero long run effects on some variable is equal to the number of common trends. Since all 8 domestic shocks have long run effects in the $\psi = 1$ case discussed in Section 2.2 and the number of domestic common trends is 4, it follows that this assumption is violated. Instead, we shall use the results in Vlaar (1998) for imposing exactly identifying (and overidentifying) restrictions on the parameters of the structural cointegrated VAR. This approach allows us to identify the shocks without having to rely on the assumption that some shocks only have temporary effects on nonstationary variables. At the same time, the Vlaar approach can be used to impose identifying restrictions on the contemporaneous and the long run responses of the endogenous variables to the domestic shocks.[18]

In Table 7, we have listed all long run and contemporaneous identifying restrictions we consider. Panel A lists the restrictions we place on the long run coefficients. The total number of identifying restrictions is equal to 17 where we take into account that the individual shocks to the fourth

[17] It may be noted that if we use these parameters instead of the counterparts in the benchmark parameterization of the theoretical model, then there does *not* exist a unique solution. In that sense, the parameters may not be viewed as reasonable, but if we were to allow for additional dynamics in the theoretical setup, then we may not run into a uniqueness problem.

[18] The Vlaar approach does not allow us to estimate a VAR model with all the long run and contemporaneous restrictions on monetary policy shocks shown in Table 7. The reason is that conditional on β, the structural form, represented by the matrix with contemporaneous effects B, is estimated also conditional on the Γ_i, δ_i, α, Ψ_i, and $\Phi_{1,i}$ parameters from the VEC model in equation (11). For example, in the case of the 2 overidentifying restrictions in Panel C, i.e. that unemployment and import prices do not react contemporaneously to monetary policy shocks, the Vlaar approach implies that the covariance matrix for u_t, the reduced form residuals, is singular. To avoid this problem, one would need to estimate α, Γ_i, etc. simultaneously with B.

common trend (the nominal trend) should have zero long run effects on output, unemployment, and employment. In terms of the theoretical model this means that we assume that nominal neutrality holds in the long run ($\xi_1 = 1$), yielding 15 identifying restrictions. The remaing 2 zero restrictions on the long run responses of labor productivity from labor supply and terms-of-trade shocks are consistent with constant returns to scale in production ($\kappa = 1$). In other words, the technology shock is identified by assuming that it is the only shock which has a long run effect on labor productivity. Similarly, the labor supply and terms-of-trade shocks are distinguished from the remaining 5 shocks by assuming that nominal neutrality holds in the long run.[19]

With all structural shocks being uncorrelated and having their variances normalized to unity (36 restrictions), it is necessary to impose 11 more restrictions to achieve exact identification. We have chosen to focus on the contemporaneous responses and the restrictions are again based on predictions from our theoretical model in Table 2, Panel B. Since the solution to the model that we discussed in Section 2.2 involves more than 11 restrictions on the contemporaneous responses to the domestic shocks, we have organized the constraints into two groups. The first, given in Panel B of Table 7, gives 11 restrictions which lead to exact identification. The remaining 32 overidentifying restrictions that our model suggests are presented in Panel C.

To distinguish the labor supply shock from the terms-of-trade shock, we use one of the restrictions on the contemporaneous responses that follow from assuming that import prices are sticky ($t_m = t - 1$). Namely, that the labor supply shock is only allowed to have an impact on import prices with a lag. Since the theoretical model, by assumption, implies that the terms-of-trade shock influences import prices contemporaneously, it follows that this zero restriction uniquely determines both shocks.

What remains to be done is to impose restrictions such that the monetary policy shock and the other shocks behind the nominal trend can be separated from each other. Assuming that wages are sticky ($t_W = t - 1$) identifies not only the wage shock, but also monetary policy and price shocks. The reason is that apart from wage and price shocks, only monetary policy shocks and domestic price shocks can influence the domestic interest rate simultaneously, whereas aggregate demand and financial market shocks cannot. This, in turn, depends on the specifications of the price setting equation (4) and the monetary policy rule (9). Specifically, the domestic interest rate responds to inflation and the wage share. Under constant returns to scale, sticky wages, and sticky import prices, neither consumer price inflation, nor the wage share, responds to aggregate demand and financial market shocks. Note also that the price shock can be separated from the monetary policy shock irrespective of whether domestic prices are sticky ($t_p = t - 1$) or not ($t_p = t$). Under either assumption, monetary policy shocks have no contemporaneous effect on the domestic price.

[19] These identifying assumptions take cointegration into account as follows. If we label the first 4 columns of the long run responses of the endogenous variables to the domestic shocks by $\beta_{1\perp}, \ldots, \beta_{4\perp}$, where $\beta'[\beta_{1\perp} \cdots \beta_{4\perp}] = 0$, then 5 identifying assumptions are imposed on these vectors. By letting the remaining 4 columns of the long run responses be proportional to the fourth column (i.e. by letting each one of these columns be given by $k_i\beta_{4\perp}$, $i \in \{5, \ldots, 8\}$, where the k_i's are free parameters) this provides us with 12 additional identifying restrictions. In the common trends model analysed by King et al. (1991) and Warne (1993), the 4 k_i parameters would be set equal to zero to ensure that $r = 4$ shocks have no long run effect on the endogenous variables. In the present setting, this would be inconsistent with the theoretical predictions.

To discriminate between aggregate demand and financial market shocks, the assumption that $t_y = t - 1$ in the aggregate demand relation is sufficient.

In the empirical model, we distinguish between the shocks in the nominal trend by using a subset of all the restrictions implied by the discussion above. Monetary policy shocks are assumed to have zero immediate effect on output, employment, the nominal wage, and the price level. The remaining 4 shocks are identified by assuming that (i) wage, domestic price, and financial market shocks do not change output simultaneously, that (ii) the financial market shock does not influence the domestic interest rate directly, and that (iii) the nominal wage does not react immediately to price and financial market shocks. We have selected this subset of the theory consistent restrictions for exact identification on the basis that (a) monetary policy shocks should influence the domestic interest rate, (b) aggregate demand shocks should influence output, (c) price shocks should influence the price level, and (d) wage shocks should influence the nominal wage.

4.2. *The Effects of Monetary Policy Shocks Identified from VAR Residuals*

Using the restrictions in Table 7 A and B, we can calculate the effects of the structural shocks (the ε's) based on the estimated VAR models. In effect, we will do so for two VAR models; one with the theoretical restrictions imposed on the cointegration vectors, and one with the unrestricted cointegration vectors (for $r = 4$). Although the theoretical restrictions are not supported by the data we think they enhance our understanding of shock effects. It is equally reasonable to study the effects in an "unconstrained" case, i.e. let the data speak more freely. Two models, eight endogenous variables, and hence eight shocks, yield in total 128 impulse responses. We will limit the presentation here to the broad picture.[20]

First, and hardly surprising, the estimated effects are in general associated with large uncertainty (the confidence level for the intervals is 67 percent). Using the estimates of the theoretical cointegration vectors in Table 6, less than half of the impulse response functions show any significant effects. Applying the unrestricted estimates of the cointegration vectors instead, we find most of the impulse responses to be significant. But this comes at the price of some estimates of long run effects being inconsistent with our theoretical model (since no theoretical restrictions on steady state relations have been imposed). Nevertheless, most of the estimated effects from various shocks that are significant are also consistent with our theoretical prejudices.

A second interesting result from these exercises is that the estimated impulse responses are much less "hump shaped" than what is typically assumed in theoretical models. The maximum shock effects are often recorded almost immediately. There are more hump shaped impulse responses in the VAR based on the estimated theoretical cointegration vectors than for the unrestricted case. Using the unrestricted vectors, we find that most variables have reached their long run levels around 2 years after a shock.

Since the focus in this paper is on the effects of monetary policy shocks, we show the interest rate and price level effects of monetary policy in Figure 5, using both restricted (Panel A) and unrestricted (Panel B) cointegration vectors. Imposing the estimates of the theoretical vectors in Table 6, we get

[20] All results are available upon request.

a price puzzle. A temporary increase in the nominal interest rate (initially by about 0.2 percentage points), results in a *higher* price level. After a small (and insignificant) disinflation during the first and second quarter after the shock, there is more inflation than in the base line no shock case for a period of 2 years. The steady state price level increases by 1.4 percent and this level is almost reached within 2 years. With unrestricted cointegration relations, we get rid of this price puzzle. The consumer price level falls in the first quarter after the shock and reaches its new equilibrium level almost immediately thereafter.

Another way to get rid of the (long run) price puzzle reported in Panel A, while keeping the cointegration restrictions, is to choose another subset of exactly identifying restrictions from Table 7. In Panel C of Figure 5 we present impulse responses for the price level and the domestic interest rate when 2 of the long run restrictions for monetary policy shocks have been replaced by 2 contemporaneous restrictions. That is, we have dropped the zero long run effects on unemployment and employment, and imposed the zero contemporaneous effect on unemployment and import prices for monetary policy shocks. While the price puzzle does not disappear completely in the 1-5 year horizon, consumer prices fall after one quarter as well as in the long run. When the same identifying restrictions are used with empirical cointegration relations, we again obtain a price puzzle.

Hence, the implications of the identified VAR models give mixed results regarding the conventional wisdom about the effects of monetary policy shocks. Under certain, apriori reasonable identifying assumptions, we get the same price puzzles that have been found in other open economy VAR models (cf. Figure 2). On the other hand, it is possible to choose among the set of identifying restrictions such that the responses to a monetary policy shock look more like the conventional wisdom. However, the results depend on both long run and contemporaneous restrictions on the effects of monetary policy in ways that, to our knowledge, have not been investigated in the earlier literature.

These results suggest a number of directions that empirical models of the effects of monetary policy could take. One is to estimate fully structural models.[21] One well know problem with these models is, however, that they hardly can account for the influence of lagged variables that one finds in empirical studies. A second possibility is to use sign restrictions of the type suggested by e.g. Canova (2001) and Uhlig (1999, 2001). A third direction is to supplement the VAR model with some information about the nature of policy shocks that can be gathered outside the VAR system. We will pursue this route in the next section.

5. Alternative Analyses of Monetary Policy

Rudebusch (1998), among others, has questioned the appropriateness of the assumptions typically applied in the VAR literature and suggests an alternative approach to monetary policy analysis by extracting shocks using futures markets for federal funds rates. In the same vein, we propose to analyse alternative sets of monetary policy shocks in this section. First, in Section 5.1, we estimate a

[21] It deserves to be emphasized that not all theoretical models suggest that the so called price puzzle really is a puzzle; see Barth and Ramey (2000).

reaction function for Sveriges Riksbank during the inflation targeting regime, using the Riksbank's own forecasts for inflation and GDP growth. The residuals from this regression are taken as our maintained policy shocks and their relation to the residuals from the VEC model is used to compute new sets of impulse response functions in Section 5.2. Similar approaches have earlier been applied by Blanchard and Watson (1986), to identify the effects of shocks to U. S. fiscal policy, and by Bagliano et al. (1999), to derive the effects of shocks to U. S. monetary policy. In Section 5.3 we will conduct a similar exercise by letting a number of currency devaluations act as monetary policy shocks very much in the same vein as Eichenbaum et al.'s (1998) VAR modeling of the exogenous, "narrative" shocks to U. S. government spending defined in Ramey and Shapiro (1998).

5.1. *Policy Shocks Estimated using an Instrument Rule*

A credible benchmark for evaluating measures of systematic and surprise movements in monetary policy requires information about policy makers' objectives and rules. The Riksbank Inflation Report in October 1999, p. 58, provides the following policy characterization:

> Monetary policy is sometimes described with a simple rule of thumb: if the overall picture of inflation prospects (based on an unchanged repo rate) indicates that in twelve to twenty-four months' time inflation will deviate from the target, then the repo rate should normally be adjusted accordingly.

Our interpretation of this quote is a reaction function of the form:

$$r_t = a + br_{t-1} + c(\pi^{\mathrm{F}}_{t,1} - 2) + d(\pi^{\mathrm{F}}_{t,2} - 2) + \epsilon_t, \tag{12}$$

where r_t denotes the policy (repo) rate in quarter t, $\pi^{\mathrm{F}}_{t,j}$ the bank's period t real time forecast of inflation j years ahead (conditional on r_{t-1}), and ϵ_t the monetary policy surprise (shock) in quarter t.

The results from regressing the nominal policy rate[22] on the Riksbank's 1 and 2 years ahead forecasts of inflation, and including an autoregressive term, are as follows:[23]

$$\begin{aligned} r_t &= \underset{(0.25)}{1.31} + \underset{(0.05)}{0.65} r_{t-1} + \underset{(0.21)}{0.40}(\pi^{\mathrm{F}}_{t,1} - 2) + \underset{(0.20)}{0.38}(\pi^{\mathrm{F}}_{t,2} - 2) + \hat{\epsilon}_t \\ R^2 &= 0.98, \quad t = 1993:1, \ldots, 2001:2, \end{aligned} \tag{13}$$

where the numbers within parentheses are estimated (asymptotic) standard errors if we assume that all variables are stationary. Since some observations of $\pi^{\mathrm{F}}_{t,2}$ are missing, we have approximated these by the most recent available forecasts. The residual diagnostics indicate that the estimated policy shocks $\hat{\epsilon}_t$ — displayed in Figure 6 — are close to Gaussian white noise. The estimates suggests relatively marked negative policy shocks in 1996 and in 2000, and positive shocks in the period 1994–95 and in 1998.

[22] Note that $r_t = 100 \ln(1 + R_t/100)$, where R_t is the repo rate in percent at the end of quarter t. That is, the repo rate has been transformed in the same way as the short term rates in the VAR model.

[23] Estimating a rule such as (12) requires access to the (real time) forecasts $\pi^{\mathrm{F}}_{t,j}$. Sveriges Riksbank has recently published every forecast made since the introduction of the inflation target in 1992. For the precise sources of the data and a discussion of various problems, see Jansson and Vredin (2001).

To examine the robustness of the results we have undertaken a number of sensitivity tests. These involve (i) measuring the policy rate as a quarterly average rather than as an end-of-quarter value; (ii) using forecasts of so called underlying inflation instead of headline inflation;[24] (iii) using consecutive 12-monthly inflation figures rather than end-of-year figures; (iv) using "truncated" samples and a different dating of the arguments in the rules; and (v) using (where appropriate) instrumental variables estimates rather than least squares. Our conclusion is that the results, as depicted in Figure 6, are robust to these alternative specifications.

5.2. *Impulse Response Experiments Based on Shocks from the Instrument Rule*

Studying the effects of the alternative measure of the policy shock within the VEC model (11) essentially boils down to ascertaining the contemporaneous relationship between the reduced form residuals of the VEC, u_t, and the policy shocks. Once this relationship has been determined, impulse response functions can be computed in the standard manner by tracing out the dynamic effects that shocks to this particular relationship imply.

Briefly, the procedure we use is as follows. First, we regress the policy shocks on the VEC residuals for the subsample 1993:1–1999:4. The coefficients on the VEC resiuduals are consistently estimated by least squares provided that the "measurement errors" (the least squares residuals in this regression) are uncorrelated with the VEC residuals. Next, we postmultiply the covariance matrix of the VAR residuals, Σ, by (a linear function of) this parameter vector.[25] Under the conditions at hand, this quantity gives an estimate of the contemporaneous effects of a (unit or one standard deviation) policy shock on the endogenous variables in the VEC.

The top 4 graphs in Figure 7 show the impulse responses for the price level and the domestic interest rate from policy shocks evaluated in the VEC model with theoretical and empirical cointegration vectors, respectively, when no contemporaneous restrictions have been imposed. In both cases the outcome is a price puzzle. Using shocks derived from a reaction function consistent with the official Riksbank monetary policy and imposing no additional restrictions on the impulse responses is thus not sufficient to dissolve the price puzzle. If we estimate the relationship between the policy shocks and the VEC residuals conditional on the restrictions that policy shocks should only influence the domestic interest rate and the nominal exchange rate contemporaneously (bottom 4 graphs), we find that the price puzzle is still present for the empirical cointegration vectors, but has disappeared when theoretical cointegration vectors are used. Hence, our empirical evidence on the effects from monetary policy shocks, whether purely VAR based or derived using external

[24] This may be interpreted as an alternative (or complementary) way of acknowledging the possibility that not only the inflation target in terms of the CPI matters for monetary policy. Because measures of underlying inflation are smoothed, they imply that certain movements in headline CPI are not counteracted by policy measures. In so far as these movements are not due to shocks to aggregate demand, a policy taking account of underlying rather than headline inflation will implicitly also stabilize the real economy.

[25] Let g denote the linear combination (an 8×1 vector) of the VAR residuals which has the largest correlation with the estimated policy shock. To ensure that the policy shock in the impulse response experiment has unit variance, we transform g by multipying it with the square root of the inverse of $g'\Sigma g$. Letting f denote this transformation, we indeed find that the policy shock given by fu_t has variance 1. The contemporaneous response to this shock, under the assumption that all other structural shocks are orthogonal to fu_t, can be shown to be given by Σf.

information about the shocks, is mixed regarding the conventional wisdom about the effects of monetary policy shocks.

5.3. *The Effects of Exchange Rate Devaluations*

Finally, in our endeavor to identify effects of monetary policy shocks in Sweden, we will now consider the case of currency devaluations. In our empirical models we take account of devaluations by including deterministic dummy variables. This is by no means an optimal model of devaluation effects. Nevertheless, we believe that these interventions are, at least in part, exogenously driven, and that the coefficients in the VAR model contain information about the effects of such shocks.

Figures 8, 9, and 10 show impulse responses from the policy interventions of December 1992, October 1982, and September 1981. In September 1981 and in October 1982 the Swedish Krona was devalued (against a trade weighted currency basket) by 10 percent and 16 percent, respectively. In December 1992 the pegged exchange rate policy was abandoned in favor of a floating exchange rate regime. In this case, the Krona initially depreciated by some 20 percent; cf. Figure 4. The impulse response function in Figures 8–10 are computed for the VEC model with the theoretically restricted cointegration vectors in Table 5.[26]

For the December 1992 devaluation, in Figure 8, we can see that the initial and long run effect is a nominal depreciation of the Krona of around 7 percent, with some overshooting in the short to medium term. The interest rate drops immediately. According to the estimated impulse response function, this drop is soon followed by an increased nominal interest rate, relative to the no devaluation case, which is consistent with most interpretations of that episode in Swedish monetary policy history. The depreciation of the Krona was at the time considered a threat to the (recently introduced) inflation target, hence a relatively contractionary monetary policy was pursued in Sweden during at least the first half of the 1990s. But the initial depreciation of the Krona in 1992 and 1993 must be viewed as an “expansionary” monetary policy shock. In a theoretical model, a devaluation would correspond to a permanent increase in the central bank’s target for the price level, or a one time increase in the level of the money stock (depending on the monetary policy regime at hand). As can be seen in Figure 8, the effect is an increase in the price level. The largest effect, in terms of inflation, is recorded immediately, but inflation continues to raise the price level for several years (except for a brief disinflation after 2 years). The long run effect is an increase in the CPI price level of 2 percent. Import prices go up slightly less. Furthermore, GDP initially declines, and then stays above the baseline level for a long time, but is ultimately not affected. Employment goes up and unemployment falls even in the long run, however. The nominal wage increases, and the real wage increases in terms of consumer prices, but falls in terms of producer prices.[27]

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[26] The results from applying the unrestricted cointegration vectors instead are somewhat different (and available upon request), but the differences are similar to those reported in Section 4.2 above. The impulse responses using unrestricted cointegration vectors are less hump shaped and converge to their new steady state levels much faster, but imply reasonable long run effects.

[27] Since import prices increase in the long run by 1.5 percent and consumer prices by 2.25, domestic producer prices must go up by more than 5 percent (holding the indirect taxes constant and using $\delta = 1/5$).

The depreciation effects in Figure 8 lend more support to the conventional wisdom about the effects of monetary policy shocks than for the structural shocks in Sections 4.2 and 5.2. In particular, the effects on the price level follow almost exactly the same pattern as in Svensson's model; cf. Figure 1. There are a couple of differences between Figure 8 and typical theoretical models, however. First, the largest effects on most growth rates occur immediately, i.e., the empirical impulse responses are not as hump shaped as the theoretical ones. Second, the monetary policy intervention in Sweden in 1992 seems to have had a permanently positive effect on employment (and a negative one on unemployment). The latter result can arise because the VEC model has not been subject to all long run restrictions suggested by our theoretical model in Section 2 (only the estimated cointegration relations, but not any of the overidentifying restrictions in Panel C of Table 7). Quantitatively, according to Figure 8 monetary policy may have a larger effect on inflation than what is commonly asserted in theoretical models; an initial decline of the nominal interest rate by 1 percentage point raises the consumer price level by around 1 percent within a couple of quarters. This may be because a devaluation is perceived as an easily observed and credible shock to the price level, resulting in a distinct and rapid effect.

The picture is almost identical when we turn to the estimated effects of the devaluation in October 1982 in Figure 9. In this case, the effect on the nominal exchange rate, both in the short and in the long run, is almost exactly equal to the devaluation (16 percent). After the first quarter, the domestic price level increases slowly towards its higher steady state level, around 4 percent above the baseline, no devaluation case. Import prices eventually stabilize at a 3 percent higher level than initially, although they overshoot in the short run. Both qualitatively and quantitatively, these effects are very close to those in Figure 8. The effects on the nominal interest rate, however, appear to have been somewhat, but not much, larger than after the devaluation in 1992. That is, if we want to measure the monetary policy shock in terms of the implicit elasticity of inflation with respect to the interest rate, then the shock in 1982 had a somewhat smaller effect on inflation than the shock in 1992. Effects on output, employment, unemployment, and wages are also similar in Figures 8 and 9.

The effects of the devaluation in September 1981, according to the impulse responses in Figure 10, were quite different. The effects on the nominal interest rate, employment, and unemployment look similar to the effects after the interventions in 1982 and 1992, but the effects on prices and the exchange rate are distinctly different. In particular, the exchange rate *appreciated* after the devaluation in 1981, relative to its baseline no devaluation level. This picture is consistent with most interpretations of the 1981 and 1982 devaluations. Whereas the former was partly anticipated (perhaps not a shock at all) and raised the real exchange rate only slightly above the level in the late 1970's (cf. Figure 3), the 1982 devaluation was viewed as a surprisingly "aggressive" and expansionary monetary policy shock. This story makes sense also if we compare the effects between the 1982 and 1992 shocks (Figures 9 and 8, respectively). In 1982 the shock to the exchange rate was about the size of the devaluation itself, whereas in 1992 the exchange rate shock was smaller than the depreciation, i.e., the latter policy intervention was partly anticipated.

These results suggest that it is possible to achieve measures of the effects of monetary policy shocks from VAR models that are consistent with the predictions of conventional theoretical models. But it seems pertinent to concentrate on periods when policy shocks have been large enough to be distinguishable from the many other shocks that hit the economy. The identification scheme, however, tells a different story about the lags in the effects of monetary policy shocks than typical theoretical models. The largest effects of monetary policy shocks on inflation do not occur after a couple of years, but almost immediately.

6. Conclusions

Theoretical predictions about the effects of monetary policy shocks have, at least in part, received empirical support in studies of the U. S. economy. However, empirical studies of *open* economies typically do not confirm theory. In particular, open economy VAR models often yield "price puzzles" such that a contractionary monetary policy shock leads to an increase of the price level.

This paper makes an effort to bridge the gap between theoretical and VAR open economy models. Although there are probably many — and interacting — reasons for this misalignment of theory and measurement, a few candidates emerge. On the one hand, the theoretical models may neglect important real world mechanisms. On the other hand, the VAR applications may be at fault, either because they are simply too poor approximations of the "true" data generating process, or because there is a mismatch between the identifying restrictions being used and the relevant theoretical predictions. Or possibly, there may exist a fundamental problem with the VAR approach to modeling monetary policy shocks in that the identified shocks do not convey strong enough signals for the impulse response analysis to be meaningful. This argument is supported by the very wide confidence intervals often recorded for impulse responses.

Our point of departure is a theoretical model — consistent with standard open economy macro models — from which we derive a VAR representation on common stochastic trends form. That is, we impose contemporaneous and long run theory consistent restrictions on our empirical VAR. Nevertheless, our structural VAR model identified monetary policy shocks often result in price puzzles. Some restrictions, that are not unreasonable from a theoretical point of view, do yield responses to monetary policy shocks that support the conventional wisdom. But there are other, also plausible, identifying schemes that yield results in line with earlier VAR models of open economies, i.e., they give rise to price puzzles. This makes it important to look for results that are less sensitive to identifying restrictions.

A conceivable alternative is to try identification of monetary policy shocks using additional and VAR model external information about monetary policy. We take estimated deviations from an instrument rule as our monetary policy shocks and feed these into the empirical VAR for impulse response analyses. However, the price puzzles prevail, unless further restrictions are imposed. Next, we looked at the Swedish currency devaluations, which arguably should contain elements of monetary policy shocks. It turns out that the effects of the Swedish devaluations in 1982 and 1992 are very similar to the effects suggested by standard open economy models (e.g., Svensson, 2000).

The devaluation derived monetary policy shocks give rise to effects that support conventional wisdom from theoretical models. Hence, the “puzzling” impulse responses should not reflect a misspecified VAR model, since our VAR model produces price puzzles for certain identifications of monetary policy shocks, but not for other. This also implies that there is no fundamental weakness about the VAR approach as such. The problem, as we see it, is that the macroeconomic times series that are typically used to analyse these questions may contain too little information about significant monetary policy shocks. VAR residuals are linear combinations of many structural shocks, many more than there are variables in the model. The restrictions that standard theoretical models come up with may not be adequate to distinguish ordinary monetary policy shocks from other kinds of shocks. Note that a monetary policy shock in a theoretical model is typically assumed to be observed by the households and firms. Hence, they have perfect information about the model they are part of and can therefore form model consistent expectations about the effects of shocks. It is difficult to conduct this experiment using a small set of macroeconomic time series.

In the case of devaluations, however, there is little question that the economy was hit by a monetary policy shock. Households, firms, and econometricians can correctly date these shocks, and using very simple macro models they can evaluate the expected effects on the economy.

Appendix: Solving the Theoretical Model

In this appendix we will discuss how the model in Section 2 can be rewritten in a format consistent with the system of linear rational expectations models that is studied by Sims (2000). Given that a solution to Sims's system exists and is unique, we present how one can transform that system back into the original variables.

Since existence and uniqueness of Sims's formulation requires that the system is stable (no unit or explosive roots), we first need to transform the theoretical model into a stable system. This means that we first need to "guess" how many unit roots there are in the model. That is, a guess about how many unit roots that are due to modelled (domestic) stochastic trends, and an assumption about which transformations of the exogenous variables that are stationary. For example, such transformations involves specifying which exogenous variables are I(1) and which linear combinations of the levels of these that are I(0).

Transforming the Model to the Sims form

Let c_t denote a vector of (guessed) stationary linear combinations, where each vector contains at least one endogenous variable. Also, let Y_t be the n dimensional vector with endogenous variables, where we define two subvectors, denoted by $Y_{1,t}$ and $Y_{2,t}$, from $Y_t = (Y_{1,t}, Y_{2,t})$. The processes $\varepsilon_{1,t}$ and $\varepsilon_{2,t}$ are determined from the $Y_{1,t}$ and $Y_{2,t}$ sets of equations, respectively, and they are stationary (although possibly autocorrelated) with zero mean. Let $Y_{3,t}$ be a vector with the exogenous I(1) variables, and $Y_{4,t}$ a vector with the exogenous I(0) variables. The dimension of the $Y_{3,t}$ (given by s_3) and of the $Y_{4,t}$ (equal to s_4) vectors sum to the number (s) of original exogenous stochastic variables,[28] while we denote the d deterministic time varying variables by d_t.

The vector Y_2 is of dimension equal to the number of unit roots not accounted for by Y_3, while the dimension of Y_1 is equal to that of c. Moreover, let ν_t denote the m dimensional vector of (guessed) stationary expectations variables in the model, where

$$\Phi_2 \Delta Y_{2,t} = \nu_{t-1} + \varepsilon_{3,t}. \tag{A.1}$$

Here, $\varepsilon_{3,t}$ is a vector of expectational errors, satisfying $E_t \varepsilon_{3,t+1} = 0$ for all t. If the model can be solved uniquely, then the solution will determine what these errors are in terms of e.g. $\varepsilon_{1,t}$ and $\varepsilon_{2,t}$. The matrix Φ_2 selects the Y_2 variables whose expectations are given by the ν vector elements. In the theoretical model, $\nu_t = (E_t \Delta s_{t+1}, E_t \Delta p_{t+1})$, where Δ is the first difference operator. Accordingly, $\Phi_2 \Delta Y_{2,t} = (\Delta s_t, \Delta p_t)$.[29]

The c_t vector is r dimensional and can be written as a linear combination of $Y_{1,t}$, $Y_{2,t}$, and $Y_{3,t}$. Let this linear combination be given by:

$$c_t \equiv B_1 Y_{1,t} + B_2 Y_{2,t} + B_3 Y_{3,t}. \tag{A.2}$$

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[28] For example, if the the theoretical model includes only the foreign short and long term interest rate and we assume that both are I(1) with a stationary spread, then Y_3 would be equal to one of the interest rates and Y_4 would be equal to the spread.

[29] In principle, we can also consider expectations variables involving the cointegration relations, but since the theoretical model does not include any such variables they are omitted here.

Since the theoretical model is specified with a minimum of dynamics, it suffices to express the c_t subsystem of the model as:

$$\begin{aligned}\Psi_c c_t + \Psi_1 \Delta Y_{1,t} + \Psi_2 \Delta Y_{2,t} + \Psi_\nu \nu_t &= \Psi_{c,1} c_{t-1} + \Psi_{1,1} \Delta Y_{1,t-1} + \Psi_{2,1} \Delta Y_{2,t-1} + \Psi_{\nu,1} \nu_{t-1} \\ &+ \Psi_3 \Delta Y_{3,t} + \Psi_{3,1} \Delta Y_{3,t-1} + \Psi_4 Y_{4,t} + \Psi_{4,1} Y_{4,t-1} + \Psi_d d_t + k_1 + \Upsilon_1 \varepsilon_{1,t},\end{aligned} \tag{A.3}$$

where k_1 contains the means of the stationary $\tau_{i,t}$ processes, and c_t can typically be ordered such that Ψ_c has unit elements on its diagonal.

The remaining subsystem of the model can be written as:

$$\begin{aligned}\Lambda_c c_t + \Lambda_1 \Delta Y_{1,t} + \Lambda_2 \Delta Y_{2,t} + \Lambda_\nu \nu_t &= \Lambda_{c,1} c_{t-1} + \Lambda_{1,1} \Delta Y_{1,t-1} + \Lambda_{2,1} \Delta Y_{2,t-1} + \Lambda_{\nu,1} \nu_{t-1} \\ &+ \Lambda_3 \Delta Y_{3,t} + \Lambda_{3,1} \Delta Y_{3,t-1} + \Lambda_4 Y_{4,t} + \Lambda_{4,1} Y_{4,t-1} + \Lambda_d d_t + k_2 + \Upsilon_2 \varepsilon_{2,t},\end{aligned} \tag{A.4}$$

where k_2 represents the means of the stationary and the drift terms of the nonstationary $\tau_{j,t}$ processes ($i \neq j$). For those equations where $\tau_{j,t}$ is a random walk (with drift), they are expressed in first difference form in (A.4).[30]

To complete the transformation of the model we need to substitute for ΔY_1 in equations (A.3) and (A.4). To do so, we will use the cointegration relations in (A.2). If the matrix B_1 is invertible, then

$$\Delta Y_{1,t} = B_1^{-1} \Delta c_t - B_1^{-1} B_2 \Delta Y_{2,t} - B_1^{-1} B_3 \Delta Y_{3,t}.$$

Using this relationship, we find that equation (A.3) can be rewritten as:

$$\begin{aligned}[\Psi_c + \Psi_1 B_1^{-1}] c_t + [\Psi_2 - \Psi_1 B_1^{-1} B_2] \Delta Y_{2,t} + \Psi_\nu \nu_t &= [\Psi_{c,1} + (\Psi_1 + \Psi_{1,1}) B_1^{-1}] c_{t-1} \\ &+ [\Psi_{2,1} - \Psi_{1,1} B_1^{-1} B_2] \Delta Y_{2,t-1} + \Psi_{\nu,1} \nu_{t-1} - \Psi_{1,1} B_1^{-1} c_{t-2} \\ &+ [\Psi_3 + \Psi_1 B_1^{-1} B_3] \Delta Y_{3,t} + [\Psi_{3,1} - \Psi_{1,1} B_1^{-1} B_3] \Delta Y_{3,t-1} \\ &+ \Psi_4 Y_{4,t} + \Psi_{4,1} Y_{4,t-1} + \Psi_d d_t + k_1 + \Upsilon_1 \varepsilon_{1,t},\end{aligned} \tag{A.5}$$

[30] For example, if we assume that $\rho_T = 1$ so that the technology process follows a random walk, then this equation must be transformed as

$$\Delta y_t - \kappa \Delta e_t = \mu^T + \varepsilon_t^T,$$

before it is included in equation (A.4). Similarly, with $\rho_L = 1$, i.e. labor supply has a stochastic trend, we can express equation (2) as

$$\Delta(l_t - e_t) + \Delta e_t - \eta \Delta w_{t_L} + \eta \Delta \nu_{y,t_L} + \eta \Delta p_{t_L} = \mu^L + \varepsilon_t^L,$$

where $t_L \in \{t, t-1\}$ as before.

Furthermore, for the benchmark solution discussed in Section 2.2, i.e. when $\psi = 1$ and the guessed 4 cointegration relations can be determined from the wage setting, the price setting, and the aggregate demand equation, we have that a stationary domestic interest rate leads to the choice of either including the financial markets equilibrium relation or the monetary policy relation in equation (A.4); the other relation will then be included in equation (A.3). With $c_{2,t} \equiv p_t - (1-\delta) p_t^m - \delta(w_t + \nu_{p,t} + e_t - y_t) - \nu_{i,t}$ and $c_{3,t} \equiv i_t$, the monetary policy equation can be written as:

$$c_{3,t} + \frac{\phi_2}{\delta} c_{2,t} - \phi_1 \Delta p_t = \phi_3 c_{3,t-1} + \phi_4 d_{\mathrm{mp},t} + \frac{\mu^M}{1-\rho_M} + \tilde{\tau}_t^M,$$

where $\tilde{\tau}_t^M = \tau_t^M - \mu^M/(1-\rho_M)$ is a stationary process ($|\rho_M| < 1$) with zero mean. Similarly, with $\nu_{s,t} \equiv E_t \Delta s_{t+1}$, the financial markets equilibrium relation can be expressed as:

$$c_{3,t} - \nu_{s,t} = i_t^f + \zeta d_{\mathrm{fm},t} + \frac{\mu^I}{1-\rho_I} + \tilde{\tau}_t^I,$$

where $\tilde{\tau}_t^I = \tau_t^I - \mu^I/(1-\rho_I)$ is a stationary process ($|\rho_I| < 1$) with zero mean. It follows that one of the processes $\tilde{\tau}_t^M$ and $\tilde{\tau}_t^I$ is included in $\varepsilon_{1,t}$, while the other is included in $\varepsilon_{2,t}$.

while equation (A.4) becomes:

$$
\begin{aligned}
[\Lambda_c + \Lambda_1 B_1^{-1}]c_t + [\Lambda_2 - \Lambda_1 B_1^{-1} B_2]\Delta Y_{2,t} + \Lambda_\nu \nu_t = {} & [\Lambda_{c,1} + (\Lambda_1 + \Lambda_{1,1})B_1^{-1}]c_{t-1} \\
& + [\Lambda_{2,1} - \Lambda_{1,1} B_1^{-1} B_2]\Delta Y_{2,t-1} + \Lambda_{\nu,1}\nu_{t-1} - \Lambda_{1,1} B_1^{-1} c_{t-2} \\
& + [\Lambda_3 + \Lambda_1 B_1^{-1} B_3]\Delta Y_{3,t} + [\Lambda_{3,1} - \Lambda_{1,1} B_1^{-1} B_3]\Delta Y_{3,t-1} \\
& + \Lambda_4 Y_{4,t} + \Lambda_{4,1} Y_{4,t-1} + \Lambda_d d_t + k_2 + \Upsilon_2 \varepsilon_{2,t}.
\end{aligned}
\tag{A.6}
$$

The assumption that B_1 has full rank is not restrictive. If the guess about the number of unique cointegration relations is correct, then we should be able to order those endogenous variables which are not included in the vector $\Phi_2 Y_{2,t}$ appropriately, i.e. into either Y_1 or Y_2. However, the choice of Y_1 is not independent of the parameter values in the cointegration relations. This means that a given B_1 will typically not be invertible for all points in the parameter space. Accordingly, for those points where a given B_1 is singular, the grouping of endogenous variables into Y_1 and Y_2 will have to be different, thus producing a new B_1 matrix (and B_2). If there does indeed exist some point in the parameter space where all possible B_1 matrices are singular, then that point corresponds to a case when the number of cointegration relations is less than the number we have assumed. For such a point it is necessary to alter the dimensions of Y_1 (smaller) and Y_2 (greater) to reflect the singularity of the original B_1 matrix. Thereby, we are able to construct a new nonsingular B_1 matrix for that point.

The system given by (A.1), (A.5), and (A.6) can now, by also using 3 identities, be cast in the form studied by Sims (2000):

$$\Gamma_0 X_t = C + \Gamma_1 X_{t-1} + \Psi Z_t + \Pi \eta_t, \tag{A.7}$$

where $X_t = (x_t, x_{t-1})$, $x_t = (c_t, \Delta Y_{2,t}, \nu_t)$ is an $n+m$ dimensional vector, while the vectors Z_t and η_t are given by $Z_t = (z_t, z_{t-1})$, where $z_t = (\Delta Y_{3,t}, Y_{4,t}, d_t, \varepsilon_{1,t}, \varepsilon_{2,t})$, and $\eta_t = \varepsilon_{3,t}$. The matrices Γ_0 and Γ_1 are both $2(n+m) \times 2(n+m)$, Ψ is $2(n+m) \times 2(s+d+n)$, and Π is $2(n+m) \times m$.

Transforming the Sims Format Solution to the Original Variables

The conditions for the existence of a unique solution to (A.7) is analyzed by Sims (2000). Assuming these conditions are satisfied, then the solution can be expressed as:

$$X_t = \Theta_1 X_{t-1} + \Theta_c + \Theta_0 Z_t + \Theta_y \sum_{i=1}^{\infty} \Theta_f^{i-1} \Theta_z E_t Z_{t+i}, \tag{A.8}$$

where the largest eigenvalues of Θ_1 and of Θ_f are both less than unity. We can then solve for X_t recursively to obtain:

$$X_t = \left(I - \Theta_1\right)^{-1} \Theta_c + \sum_{j=0}^{\infty} \Theta_1^j \Theta_0 Z_{t-j} + \sum_{j=0}^{\infty} \Theta_1^j \Theta_y \sum_{i=1}^{\infty} \Theta_f^{i-1} \Theta_z E_{t-j} Z_{t-j+i}. \tag{A.9}$$

To rewrite the solution for X_t in terms of current, lagged, and expected future z's, let J_z and $J_{z\perp}$ be defined from $z_t = J_z Z_t$ and $z_{t-1} = J_{z\perp} Z_t$, respectively. Some algebra then gives us that

$$\begin{aligned} X_t = & \left(I - \Theta_1\right)^{-1} \Theta_c + \left(\Theta_0 J_z' + \Theta_y \Theta_z J_{z\perp}'\right) z_t \\ & + \sum_{j=1}^{\infty} \left(\Theta_1^j \left(\Theta_0 J_z' + \Theta_y \Theta_z J_{z\perp}'\right) + \Theta_1^{j-1} \Theta_0 J_{z\perp}'\right) z_{t-j} \\ & + \sum_{j=0}^{\infty} \Theta_1^j \Theta_y \sum_{i=1}^{\infty} \left(\Theta_f^{i-1} \Theta_z J_z' + \Theta_f^i \Theta_z J_{z\perp}'\right) E_{t-j} z_{t-j+i}. \end{aligned} \tag{A.10}$$

In the event that z_t is white noise, then the last term on the right hand side is always equal to zero.

Let $W_t = (Y_t, v_t)$, then X_t is related to W_t through the transformation:

$$J_x X_t \equiv \Delta(L) M W_t + N Y_{3,t}, \tag{A.11}$$

where $J_x = [I_{n+m} \quad 0]$ is an $(n+m) \times 2(n+m)$ matrix and

$$\Delta(L) = \begin{bmatrix} I_r & 0 & 0 \\ 0 & \Delta I_{n-r} & 0 \\ 0 & 0 & I_m \end{bmatrix}, \quad M = \begin{bmatrix} B_1 & B_2 & 0 \\ 0 & I_{n-r} & 0 \\ 0 & 0 & I_m \end{bmatrix}, \quad N = \begin{bmatrix} B_3 \\ 0 \\ 0 \end{bmatrix}.$$

Similarly, let $D(L) = I - DL$ be defined from:

$$D = \begin{bmatrix} I_r & 0 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & I_m \end{bmatrix},$$

so that $D(L)\Delta(L) = \Delta I$. By premultiplying both sides in equation (A.11) by $M^{-1}D(L)$ and rearranging terms, we find that

$$\Delta W_t = M^{-1} D(L) J_x X_t - M^{-1} N K z_t, \tag{A.12}$$

where $\Delta Y_{3,t} = K z_t$ defines the $s_3 \times (s + d + n)$ matrix K. Substituting for X_t from equation (A.10) into (A.12), we have found the solution for W_t, and thus for Y_t, as a function of current, past, and expected future values for the exogenous variables z_t.

TABLE 1: Parameter values for the benchmark solution of the theoretical model.

Relation	parameter	value
Production	κ	1
	ρ_T	1
Labor supply	η	0.8
	ρ_L	1
Wage setting	θ_1	1.0
	θ_2	0.2
	θ_3	1/3
	ρ_W	0
	t_W	$t-1$
Price setting	ψ	1
	t_p	$t-1$
Import prices	ξ_1	1
	ξ_2	0.1
	ξ_3	0.1
	ρ_m	1
	t_m	$t-1$
Aggregate demand	γ_1	0.8
	γ_2	0.5
	γ_3	0.5
	γ_4	1
	t_y	$t-1$
Financial markets	ζ	0.25
	ρ_I	0
Monetary policy	ϕ_1	1.5
	ϕ_2	0.5
	ϕ_3	0
	ϕ_4	0.25
	ρ_M	0
CPI	δ	1/3

TABLE 2: Benchmark solution to the theoretical model.

(A) *Long Run Responses to Unit Domestic Shocks*

variable	ε_t^T	ε_t^L	ε_t^m	ε_t^M	ε_t^y	ε_t^p	ε_t^I	ε_t^W
y	0.356	0.198	-0.382	0	0	0	0	0
u	1.245	0.692	0.329	0	0	0	0	0
w	0.651	-0.115	0.214	-0.263	0.043	0.214	0.230	0.142
i	0	0	0	0	0	0	0	0
e	-0.644	0.198	-0.382	0	0	0	0	0
p^m	0.025	0.093	0.313	-0.263	0.043	0.214	0.230	0.142
p	-0.100	0.024	0.280	-0.263	0.043	0.214	0.230	0.142
s	-0.011	0.073	-0.648	-0.263	0.043	0.214	0.230	0.142
$s + p^f - p$	0.089	0.049	-0.928	0	0	0	0	0

(B) *Contemporaneous Responses to Unit Domestic Shocks*

y_t	0	0	0	0	1	0	0	0
u_t	1	1	-0.533	0	-1	-0.267	0	0.800
w_t	0	0	0	0	0	0	0	1
i_t	-0.500	0	1	1	0	0	0	0.500
e_t	-1	0	0	0	1	0	0	0
p_t^m	0	0	1	0	0	0	0	0
p_t	0	0	0.667	0	0	0.333	0	0
s_t	0.313	0.095	-1.176	-0.737	-0.043	0.286	0.770	-0.142
$s_t + p_t^f - p_t$	0.313	0.095	-1.843	-0.737	-0.043	-0.047	0.770	-0.142
$E_t \Delta s_{t+1}$	-0.500	0	1.000	1.000	0	0	-1.000	0.500
$E_t \Delta p_{t+1}$	-0.125	0.063	-0.784	-0.491	0.038	-0.143	0.513	0.239

TABLE 3: Asymptotic and bootstrapped p-values (in percent) for multivariate specification tests.

(A) *Serial Correlation Tests*

	Portmanteau		$LM(1)$		$LM(4)$	
Model	Asymp	Boot	Asymp	Boot	Asymp	Boot
$r = 1, p = 2$	0.00	4.24	0.07	0.10	5.27	8.46
$r = 2$	0.00	5.03	0.16	0.22	3.57	7.02
$r = 3$	0.00	2.40	0.41	0.93	16.34	27.63
$r = 4$	0.00	1.47	0.59	1.48	3.32	8.17
$r = 5$	0.00	1.60	7.02	16.72	2.22	7.30
$r = 6$	0.00	1.42	2.88	11.15	1.11	5.10
$r = 7$	0.00	0.90	1.46	7.48	1.63	7.67
$r = 8$	0.00	1.36	1.09	7.84	1.06	6.22
$r = 1, p = 3$	0.00	19.96	1.05	1.42	43.06	58.39
$r = 2$	0.00	20.24	3.38	4.82	11.77	24.55
$r = 3$	0.00	39.34	17.61	26.78	3.03	10.04
$r = 4$	0.00	35.31	3.95	9.89	6.46	20.23
$r = 5$	0.00	55.74	4.22	13.32	7.84	23.55
$r = 6$	0.00	54.42	10.38	31.91	17.10	43.35
$r = 7$	0.00	65.71	24.06	58.36	13.52	41.37
$r = 8$	0.00	81.54	23.30	62.06	16.80	49.42
$r = 1, p = 4$	0.00	22.02	14.22	17.70	4.12	17.40
$r = 2$	0.00	7.81	2.44	4.76	14.52	43.04
$r = 3$	0.00	6.53	3.37	8.72	12.64	43.38
$r = 4$	0.00	20.24	8.26	25.00	12.54	46.78
$r = 5$	0.00	35.48	16.56	47.04	1.20	12.61
$r = 6$	0.00	29.59	12.13	46.47	0.48	8.38
$r = 7$	0.00	50.14	14.07	55.91	1.24	14.01
$r = 8$	0.00	61.53	14.70	61.55	0.45	7.57

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 3: Continued.

(B) *Normality and ARCH Tests*

	Omnibus		Skewness		Kurtosis		ARCH	
Model	Asymp	Boot	Asymp	Boot	Asymp	Boot	Asymp	Boot
$r = 1, p = 2$	0.00	0.04	12.97	40.38	0.00	25.50	14.01	14.84
$r = 2$	0.00	0.36	17.60	47.99	0.00	17.13	9.34	11.41
$r = 3$	0.04	1.46	30.10	62.58	0.00	36.48	23.49	19.57
$r = 4$	0.00	0.62	20.21	53.00	0.00	39.16	100.00	36.67
$r = 5$	0.04	1.46	32.30	65.89	0.00	52.26	24.12	19.53
$r = 6$	0.00	0.63	21.89	55.49	0.00	58.66	30.82	23.00
$r = 7$	0.00	0.61	20.44	53.95	0.00	40.92	28.45	22.44
$r = 8$	0.00	0.47	21.09	54.54	0.00	41.95	10.37	12.99
$r = 1, p = 3$	0.00	0.25	14.52	53.84	0.00	26.35	100.00	71.77
$r = 2$	0.00	0.52	26.71	69.09	0.00	43.55	100.00	70.16
$r = 3$	7.02	28.66	24.26	68.53	0.00	40.24	100.00	67.17
$r = 4$	21.62	50.26	32.12	76.18	0.00	29.69	100.00	71.52
$r = 5$	12.98	40.72	26.19	71.97	0.00	13.31	100.00	86.87
$r = 6$	17.51	47.25	16.70	62.19	0.00	23.46	100.00	84.24
$r = 7$	26.81	56.95	20.69	65.73	0.00	10.04	100.00	77.81
$r = 8$	22.48	52.34	21.34	66.13	0.00	11.36	100.00	79.47
$r = 1, p = 4$	0.49	10.67	54.80	90.36	0.00	5.48	100.00	50.57
$r = 2$	2.24	22.64	64.14	94.33	0.00	4.78	25.98	31.00
$r = 3$	16.69	52.72	28.82	80.27	0.00	5.42	15.99	26.97
$r = 4$	25.04	64.08	59.63	94.56	0.00	58.12	100.00	54.29
$r = 5$	12.11	49.65	39.34	87.90	0.00	45.53	100.00	71.20
$r = 6$	16.36	56.83	16.85	72.88	0.00	39.92	100.00	61.67
$r = 7$	12.37	50.99	5.41	51.23	0.00	36.19	100.00	49.81
$r = 8$	15.42	53.81	4.73	48.95	0.00	54.11	100.00	63.93

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. Furthermore, 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 4: Bootstrapped distributions for the *LR* cointegration rank (trace) test.

r	$n-r+q_1$	80 %	90 %	95 %	97.5 %	99 %	LR
2 lags							
0	13	313.78	321.12	331.21	341.05	347.54	457.96
1	12	247.75	258.83	268.45	277.46	283.03	335.14
2	11	198.31	209.09	216.05	224.39	229.07	225.16
3	10	156.99	168.03	173.92	182.22	190.26	147.72
4	9	115.95	125.00	132.09	138.87	145.54	84.52
5	8	79.58	86.02	91.75	96.48	106.47	55.42
6	7	44.72	49.80	53.93	57.81	62.23	29.49
7	6	19.69	23.25	26.01	28.61	30.99	13.34
3 lags							
0	13	369.44	385.87	396.05	415.20	426.56	504.75
1	12	303.13	317.08	328.51	345.13	353.78	370.54
2	11	239.43	251.26	262.47	268.44	284.41	264.87
3	10	186.33	200.59	210.77	221.69	227.22	168.00
4	9	133.65	144.44	155.13	164.51	171.11	100.59
5	8	93.37	100.97	107.33	113.83	120.95	55.82
6	7	50.04	57.32	62.16	67.69	72.65	29.22
7	6	20.48	24.08	27.97	30.54	33.64	9.22
4 lags							
0	13	423.75	448.58	469.02	478.02	505.97	529.13
1	12	341.14	363.80	379.86	391.70	414.40	397.51
2	11	266.06	282.98	294.95	310.41	324.52	277.45
3	10	207.99	221.41	232.66	244.43	254.57	197.44
4	9	147.77	160.56	170.56	177.12	188.34	123.32
5	8	99.25	109.35	116.66	124.27	132.97	68.40
6	7	59.07	67.71	73.01	77.79	83.86	31.72
7	6	28.98	34.30	39.84	44.85	49.76	6.06

Table 5: Hypotheses about cointegration.

j	β_j	y	i	p	s	e	p^m	w	u	y^f	i^f	p^f	ν_y	ν_p	ν_i
	Case: $r = 4$, $\psi = 1$.														
	Cointegration relations														
1	Aggr. demand	1	0	$(\gamma_2 + \gamma_3)/\delta$	$-\gamma_3$	0	$-[\gamma_2 + (1-\delta)\gamma_3]/\delta$	0	0	$-\gamma_1$	0	$-\gamma_3$	0	0	$-(\gamma_2 + \gamma_3)/\delta$
2	Price setting	δ	0	1	0	$-\delta$	$-(1-\delta)$	$-\delta$	0	0	0	0	0	$-\delta$	-1
3	Wage setting	$-\theta_1$	0	-1	0	θ_1	0	1	θ_2	0	0	0	$-\theta_3$	$-\theta_3$	$-\theta_3$
4	Fin. markets	0	1	0	0	0	0	0	0	0	0	0	0	0	0
	Noncointegration relations														
5	Production	1	0	0	0	$-\kappa$	0	0	0	0	0	0	0	0	0
6	Labor supply	0	0	η	0	1	0	$-\eta$	1	0	0	0	η	0	0
7	Import prices	$-\xi_2$	0	0	$-\xi_1$	0	1	0	0	ξ_3	0	$-\xi_1$	0	0	0

TABLE 6: Estimated theoretical parameters of the cointegration vectors.

Estimated parameters (under $\delta = 1/5$, $\gamma_2 = \gamma_3 = 1/20$)

Parameter	value	Cond. std. error
γ_1	0.744	0.034
θ_1	0.406	0.079
θ_2	0.808	0.336
θ_3	0.721	0.101

TABLE 7: Identification of structural shocks from long run and contemporaneous restrictions.

(A) *Long Run Restrictions*

variable	ε^T	ε^L	ε^m	ε^M	ε^y	ε^p	ε^I	ε^W
y		a	b	0	0	0	0	0
u				0	0	0	0	0
w								
i								
e		a	b	0	0	0	0	0
p^m								
p								
s								

(B) *Contemporaneous Restrictions*

y				0	$\times$	0	0	0
u		$\times$						
w				0		0	0	$\times$
i				$\times$			0	
e				0				
p^m		0	$\times$					
p				0		$\times$		
s								

(C) *Overidentifying Contemporaneous Restrictions*

y	0	0	0		c			
u	d			0	$-c$		0	
w	0	0	0		0			
i	$-f$	0			0			f
e	$-d$	0	0		c	0	0	0
p^m	0		g	0	0	0	0	0
p	0	0	$(1-\delta)g$		0		0	0
s								

NOTES: The $\times$ character denotes a shock that has a nonzero effects on a variable according to the theoretical model.

FIGURE 1: The effects of monetary policy shocks to the short term interest rate, consumer prices, and the price level in Svensson (2000).

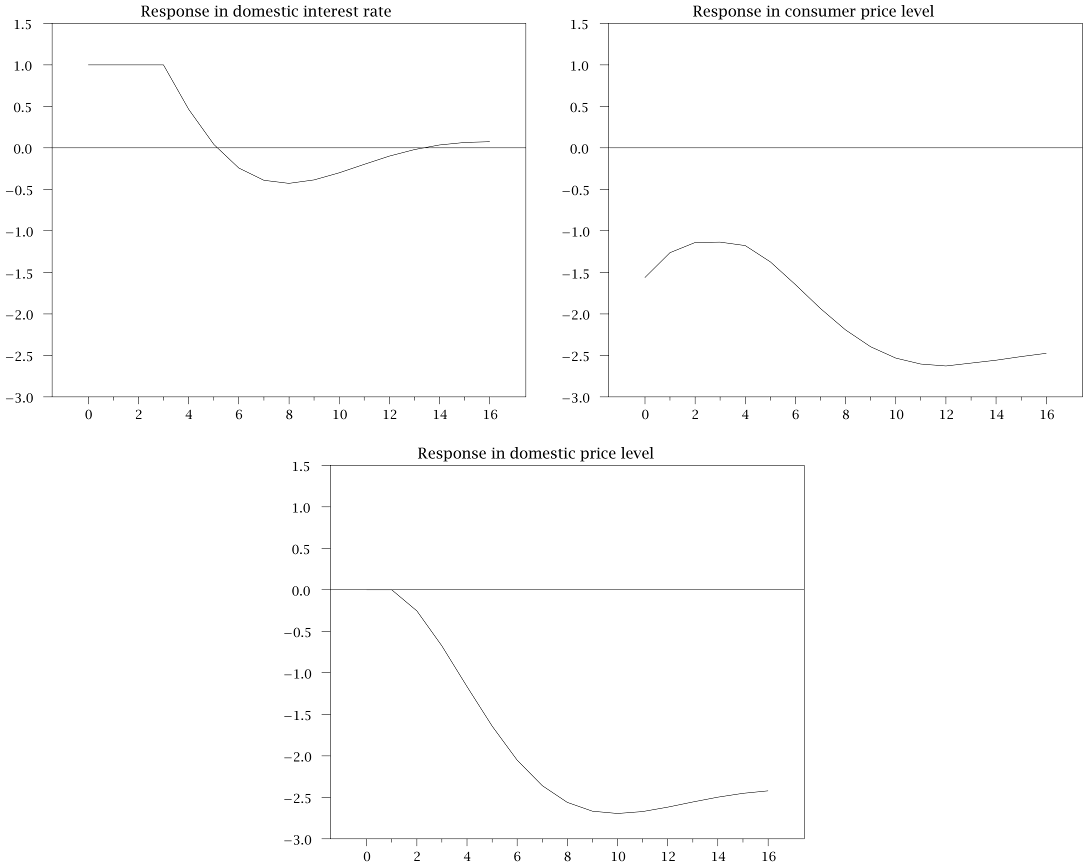


FIGURE 2: The effects on the price level and the short term interest rate in the U. S., the U. K., Italy, Germany, France, and Canada.

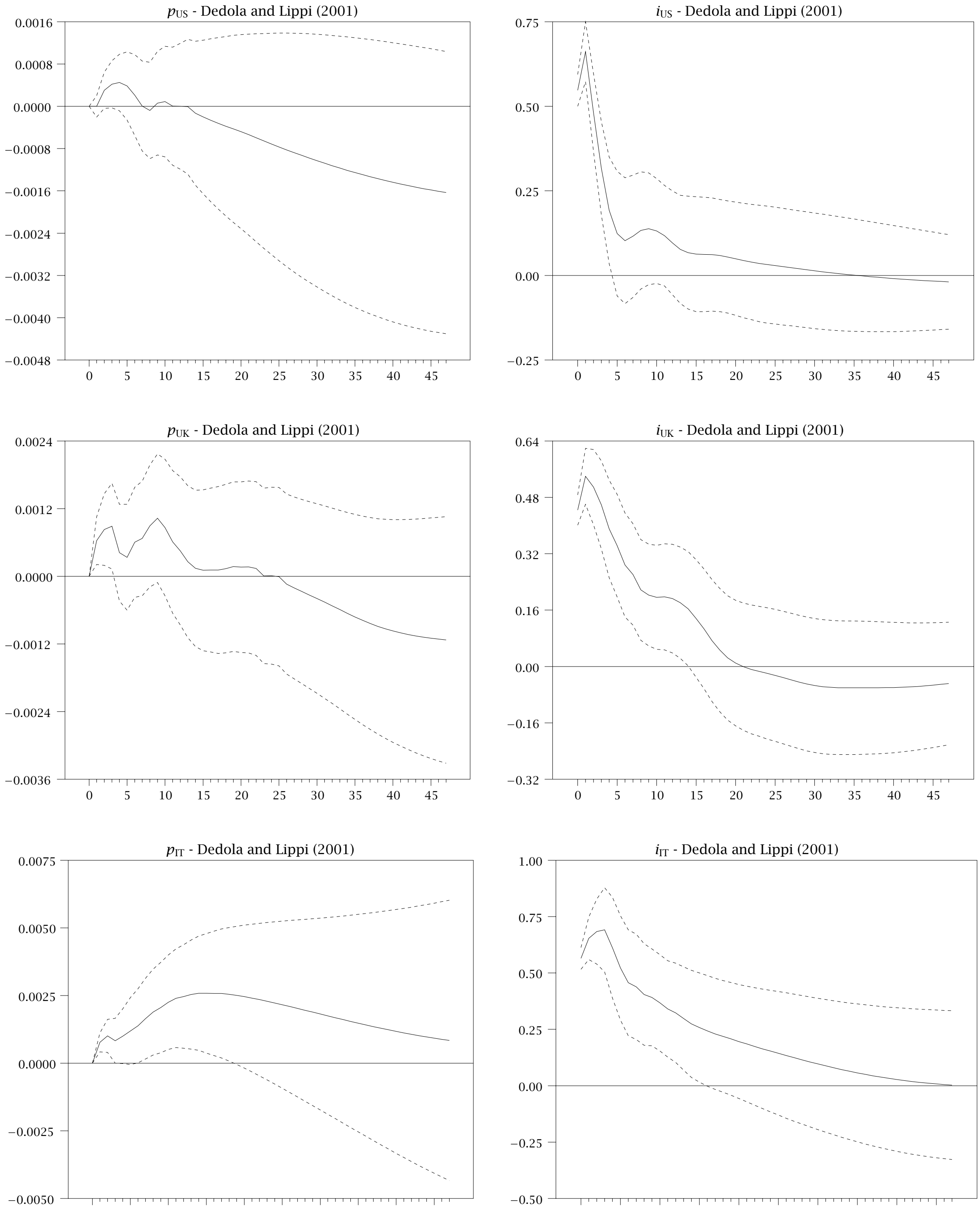


FIGURE 2: Continued.

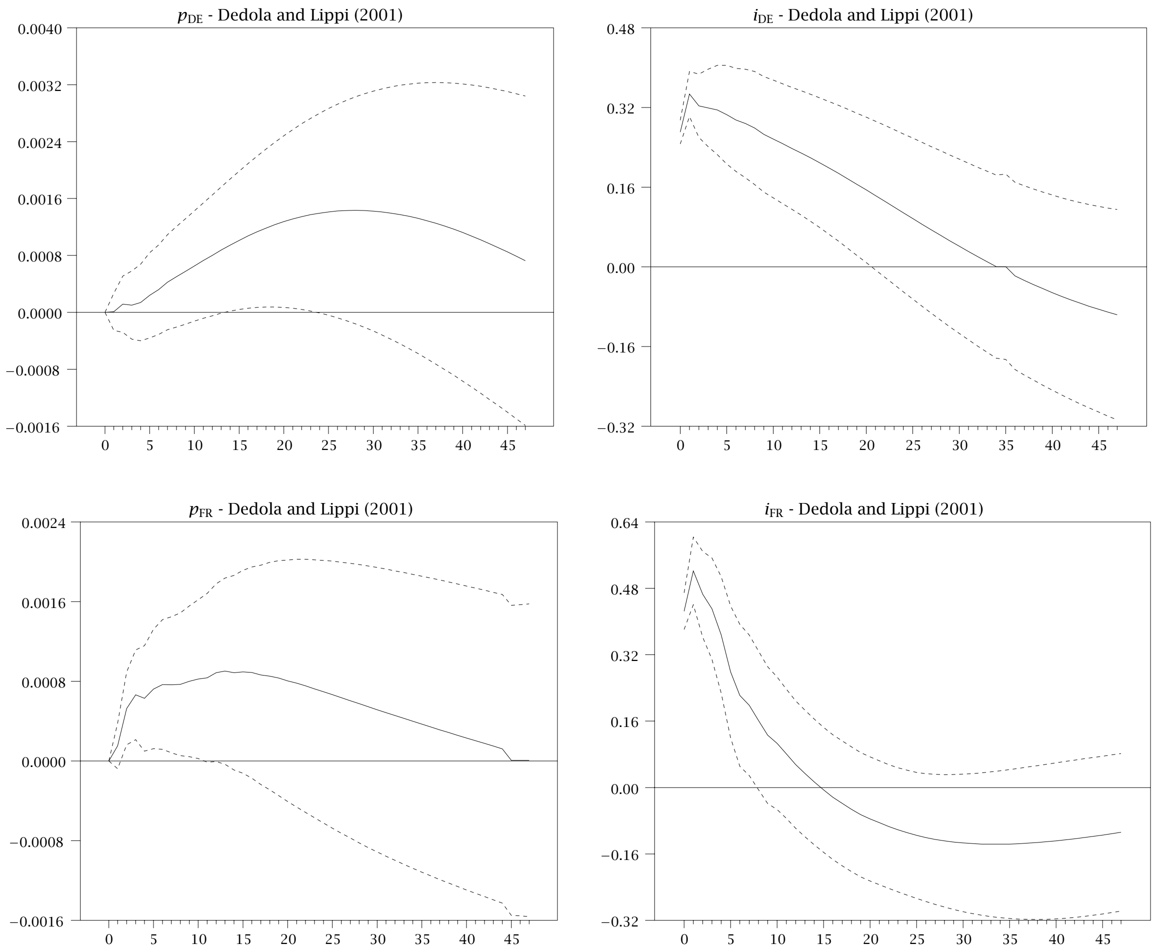


Figure 2: Continued.

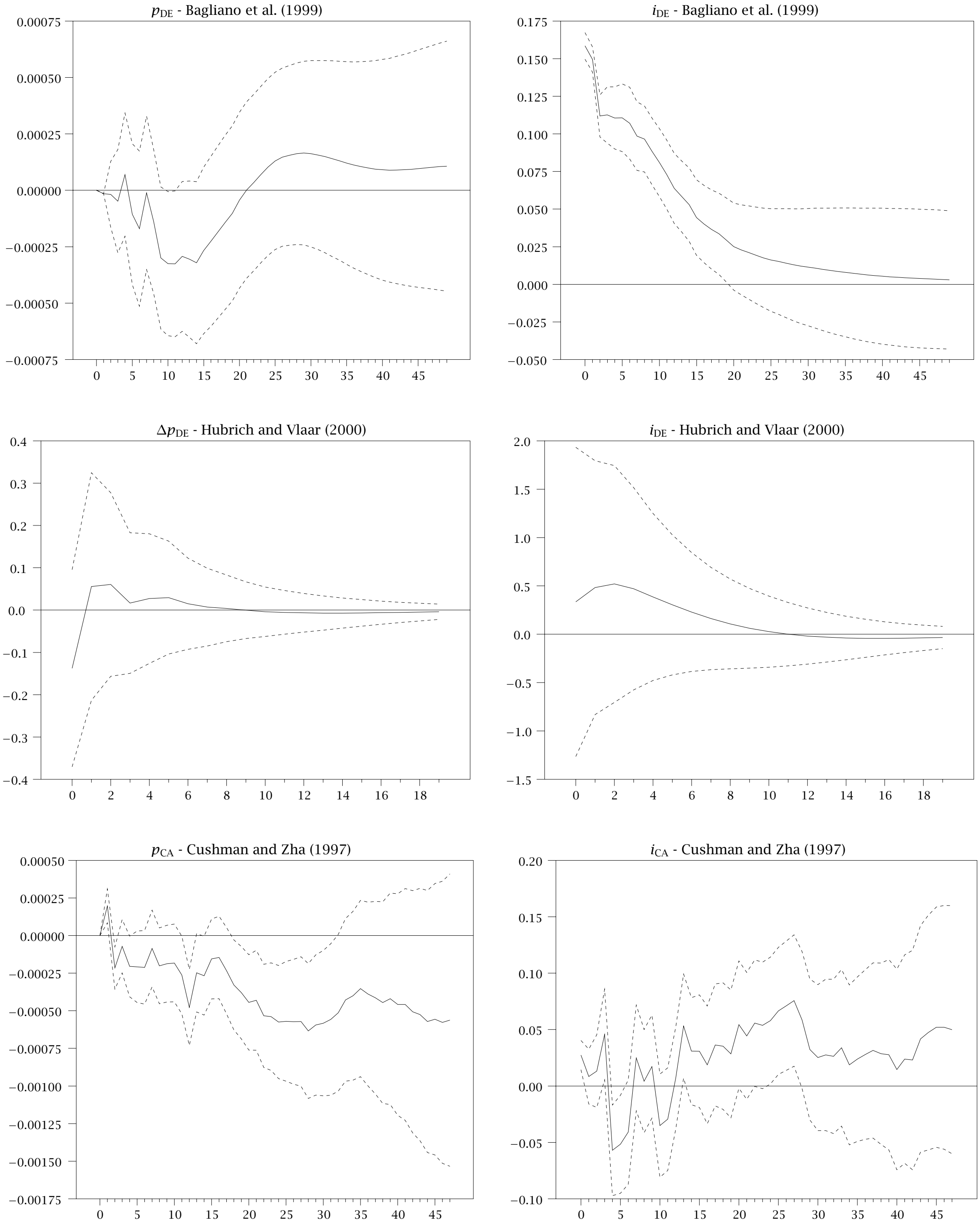


FIGURE 3: Time series observations of wage shares, tax measures, and terms of trade for the period 1970:1–1999:4.

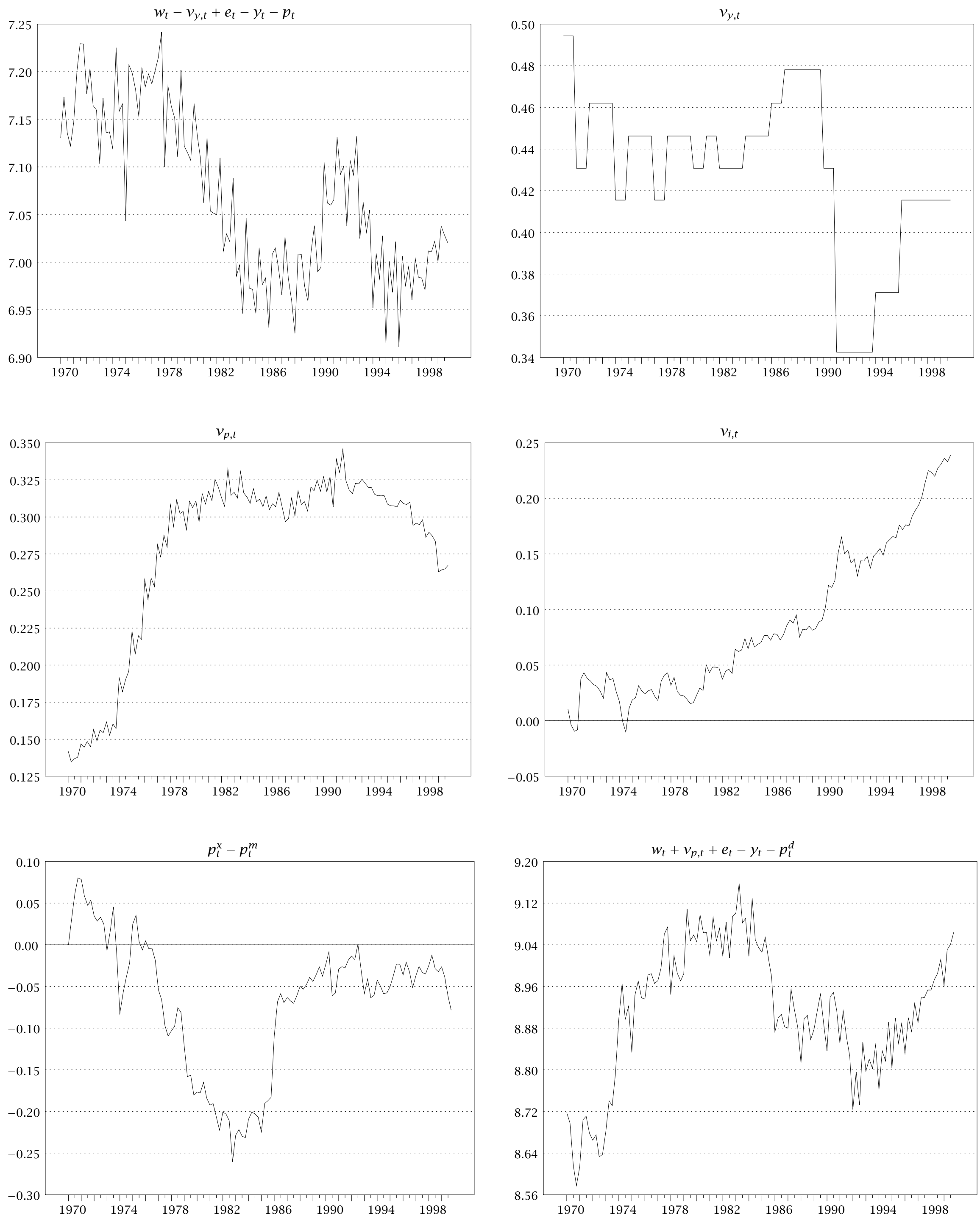


FIGURE 4: Time series observations of the 8 endogenous variables in the VAR model for the period 1970:1–1999:4.

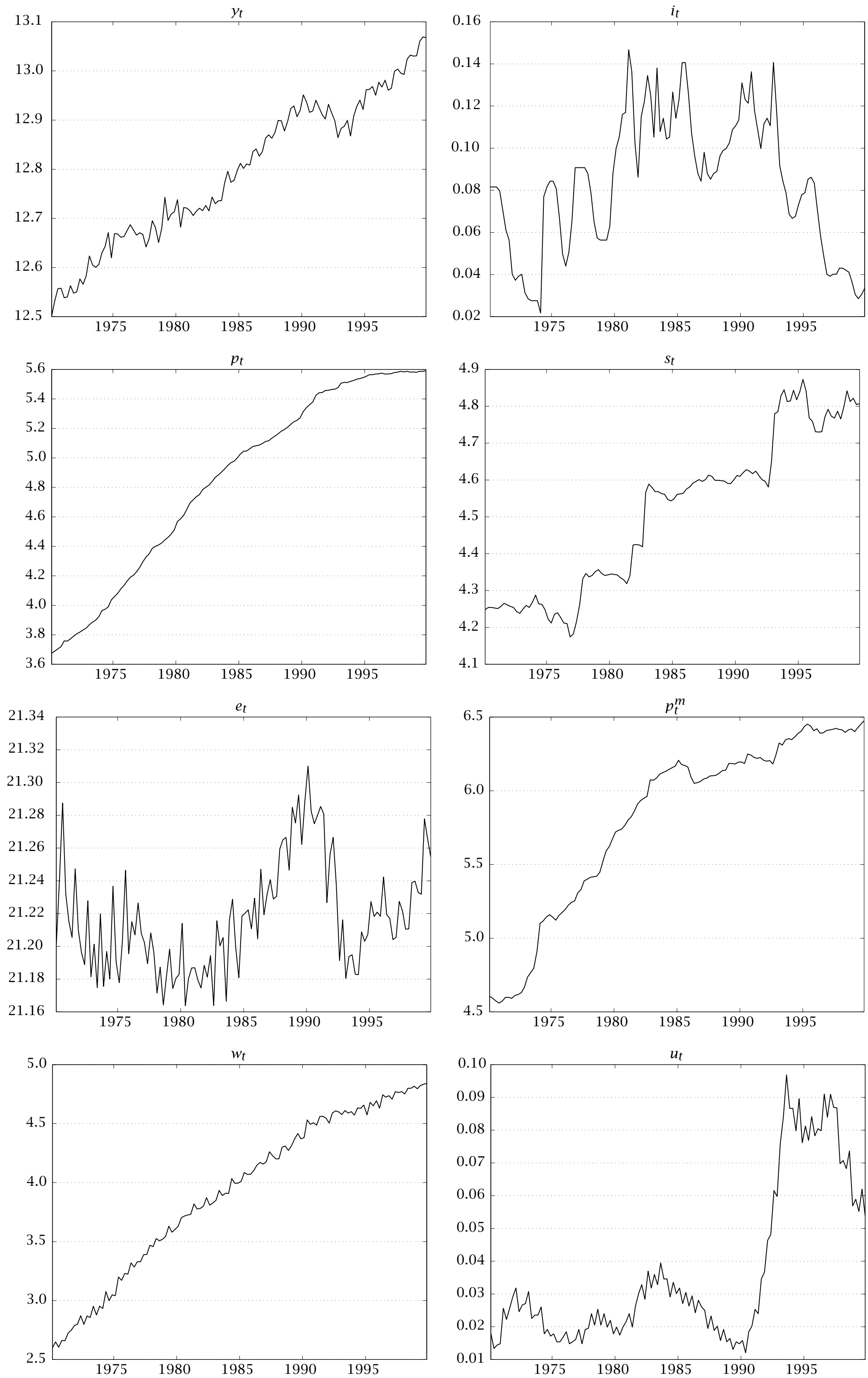


FIGURE 5: The responses in the consumer price level and the domestic interest rate from shocks to monetary policy in the structural VEC models.

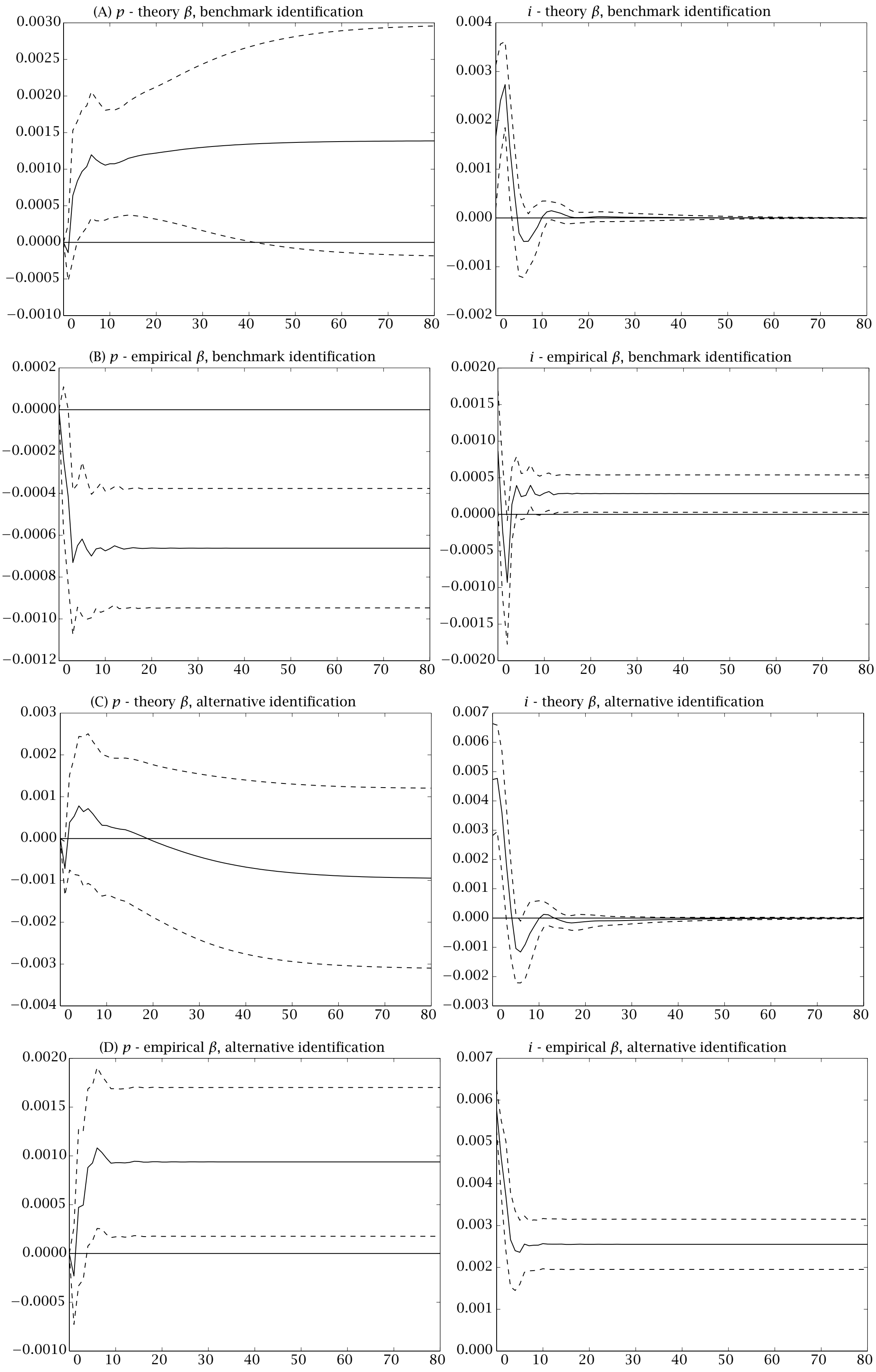


FIGURE 6: Time series plots of estimated monetary policy shocks using an instrument rule.

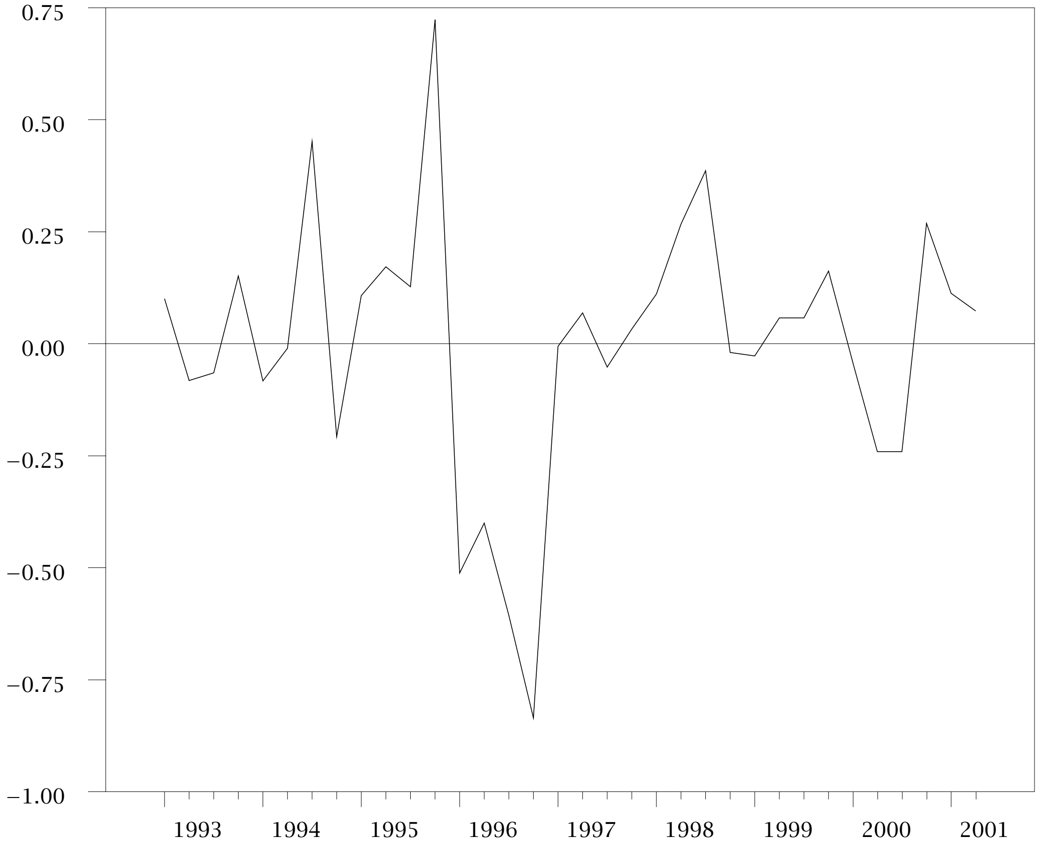


FIGURE 7: The responses in the consumer price level and the domestic interest rate from shocks to monetary policy using shocks from the estimated instrument rule.

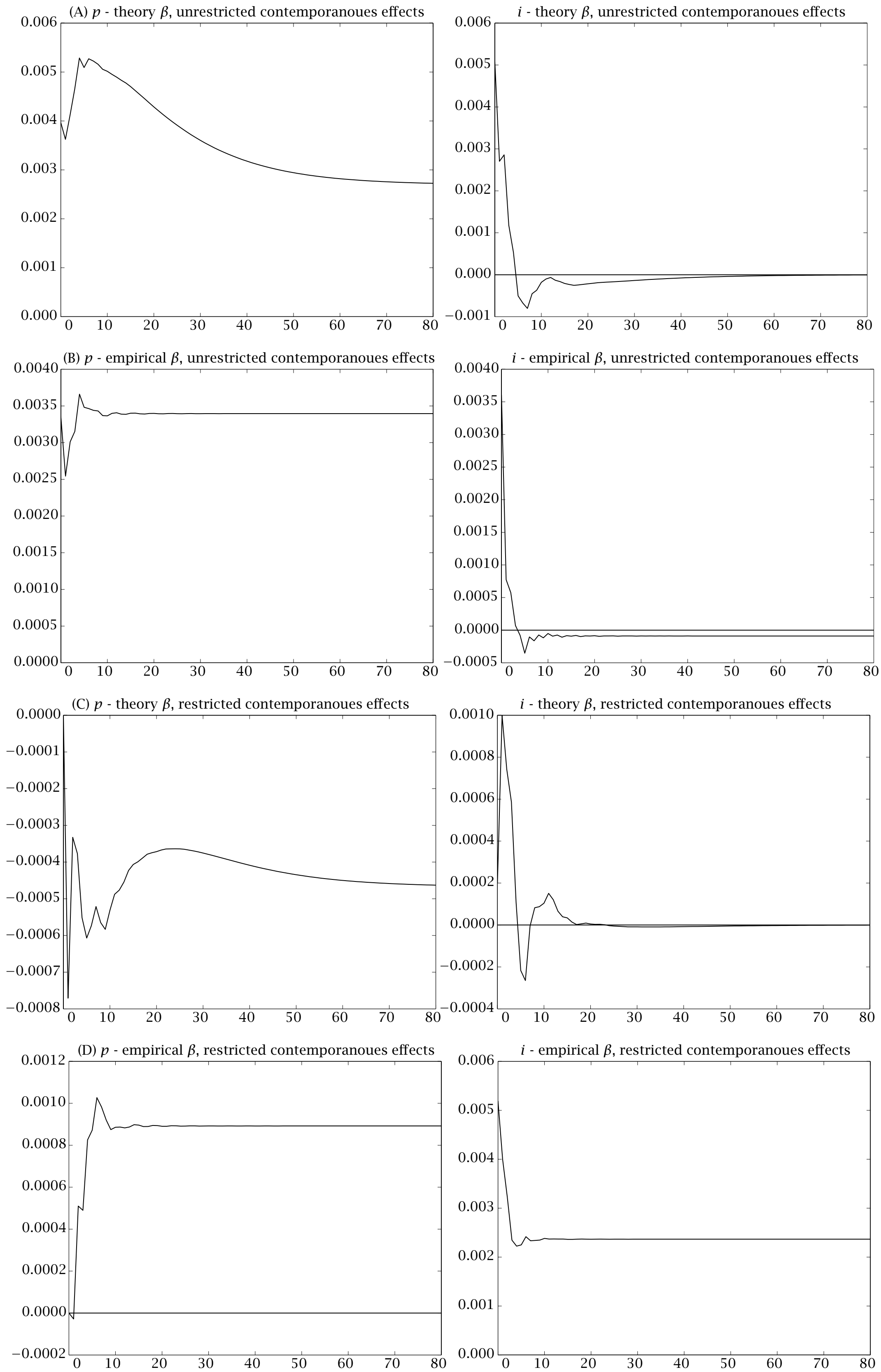


FIGURE 8: Impulse response function from a unit shock to the 1992:4 dummy using theoretical cointegration vectors.

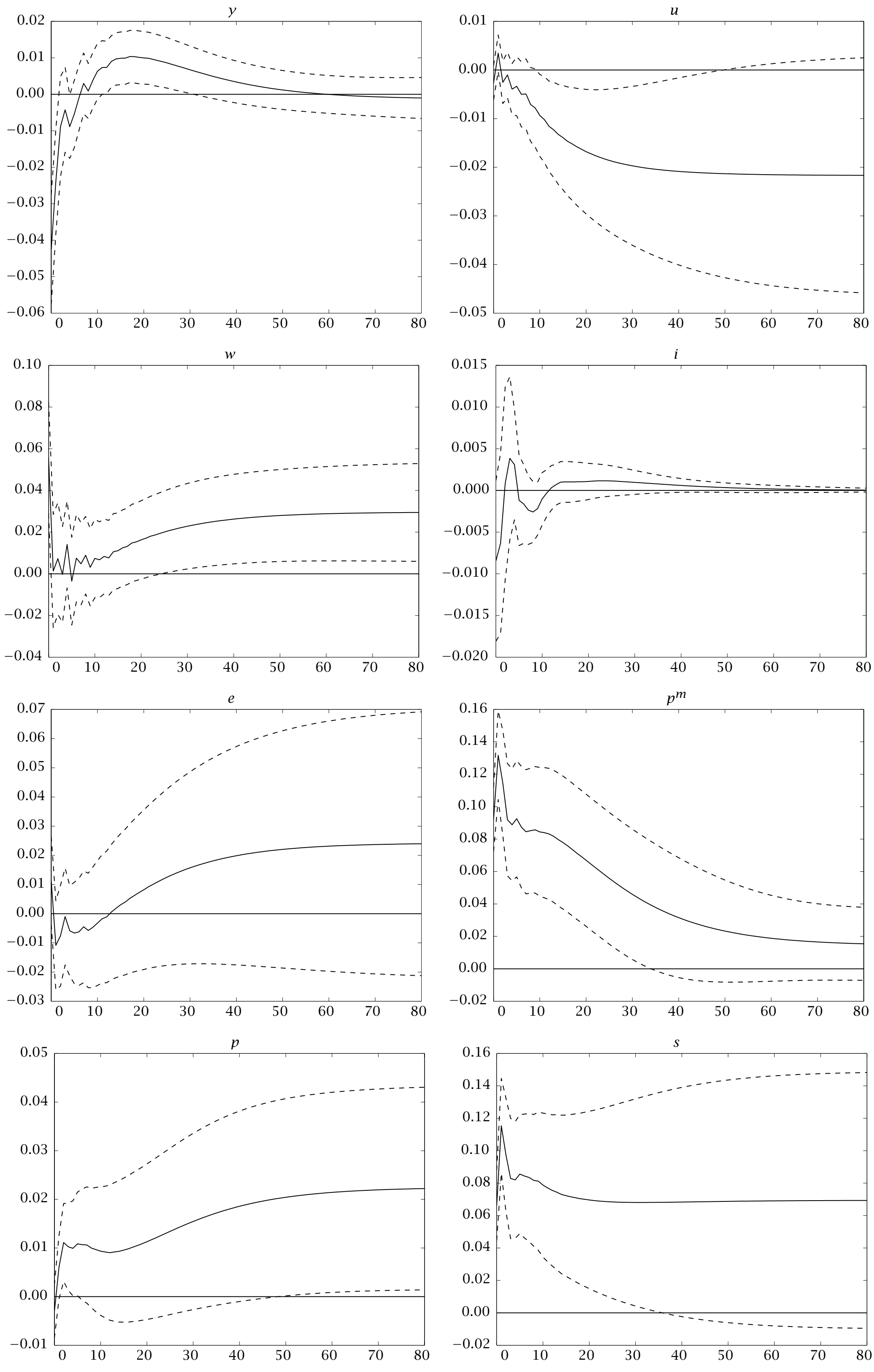


FIGURE 9: Impulse response function from a unit shock to the 1982:4 dummy using theoretical cointegration vectors.

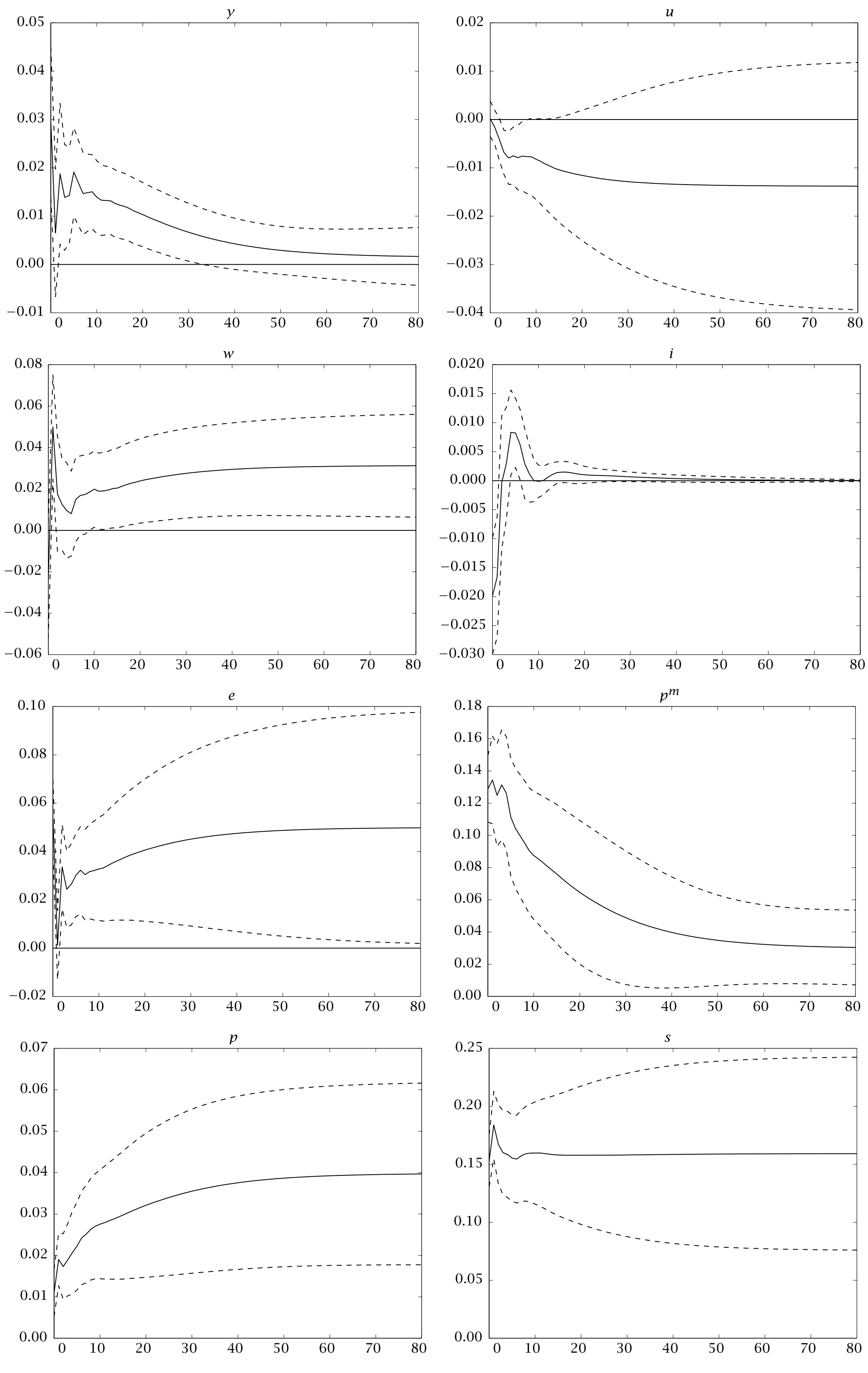
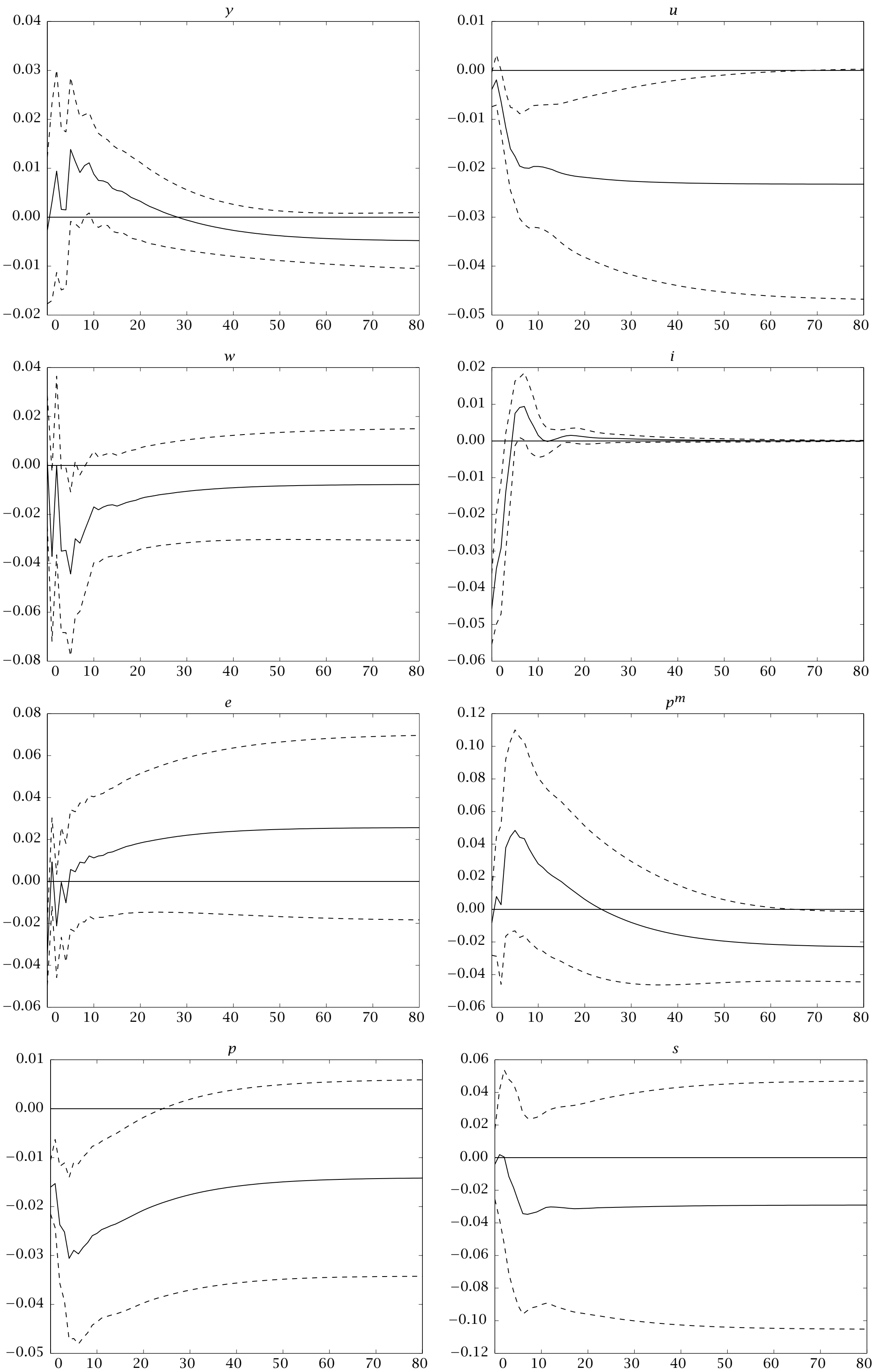


FIGURE 10: Impulse response function from a unit shock to the 1981:3 dummy using theoretical cointegration vectors.



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