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Monetary disturbances matter for business fluctuations in the G-7 ☆

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Abstract

This paper examines the importance of monetary disturbances for cyclical fluctuations in real activity and inflation. It employs a novel identification approach which uses the sign of the cross-correlation function in response to shocks to assign a structural interpretation to orthogonal innovations. We find that identified monetary shocks have reasonable properties; that they significantly contribute to output and inflation cycles in all G-7 countries; that they contain an important policy component, and that their impact is time varying.

JEL classification: C68; E32; F11

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La pietá del vecchio padre, né il debito d'amore che doveva Penelope far lieta … ma misi me per l'alto mare aperto … In fin che'l mar fu sopra noi rinchiuso.
Dante Alighieri

1. Introduction

The high correlation between monetary and real aggregates over the business cycle has attracted the attention of macroeconomists for at least 40 years. Friedman and Schwartz (1960) were among the first to provide a causal interpretation of this relationship: they showed that the comovements of money with output were not due to the passive response of money to the developments in the economy, and argued that rates of change in money were good approximations to monetary policy disturbances. Since then generations of macroeconomists have tried to empirically refute Friedman and Schwartz's interpretation. In particular, the literature has documented that unforecastable movements in money produce responses in interest rates that are difficult to interpret—i.e. they generate the so-called liquidity puzzle (see Leeper and Gordon, 1992). To remedy these problems, Sims (1980) and Bernanke and Blinder (1992) suggested the use of short-term interest rate innovations as indicators of monetary policy disturbances. In this case it is the response of the price level to policy disturbances that is hard to justify (see Sims, 1992). As a consequence of these difficulties, the last 10 years witnessed a considerable effort to identify monetary policy disturbances using parsimoniously restricted time series models (see Gordon and Leeper, 1992; Christiano et al., 1996; Leeper et al. 1996; Bernanke and Mihov, 1998).

The methodology used in these exercises involves three steps: run unrestricted VAR models; identify monetary policy shocks by imposing exclusion restrictions, justified by economic theory, informational delays, and other informal constraints; and measure the contribution of identified monetary policy shocks to output fluctuations at different horizons. On this last issue, the consensus view is that the contribution of monetary policy to output fluctuations in the post-World War II era is modest (see e.g. Uhlig, 1999; Kim, 1999).

In this paper we assess the importance of monetary disturbances as sources of cyclical movements in economic activity using a novel two-step procedure. First, we extract orthogonal innovations from a reduced form model. These innovations have, in principle, no economic interpretation, but they have the property of being contemporaneously and serially uncorrelated. Second, we study their informational content. In this second step we are guided by aggregate macroeconomic theory: we employ the sign of the theoretical comovements of selected variables in response to an orthogonal innovation to assign a structural interpretation to a VAR disturbance.

Our identification approach has a number of advantages over competing ones, and complements the methods recently proposed by Faust (1998) and Uhlig (1999). First, our procedure clearly separates the statistical problem of orthogonalizing the covariance matrix of reduced form shocks from issues concerning the identification of structural disturbances. Second, unlike structural VAR approaches, it achieves identification avoiding the imposition of zero constraints on impact responses, potentially inconsistent with the implications of a large class of general equilibrium monetary models (see Canova and Pina, 1999), or on the long-run response of certain variables to shocks, for which distortions due to measurement errors and small sample biases may be substantial (see e.g. Faust and Leeper, 1997). Third, all constraints employed are explicitly stated—so no circularity between identification and inference arises—and sensitivity analysis on the space of orthogonal decompositions consistent with the identifying restrictions is carried out systematically.

One important aspect of our exercise, which distinguishes it from the existing literature, is the international focus of the comparison (one exception is Kim, 1999). We are interested in knowing not only whether monetary disturbances are important in driving domestic cycles, but also they produce a similar pattern of responses in real and nominal variables in the G-7.

Four results emerge from our analysis. First, our approach identifies monetary disturbances in all seven countries. In general, these shocks can be classified into three broad categories: those generating liquidity effects; those generating temporary (expected) inflation effects and those linked to turbulence in international financial markets. We show that the time path of the monetary shocks we generate is interpretable, and that in the US, these shocks are significantly related to Federal Funds Future rates innovations (see Rudebush, 1998) and imply reasonable policy reaction functions. Second, shocks in each category produce responses in macroeconomic variables which are similar across countries. Third, monetary disturbances explain large portions of output and inflation fluctuations. For example, their combined explanatory power for output variability in Germany, Canada, UK and Italy exceeds 22% and for inflation variability in the US, UK, Japan and Italy exceeds 54%. Fourth, monetary disturbances are quickly incorporated into the slope of the term structure, thereby supporting the conjecture that they have an important policy component.

Our qualitative conclusions are broadly robust to sample splitting with one qualification. The number of monetary innovations that we uncover and their predictive power for the variability of output and inflation changes somewhat across subsamples. Results are also robust to the use of alternative estimation techniques.

The finding that monetary disturbances explain a large percentage of output variations in many countries is somewhat surprising, and appear at odds with some recently held views about sources of output fluctuations (exceptions are Roberts, 1993; Faust, 1998). For the US, which has been the focus of the majority of the analyses, our evidence diverges from the assessments of Leeper et al. (1996) or Uhlig (1999) in two important ways. The combined explanatory power of the monetary disturbances we identify is significant. Furthermore, the shock which is more closely related to those considered by these authors, explains in the median 38% of output variability but this percentage is dramatically increased in the post-1982 sample. For the other G-7 countries, the monetary shocks we have identified account for a percentage of output variance which is always 2–3 times larger than that found by Kim (1999).

The reminder of the paper is organized as follows. The next section presents the reduced form model and the issues connected with its specification. Section 3 discusses the basic intuition behind our identification procedure. Section 4 presents the results of our investigation. Section 5 analyzes the responses of the slope of the term structure to identified monetary shocks. Section 6 concludes.

2. The specification of the statistical model

Our reduced form model is an unrestricted VAR. We use an unrestricted VAR since it is a good approximation to the DGP of any vector of time series, as long as enough lags are included (see e.g. Canova, 1995). We use two alternative setups: single country VAR models including a measure of real activity (IP), of inflation (INF), of the slope of the term structure of the nominal interest rates (TERM) and of real balances (M/P); and a pooled VAR with country-specific fixed effect containing the same four variables for all countries. The sample we use covers monthly data from 1973:1 to 1995:7; industrial production, CPI and nominal interest rates are from the OECD database while monetary (M1) data are from IFS statistics. All series are seasonally adjusted.

Reduced form VAR models, which include real activity, inflation and measures of interest rates and money have been examined by many authors (e.g. Sims, 1980; Farmer, 1997). Here we maintain the same structure except that we employ a measure of the slope of the term structure in place of a short-term interest rate. We do this because recent results by Stock and Watson (1989), Estrella and Hardouvelis (1991), Bernanke and Blinder (1992) and Plosser and Rouwenhorst (1994) demonstrated the superior predictive power of the slope of term structure for real activity and inflation relative to a single measure of short-term interest rates in many countries. Also, the slope of the term structure has information about nominal impulses that other variables, such as unemployment or real wages, may not have. Unlike part of the literature, we use real balances, as opposed to nominal ones, for two important reasons. First, the model we present in the next section has important implication for real balances. Second, the responses of real balances allow us to distinguish monetary from other types of real demand disturbances. We have experimented with specifications including either stock returns or both a short- and a long-term nominal rate separately. The results we present are insensitive to the addition of these variables to the VAR.

In order to interpret responses to shocks as short-term dynamics around a stationary (steady) state, the VAR must be stationary, possibly around a deterministic trend. Given the relative small size of our data set, tests for integration and cointegration are likely to have low power and this may affect economic inference at a second stage. We therefore prefer to be guided by economic theory in selecting relevant variables and use that subset of them which is likely to be stationary under standard assumptions. The model we present in Section 3 generates stationary paths for linearly detrended output, inflation, term structure and real balances. Visual inspection of the linearly detrended time series for the four variables in the seven countries shows that there is no compelling evidence of non-stationarities. For VAR models with these variables, the Schwarz criterion indicates that the dynamics for all countries are well described by a VAR(1), except for Japan, where a VAR(2) is used.

Because the VAR is a reduced form model, the contribution of different sources of structural disturbances to output and inflation cycles cannot be directly computed. To obtain structural shocks we proceed as follows. First, we construct innovations from the reduced form residual having the property of being serially and contemporaneously uncorrelated. Second, we use theory to tell us whether any of the components of the orthogonal innovation vector has a meaningful economic interpretation.

Formally, let the Wold MA representation of the system be

$$Y_t = \phi + B(\ell)u_t, \quad u_t \sim (0, \Sigma), \tag{1}$$

where Y_t is a 4×1 vector and $B(\ell)$ a matrix polynomial in the lag operator. All orthogonal decompositions of a Wold MA representation with contemporaneously uncorrelated shocks featuring unit variance–covariance matrix are of the form

$$Y_t = \phi + C(\ell)e_t, \quad e_t \sim (0, I), \tag{2}$$

where $C(\ell) = B(\ell)V$, $e_t = V^{-1}u_t$ and $\Sigma = VV'$. The multiplicity of these orthogonal decompositions comes from the fact that for any orthonormal matrix $Q, QQ' = I$, $\Sigma = \hat{V}\hat{V}' = VQQ'V'$ is an admissible decomposition of Σ. One example of an orthogonal decomposition (which will not be used in this paper) is the Choleski factor of Σ, where V is lower triangular. In that case, it is well known that alternative ordering of the variables of the system (i.e. different orthogonal representations of Σ) may produce different structural systems. Another example of an orthogonal representation is the eigenvalue–eigenvector decomposition $\Sigma = PDP' = VV'$ where P is a matrix of eigenvectors, D is a diagonal matrix with eigenvalues on the main diagonal and $V = PD^{1/2}$. Under the assumption of orthogonal shocks, the impulse response of each variable to any shock is given by the coefficients of the vector of lag polynomials $C(\ell)\alpha$, where α satisfies $\alpha'\alpha = 1$.

As shown in the next section, economic theory provides important information on the signs of the pairwise dynamic cross-correlations of certain variables in response to structural shocks. The dynamic cross-correlation function of Y_{it} and Y_{jt+r}, $r = 0, \pm 1, \pm 2, \ldots$, is

$$\rho_{ij}(r) \equiv \mathrm{Corr}(Y_{it}, Y_{j,t+r}) = \frac{\mathrm{E}[C^i(\ell)e_t C^j(\ell)e_{t+r}]}{\sqrt{\mathrm{E}[C^i(\ell)e_t]^2 \mathrm{E}[C^j(\ell)e_{t+r}]^2}}, \tag{3}$$

where E indicates unconditional expectations and C^h the h row of $C(\ell)$. Hence, the pairwise dynamic cross-correlation conditional on the particular shock defined by α is

$$\rho_{ij|\alpha}(r) \equiv \mathrm{Corr}(Y_{it}, Y_{j,t+r}|\alpha) = \frac{(C^i(\ell)\alpha)(C^j(\ell + r)\alpha)}{\sqrt{[(C^i(\ell)\alpha)]^2[(C^j(\ell + r)\alpha)]^2}} \tag{4}$$

whose sign only depends on the sign of $(C(\ell)^i\alpha)(C^j(\ell+r)\alpha)$, the cross-product of the impulse responses of variables i,j at lag r to the shock. Hence, given an orthogonal representation, it is easy to check whether a shock produces the sign of the cross-correlation function required by theory.

In this paper, we want to explore the space of orthogonal decompositions to see whether for some α and certain variables i,j, $\rho_{ij|\alpha}(r)$ conforms with the predictions of economic theory. Because the space of V is uncountably large when one considers non-recursive models, two questions naturally arise. First, how to systematically search over the space of orthogonal decompositions for shocks which conform to theory. In the appendix we detail an algorithm, based on results provided by Press (1997), which we found useful for that purpose. Second, how to choose among various decompositions which recover *some* interpretable disturbance. Here we follow three general principles. First, we restrict attention to those decompositions that maximize the number of shocks exhibiting conditional correlations consistent with theory. If there is no decomposition for which all four shocks are identifiable, we concentrate on those for which only three shocks are identifiable, and so on. Second, if there is more than one decomposition that produces the same maximum number of identifiable shocks, we sequentially eliminate candidates making the sign requirements more stringent. Thus, for example, suppose that when one considers only sign restrictions at $r=0$ and obtains three candidate decompositions which identify all four shocks. Then, among these three candidates, we choose the one that satisfies the sign restrictions also at $r=\pm 1, \pm 2$, etc. Third, if this is still not enough to uniquely select a decomposition, we enlarge the vector of conditional correlations whose sign need to be matched, adding the pairwise correlation between the variables of the system and an additional one for which theory has information. For example, in the case of monetary shocks, one may use money and prices cross-correlation function to identify them. If, after having used, say, r up to 12, there is still more than one decomposition available, one may also want to look at the cross-correlation of money and interest rates to eliminate decompositions which, e.g. do not generate liquidity effects.

Although we rely on the *sign* of the theoretical cross-correlation function, one may be, at times, interested in using the *magnitude* of these correlations to identify shocks. In this case one could select the orthogonal decomposition that minimize the distance between a vector of cross-correlation functions of the model and of the data. While, as shown in the next section, sign restrictions are shared by a large class of models with different microeconomic foundations and magnitude restrictions are typically model dependent. Therefore, by taking this alternative route to identification, one has to take a firm stand on the reference model producing the correlations, and the search over orthogonal decompositions may lead to its rejection if the data are inconsistent with the magnitude restrictions imposed. Hence, using sign restrictions is equivalent to using only a minimal set of widely agreed (non-parametric) restrictions to identify shocks.

Once we have explored the space of identifications and selected a candidate, we measure their contribution to output and inflation cycles using the variance decomposition. The variance of Y_{it} allocated to α at horizon τ is

$$z^{\tau}(i,\alpha) = \frac{\sum_{s=0}^{\tau-1}(C_s^i\alpha)^2}{\sigma_{i\tau}^2}, \tag{5}$$

where $\sigma_{i\tau}^2$ is the forecast error variance of Y_i at horizon τ. We compute confidence bands for the $z^{\tau}(i,\alpha)$ numerically drawing 1000 Monte Carlo replications, ordering them and extracting the 68% band (from the 16th to the 84th percentile) as suggested by Sims and Zha (1999).

There are several differences between our approach and the one commonly used in structural VARs (SVAR). In SVAR one typically imposes "economic" or "sluggish" restrictions on impact coefficients or on the long-run multipliers of shocks and interprets the resulting long-run (short-run) dynamics. The imposition of economically or informationally motivated zero restrictions achieves two goals at once: disentangle the reduced form shocks and make them structurally interpretable. The two-step approach we propose separates the statistical problem of producing orthogonal shocks from the economic one of interpreting them (much in the spirit of Cooley and LeRoy, 1985). Furthermore, instead of identifying shocks by imposing zero restrictions on the contemporaneous impact of shocks, restrictions which may be inconsistent with a large class of general equilibrium models (see Canova and Pina, 1999), or on their long-run effects, for which small sample biases may be substantial (see Faust and Leeper, 1997), we use sign restrictions on a vector of conditional cross-correlations to assign a structural interpretation to orthogonal disturbances.

Several authors, including Leeper et al. (1996), Faust (1998) and Uhlig (1999), have pointed out that identification of a set of shocks is typically achieved using both a set of formal zero restrictions (e.g. output is not contemporaneously responding to money supply shocks) and of informal prior constraints (e.g. prices should not decline in response to a expansionary money supply shock), that only formal constraints are used to compute intervals around point estimates of the statistics of interest, and that the way informal constraints are used may render inference circular. In our approach, all constraints used are cast in the form of formal sign restrictions on the pairwise cross-correlation functions and all are used to compute confidence intervals.

Our approach shares similarities with the ones recently proposed by Faust (1998) and Uhlig (1999). Faust provides a way to examine the validity of a statement for all identification schemes which produce "reasonable" impulse responses, and constructs counterexamples if they exist. Uhlig evaluates the correctness of a statement by computing either the variance share of a particular shock for all identifications which minimize a penalty function, or the set of responses which satisfy some a priori sign restrictions. With both methods, decompositions which produce impulse responses having signs different from those assumed to be "reasonable" are penalized, explicitly with arbitrary weights, or implicitly by being discarded. We share with both authors the desire of systematically examining a variety of identification schemes and of making all restrictions formal. We differ in the function used to identify shocks (cross-correlations vs. impulse responses or variance decompositions), in the criteria used to select among orthogonal decompositions satisfying the restrictions, and in the fact that our approach allows to sequentially impose more stringent restrictions to eliminate candidate orthogonalizations.

3. The theoretical restrictions

In this section we highlight the type of sign restrictions that a particular model produces in response to structural shocks. We then argue that these restrictions are generic, in the sense that in a number of models with different microfoundations the joint dynamics of output, inflation and real balances in response to shocks have similar signs, at least contemporaneously. Therefore they can be used regardless of the confidence a researcher has in the specific model we describe here.

The economy we consider is a version of the limited participation model used by Christiano et al. (1997). The economy is populated by five types of agents: households, firms, financial intermediaries, a fiscal and a monetary authority. The households are all identical, own the firms and the financial intermediaries, and maximize the expected discounted sum of instantaneous utilities derived from consuming a homogenous good and from enjoying leisure. The timing of the decision is the following: at the beginning of period t, households carry over M_{t-1} units of money and bonds B^j_{t-1} of maturities $j = 2, \ldots, n$ and choose cash for purchases, Q_t, before observing the shocks. Then all shocks are realized, the households take their remaining financial assets ($M_{t-1} - Q_t$ and the holding of bonds) to the banks and the monetary injection, X_t, is fed into the banking system. At this point, households rebalance their portfolio of assets by purchasing bonds of maturities $j = 1, \ldots, n$ at price b^j_t from the bank, and choose the number of hours worked. The time endowment is normalized to one; capital is in fixed supply and normalized to one. At the end of production time, households collect wage payments, $W_t N_t$, and use them with the cash set aside, Q_t, to purchase goods. After goods are purchased, households receive capital income—dividends from owing the firms (D_t), the financial intermediaries (F_t) and returns from maturing bonds (B^1_t) — and pay taxes (T_t). The program solved is

$$\mathrm{Max}_{\{C_t, Q_t, N_t, M_t, B^j_t\}} \mathrm{E}_0 \sum_{t=0}^{\infty} \beta^t [(\ln(C_t)) + \gamma \ln(1 - N_t)] \tag{6}$$

subject to

$$P_t C_t \leqslant Q_t + W_t N_t, \tag{7}$$

$$\sum_{j=1}^{n} b^j_t B^j_t \leqslant M_{t-1} - Q_t + \sum_{j=2}^{n} b^{j-1}_t B^j_{t-1}, \tag{8}$$

$$M_t \leqslant F_t + D_t + B^1_t + Q_t + W_t N_t - P_t(C_t + T_t), \tag{9}$$

where M_{-1}, B^j_{-1} are given and E_0 is the expectation conditional on information at time 0.

Firms are identical and face a decreasing returns to scale technology perturbed by an exogenous technology shock v_t. Each firm maximizes profits subject to the technology and to a cash-in-advance constraint, since wages are paid before the firm collects revenues from the sales of the product. Profits at each t are measured by the difference between the receipts from selling the good, Y_t, at price P_t, and the wage costs $(1+R_t)W_tN_t$. The problem solved by the firm is

$$\mathrm{Max}_{\{N_t\}}\ P_tY_t - (1+R_t)W_tN_t \tag{10}$$

subject to

$$W_tN_t \leqslant M_{t-1} - Q_t + X_t, \tag{11}$$

$$Y_t \leqslant v_tN_t^{\alpha}. \tag{12}$$

We assume $\ln(v_t) = a\ln(t) + (1-\rho)\ln(v) + \rho\ln(v_{t-1}) + \vartheta_t$, with $\vartheta_t \sim \mathrm{iid}(0, \sigma^2_{\vartheta})$, $|\rho|<1$, $\alpha \in [0,1]$.

Financial intermediaries collect deposit from the households, $M_{t-1} - Q_t$, trade bonds with them and receive the injection X_t from the monetary authority. These funds are supplied in the loan market at the gross interest rate of $(1+R_t)$. Market clearing in the loan market requires that (11) is satisfied with equality. After repaying all maturing bonds, profits distributed to the households are equal to

$$F_t = (1+R_t)W_tN_t - B^1_t. \tag{13}$$

The fiscal authority finances consumption expenditure G_t by lump sum taxes T_t. We assume $\ln(G_t) = (1-\theta)\ln(G) + \theta\ln(G_{t-1}) + \varphi_t$, with $\varphi_t \sim \mathrm{iid}(0, \sigma^2_{\varphi})$, $|\theta|<1$.

The monetary authority issues cash at no cost and transfers it to the bank. We assume a simple policy rule, which has both an exogenous and an endogenous component, of the form $R_t = \pi_t^{0.5}Y_t^{0.1}\varepsilon_t$, where $\pi_t = P_t/P_{t-1}$ and ε_t, the policy shocks, satisfy $\ln(\varepsilon_t) = (1-\phi)\ln(\varepsilon) + \phi\ln(\varepsilon_{t-1}) + \omega_t$, with $\omega_t \sim \mathrm{iid}(0, \sigma^2_{\omega})$, $|\phi|<1$. Monetary injections are defined as $X_t = M_t - M_{t-1}$.[1]

In equilibrium all markets clear and the interest rate for one-period bonds is $1 + R_t = 1/b^1_t$, that is, the nominal return earned by the household on one-period bonds equals the return earned by the intermediaries on their one-period loans. The interest rates for bonds of longer maturities can be obtained using the standard pricing formula $1 + R^j_t = -(1/j)\log(\mathrm{E}_t\beta\lambda_{t+j}/\lambda_t)$ where λ_t is the Lagrangean multiplier on (8).

Since an analytic solution to the model cannot be computed, we log-linearize the equilibrium conditions around the steady state. We construct the slope of the term structure by taking the difference between a long-term rate and a short one ($\mathrm{SL}_t = \lim_{j\to\infty}\hat{R}^j_t - \hat{R}_t$) where a hat indicates percentage deviations from the steady state. To generate time series out of the model, we choose the time unit to be a quarter. We let $\bar{N} = 0.30$, $\alpha = 0.65$, $\bar{\Pi} = 1.0$, $\beta = 0.99$, $\bar{c}/\bar{y} = 0.8$ where $\bar{c}/\bar{y}$ is the share of

[1] Although we have assumed a specific form of the policy reaction function, none of the results we present depend on the exact form of this function (see e.g. Canova and Pina, 1999).

consumption in output, $\bar{N}$ is hours worked and $\bar{\Pi}$ is the gross inflation in the steady states, α is the exponent of labor in the production function and β is the discount factor. These parameters imply that in steady state the gross real interest rate is 1.01, output is 0.46, deposits are 0.29, real balances 0.37, the real wage 0.88, the share of leisure in utility is 0.65, and $\gamma = 1.86$, which are in line with those used in the literature. Finally, we parametrize the stochastic processes for the three shocks to all have the same persistence (0.95) and the same coefficient of variation (1/0.71).

Fig. 1 reports the pairwise cross-correlation of output, inflation and real balances in response to the three structural shocks. A technology disturbance generates S-shaped correlations between output and inflation and inflation and real balances and in both cases the contemporaneous cross-correlation is negative. On the other hand, the cross-correlation between real balances and output is positive everywhere. Government expenditure shocks produce an inverted S-shape correlation between inflation and output and the contemporaneous cross-correlation is positive. The cross-correlation between inflation and real balances has an S shape with a negative contemporaneous cross-correlation while the correlation between real balances and

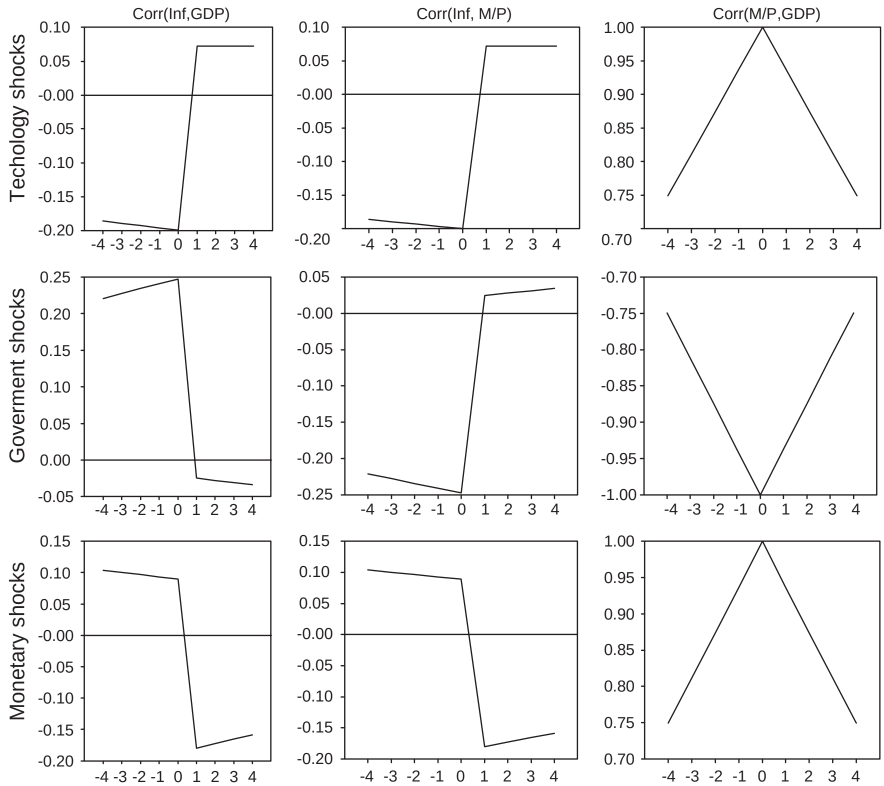

Fig. 1. Cross correlation in a limited participation model.

output is negative in the range. Finally, monetary disturbances produce positive contemporaneous cross-correlations for all pairs of variables.

The interpretation of these patterns is very simple. A surprise increase in $\hat{v}_t$ increases output and consumption on impact since government consumption is constant at its steady-state level. Because of the cash in advance constraint, an increase in consumption requires an increase in real balances to finance expenditure. With the given policy rule, short-term nominal rates must increase (the slope of the term structure declines) to make agents hold exactly the right amount of money. Since agents are richer, the wealth effect of the shock makes hours decline and leisure increases temporarily. Because labor demand by firms has increased, the real wage is higher after the shocks, making the wealth effect even stronger. In other words, as agents become more productive, they devote more time to leisure and less to production. Also, because the nominal rate increases and the inflation rate declines, real balances and the ex post real rate increase substantially after the shock.

A unitary surprise increase in $\hat{\varphi}_t$ makes private consumption decline and, because of a wealth effect, labor supply and output increase. Since aggregate demand increases, prices go up on impact. Since consumption declines, money demand also declines and the short-term rate decreases (the slope of term structure increases) to induce agents to hold the existing stock of money. As a consequence, leisure declines to maintain the time constraint satisfied. Real balances and ex post real returns also decline, as the nominal rate decreases while inflation has increased on impact.

Finally, a unitary surprise increase in $\hat{\omega}_t$ decreases the cost of production for firms and this increases their labor demand. Hence both wages and hours increase, leading to an increase in output and consumption. As money increases are larger than output increases, there will be inflation. However, since the increase in inflation is smaller than the increase in money creation, real balances increase. Since the liquidity effect dominates the expected inflation effect, a positive monetary shock decreases nominal short-term rates at impact and rises the slope of the term structure.

In sum, the three types of (temporary) disturbances we consider produce joint comovements of output, inflation and real balances of different signs. One may be curious as to whether these restrictions are specific to the model, in which case the analysis of the next few sections is relevant only to the extent that the model is a credible description of the data, or whether they are shared by a large class of economies, in which case the analysis can be conducted without any reference to a specific member in this class and the characterization we provide more robust.

The contemporaneous sign restrictions that the model produces are very generic, in the sense that the class of models where innovations move output, inflation and real balances in the way we have described is relatively broad and includes economies with different microfoundations and frictions. For example, in Lucas (1972) misperception model, where agents cannot distinguish shocks to relative prices from shocks to the aggregate price level, demand (monetary) and supply (technology) disturbances produce comovements in output, inflation and real balances with the required sign characteristics. New-keynesian models with menu costs or sticky prices and monopolistic competition of the type examined by Mankiw

(1985) or Gali (1999) or indeterminacy models of the type described in Farmer (1999) generate a similar pattern of comovements in response to demand, monetary and technology disturbances, even though the quantitative features of inflation and output responses in the short run will be different from those produced by Lucas' model. Finally, also a static undergraduate textbook model, depicting downward sloping aggregate demand curve, an upward sloping short-run aggregate supply curve and a vertical long-run aggregate supply curve in the inflation–output plane (see e.g. Abel and Bernanke, 1995, p. 382) has the feature that technology, government and monetary shocks generate the required sign restrictions on the responses of output, inflation and real balances. Because of its static nature, this model has not much to say about the exact timing of these comovements. Common sense suggests that if prices are flexible, the majority of the adjustments should occur almost contemporaneously, in which case the pairwise contemporaneous cross-correlation of these three variables can be used to identify the informational content of shocks. On the other hand, if prices are sticky or there is sluggishness in output adjustments, propagation may take time so that leads and lags of the pairwise cross-correlation function contain the information needed to identify structural disturbances. Clearly, one can build examples where the responses of these three variables deviate from the characterization we have provided here. Sign restrictions can also be used as identification devices for this alternative class of models. By comparing the responses of interesting variables to shocks one can then discard one class of models and retain another one.

4. The results

While in the previous section we have described the restrictions on output, inflation and real balances implied by three shocks, here we focus the discussion entirely on monetary disturbances. Canova and De Nicoló (2000) use the same machinery to study the relative contribution of demand and supply shocks to business cycle fluctuations in the G-7.

Before describing the results in detail, it is worth discussing two features of the approach which may be puzzling to the reader. First, it may be the case that monetary shocks are not identifiable, that is, the sign restrictions we impose may be inconsistent with the data. This may occur if the statistical features of the shocks we attempt to identify are misspecified (e.g. we require the disturbances to be transitory but some shocks may have permanent characteristics) or if the set of variables we use does not have clear informational content (e.g. labor market variables may be capturing monetary shocks better than industrial production). Second, it may be the case that we identify more than one monetary shock. This does not typically occur in standard VARs because the number of variables used matches the number of structural shocks one wants to identify and conventional names are given to shocks which lack clear economic content. In large systems one could use other sign restrictions to disentangle the information content of multiple monetary shocks (for example, distinguish those generated in credit markets from those generated in federal funds markets, from discount window shocks, etc.). However, in small systems this is not possible and one has to use informal devices to achieve this scope. The exercises we conduct in this and the next section are designed to understand better the nature of these multiple shocks.

4.1. Identifying US Monetary disturbances

To illustrate how the identification procedure works, the type of disturbances it generates and the policy functions it produces, we examine the case of the US in detail. Figs. 2, 3 and 4 present, respectively, the estimated cross-correlation function for inflation and industrial production, inflation and real balances and real balances and industrial production, conditional on the orthogonalized VAR innovations for $r = -4, \ldots, 0, 1, \ldots, 4$; the impulse response of the variables of the system to orthogonal innovations; and the time path of the disturbances.

Fig. 2 shows that the first orthogonal shock generates positive pairwise contemporaneous cross-correlation functions in the relevant range, and therefore qualifies as "monetary" disturbances. The fourth orthogonal shock also produces

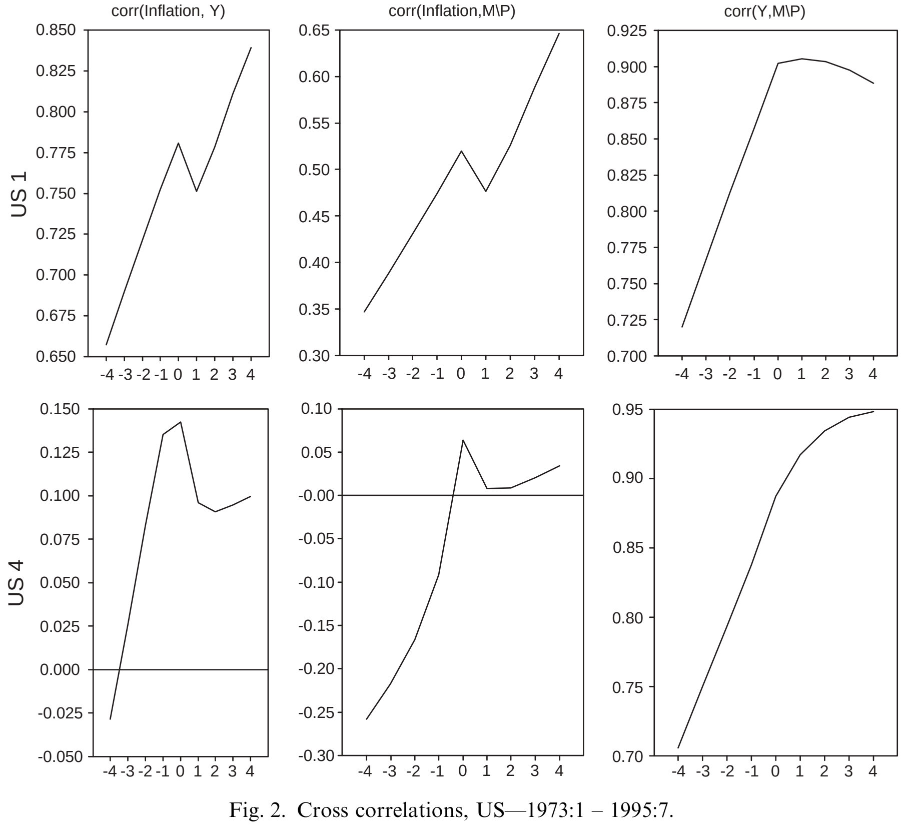

Fig. 2. Cross correlations, US—1973:1 – 1995:7.

positive values for all three cross-correlations at $r = 0$. However, the correlation between real balances and inflation is insignificant for a range of values of r. Hence, although this shock fits the prototype of monetary disturbance we have described, some care must be exercised in labelling it.

Fig. 3 indicates that the two monetary disturbances have distinct effects on real activity, inflation and the slope of the term structure. The first monetary shock produces sizable responses of industrial production and increases in real balances are associated with temporary but small increases in inflation and a decline in short-term rates relative to the long ones (the slope increases). This pattern is consistent with a standard liquidity interpretation of the shock. Note also that the response of real balances is almost synchronized with that of industrial production, suggesting that velocity may be nearly constant in responses to this shock. The second disturbance has negligible short-run real effects, but the impact response of inflation is strong and the slope of the term structure declines considerably for about two years after the shock. Since also output declines over this period, it may be reasonable to suspect that short-term rates have increased relative to long-term ones. This pattern is

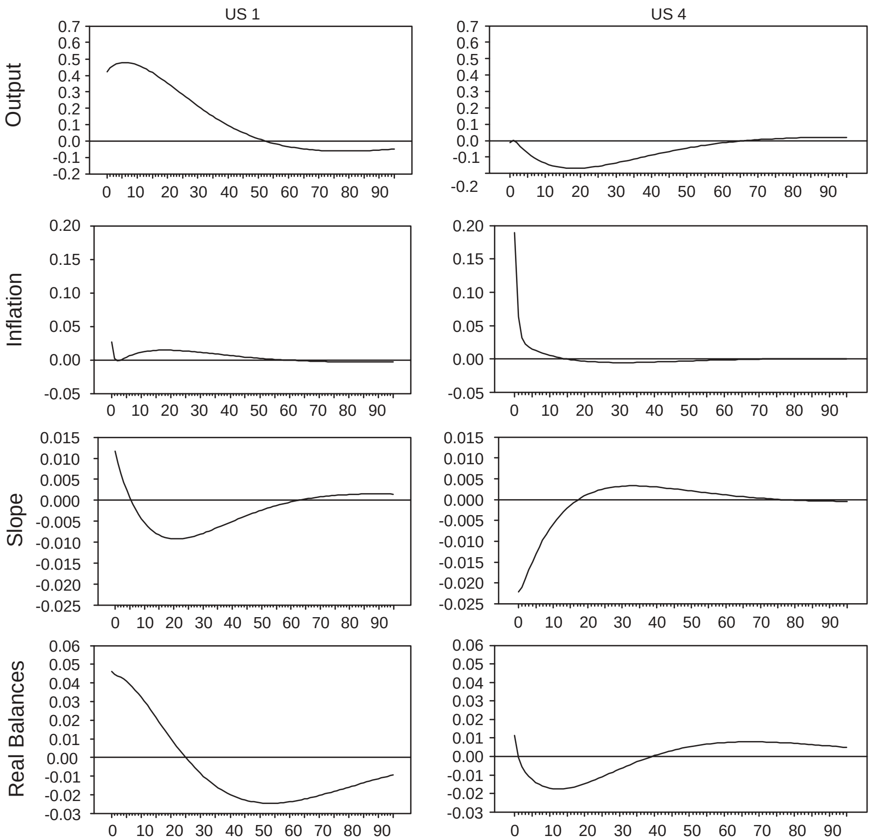

Fig. 3. Responses to monetary shocks, US (sample 1973:1–1995:7).

consistent with the idea that this type of shocks occur close to full employment and produce a temporary inflation effect that dominates the liquidity effect.

Two restrictions typically employed in high-frequency structural VARs are that prices (inflation) and production do not contemporaneously react to monetary shocks. The presumption is that there is sluggishness in the way prices are determined and that monetary shocks take time to produce real effects. Since our approach employs alternative identifying assumptions, we are in the position to verify whether these restrictions hold in the data or not. Fig. 3 indicates that neither of the two monetary disturbances has negligible instantaneous effects on both output and inflation. Hence, the identifying restrictions typically employed in VARs may be dubious and inference possibly flawed.

Fig. 4 shows that the volatility of the first monetary shock is approximately constant over the sample except for two large spikes around 1987–89. This shock also displays significant negative movements in 1974, 1979 and around the so-called Romer and Romer dates. The second monetary shock displays periods of high volatility in 1973–75 and 1979–82. Also, after 1982 its volatility declines, and there

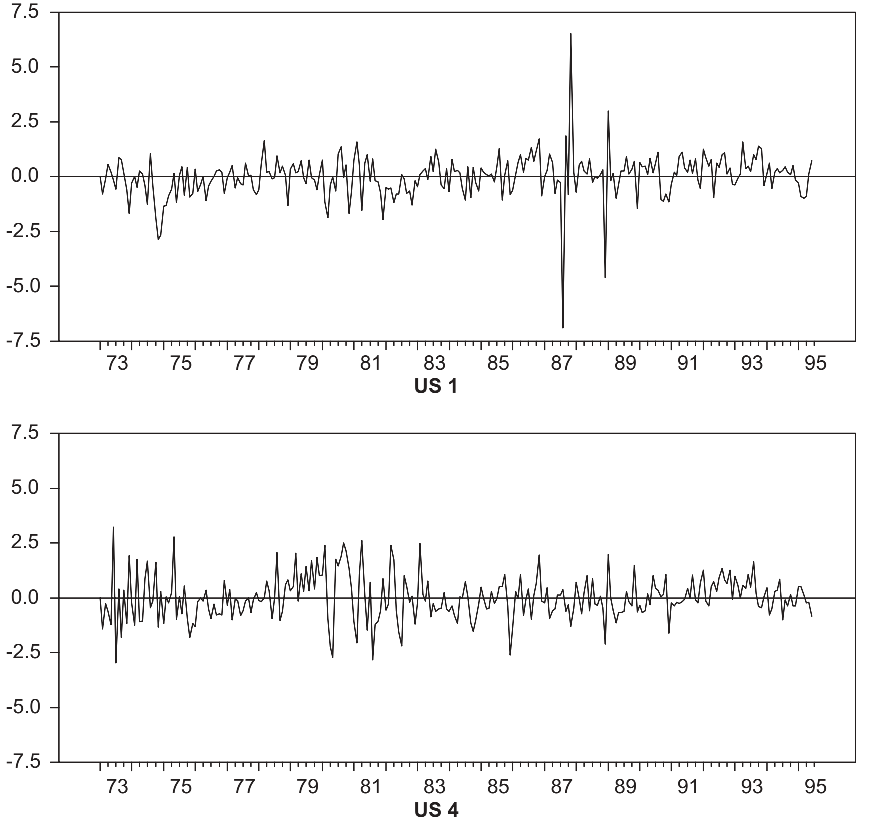

Fig. 4. Time path of nominal shocks, US — 1973:1–1995:7.

are only two episodes of significant negative disturbances: in correspondence with the Plaza Agreement (end of 1985) and at the end of 1988.

4.2. Are identified monetary shocks reasonable?

Before proceeding with the analysis, it is worth examining whether the two monetary disturbances we have identified are reasonable using other two criteria.

Rudebush (1998) has forcibly argued that structural shocks recovered with standard VAR identifications are unreasonable as they are unrelated to market perceptions of what monetary policy shocks are. He argued that innovations in the Federal Fund Future (FFF) rate carry this information and shows that structural VAR shocks are a poor proxy for these innovations. How are our monetary shocks related to innovations in FFF rate? We construct innovations in 1-month FFF rates by regressing the series on a lag of itself and three lags of the industrial production index and the slope of the term structure. The regression of our two monetary shocks on these innovations give the following results (t-statistics in parentheses):

$$\begin{array}{lllll} \mathrm{US}_{1t} = & -0.09 & + & 0.797\mathrm{FFF}_t, & R^2 = 0.04, \\ & (0.07) & & (1.84) & \\ \mathrm{US}_{4t} = & 0.02 & + & 0.608\mathrm{FFF}_t, & R^2 = 0.02. \\ & (0.07) & & (1.60) & \end{array} \tag{14}$$

Hence, FFF rate innovations are positively correlated with both monetary innovations; the regression coefficient is high and significant at 10% in both cases. However, since the volatility of the shocks we extracted is much larger than the volatility of innovations in FFF rate, the R^2 of the regressions is low. Overall, our monetary innovations appear to fare much better than the monetary policy innovations obtained with the VARs examined by Rudebusch.

Following Taylor (1993) several authors have claimed that a rule with stronger feedback from inflation than output is a good representation for the monetary policy conduct in the US for the last 20–30 years. Leeper et al. (1996) have argued that when monetary policy decisions are made, contemporaneous values of prices and output are typically unavailable and suggest that a partial accommodative rule, relating a monetary aggregate to a nominal interest rate, could do as well. One question of interest is therefore whether the rules produced by the two monetary shocks fit in one of these categories and have reasonable coefficients. Table 1 presents these rules, which are normalized, for convenience on the slope of the term structure together with the rules obtained using a Choleski decomposition with IP, Inflation, Slope and Real Balances in that order (as in Christiano et al., 1996) and forcing the slope to contemporaneously react to real balances (in the spirit of Gordon and Leeper (1992) or Leeper et al. (1996)). The first monetary shock produces a rule which resembles a partial accommodative rule with the addition of a modest feedback from output to the slope. The second rule is less easily interpretable in the sense that while the slope declines when inflation increases, as common sense would suggest, there is a positive relationship between real balances and the slope of

Table 1
Monetary policy rules

Sample 1973:1–1995:7	
US shock 1	Term $= -0.15\text{IP} - 2.97\frac{M}{P}$
US shock 4	Term $= -0.25\text{INF} + 0.33\frac{M}{P}$
Choleshi	Term $= -0.0009\text{IP} - 0.0006\text{INF}$
LSZ	Term $= 0.25\frac{M}{P}$

Notes: In Choleski the order is IP, Inflation, Term, Real Balances. In Leeper, Sims, Zha (LSZ), the order is the same but Term reacts only to Real Balances.

the term structure, probably due to the presence of (expected) inflationary effects. In both cases, the magnitude of the estimated coefficients is reasonable. In comparison, a Choleski decomposition produces, approximately, a slope targeting rule while the other specification also implies a positive trade-off between real balances and the slope.

In conclusion, identified monetary disturbances appear to be sound according to all criteria used. The first shock is generated with a simple partial accommodative rule, is significantly related to FFF rate innovations, produces liquidity effects, a sluggish response of inflation and a hump-shaped response in output. The second monetary shock is generated by a rule with a somewhat more difficult interpretation, but it is also significantly related to FFF rate innovations; has sluggish effects on output but strong contemporaneous effects on inflation.

4.3. *Identifying monetary disturbances in the other G-7 countries*

Table 2 reports the number of monetary shocks we have identified in the seven countries. Note that there is at least one monetary disturbance in all countries, and in Japan, Italy and UK three orthogonal shocks appear to be of monetary type.

Identified monetary disturbances fit three broad patterns. First, in five countries (Germany, France, Italy, Japan and Canada) at least one shock produces responses which are similar to those generated by the first US monetary shocks and fit our a priori idea of what a monetary *policy* disturbance does, i.e. when contractionary, such a shock should reduce nominal balances, decrease output, either on impact or with a short lag, contract inflation, make real balances decline and the short nominal interest rate increase relative to the long one. In all these instances the joint behavior of the four variables is consistent with the presence of a liquidity effect and the absence of the so-called "price puzzle" (see Sims, 1992).

Second, there is a group of monetary shocks which has perverse output effects. Expansionary disturbances of this type produce responses qualitatively similar to those of the second US monetary shock: nominal balances increase, output decreases on impact or with a short lag; instantaneous inflation responses are positive followed by a decline; the response of the slope term structure is positive and humped shaped. As shown in Fig. 5, there are disturbances in Germany, UK and Japan with these

Table 2
Identification

Country	Rotation	θ	Shock 1	Shock 2	Shock 3	Shock 4
Sample 1973:1–1995:7						
US	8	0.94	Monetary			Monetary
Germany	5	0.47	Monetary		Monetary	
Japan	8	1.53	Monetary	Monetary		Monetary
UK	1	0.31		Monetary	Monetary	Monetary
France	6	1.09		Monetary		
Italy	1	0.31		Monetary	Monetary	Monetary
Canada	1	0.62	Monetary			
Pooled	7	0.47		Monetary	Monetary	
Sample 1973:1–1982:10						
US	4	0.62	Monetary	Monetary	Monetary	
Germany	9	0.94			Monetary	
Japan	1	0.00		Monetary	Monetary	Monetary
UK	3	0.47		Monetary		Monetary
France	3	0.00		Monetary	Monetary	
Italy	1	0.47		Monetary	Monetary	Monetary
Canada	4	1.09			Monetary	Monetary
Pooled	1	0.62	Monetary		Monetary	
Sample 1982:11–1995:7						
US	2	0.31	Monetary		Monetary	Monetary
Germany	1	1.25		Monetary		Monetary
Japan	5	1.09		Monetary		
UK	4	1.25			Monetary	
France	2	0.62			Monetary	Monetary
Italy	7	0.31	Monetary			
Canada	7	1.41		Monetary		
Pooled	7	0.94				Monetary

Notes: In the rotation column, 1 indicates that the first two elements of the standardized eigenvalue–eigenvector decomposition matrix are rotated; 2 indicates that elements one and three of this matrix are rotated; 3 indicates that elements one and four of this matrix are rotated; 4 indicates that elements two and three of this matrix are rotated; 5 indicates that elements two and four of this matrix are rotated; 6 indicates that elements three and four of this matrix are rotated; 7 indicates that elements one and two, and three and four of this matrix are contemporaneously rotated; 8 indicates that elements one and three, and two and four of this are contemporaneously rotated; 9 indicates that elements one and four, and two and three of this matrix are contemporaneously rotated. θ measures the angle of rotation.

features. Once again this pattern is consistent with the idea that a surprise increase in nominal balances makes real balances decline on impact, probably because these shocks occur close to full employment. Output then declines either because demand has declined or because high inflation has increased costs of production. When these effects are persistent, increases in expected inflation translate into an increase in the long-term interest rates relative to short-term ones over the medium run.

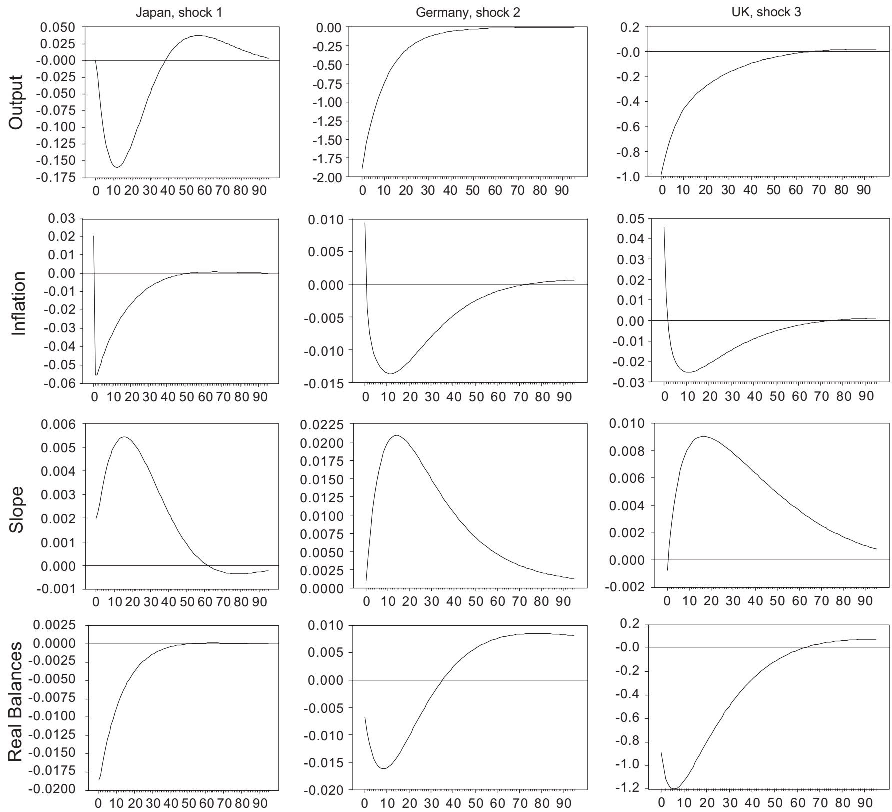

Fig. 5. Responses to monetary shocks, G-7—1973:1–1995:7.

The final typical pattern characterizing identified monetary shocks across countries appears to be linked to international factors. That is, the volatility of some of the disturbances increases at times of speculative pressure in international currency markets and, for European countries, at time of realignment of their exchange rates within the EMU. In Fig. 6 we report the time path of two such shocks, one for Germany (spikes in 1984, 1992 and 1994) and one for Italy (spikes in 1979, 1989 and 1992). In general, the volatility of these shocks increases at times of speculative pressure in international currency markets. Positive realizations of this type of disturbances generate strong expected inflation effects and produce positive hump-shaped responses in the term structure.

4.4. The explanatory power of monetary disturbances

Next, we calculate the contribution of monetary shocks to output and inflation cycles. What we compute here Japan are lower bounds, because there are orthogonal innovations without an informational content. These innovations may also contain

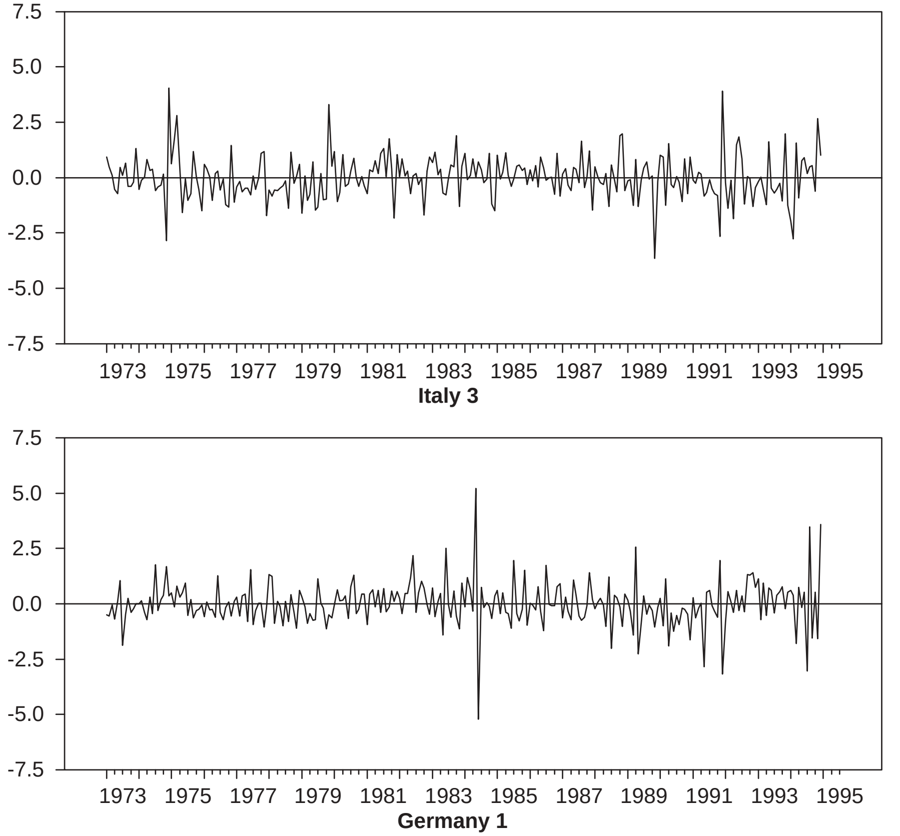

Fig. 6. Selected monetary stocks, G-7—1973:1–1995:7.

components which may be monetary in nature, distinct from and uncorrelated from the ones we measure, so that the total percentages we present here could be augmented if, by means of other variables or additional constraints, we could uncover what drives the remaining unnamed innovations. Table 3 presents 68% bands for (i) the contribution of each monetary shock, (ii) the total contribution of monetary shocks to the forecast error variance of output and inflation at 24-step horizon. Varying the forecasting horizon between 12 and 48 steps has no effects on the results.

The table displays four important features. First, except for Germany and Canada, no single monetary shock explains extreme portions of output variability. Hence, the disturbances we have identified are not of the type extracted by Faust (1998). The case of Germany is special and appears to be due to the large break in the real balance series following unification, while for Canada the disturbance that explains a large portion of output variability is highly volatile when financial markets turbulence increases. Second, the combined contribution of monetary disturbances

Table 3
Percentage of the 24-month forecast error variance of industrial production and inflation explained by monetary disturbances

	Variance of industrial production					Variance of inflation				
	Shock 1	Shock 2	Shock 3	Shock 4	Sum	Shock 1	Shock 2	Shock 3	Shock 4	Sum
Sample 1973:1–1995:7										
USA	10–56			1–11	16–60	3–11			44–57	54–64
Germany	69–82		15–26		95–99	3–15		10–19		25–32
Japan	14–41	0–8		0–5	22–45	0–5	4–12		78–89	87–99
UK		9–51	5–24	1–10	37–77		2–9	64–87	0–3	75–98
France		0–6					16–19			
Italy		3–16	9–18	3–12	25–45		78–86	0–3	9–12	85–95
Canada	67–87					1–7				
Pooled		15–20	7–43		23–55		50–65	3–15		51–80
Sample 1973:1–1982:10										
USA	20–58	4–22	18–51		68–94	3–17	35–51	6–23		76–97
Germany			28–50					1–7		
Japan		1–9	34–58	4–18	55–76		74–91	2–7	0–4	73–98
UK		18–59		12–14	31–66		2–10		0–3	3–10
France		8–13	30–52		39–60		76–92	10–13		86–95
Italy		7–25	19–32	1–8	39–61		63–74	2–9	19–22	76–92
Canada			17–43	18–30	34–70			2–11	5–8	4–15
Pooled	19–71		11–69		74–91	1–19		12–28		33–63
Sample 1982:11–1995:7										
USA	53–89		4–37	0–8	56–99	2–13		0–3	8–11	6–16
Germany		0–10		0–6	0–14	18–23		60–80	88–97	
Japan		2–17						19–23		
UK			31–63					5–23		
France			14–44	1–27	16–55			7–60	4–29	13–83
Italy	56–84					1–9				
Canada		0–11					2–3			
Pooled				0–17					53–61	

Notes: The forecast error variance is computed using a 4 variable VAR model. The table shows the 68% error band for the 24-month forecast error variance in the variable explained by sources of structural innovations. Bands are computed using Monte Carlo replications. Sum reports the standard error bands for the total contribution of monetary disturbances to the variability of output and inflation.

for real fluctuations is large, except in France. In the US, for example, monetary shocks explain between 16% and 60% and in the UK between 37% and 77% of the variance of output. Third, one monetary disturbance accounts for a large portion of inflation variability in US, Japan, UK and Italy. Fourth, the combined contribution of monetary disturbances to inflation variability exceeds 50% in four of the seven countries.

How do these results relate to those present in the literature? Most analyses have concentrated on the US and the consensus view appears to be that a small portion of output variability (between 0% and 20% or 15% and 35%, depending on the estimates) is due to monetary disturbances. The first monetary shock, whose characteristics are most similar to those examined in the literature, has a much larger median impact (38%) but the estimated bands are large and include, to a large extent, existing estimates. Kim (1999) examined monetary disturbances in the G-7 using a standard identification approach and has found them to be of negligible importance for output variability. Our results indicate, on average, a much larger role for monetary shocks in all countries, and this is true even when we restrict attention to disturbances which generate liquidity effects.

4.5. Sub-sample analysis

The domestic and international portions of monetary markets of the G-7 countries have undertaken substantial changes over the sample. For example, capital controls and restrictions on domestic holdings of foreign currencies have been gradually eliminated during the 1980s. Domestic banking constraints, e.g. regulation Q in the US or quotas on the portfolio of banks in European countries, have also been scrapped in favor of more market-oriented policies. These changes may have affected the way monetary disturbances are transmitted to the real economy and the lag needed for prices and quantities to fully adjust to these disturbances.

In this subsection we report evidence obtained from two subsamples (73:1–82:10 and 82:11–95:7) in order to check whether instabilities or regime shifts change the essence of the results. It should be kept in mind that by breaking the sample we avoid to mix periods with different structural characteristics, but estimates of the cross-correlation functions are more likely to be imprecise, and the informational content of orthogonal innovations more difficult to detect. We chose 1982:10 as common break point following the existing literature (see e.g. Kim, 1999). While there are arguments in favor of choosing a unique sample break for all countries, it is also the case that, at least in Europe, there are episodes which may require further subdivisions (the German unification in 1990, the breakdown of the monetary snake in 1979 and of the EMS in 1992, and so on). We do not investigate these additional potential breaks as the sample size becomes too short to make sense of the estimates of the cross-correlation function. The time path of the identified disturbances suggests that these episodes are better characterized as outliers than as structural breaks with changing dynamics.

For the first subsample, the results mirror those obtained for the full sample, but some differences also emerge. For example, the number of identified monetary disturbances changes: in the US, Japan and Italy we recover three of these shocks; in UK, France and Canada two while in Germany only one shock is monetary in nature. Despite these differences, the general conclusions remain: within each country there are shocks which generate liquidity effects and others which generate inflation effects; the combined contribution of monetary disturbances to the variability of industrial production is large and significant; monetary shocks are the dominant source of variability in inflation in four of the seven countries.

In the second subsample, we recover at least one monetary disturbance in all countries but the relative importance of these shocks for industrial production and inflation fluctuations changes. In Japan, the UK and France, the total contribution of monetary disturbances to real fluctuations declines to a more modest level while in the US and Italy it increases. Similarly, the combined contribution of monetary disturbances for inflation variance declines in US, Canada, Japan and Italy while it increases in Germany and the UK. Also in this subsample, identified shocks fall into the three broad categories we have found for the full sample. However, inspection of the time path of the estimated disturbances indicates that the shock which contributed most to the variability of inflation in France, Germany the UK is highly volatile at times of realignments and/or disruptions of the European Monetary System and at German unification. Hence, an "EMU" shock, more related to turbulence in international money markets than to domestic (policy) changes, may be present in this subsample.

4.6. A pooled VAR

Instead of asking how important are monetary disturbances in explaining output and inflation cycles in each of the G-7, one may be interested in knowing what is the "typical" effect of a monetary innovation in an average G-7 country. To investigate this question we estimate a pooled VAR model with a country-specific intercept.

A pooled model correctly recovers the average informational content of orthogonal innovations if the DGP of the actual data were the same for all countries, apart from a level effect. When this is the case and the time-series dimension of each sample is short, we can obtain more precise estimates of the cross-correlation function by pooling together the seven data sets. In practice, this means that inference may be more accurate since the mechanism driving output and inflation fluctuations may have been operating in a larger number of instances. For example, one should a priori expect monetary shocks in the European countries to have a common component with the differences previously noted due to small sample sizes. By pooling data together one hopes that this commonality will translate in repeated observations on either the same source or the same propagation mechanism, therefore providing a more accurate representation of the forces at work.

The drawbacks of pooling are well understood. Neglecting heterogeneity in the dynamics produces inconsistent estimates of the parameters and biases structural inference, i.e. we get more precise estimates of the possibly wrong source of disturbance. Under the assumption that short-term dynamics are the same across countries and the samples are large enough, single country VARs and the pooled VAR will give identical information on the structural sources of disturbances.

For a pooled VAR we identify two monetary disturbances in the full sample (see Table 1). The qualitative similarities between the pooled and the US model are remarkable: not only identified shocks produce the same dynamics but also their relative importance for industrial production and inflation variability is similar (see Table 2). For the first subsample, we also identify two monetary disturbances. Jointly, these shocks account for a large portion of the variability of industrial production and they explain between one-third and two-thirds of the variability of inflation. For the second subsample, we identify only one monetary shock but its contribution to industrial production fluctuations is negligible.

To summarize, the cross-country dynamics following orthogonal VAR innovations are sufficiently homogeneous for the full and the first subsample to make pooled estimates of the cross-correlation function meaningful. In these samples, the results once again emphasize the important role that monetary disturbances play for output and inflation cycles. For the second subsample, heterogeneities appear to be important, so the misspecification present in the pooled VAR prevents us from drawing useful conclusions regarding the importance of monetary shocks.[2]

5. The variability of the slope of the term structure

Implicit in our identification scheme is the idea that monetary disturbances are policy driven, i.e. they are expected to represent disturbances that move the supply of funds. However, it may be the case that under certain policy design (for example, an interest rate targeting) identified monetary shocks represent money demand disturbances. One way to disentangle these two possible interpretations is to examine how the slope of the term structure responds, and measure the time needed for these disturbances to be fully incorporated in this variable.

Liquidity theories of monetary policy (see e.g. Christiano et al., 1997) stress that the magnitude of the real effects crucially depends on how quickly financial markets adjust to monetary disturbances. For example, Evans and Marshall (1998) have shown that, at least for the US, contractionary monetary policy shocks produce a contemporaneous positive response of the slope of the term structure, and that this response changes sign in the medium run when expected inflation effects become important. Since for part of the sample several Central Banks followed monetary rules that implicitly or explicitly gave heavy weights to interest rates, we should expect a speedy reaction of the slope of the term structure in many of the G-7 countries, if the disturbances we have identified are truly policy shocks. In this situation, disturbances that move money demand should leave the term structure of interest rates unaffected—exactly if the central bank follows a fully accommodative rule, approximately if interest rate smoothing policies are implemented.

Our discussion in Section 4.1 has already pointed out that the slope of the term structure in the US quickly responds to both monetary disturbances, and that the shape of the response depends on the relative importance of liquidity and expected inflation effects. The remaining G-7 countries display similar features. Out of the 13

[2] We have also examined the typical dynamics obtained by averaging the relevant statistics over the seven countries as suggested by Pesaran and Smith (1995). The results obtained are mixed and the procedure is unable to provide any sharp conclusion about sources of output and inflation cycles.

monetary disturbances we identified for the full sample, eight shocks, if contractionary, produce an instantaneous increase in the slope of the term structure and, in five cases, hump-shaped responses in the medium run. For the remaining five cases a shock, if contractionary, produces first a decline and then an increase in the slope, but in all cases we observe hump-shaped responses in the medium run. One reason for the presence of these two different patterns may be related to the credibility that different central banks may have gained over the period. Another may have to do with the fact that some monetary shocks are related to disturbances in international financial markets, thereby producing strong expected inflation effects.

Table 4, which reports the total percentage of variance in the slope of the term structure jointly accounted for by monetary shocks at 3- and 24-month horizons, supports of the hypothesis that identified disturbances contain a large policy component. For the full sample, monetary disturbances account for 95–99% of the fluctuations in the slope in Japan, UK, Germany and Italy at the 3-month horizon and this percentage is also large at the 24-month horizon, except for Germany. For France and the US, the percentage is slightly smaller at the 3-month horizon (66–84% and 25–58%, respectively) and the percentage at the 24-month horizon is approximately the same. Finally, for Canada there is little evidence that monetary shocks are responsible for variations in the term structure at both horizons.

As it was the case for fluctuations in industrial production and inflation, the relative importance of monetary shocks for the variability of the term structure changes over time. For the 1973–82 period, monetary shocks account for 70–99% of the variance of slope of the term structure at the 3-month horizon in five countries; in Germany the percentage is smaller and in France it is nil. For the 1982–95 period, monetary shocks are the overwhelming source of term structure variability in the US, UK and France at the 3-month horizon; they are important in Japan and Germany, and have no influence in Canada and Italy.

6. Conclusions

This paper examined the importance of monetary disturbances for output and inflation fluctuations using a novel two-step identification approach. The proposed procedure is advantageous for several reasons: it uses widely agreed sign restrictions derived from economic theory to identify shocks; it clearly separates the statistical issue of obtaining contemporaneously uncorrelated innovations from that of identifying their informational content; and it allows us to explore the space of identifications systematically.

The consensus view about the contribution of monetary disturbances to output fluctuations seems to be that these shocks have, at most, a modest importance (see Sims (1998) or Uhlig (1999)). This view has been challenged by Roberts (1993) and, more recently, by Faust (1998), who claim that there are identification schemes which produce reasonable dynamics, where monetary disturbances account for a

Table 4
Percentage of the forecast error variance of the slope of the term structure explained by monetary disturbances

	3-month horizon	24-month horizon
Sample 1973:1–1995:7		
USA	25–58	25–57
Germany	98–99	26–54
Japan	98–99	73–91
UK	95–99	77–95
France	66–84	49–72
Italy	98–99	95–99
Canada	0–3	18–34
Pooled	1–67	14–55
Sample 1973:1–1982:10		
USA	69–96	58–95
Germany	19–46	10–26
Japan	97–99	78–94
UK	93–98	61–83
France	3–11	13–39
Italy	91–99	83–99
Canada	89–98	55–87
Pooled	86–97	76–87
Sample 1982:11–1995:7		
USA	97–99	90–99
Germany	20–39	34–56
Japan	35–63	21–47
UK	91–96	4–14
France	93–99	77–96
Italy	0–3	0–7
Canada	0–2	27–52
Pooled	3–69	18–45

Notes: The forecast error variance is computed using a 4 variable VAR model. The table shows the 68% error band computed using Monte Carlo replications for the total contribution of monetary disturbances to the variability of the slope of the term structure.

large portion of output variability in the US. Our results reinforce these challenges in several ways.

First, we find that monetary disturbances are at work in all countries and in all samples. We show that identified monetary disturbances generate either liquidity or inflation effects, which appear to be linked either to domestic or to international factors and that they produce similar responses across countries. Second, we show that these shocks play a major role in driving output, inflation and term structure variability in most of the G-7 countries. Third, for every country and all subsamples, we show that the slope of the term structure quickly reacts to these disturbances.

This last result, together with the observation that, in the US, monetary disturbances are related to the FFF rate innovations and generated reasonable policy functions, leads us to conclude that the shocks we recover are theoretically sound and have important policy components.

Our results also provide empirical support to the recent resurgence of interest in theoretical models where monetary shocks are the engine of the business cycle and suggest that a careful study of the nature of these shocks may shed important light on mechanics of propagation across various markets within and across countries.

Appendix

In this appendix we describe how we explore the space of orthogonal decompositions. It is well known that, if we exclude the case of recursive models, the set of possible identifications is uncountable and it is difficult to search effectively. The algorithm we employ makes use of the following result which is contained in Press (1997).

Result. Let P be the matrix of eigenvectors and D the matrix of eigenvalues such that $\Sigma = PDP'$. Then $P = \prod_{m,n} Q_{m,n}(\theta)$ where $Q_{m,n}(\theta)$ are rotation matrices of the form

$$Q_{m,n}(\theta) = \begin{pmatrix} 1 & 0 & 0 & \dots & 0 & 0 \\ 0 & 1 & 0 & \dots & 0 & 0 \\ \dots & \dots & \dots & \dots & \dots & \dots \\ 0 & 0 & \cos(\theta) & \dots & -\sin(\theta) & 0 \\ \vdots & \vdots & \vdots & 1 & \vdots & \vdots \\ 0 & 0 & \sin(\theta) & \dots & \cos(\theta) & 0 \\ \dots & \dots & \dots & \dots & \dots & \dots \\ 0 & 0 & 0 & 0 & 0 & 1 \end{pmatrix},$$

where $0 < \theta \leqslant \pi/2$ and the subscript (m, n) indicates that rows m and n are rotated by the angle θ.

To translate this result into an algorithm that searches the space of orthogonal decompositions, note first that in a system of N variables there are $(N(N-1)/2)$ bivariate rotations and $(N(N-1)/4)$ combinations of bivariate rotations of different elements of the VAR, for a fixed θ. Hence, for $N = 4$ there are 9 possible rotation matrices. Second, since $Q_{m,n}(\theta)$ are orthonormal $\Sigma = \hat{V}\hat{V}' = VQ_{m,n}(\theta)Q_{m,n}(\theta)'V' = PD^{0.5}Q_{m,n}(\theta)Q_{m,n}(\theta)'D^{0.5}P'$ is an admissible decomposition. Hence starting from an eigenvalue–eigenvector decomposition we can "decuple" it in one direction or another, for each θ. Third, we grid the interval $[0, \pi/2]$ into M points, and construct $9M$ orthogonal decompositions of Σ. This last step transforms an uncountable into a large but finite search.

In practice, one needs to choose both how many r to include and how many points the grid should have. To maintain computations feasible, we start the process by choosing $r = 0$ and/or $r \pm 1$. Since theory typically has strong prediction for either contemporaneous or one-period lagged correlations (e.g. in model with sticky prices), this starting point is not restrictive. Moreover, as mentioned, the number of sign correlation restrictions considered can be increased if multiple candidates satisfy the restrictions. In the specific application of this paper, matching the sign of cross-correlations at $r = 0, 1$ was sufficient to select a unique candidate. Also, to keep computation manageable, we limit the number of grid points to be less than 500. Depending on the country, between 30 and 500 points for each rotation matrix were sufficient to cover the interval effectively.

The algorithm described is related to the one employed by Uhlig (1999, Proposition 1, Appendix A). In his application the vector α, which defines the selected identification, is a particular transformation of sine and cosine functions. Here the matrix $Q_{m,n}(\theta)$ defines the selected identification and is an explicit function of sines and cosines of an angle.

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