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Ex-ante dynamics of real effects of monetary policy: Theory and evidence for Poland and Russia, 2001-2003

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Abstract

The paper introduces an indicator of expected real effects of an inflation-targeting policy. It compares some measures of the expected and output-neutral inflations. It is shown that the indicator, REIT, can be computed with a use of a simple vector autoregressive model of inflation and output gap. If the monetary authority has some discretion regarding the timing of monetary actions, REIT can be used to identify the optimal times for such actions. A simulation experiment illustrates its rationale. REIT has been used by the Polish Monetary Policy Council since 2001 in it's inflation targeting and might contribute to a substantial decline in Polish inflation in 2003 and an increase in output growth in 2004. A similar indicator computed for Russia and used as a mean of monitoring monetary policy confirms that the expansionary policy in 2002 - 2003 might stimulate economic growth in Russia in 2004 – 2005.

1. Introduction

The importance of good timing of monetary policy measures in terms of output effects has long been recognized (see e.g. statements by the executive board member of the European Central Bank, Issing 2001, 2002, Governor of the Bank of Canada, Dogde 2001, Mankiw 2001, van Gaasbeck 2001, Mankiw and Reis 2003). However, little has so far been done in practice to develop gauges and indicators which would help to determine the proper timing of the monetary actions, especially in the context of direct inflation targeting. In practice, measures aimed at keeping inflation within target bounds can often be undertaken with some time discretion. Pure intuition tells us that, *ceteris paribus*, such measures should be undertaken in periods when their real effects would be the most desirable (for minimising output fluctuations or output loss). This paper proposes a simple indicator which could help to evaluate in advance whether an anti-inflationary measure would also cause a minimal output distortion. Such an indicator can be computed from a vector autoregressive output-inflation model using decompositions analogous to those associated with models of output-neutral (core) inflation. The general idea of the indicator, denoted *REIT* (*real effect of inflation targeting*) is outlined in Section 2. Section 3 describes its further development within the framework of vector autoregressive modelling. Section 4 analyses the performance of *REIT* in a series of numerical experiments. Section 5 shows how *REIT* has been used by the Monetary Policy Council of Poland for inflation targeting in 2001-2003**.** It is claimed here that the use of *REIT* might have improved the decisions of the Council in 2001, resulting in a significant reduction in inflation in 2003 and an increase in output in 2004. A similar indicator computed for Russia, where it was not reported to the monetary authorities, shows that the actions taken reduce interest rates in 2002 and 2003 might have positively affected Russian economic growth in 2004 and 2005. However, further actions regarding the interest rate, undertaken in 2004, are likely to prove ineffective.

2. The basic model

The problem is illustrated by a simple generalisation of a typical aggregate supply function:

$$y_t = \overline{y}_t + q(p_t - p_t^n) \quad , \tag{1}$$

where y_t denotes output in time period t, $\overline{y}_t$ its natural level, p_t the aggregate price index, p_t^n the output-neutral price. All these variables are in logs. Interpretation of the parameter q varies depending on the particular microeconomic foundations of the supply function. In purely neoclassical models $p_t^n = p_t^e$, i.e. output-neutral price is equal to that expected at *t*-1 for time *t*. The short-run representation of (1) is:

$$\tilde{y}_t = q\,(p_t - p_t^n) \quad ,$$

where $\tilde{y}_t$ is the output gap defined as the difference between the logs of the actual and natural output levels, headline (observed) inflation is defined as $p_t = p_t - p_{t-1}$ and output-neutral inflation as $p_t^n = p_t^n - p_{t-1}$. In a purely neoclassical model, $p_t^n = p_t^e$. Nevertheless, there might exist an economy with less-than-perfectly-flexible prices, where some individual relative prices cannot be fully adjusted after a shock and hence could have long-lasting effects on output, even if fully expected.

Suppose that, at time *t*-1, the monetary authority is making its decision regarding the control of inflation on the basis of all available information. Such a decision is unexpected for other agents, for whom the relationship between the price expected at *t*-1 and observed at *t* is:

$$p_t = p_t^e + u_t \quad , \tag{2}$$

where u_t is a shock unexpected at *t*-1. This shock can be interpreted as a composition of a supply shock and an unexpected outcome of a monetary action. Obviously, for all agents in the economy except the central bank (CB), $E_{t-1}u_t = 0$, where E_t denotes the expected value based on information available at time *t*. The CB, albeit equally ignorant regarding a possible supply shock, is in the privileged position of having its own evaluation of the possible inflationary impact of the monetary measure:

$$E_{t-1}^B u_t = m_t \quad ,$$

where E_t^B denotes the CB's bankers' expectation at time *t*. Another decomposition of p_t is:

$$p_t = p_t^n + w_t \quad , \tag{3}$$

where w_t is the non-neutral component of inflation. The evaluation of p_t^n is also based on information available at time *t*-1 and is known at that time. Referring to the seminal literature on inflation decomposition, p_t^e is similar to core inflation in the sense of Eckstein (1981), i.e. the systematic (predictable) component of the increase in production costs. In turn, p_t^n is analogous to core inflation in the sense of Quah and Vahey (1995), i.e. the component of expected inflation which does not cause a real effect in the medium and long-run.

From (2) and (3) we obtain w_t as

$$w_t = p_t^e - p_t^n + u_t \quad ,$$

which yields $E_{t-1}^B w_t = p_t^e - p_t^n + m_t$. Hence the CB's conditional expected value of an increase in output is

$$E_{t-1}^B \tilde{y}_t = q \cdot E_{t-1}^B w_t = q \cdot (p_t^e - p_t^n + m_t) \quad . \tag{4}$$

On the basis of (4) one can derive an interesting indicator for proper timing of a monetary measure. The sign and magnitude of the difference between p_t^e and p_t^n indicate the possible output effects of a monetary action undertaken at time t-1. Let us write

$$REIT_t = p_t^e - p_t^n \quad ,$$

where $REIT_t$ denotes the *real effect of inflation targeting*. Suppose that, in order to keep inflation within target, the central bank wants to undertake a contractionary measure, that is expected to change inflation by $m_t < 0$. Clearly, the expected non-neutral component of inflation, proportional to the expected real effect is given by (4). Hence the output loss generated by such a contractionary action will be smaller if the action is undertaken while $p_t^e > p_t^n$, i.e. when $REIT_t$ is positive, rather than when $REIT_t \leq 0$. Similarly, it is expected (by the CB) that an expansionary policy ($m_t > 0$) will be relatively effective with positive $REIT_t$, since in this case the expected output gain is greater than with a non-positive $REIT_t$.

Hence it is conjectured that the CB should pay particular attention to monitoring the differences between p_t^e and p_t^n. Suppose, for instance, that the CB has some time to make a decision aimed at reducing inflation, say between times T and T+k. If the secondary objective is to minimise the loss in output, then the optimal time for an increase in interest rate would be

$$t_{opt} = \{t: \max(REIT_t)\} \quad , \; t = T, T+1, \ldots T+k \quad .$$

Such a decision requires sufficient knowledge about both inflation indicators, p_t^e and p_t^n, in periods T, T+1, … T+k. As mentioned above, p_t^e can be regarded as a gauge of the overall inflationary tendency over a long period of time and could be evaluated by one (or more) of the core inflation measures (in the Eckstein sense). One way of computing p_t^n is similar to that derived from the Quah - Vahey (1995) structural decomposition of a vector autoregressive model by Dewachter and Lusting (1997), Gartner and Wehinger (1998), Wehinger (2000) and Hahn (2002).

3. Further interpretation: a *VAR* model

In order to explain further the relevance of *REIT*, we consider a simple two-equation vector autoregressive (*VAR*) model with $\tilde{y}_t$ and p_t. Let us assume that both $\tilde{y}_t$ and p_t are integrated of order zero and that their cummulatives, the logs of natural levels of output and prices, are not cointegrated. Hence, the *VAR* model can be written as:

$$A(L)Z_t = K + U_t \quad , \tag{5}$$

where $Z_t = \begin{bmatrix} \tilde{y}_t \\ p_t \end{bmatrix}$, $A(L)$ is the lag polynomial operator, $K = \begin{bmatrix} k_1 \\ k_2 \end{bmatrix}$ the vector of constants, $U_t = \begin{bmatrix} u_{1t} \\ u_{2t} \end{bmatrix}$ innovations with zero expectations and variance-covariance matrix $\Sigma = \begin{bmatrix} s_{11} & s_{12} \\ s_{12} & s_{22} \end{bmatrix}$. Suppose further that the output innovation u_{1t} can be decomposed into a technological shock, w_t, and the real effect of the inflation 'surprise', u_{2t}, i.e.

$$u_{1t} = w_t + d\,u_{2t} \quad . \tag{6}$$

Consequently:

$$U_t = \begin{bmatrix} 1 & d \\ 0 & 1 \end{bmatrix} W_t \ , \tag{7}$$

where $W_t = [w_t \ , u_{2t}]'$.

A convenient derivation of such a model could be done on basis of the microfoundations independently given by Chari *et al* (2000) and Ascari (2000) for the Taylor (1979, 1980) concept of staggered prices. A comparison of the Chari e*t al* and Ascari models is given by Dixon and Kara (2005), empirical insight is analysed by Whelan (2004) and, for a critique of this approach, see Fourgère, Le Bihan and Sevestre (2004). According to this theory, and unlike familiar Calvo (1983) scheme, the probability of a price (or wage) contract expiring in specific periods of time may not be constant, but the fraction of firms setting prices at a given time covering a fixed period N is constant and equal to $1/N$. A general equilibrium approach, with maximisation of intertemporal consumers' utility functions and profit functions for final and intermediate goods (with staggered price effect explicitly formulated) and after some simplification and linearization around the steady-state yields

$$x_t = E_t\left(\sum_{i=0}^{2(N-1)} f_i \tilde{p}_{t+i} + g \sum_{i=0}^{2(N-1)} f_i \tilde{y}_{t+i} + \right) + e_t \quad , \tag{8}$$

where $\tilde{p}_t$ and $\tilde{y}_t$ are deviations of prices and output from the steady-state and e_t is the expectation error. The weights f_i are derived from the assumption that contract wages, x_t, reflect consumers' expectations regarding future prices and excess demand, as measured by the output gap. The parameter g in the original Taylor model represents the sensitivity of wages to aggregate demand policy, while in Chari *et al* (2000) it is interpreted as the elasticity of equilibrium real wage with respect to consumption. Since in the Taylor model price is a weighted average of negotiated, contemporaneous and preceding, nominal wage contracts under rational expectations, formula (8) leads to *VAR* system (5), where the lag length is *N*-1. Similar micro support for the inflation-output *VAR* model has also been used by Coenen and Wieland (2005).

The technological shock w_t and inflation surprise u_{2t} in (6) are uncorrelated, which implies that $d = s_{12} / s_{22}$, so that the covariance matrix of vector $W_t = [w_t, u_{2t}]'$ is diagonal. The vector moving average representation of (5) is

$$Z_t = M + C(L)U_t \quad , \tag{9}$$

where $M = [m_1, m_2]' = EZ_t = C(1)K$ and $C(L) = A^{-1}(L) = I + C^{(1)}L + C^{(2)}L^2 + \ldots$. Using (7), this can also be expressed as a vector moving average representation of technological shocks and inflation surprises, i.e.

$$Z_t = M + C(L)\begin{bmatrix} 1 & d \\ 0 & 1 \end{bmatrix} W_t = M + S(L)W_t \quad , \tag{10}$$

where:

$$S(L) = C(L)\begin{bmatrix} 1 & d \\ 0 & 1 \end{bmatrix} = I + S^{(1)}L + S^{(2)}L^2 + \ldots, \text{ and } S^{(i)} = C^{(i)}\begin{bmatrix} 1 & d \\ 0 & 1 \end{bmatrix} \quad . \tag{11}$$

Decomposition into the unitary innovations in (5) is given by

$$Z_t = M + \Gamma(L)\Phi_t \quad , \tag{12}$$

where: $\Gamma(L) = \Gamma^{(0)} + \Gamma^{(1)}L + \Gamma^{(2)}L^2 + \cdots$, $\Phi_t = [j_{1t}, j_{2t}]'$, and $E\Phi_t\Phi_t' = I$ (identity matrix).

The desired long-run output-neutral decomposition is defined as

$$Z_t^* = \begin{bmatrix} g_{11} & 0 \\ g_{21} & g_{22} \end{bmatrix} \times \begin{bmatrix} j_{1t} \\ j_{2t} \end{bmatrix} \quad ,$$

where g_{kj} (k,j = 1,2) are elements of the long-run matrix $\Gamma(1)$, i.e.

$$\Gamma(1) = \Gamma^{(0)} + \Gamma^{(1)} + \Gamma^{(2)} + \cdots = \begin{bmatrix} g_{11} & 0 \\ g_{21} & g_{22} \end{bmatrix} \quad . \tag{13}$$

Comparing (9) and (10) with (12) and noticing (11), we obtain the relationship between vectors U_t (or W_t) and Φ_t

$$\Gamma(1)\times\Phi_t = C(1)\times U_t = S(1)W_t \ , \tag{14}$$

so that $\Gamma(1)$ can be computed as the lower-triangular Cholesky factor of $C(1)\Sigma C(1)'$.

It is useful to note that matrix $\Gamma^{-1}(1)$ is also lower-triangular,

$$\Gamma^{-1}(1) = \frac{1}{g_{11}g_{22}} \begin{bmatrix} g_{22} & 0 \\ -g_{21} & g_{11} \end{bmatrix}, \tag{15}$$

and from (14)

$$\Phi_t = \Gamma^{-1}(1)\times C(1)\times U_t = \Gamma^{-1}(1)\times S(1)\times W_t \quad . \tag{16}$$

Denote

$$V_t = \begin{bmatrix} v_{1t} \\ v_{2t} \end{bmatrix} = S(1)\times W_t = \left(I + S^{(1)} + S^{(2)} + \ldots\right) \begin{bmatrix} w_t \\ u_{2t} \end{bmatrix} \quad . \tag{17}$$

Equations (16), (17) and (12) suggest the interpretation of v_{1t} as the cumulative shock effect on output and v_{2t} as the cumulative shock effect on inflation. Using (15), (16) and (17) we obtain

$$\Phi_t = \begin{bmatrix} j_{1t} \\ j_{2t} \end{bmatrix} = \Gamma^{-1}(1)\times V_t = \frac{1}{g_{11}g_{22}} \begin{bmatrix} g_{22} & 0 \\ -g_{21} & g_{11} \end{bmatrix} \begin{bmatrix} v_{1t} \\ v_{2t} \end{bmatrix} = \frac{1}{g_{11}g_{22}} \begin{bmatrix} g_{22}v_{1t} \\ -g_{21}v_{1t} + g_{11}v_{2t} \end{bmatrix} \quad .$$

This shows that $j_{1t} = g_{11}^{-1}v_{1t}$ is proportional to the cumulative shock effect on output and the long-run output-neutral component. The term j_{2t} can be expressed as a linear combination of the cumulative shock effect on output and inflation

$$j_{2t} = -\frac{g_{21}}{g_{11}g_{22}} v_{1t} + \frac{1}{g_{22}} v_{2t} \quad .$$

Using the orthogonality condition $E\Phi_t\Phi_t' = I$ we can show that $Var(v_{1t}) = g_{11}^{2}$ and $Cov(v_{1t}, v_{2t}) = g_{11}g_{22}^{-1}g_{21}^{2}$. Also $Var(v_{2t}) = g_{22}^{2} - g_{21}^{2} + 2g_{21}^{3}g_{22}^{-1}$. Given information available at t-1, the one-step ahead forecast of p_t can be computed from one of the following:

$$p_t^e = m_2 + \sum_{i=1}^{\infty}(g_{21}^{(i)}j_{1t-i} + g_{22}^{(i)}j_{2t-i}) \quad ,$$

where $g_{kl}^{(i)}$ are the elements of matrices $\Gamma^{(i)}$ (see (12)), or

$$p_t^e = m_2 + \sum_{i=1}^{\infty}(c_{21}^{(i)}u_{1t-i} + c_{22}^{(i)}u_{2t-i}) \quad ,$$

where $c_{il}^{(i)}$ are elements of matrices $C^{(i)}$ (see (9)), or, from (10),

$$p_t^e = m_2 + \sum_{i=1}^{\infty}[c_{21}^{(i)}w_{t-i} + (d\,c_{21}^{(i)} + c_{22}^{(i)})u_{2t-i}] \quad . \tag{18}$$

Ex-ante evaluation, of output-neutral inflation based on information from the past, is given by (see (12) and (13))

$$p_t^n = m_2 + \sum_{i=1}^{\infty}g_{22}^{(i)}j_{2t-i} \quad , \tag{19}$$

where j_{2t-i} can be obtained from (16):

$$j_{2t-i} = ([c_{21} - c_{11}\frac{g_{21}}{g_{11}}]w_{t-i} + [(c_{22} + d\,c_{21}) - \frac{g_{21}}{g_{11}}(c_{12} + d\,c_{11})]u_{2t-i} \quad . \tag{20}$$

This yields

$$p_t^n = m_2 + \frac{1}{g_{22}}\sum_{i=1}^{\infty}g_{22}^{(i)}([c_{21} - c_{11}\frac{g_{21}}{g_{11}}]w_{t-i} + [(c_{22} + d\,c_{21}) - \frac{g_{21}}{g_{11}}(c_{12} + d\,c_{11})]u_{2t-i} \ . \tag{21}$$

Finally, $REIT_t$ can be defined as

$$REIT_t = p_t^e - p_t^n = \sum_{i=1}^{\infty}g_{21}^{(i)}j_{1t-i} = \frac{1}{g_{11}}\sum_{i=1}^{\infty}g_{21}^{(i)}[c_{11}w_{t-i} + (c_{12} + d\,c_{11})u_{2t-i}] \quad .$$

In order to provide further interpretation for the possible real effects of inflation targeting, let us narrow our analysis for a *VAR*(1) model (where staggered wage contracts cannot cover more than 2 periods):

$$Z_t = K + BZ_{t-1} + U_t \quad . \tag{22}$$

This corresponds to $A(L) = I - BL$ in (5), where I is the identity matrix. Note that in the particular case of model (22) the lag operator $C(L)$ in (9) is reduced to $C(L) = I + BL + B^2L^2 + B^3L^3 + ...$ and hence the vector moving average representation of (22) can be simplified to

$$Z_t = M + U_t + BU_{t-1} + \sum_{i=2}^{\infty} B^i U_{t-i} \quad . \tag{23}$$

In order to show how *REIT* affects output, let us assume that, for period t, both p_t^e and $REIT_t$ are fixed at given levels. Identification here can be achieved by setting U_{t-1}. We introduce matrix F such that

$$FU_{t-1} = \begin{bmatrix} p_t^e - g_1 \\ REIT_t - g_2 \end{bmatrix} \quad , \tag{24}$$

where g_1 and g_2 depend only on information from periods t-2, t-3 *etc*:

$$g_1 = m_2 + \sum_{i=2}^{\infty} (c_{21}^{(i)} u_{1t-i} + c_{22}^{(i)} u_{2t-i}) \text{ , and } g_2 = \sum_{i=2}^{\infty} g_{21}^{(i)} j_{1t-i} \ .$$

Assume also that matrix F is invertible (discussion of the consequence of its non-invertibility is given further in this section). Hence the representation (23) can be re-written as

$$Z_t = M + U_t + Q\begin{bmatrix} 0 \\ REIT_t \end{bmatrix} + Q\begin{bmatrix} p_t^e - g_1 \\ g_2 \end{bmatrix} + \sum_{i=2}^{\infty} B^i U_{t-i} \ , \tag{25}$$

where $Q = BF^{-1}$. The matrix Q takes the form $Q = \begin{bmatrix} q_1 & q_2 \\ 1 & 0 \end{bmatrix}$, where q_1 and q_2 are functions of model parameters. Clearly, the element of interest here is q_2, which shows an effect of $REIT_t$ on output. It can be shown that

$$q_2 = -\frac{\det B}{\det F} \quad . \tag{26}$$

After some algebraic manipulation, the determinant of F can be expressed as

$$\det F = -\frac{(b_{22}^2 - b_{22} + b_{12}b_{21})H}{g_{11}\det(I - B)} , \tag{27}$$

where b_{ij} $(i, j = 1,2)$ are the elements of matrix B and

$$H = b_{21} g_{11} (b_{11} + b_{22} - 1) + g_{21} (b_{22}^2 - b_{22} + b_{12} b_{21}) . \tag{28}$$

The economic conditions for the staggered price model (in particular, inflation and output persistence) imply that $\det B > 0$ and the stability condition implies that $\det(I - B) > 0$. Moreover, g_{11} is also positive, being the upper-left element of the Cholesky decomposition. Consequently, the condition for $REIT > 0$ to have a positive impact on output is the negativity of the numerator in (27). Let us denote

$$G = b_{22}^2 - b_{22} + b_{12} b_{21} \quad .$$

With the inflation persistence coefficient, b_{22}, being between zero and one and b_{12} and b_{21} having the opposite signs, G is negative. In fact the sign of b_{21}, which is the parameter representing the consumption elasticity of equilibrium real wage, should be positive. Even if the sign of b_{12} is also positive, G would usually be negative, since the product of cross-effects, is small relative to the inflation persistence coefficient, which often takes values around 0.5. In such case, the condition $q_2 > 0$ depends on whether $H < 0$. From (28) it is clear that this in turn depends on the correlation matrix Σ of U_t and the matrix of coefficients B, which determine on the sign of g_{21}. However, given $H = 0$ and taking into account that g_{11} and g_{21} are both functions of the elements of matrices Σ and B, b_{12} can be expressed as an implicit function of the remaining parameters in (28). For $b_{21} > 0$, there exist such $b_{12}^* < 0$ for which $H = 0$, and thus $\det F = 0$. For $b_{12} > b_{12}^*$, its relation with q_2 is of a hyperbolic nature with the real effect of a positive $REIT$ being large for small positive differences between b_{12} and b_{12}^* and diminishing with increases in this difference. This is illustrated by plots of numerical values of q_2 against b_{12} in selected $VAR(1)$ models. Figure 1 represents a situation where parameters $b_{21} = 0.1$ and $b_{22} = 0.6$ and the covariance matrix is given by

$$\Sigma = \begin{bmatrix} 1.0625 & 0.25 \\ 0.25 & 1.0 \end{bmatrix} \quad .$$

The parameter b_{11} (output persistence) varies such that b_{11} = 0.4 , 0.6 and 0.8 and the parameter b_{12} varies from $b_{12}^* \approx -0,1625$ (numerically computed) to 0.1725. This indicates that, for a fixed expected inflation, a negative cross-effect of inflation on output (but greater than b_{12}^*) causes $REIT$ to affect output with much more force than the positive cross-effects. It also illustrates that an increase in output persistence, although positively affecting output here, does not markedly change the 'bound' value of b_{12}^{*}.

The 'bound' value of b_{12}^{*} does, however, depend on inflation persistence. Figure 2 shows the numerically computed values of b_{12}^{*} plotted against the inflation persistence parameter b_{22}, which varies the range 0.3 - 0.8. It shows that the lower bound for the b_{12} parameter is quite low for low and intermediate levels of of inflation persistence. Only for very high levels of b_{22} (approaching the limits of the stability condition of the *VAR*(1) model), must the parameter b_{12} be positive or mildly negative in order to ensure a positive real effect of a positive *REIT*.

Figure 1: Impact of *REIT* (q_2 in relation to b_{12})

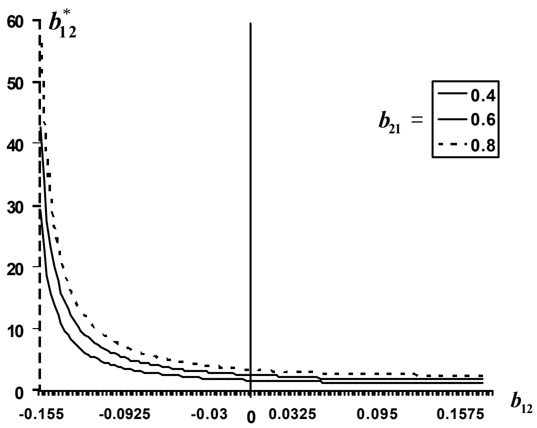
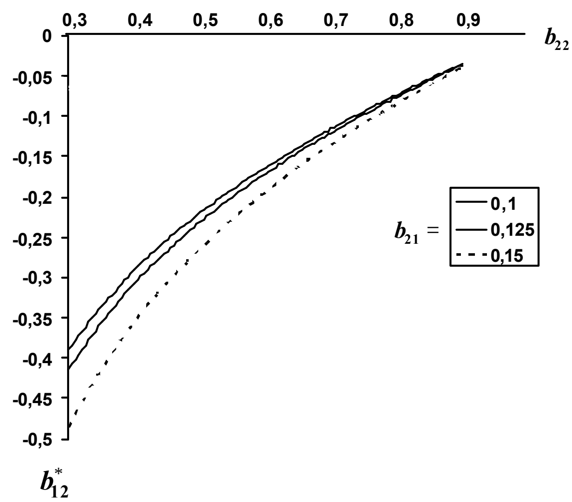


Figure 2: Bound values b_{12}^{*} and inflation persistence



4. Simulation experiment

Since we are considering an inflation-output model without policy instruments, it is not possible to directly model the impact of a monetary measure. This is not the aim of the paper, which concentrates on the climate and possible effects of monetary actions rather than on the decision-making itself (see e.g. Uhlig 2005). Hence the rationale of *REIT* can be illustrated by simulation of time-aggregated real effects of inflationary shocks on output assuming that these shocks result from monetary policy actions. In order to carry out the experiment, model (5) was simplified to a first-order *VAR* model with parameters set as follows:

$$\begin{bmatrix} \tilde{y}_t \\ p_t \end{bmatrix} = K + B \begin{bmatrix} \tilde{y}_{t-1} \\ p_{t-1} \end{bmatrix} + U_t = \begin{bmatrix} 0.1 \\ 0.1 \end{bmatrix} + \begin{bmatrix} 0.5 & 0.1 \\ 0.2 & 0.85 \end{bmatrix} \begin{bmatrix} \tilde{y}_{t-1} \\ p_{t-1} \end{bmatrix} + \begin{bmatrix} w_t + 0.05 u_{2t} \\ u_{2t} \end{bmatrix} , \tag{29}$$

with t = 1, 2, … , 100. The inflationary and technological shocks u_{2t} and w_t were generated from independent normal standard distributions. In this model the sign of q_2 in (26) is always positive, so that only *REIT* > 0 could produce a positive real effect. In approximating p_t^e from (18), the first one-hundred terms of the infinite series were used. The computation of p_t^n can be done from (19) and (20), where the $g_{22}^{(i)}$ are the lower right elements of the matrices $\Gamma^{(i)}$ obtained from (9), (12) and (14) as $\Gamma^{(i)} = C^{(i)} A(1)\Gamma(1)$ with $\Gamma(1)$ defined by (13) and $C^{(i)} = B^i$ in the *VAR*(1) model. This leads to the approximation of p_t^n using the first one-hundred terms in (19).

The simulations were organised and conducted in the following way. Let h be the simulation horizon, so that t = 1, 2, …, h (the initial observations are then indexed from -99 to 0). In the experiments described here h = 18. There are K=6 main experiments (or groups of experiments), denoted by $\mathcal{J}(k)$, k = 1, 2,…,6. Let $REIT_t^{(i)}(k)$, i = 1,2,…,100, be the time-t value of *REIT* in experiment k run i.

In experiment $\mathcal{J}(1)$, for the different runs, the values of *REIT* are fixed for t = 1 that $REIT_1^{(1)}(1) = -1$, $REIT_1^{(2)}(1) = -0.99$, *etc*., until $REIT_1^{(100)}(1) = -0.01$. This is achieved by forcing U_0 as in (24), with p_t^e set at its unrestricted level, which is generated from (18) for t = 0, 1, …, 18. All subsequent (in time) values of *REIT* are set at zero, that is $REIT_t^{(i)}(1) = 0$ for t = 2, 3, …, h and i = 1,2, …, 100. In further experiments, $\mathcal{J}(k)$, k = 2, 3, …, K, values of

$REIT_t^{(i)}(k)$ are set at zero for $t = k+1, k+2, \ldots, h$ and all i. Here the first k subsequent values of $REIT_t^{(i)}(k)$ are set between –1 and –0.01, as in $\mathcal{J}(1)$,

$$REIT_1^{(i)} = REIT_2^{(i)} = \ldots = REIT_k^{(i)} = -1 + 0.01 \times (i-1),\ i = 1,2,\ \ldots,\ 100.$$

Hence the simulation aims at mimicking a situation with k periods of an identical negative *REIT* followed by $h - k$ periods of an 'improved' *REIT*, so that

$$REIT_{k+1}^{(i)}(k) = REIT_{k+2}^{(i)}(k) = \ldots = REIT_h^{(i)}(k) = 0 \quad .$$

It might be convenient to interpret the simulation as a scheme where in periods 1, 2, …, k the CB considers an anti-inflationary action which could be delayed over periods $k + 1$, $k + 2$, …, h, if the situation (in terms of minimising output losses) were to become more favourable. Clearly, for each period after first k, the situation is indeed better, since $REIT = 0$ implies smaller output losses than does $REIT < 0$.

Let us denote by $\tilde{y}_t^{(i)}(k)$ the output gain (measured in relation to full capacity) in period t under $REIT_t^{(i)}(k)$. An increase in i indicates an increase in $REIT_t^{(i)}(k)$. Consequently, with the increase of $REIT_t^{(i)}(k)$, there will be an increase in the corresponding sums of output, ie $\sum_{t=1}^{h} \tilde{y}_t^{(i)}(k) < \sum_{t=1}^{h} \tilde{y}_t^{(j)}(k)$, for $i < j$, and $\sum_{t=k+1}^{h} \tilde{y}_t^{(i)}(k)$ does not depend on i because $REIT_t^{(i)}(k) = 0$ for all i if $t > k$.

Hence, for each $d = 1, 2, \ldots, h - k$, there exists $v = v(d,k)$ such that

$$\sum_{t=1}^{h} \tilde{y}_t^{(n)}(k) \approx \sum_{t=k+d}^{h} \tilde{y}_t^{(n)}(k) \qquad ,$$

In another words, there is a case where it is possible to delay an action affecting the monetary target until *REIT* becomes better, without sacrificing accumulated output. The value of $REIT_1^{(v)}(k)$ for given k and d is here called the *marginal REIT* and can be interpreted as the value of *REIT* in period one for which, under expected *REIT*`s of zero for periods after k, it would pay (in terms of foregone output) to delay an inflation-targeting action by d periods.

Table 1 gives the averaged (over 1,000 replications) marginal *REIT* values obtained for model (29), for $\mathcal{J}(k)$ and $d = 1,2, \ldots, h$-1. For the sake of interpreting the figures given in this table consider, for instance, the marginal *REIT* value of –0.570 obtained in $\mathcal{J}(2)$, with $d = 5$. This means that if it is expected (in time 0) that, for two subsequent periods, *REIT* will be exceed –0.570 and, starting from the third period, *REIT* will stabilise at zero, the monetary authority could delay the inflation-reducing measure for up to five periods with no a danger of creating additional loses in output. Analogously, they might delay an expansionary policy decision for up to five periods, since it would not increase output. As expected, the marginal *REIT* value diminishes (albeit not monotonically, due to the randomness of the experiment) with the lengthening of the delay. That means, not surprisingly, that with a very low *REIT* in the initial period, an anti-inflationary measure can be delayed even for a relatively long time. If negative *REITs* continue beyond the first period, and an inflation targeting action is delayed by up to *d* periods, the marginal *REIT*` will have a tendency to increase, giving rise to the delaying of such action even with a relatively moderate $REIT_1$.

5. Has *REIT* been taken seriously? Active policy in Poland and monitoring in Russia

During the period 2001 – 2003, the current *REITs* and forecasts of them were computed for Poland on a monthly basis and delivered to the Monetary Policy Council of Poland (*MPC*). The *MPC* of Poland was established in 1998 at the National Bank of Poland (central bank). The *MPC* is a fully independent body, appointed partly by Parliament and partly by the President of Poland, with the primary job of pursuing anti-inflationary policy by setting the base interest rate. Inflation is to be maintained within pre-specified bounds. At the monthly meetings of the *MPC* decisions on whether to change the current base rate (and, if so, by how much) are made by voting.

Table 1: Simulated marginal *REIT's*, averaged over 1,000 replications

d: (delay in policy action)	Number of experiment: (k: No. of initial periods with non-zero *REIT*)					
	$\mathcal{J}(1)$	$\mathcal{J}(2)$	$\mathcal{J}(3)$	$\mathcal{J}(4)$	$\mathcal{J}(5)$	$\mathcal{J}(6)$
1	-0.475	N/A	N/A	N/A	N/A	N/A
2	-0.547	-0.486	N/A	N/A	N/A	N/A
3	-0.576	-0.529	-0.486	N/A	N/A	N/A
4	-0.591	-0.560	-0.520	-0.462	N/A	N/A
5	-0.619	-0.570	-0.555	-0.504	-0.473	N/A
6	-0.644	-0.587	-0.567	-0.540	-0.500	-0.452
7	-0.647	-0.611	-0.581	-0.569	-0.517	-0.477
8	-0.667	-0.622	-0.599	-0.586	-0.533	-0.502
9	-0.667	-0.626	-0.611	-0.596	-0.546	-0.516

10	-0.674	-0.644	-0.619	-0.605	-0.558	-0.528
11	-0.681	-0.647	-0.625	-0.610	-0.570	-0.541
12	-0.687	-0.651	-0.632	-0.618	-0.583	-0.552
13	-0.673	-0.654	-0.638	-0.622	-0.588	-0.563
14	-0.674	-0.659	-0.644	-0.623	-0.592	-0.571
15	-0.674	-0.664	-0.650	-0.620	-0.605	-0.581
16	-0.688	-0.660	-0.653	-0.625	-0.615	-0.584
17	-0.696	-0.657	-0.657	-0.623	-0.621	-0.595

REITs were also computed for Russia for 2003 – 2004, in this case more as an academic exercise than as a policy tool.[1] The Russian monetary policy targets have been defined in a much less transparent way than the Polish targets. It is widely accepted that the Central Bank of Russia is primarily interested in maintaining a 'dirty float' of the domestic currency, by undertaking occasional interventions. During the period under investigation, high world oil prices (the primary source of government revenue) caused, under significant domestic inflationary pressure, real appreciation of the rouble. Such a situation calls for monetary instruments other than interest rate-oriented inflation targeting. In the attempted sterilisation, more weight must be placed on fiscal rather than monetary measures, the traditional monetary interventions being limited to controlling credit to commercial banks and setting high bank reserve requirements (see e.g. Arhend, 2004, Esanov et al 2004). Consequently, the Russian CB did not show an interest in the *REIT,* which (unlike in Poland), was not reported to the Russian monetary authorities.

The concept of *REIT*, described in earlier sections, can be of a practical relevance if two crucial components of headline inflation, p_t^e and p_t^n, can be evaluated *ex-ante*. For empirical work, however, the computing *REIT* with the required accuracy, is more complicated than a mere extrapolation of a vector autoregressive model such as (5). Although such extrapolation would provide mutually consistent estimates of the inflationary components and seems to be adequate for computing p_t^n, its application in the evaluation of p_t^e is more questionable. The limited economic information included in a two-dimensional *VAR* model is not likely to suffice for a precise forecast of inflation, and the resulting estimate might be very different from that used by economic agents.

[1] This research was carried out under the auspices of the Think Tank Partnership Project, *Timing of Monetary Policy and Inflation Monitoring in Poland and Russia*.

In light of the above considerations, out use of *REIT* methodology as an empirical tool for Poland and Russia included the calculation of expected inflation in a more complex way. Using monthly time series data on individual consumers' prices and their aggregates (such as the consumers' price index), several of predictive inflation measures were computed using a number of the limited influence estimators (percentile means, trimmed means with various trims, exclusion means), smoothed estimators (Kalman filter, Hodrick-Prescott, *ARMA*, Holt) and more recently developed methods (e.g. Arrazola and de Hevia, 2002). These methods have been applied separately in computing forecasts for 1-16 months. For each forecasting horizon, the three best methods (in terms of minimal *ex-post* root mean square error of forecast) were selected and combined, using regression weights, into an optimal mechanical forecast. These mechanical forecasts were in turn combined with external experts' evaluations to produce the values of p^e_{t+i}, $i = 1,2, \ldots, 16$ (detailed description of this inflation forecasting methodology is given in Charemza *et al*., 2006). The corresponding values of p^n_{t+i} were obtained from the estimated *VAR* model, where monthly data for the index of industrial production were used to approximate $\tilde{y}_t$. Similar data were used with the *VAR* models of the Polish and Russian inflation and output gap. The main difference here is the use of annual indices for Poland (computed monthly, ie in relation to the corresponding month of the previous year) and monthly indices (in relation to previous month) for Russia. This stems from the different ways of reporting inflation in Poland and Russia.

Figures 3a and 3b give monthly *CPI* inflation figures (annual for Poland and monthly for Russia) and *REIT*. For Poland, *REIT* was originally submitted to the *MPC* as several separate forecasts for up to 16 months. These forecasts are represented here in an aggregate way, as weighted averages where the weights are approximated from impulse responses of output (for Poland, see Lyziak 2002). For Russia the way of computing the aggregate *REIT* is similar, with weights identical to that used for Poland. Inflation and *REIT* are plotted on different axis, and months for which the *MPC* decided to change the base interest rate are indicated by arrows. The length and direction of each arrow correspond to the direction and magnitude of each change.

Figure 3a: *REIT* and inflation, Poland

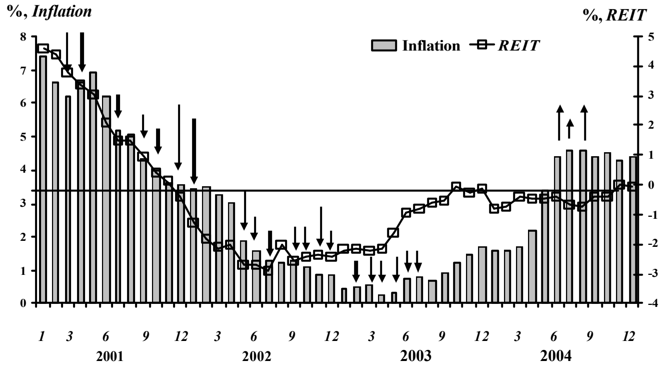
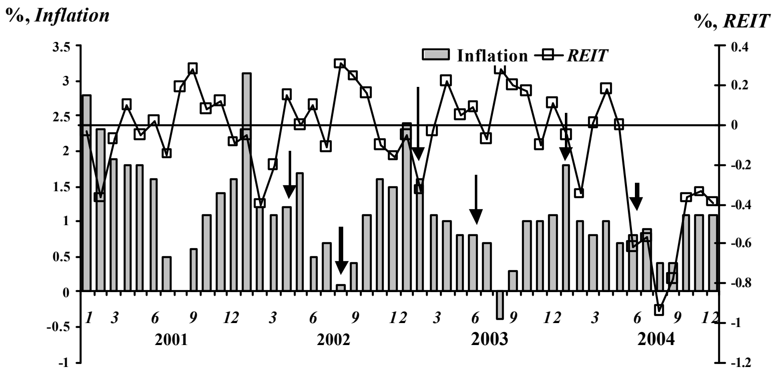


Figure 3b: *REIT* and inflation, Russia



In Poland, there was initially a period of 11 months, from January to November 2001, with positive (averaged) *REIT*. After that, *REIT* becomes negative. Within the period of positive *REITs*, the *MPC* reduced the interest rate by a total of 6 percentage points (from 21.5% to 15.5%), which gives an average reduction per month of 0.55%. During the subsequent period of negative *REITs*, the average monthly reduction in the interest rate was smaller (0.42%). In both periods the average time between interest rate changes was similar: 2.2 months in the period of positive *REITs* and 1.92 months afterwards. Obviously, it is difficult to say *ex-post* to what extent information on *REIT* supplied to the *MPC* affected their decision, but Figure 3, together with the simple statistics given above, suggest at least symptomatic relation between *REIT* and monetary decisions. It can therefore be conjectured that the period of more active expansionary monetary policy (more drastic changes in the interest rate at similar intervals) corresponds to that of the positive *REITs.*

For Russia, the estimates of *REIT* are more volatile, with greater frequency of changes in the regimes and less persistence. This is presumably due to the fact that the data here represent monthly rather than annual inflation and that the observed and neutral inflation, p_t^n, exhibits different seasonal patterns. It can be noticed that, unlike for Poland, positive *REIT*'s generally correspond to periods of low inflation. This seems to be intuitively plausible: during periods of low inflation, under the 'dirty' float, there is little danger of real appreciation of the rouble. Hence, a fall in interest rate is likely to increase competitiveness and, after a delay, positively affect output. Initially, during the period investigated, decisions regarding the CB refinancing rate, in April and August 2002 and in June 2003, coincided with positive *REITs*. Further decisions, however, in January and June 2004, were undertaken in an unfavourable climate, when *REITs* were negative.

Figures 4a and 4b give monthly *CPI* inflation and quarterly *GDP* growth, for Poland and Russia respectively. For Russia, the background shape represents growth of the non-industrial components of *GDP* (agriculture, construction, transport, retail sales, paid services to households), denoted as *NI-GDP*[2]. The reason for showing the growth of non-industrial components of *GDP* separately is to evaluate possible impacts of monetary policy on these sectors of the Russian economy, which might be less directly dependent on oil industry revenues and hence more sensitive to monetary policy. For both countries, the pattern is similar: significant *GDP* growth in 2003 and 2004, accompanied by falling inflation. The *GDP* growth was due to productivity growth in Poland and strong world demand for Russian petroleum, as unemployment was not declining in either country. Since it is generally acknowledged that the average response of output to a change of the base interest rate is in the range of 8 to 11 quarters (see Kokoszczynski *et al* 2002), one might conjecture that active monetary policy undertaken in 2001 in Poland paid off by a substantial delayed increase in *GDP*. To what extent information on *REIT* given to the Monetary Policy Council contributed to this success, of course unknown, but the positive associations are striking. For Russia, it seems to be more of a coincidence that expansionary measures were effected in 2002 during the periods of positive *REIT*. Nevertheless, as the substantial economic growth in 2004 suggests, the timing of these actions was good. The fact that further actions, in 2004, were undertaken while *REIT* was negative gives a warning signal regarding further increases in Russian *GDP* in the first half of 2006.

[2] Own computations using official *GosKomStat* data available at the Institute for Complex Strategic Studies site http://www.icss.ac.ru

Figure 4a: Inflation and *GDP* in Poland, 2001-2004

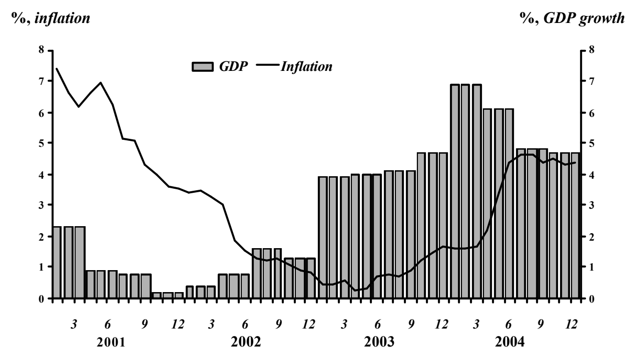
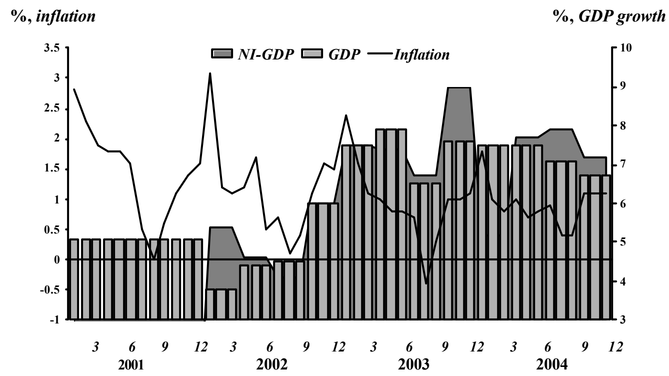


Figure 4b: Inflation and *GDP* in Russia, 2001-2004



6. Conclusions

The *REIT* concept, which is relatively simple and computationally straightforward, can be applied in many situations that require an evaluation of possible real effects of inflationary policy or a projection. For instance, it can be used to investigate the effects of possible external supply shocks on inflation. In this case the model applied would be analogous to that described in this paper, with some modification of the interrelation mechanism for the real and monetary effects. Further on, the *REIT* concept might help to reconcile the problem of inflation-output sacrifice and the problem of inflation control with minimal output loss. Some difficulties would, however arise concerning the construction of an empirically sound gauge of output-neutral inflation. The methodology applied in this paper, which involves computing the gauge from a two-variable *VAR*, is fairly simple and can be further improved. An investigation based on a larger *VAR*, possibly involving monetary policy instruments and monetary aggregates, might generate more specific results. It is also likely that more accurate estimates can be obtained from a more disaggregated price system, presumably by evaluating large disaggregated panel data systems for individual prices and outputs.

References

Ahrend, R. (2002), 'Accounting for Russia's post-crisis growth', Economics Department Working Paper No. 404, OECD.

Arrazola, M. and J. de Hevia (2002), 'Alternative measure of core inflation', *Economics Letters* **75**, pp. 69-73.

Ascari, G. (2000), 'Optimising agents, staggered wages and persistence in the real effects of money shocks', *Economic Journal* **110**, pp. 664-686.

Calvo, G. A., (1983), 'Staggered prices in a utility-maximizing framework', *Journal of Monetary Economics* **12**, pp. 383-398.

Charemza, W.W., P. Jelonek and S. Makarova (2006), 'Probabilistic inflation forecasting for Poland'. National Bank of Poland, in preparation.

Chari, V.V., P.J. Kehoe and E.R. McGrattan (2000), 'Sticky price model of the business cycle: can the contract multiplier solve the persistence problem?`, *Econometrica* **68**, pp. 1151-1179.

Coenen, G. and V. Wieland (2005), 'A small estimated Euro area model with rational expectations and nominal rigidities', *European Economic Review* **49**, pp. 1081-1104.

Dewachter, H. and H. Lusting (1997), 'A cross-country comparison of CPI as a measure of inflation', CES Discussion Paper Series, No 97 06, Department of Economics, Katholieke Universiteit Leuven.

Dixon, H. and E. Kara (2005), 'Persistence and nominal inertia in a generalised Taylor economy: how longer contracts dominate shorter contracts', University of York, mimeo.

Dodge, D. (2001), 'Monetary policy and inflation targets in Canada', speech of the Governor of the Bank of Canada before the Standing senate Committee on Banking, Trade and Commerce, Ottawa.

Eckstein, O. (1981), *Core inflation*, Prentice-Hall, Englewood.

Esanov, A. C. Merkl and L.V de Souza (2004), 'Monetary policy rules for Russia', BOFIT Discussion Paper No 11, BOFIT, Helsinki.

Fourgère, D., H. Le Bihan and P. Sevestre (2004), 'Calvo, Taylor and the estimated hazard function for price changes', presented at the conference *Inflation Persistence in Euro Area*, European Central Bank, Frankfurt am Mein.

Gartner, C. and G.D. Wehinger (1998), 'Core inflation in selected European Union countries', Working Paper No. 33, Oesterreichische Nationalbank, Vienna.

Hahn, E. (2002), 'Core inflation in the Euro area: evidence from the structural VAR approach', CFS Working Paper No. 2001/09, Johann Wolfgang Goethe Universitat.

Issing, O. (2001), 'The challenge for monetary policy: a European perspective', speech at the German-British Forum, London.

Issing, O. (2002), 'The role of monetary policy in managing economic risks', speech at the 18th Annual Conference of the National Association for Business Economics, Washington.

Kokoszczynski, R., T. Lyziak, M. Pawlowska, J. Przystupa, and E. Wróbel (2002), `Mechanizm transmisji polityki pienieznej', Materialy i Studia No. 151, National Bank of Poland, Warsaw.

Lyziak, T. (2002), `Monetary transmission mechanism in Poland: the strength and delays', Materialy i Studia No. 26, National Bank of Poland, Warsaw.

Mankiw, N.G. (2001), 'U.S. monetary policy during 1990's', NBER Working Paper No. w8471.

Mankiw, N.G. and R. Reis, (2003), 'What measure of inflation should a central bank target?', *Journal of the European Economic Association* **1**, pp. 1058-1086.

Quah, D., S. Vahey (1995), 'Measuring core inflation', *Economic Journal* **105**, pp. 1130-1144.

Taylor, J.B. (1979), 'Staggered wage setting in a macro model', *American Economic Review* **69,** *Papers and Proceedings*, pp. 108-113.

Taylor, J. B. (1980), 'Aggregate dynamics and staggered contracts', *Journal of Political Economy* **88**, pp. -24.

Uhlig, H. (2005), 'What are the effects of monetary policy on output? Results from an agnostic identification procedure', *Journal of Monetary Economics* **52**, pp. 381-419.

van Gaasbeck, K.A., (2001), 'Is time everything? Investigating U.S. monetary policy with interest rate runs', University of California, Davis, mimeo.

Wehinger, G.D. (2000), 'Causes of inflation in Europe, the United States and Japan: some lessons for maintaining price stability in the EMU from a structural *VAR*', *Empirica* 27, pp. 83-107.

Whelan, K. (2004), 'Staggered price contracts and inflation persistence: some general results', Central Bank and Financial Services Authority of Ireland, Dublin (mimeo).