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Estimating the Effect of Hungarian Monetary Policy within a Structural VAR Framework

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Abstract

A standard approach in measuring the effect of monetary policy on output and prices is to estimate a VAR model, characterise somehow the monetary policy shock and then plot impulse responses. In this paper I attempt to do this exercise with Hungarian data. I compare two identification approaches. One of them involves the 'sign restrictions on impulse responses' strategy applied recently by several authors. I also propose another approach, namely, imposing restrictions on implied shock history. My argument is that in certain cases, especially in the case of the Hungarian economy, the latter identification scheme may be more credible.

In order to obtain robust results I use two datasets. To tackle possible structural breaks I make alternative estimates on a shorter sample as well.

The main conclusions are the followings: (1) although the two identification approaches produced very similar results, imposing restrictions on history may help to dampen counterintuitive reaction of prices; (2) after 1995 a typical unanticipated monetary policy contraction (a roughly 25 basis points rate hike) resulted in an immediate 1 per cent appreciation of the nominal exchange rate (3) followed by a 0.3% lower output and 0.1-0.15% lower consumer prices; (4) the impact on prices is slower than on output; it reaches its bottom 4-6 years after the shock, resembling the intuitive choreography of sticky-price models; (5) using additional observations prior to 1995 makes identification more difficult indicating the presence of a marked structural break.

JEL classification: C11, C32, E52

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1 Introduction

During the 1990s many researchers attempted to estimate the effect of monetary policy on output and prices using the structural VAR approach. The purpose of the research was often to find the monetary general equilibrium model most consistent with the data. Despite the effort devoted to this issue there remained some unresolved problems, although some consensus results also emerged.

From the central banker's point of view, especially if he is an inflation targeter, the most important thing is perhaps the behaviour of prices in the wake of a monetary policy action. Unfortunately, the reaction of prices has shown the largest variability across models. Nevertheless, as Christiano, Eichenbaum and Evans (1998) (henceforth CEE) claim, the impulse response of some other variables, like that of output, had proven to be very robust to specification.

Even if we have a consensus view about the impulse responses, it is useful to clarify first what we can learn from structural VARs about the effect of monetary policy. The standard SVAR approach involves identifying monetary policy shocks and quantifying their consequences. These shocks are unexpected deviations from the systematic behaviour of monetary policy, from the so-called 'monetary policy rule'. But since these deviations usually explain a small part of the policy instrument's variation, the question naturally arises: are these shocks important in understanding the transmission of monetary policy? Why do we not simply regress changes in output and prices on changes in the policy instrument, for example, in the short interest rate?

The answer is because interest rate changes are mainly endogenous (i.e. consistent with the policy rule) reactions of monetary policy to other types of shocks coming from the economy. If we trace the development of prices following that particular change in interest rate, we only get a picture about the consequences of that particular shock which, among others, caused the interest rate movement. Clearly, the endogenous reaction of monetary policy is only one channel through which disturbances exert their influence on prices. It is therefore crucial to separate autonomous disturbances coming from monetary policy from other types of shocks.

Even if we are aware of the advantages of identifying pure monetary policy shocks, their interpretation is not yet straightforward. Some possible versions are listed in CEE (1998). I would like to cite two of them here. The first is perhaps the most often used 'exogenous shift in preferences' term. Since shocks are one-off deviations from the rule, and the rule can be derived from the decision maker's preferences, this explanation may not be very convincing. If one would like to model changing monetary policy preferences, policy rule with time-varying coefficients may better describe the actual behaviour.

Another approach involves saying that those shocks are due to imprecise measurement, lack of reliable real-time data, or statistical error. Although this seems to undermine the claimed usefulness of monetary policy shocks at first glance, I would prefer the latter interpretation in a linear modelling environment. Despite being small and unintended, these 'errors' help us to unveil the reaction of macro variables when the only source of the disturbance is the monetary policy. When a decision maker has an erroneous picture about the state of the economy, he or she deviates from systematic behaviour involuntarily and makes the economy reveal the difference in its response from the 'normal' course. Of course, these errors are small relative to the predictable actions. Put in another way, the investigation of monetary policy shocks does not help much to characterise the monetary policy, but rather the response of variables to monetary policy, the transmission mechanism.

The identification of these monetary policy shocks is not straightforward. Special care should be taken in choosing the appropriate approach when working with data such as our Hungarian time series. The bulk of the literature has dealt with large, closed economies with stable institutions; hence the adoption of known methods to small open economies just having undergone some transition processes should be coupled with critical modifications. Two principles are recommended: (1) one should seek the identification that uses the least structural knowledge about the economy and (2) one should check the robustness of the results by using alternative approaches, too. These recommendations are not orthogonal to each other.

This paper tries to meet these requirements. The first is taken into account by imposing sign, or more generally, inequality restrictions instead of concrete values. The second is fulfilled by using two independent sets of assumptions: in the baseline identification I impose sign restrictions on impulse responses similarly to Uhlig (2004) and Jarocinski (2004). The alternative strategy is to some extent related to the 'narrative approach' of Romer and Romer (1989) and to the approach of Rudebusch (1998) and Bagliano and Favero (1997, 1999). The basic idea in all these papers is to use historical evidence regarding monetary policy shocks. My identification scheme, however, is more liberal, since I only specify the date of the largest contractionary and loosening monetary policy shocks, and I do that using only inequality restrictions.

One of the main conclusions drawn after having experimented with several specifications is that using data between 1995 and 2004 provides results more comparable to the consensus of the SVAR literature than the longer sample beginning in 1992. The second important technical observation is that the results are quite robust to the identification strategy.

As far as transmission mechanism is concerned, a typical monetary policy shock during the past nine years caused roughly an immediate 25 basis points short interest rate rise and a one per cent appreciation of the nominal exchange rate. The output declines very quickly after the shock, reaching its minimum at -0.3% within the first 3 years. The reaction of consumer prices is much more protracted, but somewhat smaller: the maximum reduction is 0.1-0.15% between the 4th and 6th years after the shock.

The structure of the paper is as follows: in section 2 the issue of identification is discussed. In sections 3 and 4 the baseline and the alternative estimates are presented. The last section concludes. At the end of the paper the reader can find several charts, some of them not referred to in the text but nevertheless conveying interesting information.

2 Identification of monetary policy shocks

In the following subsections typical identification schemes are outlined[2]. Identification of structural shocks, such as monetary policy shocks, involves the imposition of some restrictions. These strategies can be more or less classified by the statistics that are restricted or by the precision of the restrictions, namely whether they require the target parameter to equal some real number, or just to be greater or less than certain values. This section follows the latter grouping but in subsection 2.4 the former aspect is also touched on. Subsection 2.5 justifies the identification strategy adopted in this paper.

2. 1 Major aspects of identification

Within a structural VAR framework one estimates first the reduced form model, which is approximated by a vector-autoregressive specification:

$$y_t = A_1 y_{t-1} + A_2 y_{t-2} + \ldots + A_p y_{t-p} + Bz_t + u_t$$

where y stands for the vector of n endogenous variables, z contains intercept, deterministic trend and other exogenous variables, p is the number of lags included, and vector u is the unexplained part of the vector process. B is the matrix of coefficients of exogenous variables, while $A_1 \ldots A_p$ are n×n coefficient matrices of lagged endogenous variables.

The estimated residuals $\hat{u}_t$ are historical shocks to the corresponding endogenous variable. If, for example, the residual of price level equation in Q1:97 is .01, we claim that one percentage point of the change in price level was unexpected, at least as far as our specification contains all relevant information market participants possessed in Q4:96. However, the source of that disturbance is not yet identified.

We are usually not interested in estimating price level or output innovations, but rather economically meaningful, i.e. supply, demand, etc. shocks, and particularly their dynamic effect on some variables. If, in our example, the output grew unexpectedly in the same period, i.e. it has a positive residual in Q1:97, we can suspect the presence of demand side pressure.

The main task after having estimated the VAR model is to decompose residuals into these structural shocks. This corresponds to finding the contemporaneous relationship between structural and reduced form innovations, or finding matrix *C* in the equation

$$u_t = Ce_t$$

where u denotes the vector of estimated residuals (output, price level, etc.) and e the vector of structural shocks (technology, demand, etc.). It is assumed that structural shocks are orthogonal to each other, while the same is not necessarily true for VAR residuals. Matrix *C* contains the contemporaneous impact of structural disturbances on endogenous variables. The element in the i-th row and j-th column

[2] For a more detailed introduction, but another classification scheme, see CEE (1998).

is the magnitude by which the j-th structural shock affects i-th variable simultaneously. With this formulation the residuals of the reduced form are derived as linear combinations of structural shocks.
Unfortunately, this matrix is not unique, which means that there is more than one structural model that has the same reduced form. We have to add some additional information in order to obtain the results we are searching for. Providing this information is the identification of structural shocks. It can be shown that in order to achieve full or exact identification one needs to impose $n(n-1)/2$ restrictions on C - in addition to n normalisation. When working with fewer restrictions (underidentified system) the point estimates of the parameters we are interested in (e.g. the response of output to one standard deviation monetary policy shock in periods 1, …, 8) broaden to intervals. In the overidentification case we have more assumptions than required for exact identification. The logic of estimation is then somewhat different: one weights the deviations from the restrictions and optimises.
Identification is the most sensitive part of the estimation procedure. We have to assume something about the structure we are investigating. Results from identified VARs usually take the form of a conditional statement. In particular, monetary transmission SVARs usually produce findings that sound like this: 'assuming monetary policy shocks' effect on x,y,… has the property …, the effect of monetary policy on z,w,… can be characterised as follows:…'.
Accordingly, identifying assumptions optimally represent our least disputable prior knowledge about that particular mechanism. This is important in order to obtain credible results. There are, however, two difficulties in finding the appropriate set of restrictions: (1) we have to impose enough restrictions in order to obtain a clear result and to avoid 'informal identification'; (2) we have to impose few enough restrictions in order to have a convincing identification strategy. While (2) is in accordance with the above-mentioned logic, the former criterion may require further explanation.
Let us consider the example of identifying monetary policy shocks – the purpose of this paper. Monetary policy shocks share some common features with other shocks. Autonomous monetary tightening, for instance, may be similar to a positive demand shock in its contemporaneous effect on the interest rate: in both cases one expects a higher policy rate in the period the shock hits the economy. The reasons are, however, different. Whereas in the first case this is an unexpected deviation of monetary policy from its rule, in the second case the higher interest rate is a consequence of systematic monetary policy that reacts immediately to inflationary pressure. In order to distinguish between the two disturbances further assumptions are needed. Assuming that an autonomous monetary contraction appreciates the exchange rate may, for example, disentangle it from demand shock.
Another reason for having a rich restriction set arises from realising that sometimes implicit assumptions are applied during the model selection procedure. The econometrician usually has a high degree of freedom. Within the SVAR framework, selecting the number of variables and the variables themselves (e.g. GDP vs industrial production as a measure of real output) included in VAR, the choice of sample, lag length, etc. are subject to decision, even if one relies on some model selection criteria. Typically, the researcher estimates several models and compares their outputs. He is inclined to keep the specification that meets some expectations not made explicit prior to estimation. Put in another way, specifications producing more appealing impulse responses are preferred to other set-ups, even if they all meet formal identifying restrictions to the same extent.

This model selection mechanism uses informal or implicit identifying restrictions. Distaste for 'price puzzle' is a good example. We call price puzzle the observed perverse behaviour of the price level following a monetary policy shock, that is rising prices after unanticipated monetary contraction. Suppose we have two sets of impulse responses triggered by one standard deviation monetary policy shock, both obtained from VAR imposing the same identifying restrictions. One of them exhibits price puzzle, the other does not. It is difficult then to resist the temptation of keeping the well-behaving specification while rejecting the other, which means imposing ex post additional restrictions.

A transparent identification strategy should avoid such steps by making those assumptions explicit. Nonetheless, even a priori exclusion of price puzzle is hard to justify as long as our aim is to estimate the effect of monetary policy shocks on prices, since we have then no chance to answer the question: 'Is the price puzzle a reality or just an identification failure?[3]

2. 2 Point (zero) restrictions

The most popular identification approach is to restrict some elements of matrix C to be zero. This strategy has the advantage that a structure of contemporaneous impacts like that can be translated to delayed reaction. Identification of monetary policy shocks is usually based partly on assuming no immediate effect on output and prices.

As a special case, so-called recursive identification involves an ordering of the variables. In this specification structural innovations affecting some variable do not appear contemporaneously in the residuals of variables ordered before. The matrix C becomes lower triangular and can be obtained by a Cholesky decomposition of the VAR's covariance matrix.

If we believe that the source of all nominal shocks is the monetary policy, and that monetary policy shocks do not affect output and prices contemporaneously, a 3-variable VAR, including output, prices and interest rate together with a Cholesky decomposition with the innovations in the interest rate ordered last is a good minimal workhorse, which is especially appropriate for international comparison – see, for example, Gerlach and Smets (1995). Most authors use larger models in order to include as much information as possible supposed to be available to monetary policy makers when making decisions, while maintaining the recursiveness assumption – for example CEE (1998), Peersman and Smets (2001).

Faust et al (2003) estimate first on high frequency data the contemporaneous impact of monetary policy shock and then use the coefficients in their monthly VAR. Although their identification is more sophisticated and fits better the theme of the next subsection, this is an example of using non-zero point restriction. Similarly, Smets (1997) estimates the contemporaneous impact of monetary policy and exchange rate shocks on interest rate and exchange rate outside the VAR and uses those estimates in his transmission VAR identification. In both cases the two step approach is necessary because of the supposed simultaneity between financial variables, thus the invalidity of recursiveness assumption.

[3] What kind of relevance might be attributed to statements such as the following: 'assuming that a contractionary monetary policy shock causes lower prices for one year, we get the result that the response of price level to one standard deviation monetary contraction is…'? In this conditional statement the condition and the statement is mixed up. This problem, however, refers to the aspect of identification credibility.

This point is crucial regarding estimation with Hungarian data. As I argue later, in addition to monetary policy, the risk assessment of forint denominated assets must have been the main force influencing the nominal interest rate and exchange rate during the past decade. Due to the quick reaction of monetary policy to exchange rate movements and the exchange rate to monetary policy surprises, the simultaneity problem seems to be highly relevant, ruling out a priori the adoption of recursive identification.
Another strategy is based on the assumed long-run neutrality of monetary policy. In practice, this means that monetary policy has only a temporary effect on real variables, such as output. Such restriction was applied by Clarida and Gali (1994), Gerlach and Smets (1995), and with Hungarian data Csermely and Vonnák (2002), among others. Note that imposing zero long-run effect is also a point (zero) restriction. For the shortcomings of such restrictions see Faust and Leeper (1994). Perhaps their most important criticism is that in finite samples the long-run effect of shocks is imprecisely estimated and the inferences regarding impulse responses are biased.

2. 3 Interval (sign) restrictions

The risk of imposing too disputable restrictions can be reduced by being less ambitious and letting parameters (response in certain periods, cross-correlations, etc.) lie in an interval instead of requiring them to take a certain value. This approach can be considered as a robustness check of identification trying to answer the implicit question: how stable are our results if we perturb the parameters of our assumption set? This was the original idea behind Faust's (1998) approach.
Some authors impose their restrictions on impulse responses. Faust (1998) considers only the immediate effect. Uhlig (2004) requires the restrictions to hold throughout a longer period of time. He also does a robustness check with respect to the length of that period.
Canova and De Nicolo (2002) adopt another approach. They first calculate dynamic cross-correlations of variables following a monetary policy shock with the help of a theoretical model. They then identify monetary policy shocks by demanding it reproduce the sign of those cross-correlations as much as possible.
Finally, the other strategy applied in earlier versions of Uhlig's 2004 paper is also worth mentioning. He gave room for his sign preference by minimising a loss function that penalises the deviation of impulse response from the restrictions continuously. In this way he ended up with an exactly identified system, but at the cost of constructing a penalty function that inevitably contains some arbitrariness.

2. 4 Restrictions on implied structural shock series

As mentioned in the previous subsection, identification can be based not only on responses of individual variables but also on, for example, cross-correlation functions, as Canova and De Nicolo (2002) did. Another plausible strategy is to focus on the history of shocks. One can make use of an additional information set in identifying the historical development of shocks. Using these estimates in the VAR, it is easy to plot impulse responses or to calculate other statistics related to those monetary policy shocks.

Romer and Romer (1989) apply a so-called narrative approach. They created a dummy variable that took the value of 1 in periods when the Fed was deemed to be excessively contractionary. The assessment was based on historical evidence, or more precisely, on their reading of Federal Reserve documents. They used the dummy in a univariate regression. The response to change in that variable was regarded as the effect of monetary policy shocks.

Rudebusch (1998) as well as Bagliano and Favero (1997, 1999) estimate historical monetary policy shocks from financial market data. They do this by comparing expectations reflected in futures or implied forward rates with the actual short-term interest rate one period later. They plug the difference into their VAR as an exogenous variable.

The Romers' dummy variable is subject to the criticism that it is not orthogonal to other important shocks and thus involves a mixture of monetary policy and other disturbances when interpreted as structural innovations, as Leeper (1996) points out. In CEE (1998) this problem is remedied by using VAR and giving room for other types of shock to appear implicitly in the residuals but orthogonally to the exogenous monetary policy shocks, similarly as in Bagliano and Favero (1997, 1999) or in Rudebusch (1998).

Sims (1996) criticised Rudebusch's approach by pointing out that identification based on shock series may be much less reliable than other strategies. His argument is that identification schemes producing similar impulse responses can produce quite different shock series due to omitting some variables from the policy rule part of the specification.

2. 5 The approach of this paper

Based on historical evidence of the nineties, there is a strong prior belief that risk premium (exchange rate) shocks played a predominant role in shaping Hungarian interest rate and exchange rate development. Thus it is necessary to have a model that can distinguish between two types of nominal shocks, which involves the inclusion of at least two financial variables. On the other hand, short time series constrain our possibilities to construct a model with many variables. To balance these requirements I chose a 4-variable VAR, adding nominal exchange rate to the minimal variable set of output, price level and short interest rate.

The 4-variable set-up and the supposed importance of both monetary policy and risk premium shocks make the identification difficult. The Magyar Nemzeti Bank (MNB) has always paid special attention to the exchange rate, due to its prominent role in monetary transmission mechanism. In the crawling narrow band regime it was the legal duty of MNB to keep the exchange rate within a ±2.25 cent neighbourhood of the continuously devaluated central parity. Even later, after widening the fluctuation band, the exchange rate remained an important device in coordinating expectations or, at least, in indicating the commitment of monetary policy to disinflation. Sometimes this was manifested in a very quick reaction of interest rate policy to considerable exchange rate movements, mostly due to sudden shifts in risk assessment of forint investments, i.e. risk premium shocks.

On the other hand, being an asset price with a relatively efficient market, the nominal exchange rate of the forint reacts immediately to unexpected shifts of monetary policy. Therefore, the recursive identification approach is not appropriate for an econometrician trying to isolate monetary policy shocks using recent Hungarian data. The simultaneity of the nominal variables with respect to both nominal shocks calls for some alternative identification scheme. The only exception when I applied contemporaneous zero restriction is the case of industrial production in the monthly dataset.

In Csermely and Vonnák (2002) we tried to separate monetary policy shocks from risk premium shocks by assuming that among all possible nominal shocks these two induce the largest immediate appreciation and depreciation of the exchange rate. We admitted that although the impulse responses to risk premium shock met our expectations, in the case of monetary policy the results were not convincing.

As a refinement of that paper's strategy, I assumed here that a contractionary monetary policy shock results in appreciation of nominal exchange rate, that is, I imposed a sign restriction on impulse response function. The same strategy was applied in Jarocinski (2004). If I had also pursued identifying risk premium shocks, I would have imposed a similar restriction but with the opposite sign on exchange rate.

In order to obtain credible results and to reduce the risk of identifying a mixture of several shocks instead of pure monetary policy shock, I also applied a completely different approach. Partly in the spirit of the "narrative approach" and of the Rudebusch-Bagliano-Favero type "identification based on financial market data" strategy, I identified monetary policy shocks by fixing the dates of the biggest unexpected monetary contraction and easing. Both episodes can be associated with an important, and at least in our sample, unique shift in monetary policy stance. I expect my strategy to gain special credibility from the fact that among economists familiar with the past decade of Hungarian economic policy there is not much debate about the two extreme points of monetary policy shocks. Note that in contrast with Romer and Romer (1989), Rudebusch (1998) and Bagliano and Favero (1997, 1999), my second identification is also an example of interval (or inequality) restrictions.

An important feature of this approach is worth mentioning. I identify only monetary policy shocks as Bernanke and Mihov (1996), Uhlig (2004), and Jarocinski (2004) did. In this way I am relieved of the duty of specifying all relevant shocks and searching for further credible identification assumptions. On the other hand, some monetary policy shock vectors may be inconsistent with an implicit structure of the unexplained part of the covariances. When a shock vector is accepted as a monetary policy shock, there is no check whether a reasonable and complete decomposition of VAR residuals could be achieved including that particular shock vector. I assess the costs of my approach to be much lower compared to the benefits from not identifying a full structure.

Later in this paper the near equivalence of both identification approaches is demonstrated[4]. A natural consequence would then be to combine these strategies and impose all restrictions simultaneously. However, I do not present results from a combined identification, since it would not alter the main conclusions.

[4] At least on this dataset.

3 Baseline estimation on Hungarian data

In this section I present the results from quarterly VAR estimated on the longest available data series. Although this specification is a natural starting point of the research, later I argue that we can obtain more appealing results from alternative specifications.

3. 1 Data and VAR specification

For the baseline estimate I used quarterly series of Hungarian data: logarithm of real GDP, CPI, nominal effective exchange rate and logarithm of 1+(3-month treasury bill yields)[5]. The frequency of the latter three was converted by taking the period average. An increase in exchange rate corresponds to depreciation. Since quarterly GDP data prior to 1995 is not provided by the Central Statistical Office, estimates of Várpalotai (2003) were used for that period. The series cover the period Q2:1992-Q4:2003. GDP and CPI are seasonally adjusted.

Following several authors – e.g. Uhlig (2004), Peersman and Smets (2001) – I estimated the VAR in levels. The reader interested in the debate surrounding the question how to make inference and how to interpret results when the data is likely to contain some unit roots should refer to Sims (1988), Sims and Uhlig (1991), Phillips (1991) and Uhlig (1994), among others. Following Uhlig (2004), I make inference in a Bayesian manner and interpret results using Bayesian terminology; thus the difficulties which arise when attempting to construct classical confidence bands in the presence of near unit root regressors can be avoided.

Three lags were enough to produce unautocorrelated residuals, based on the evidence of the multivariate LM-test. The Akaike information criterion also suggested 3-lags specification. An intercept was also included in the VAR.

3. 2 Estimation and inference

The estimation procedure applied here and the presentation of the results is almost the same as in Uhlig (2004) with the exception of the case when monetary policy shocks were identified by imposing restrictions on shock series.

First, the coefficients and the covariance matrix of the residuals were estimated by OLS. I then used Normal-Wishart prior distribution parameterised by the VAR's coefficient and covariance matrices. As shown in Uhlig (1994), the posterior distribution will then also be Normal-Wishart. My approach differs from Uhlig (2004) in that I excluded the possibility of explosive dynamics by truncating the posterior.[6]

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[5] If, for example, the annual yield is 8 per cent, the corresponding data point is ln(1.08).

[6] Technically it was carried out by calculating the largest eigenvalue for each random draw from the VAR posterior. If the modulus was not greater than one I proceeded with that draw, otherwise I dropped it. As far as the number of draws is concerned, this truncation seemed to be effective in the sense that the procedure was often repeated because of too large eigenvalues. Interestingly, this did not influence the shape of median impulse responses much, with the only exception of price level: excluding explosive roots decreased the relative frequency of 'price puzzle' type responses in the set of all possible responses. Not surprisingly, the posterior distributions became more focused on the median value, allowing for 'more significant' results.

For each draw from the VAR posterior I randomly chose a candidate monetary policy shock, which is in the form of a 4x1 vector comprising the immediate effect on the variables. Depending on where to impose identifying restrictions, I calculated the relevant impulse responses or the shock series implied by the particular shock vector. If the impulse responses or shock series met the expectations, the draw was kept, otherwise dropped. This procedure corresponds to having an implicit flat prior on the part of the 4-dimensional unit sphere that contains 'credible' monetary policy shocks and represents our identification scheme.

As a consequence, if we interpret this procedure as a Bayesian estimation, our prior is formulated on the parameter space consisting of the subspaces of VAR coefficients, covariances and monetary policy shock vectors. As Uhlig (2004) points out, our procedure is a re-estimation of the VAR model, since, depending on how many draws from the 'monetary policy shock space' satisfy the identifying conditions, some parts of VAR coefficient prior will be overrepresented while others underrepresented.

The quantiles of posterior distributions for impulse responses and other outputs reported in the Appendix are calculated from the set of successful draws that in each case contained more than 2000 elements[7].

3. 3 Impulse responses from sign restriction approach

In the first experiment I identified monetary policy shocks by imposing restrictions on the sign of impulse responses. In particular, it was assumed that an unanticipated monetary policy tightening results in more appreciated exchange rate (negative response) and higher interest rate (positive response). I chose the length of the restriction to be 4 periods, but all the results are robust to changes in the length of restriction. This identification scheme is similar to that of Jarocinski (2004) with the exception that I did not restrict the immediate output response to be zero.[8]

Whereas a monetary policy shock should behave as we prescribed, it is not clear how we can exclude other sources of disturbances which produce the same initial responses. The answer is that we can never be sure. The same applies, however, to other identification strategies irrespective of whether our prior belief is formulated as point or interval restrictions. Researchers using the SVAR approach usually assume that the number of endogenous variables equals the number of relevant shocks. In addition, the looseness of interval restrictions (in other words, the underidentification) can make this problem more serious and the resulting picture more blurred – relative to an exactly (or over-) identified system with point restrictions. Nevertheless, this is the price we have to pay for greater credibility of our identification. In this way we end up with less significant results, but all those results that are significant will have more convincing power.

Figure 2 shows the resulting impulse responses with the error bands created as quantiles of the posterior distributions for each period. The shape of the consumer price level response suggests that we

[7] As Sims and Zha (1998) demonstrate, this reporting technique is not optimal, since it may convey a misleading picture about the shape uncertainty of impulse response. In order to understand their intuition, it may be enough here to note that the median impulse response plotted as thick line is the interlacement of points of different impulse responses, and usually is not a plausible impulse response itself. As Figure 1 suggests, a presentation of shape uncertainty in the spirit of the above-mentioned paper might be useful here.

[8] This is true only in the quarterly series case. When using monthly data I used that restriction – see the next subsection.

probably mixed too many types of shocks under the label 'monetary policy'. The quite significant[9] increase 3 after the shock is the well-documented price puzzle. The usual interpretation is that another shock is identified as monetary policy shock, namely a shock to future inflation – see, for example, Sims(1992). This is anticipated by the monetary policy-maker, who therefore tightens monetary conditions. The usual remedy to this problem is to include some variables in the VAR that play the role of leading indicators of inflation, typically commodity prices – see Sims (1992) and CEE (1998). Uhlig (2004) excludes this puzzle by using the condition of negative price response to contractionary monetary policy shock as an identifying restriction. His estimation focuses on the response of output, therefore his approach could be justified[10]. In our case, however, it is the response of prices, among others, we are interested in, and thus it would not be appropriate to impose restrictions on price level impulse response.

The responses of the interest rate and the exchange rate help us to imagine the size of the shock. The 3-month TB-yield increases by 60 basis points immediately, while the nominal exchange rate appreciates by almost 0.7 per cent. Note that since we restricted the sign of both impulse responses for the first four quarters, the entire posterior distribution is above and below zero for nominal interest and exchange rate, respectively.

The output responds moderately to unanticipated monetary tightening. The immediate effect is virtually zero. This observation suggests that we indeed identified a nominal shock. The level of output declines gradually and reaches its minimum in the third year after the shock. The size of the decrease at its bottom is not particularly huge, 0.2 per cent, but it is worth noting that roughly 95 per cent of the posterior distribution is below zero during the third year; thus we can consider this effect as significant.

An interesting feature of our results is the relative sharpness of the real exchange rate impulse response. The width of the middle two-thirds of the posterior distribution is only 0.3-0.4 of a percentage point in the fifth year after the shock, which is three times wider in the case of nominal exchange rate. This is due to the fact that identification uncertainty is highly correlated between prices and nominal exchange rate. If, for example, some plausible (i.e. meeting identifying criteria) monetary policy shock vector generates rising prices after a contractionary shock (price puzzle), it is likely to generate a more depreciated nominal exchange rate for the same period.[11] Put differently, our data and identifying assumptions have very stable consequences regarding the response of the real exchange rate to monetary policy shocks, but not regarding the price level and nominal exchange rate.[12]

In the next subsection we compare these results to those obtained from an alternative identification strategy.

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[9] The word 'significant' may be a bit misleading here. Since we apply Bayesian inference philosophy, the probability coverage terminology is more appropriate. The right interpretation is that 'with probability x the response is above (or below) zero'. In our case we can claim that with probability more than 84% the response of the price level is positive 3–6 years after the shock, conditioned on the data.

[10] Nevertheless, see footnote 2 regarding this issue.

[11] This fact has the consequence that we could have mitigated the undesired price puzzle by lengthening the restriction period of nominal exchange rate. However, my identification philosophy was to impose explicitly all features about which we have firm prior belief. Following this logic the only legitimate way to fight against price puzzle would have been to require a negative price level response to monetary tightening.

[12] The same will be true for all the other estimation strategies to be introduced later in this paper with the exception of shorter sample experiences.

3. 4 Impulse responses from restrictions on shock history

As advocated in section 2, identifying restrictions imposed on implied shock history may sometimes have a communication advantage over restrictions on impulse responses. In Hungary during the past 10 years one of the largest monetary loosening was the austerity package of financial minister Bokros, which contained a surprise depreciation of the forint in order to balance the government budget and the current account in March 1995. On the other hand, the widening of the narrow exchange rate band in May 2001 and the following appreciation surprised the market into the opposite direction. In both episodes monetary policy deviated to a considerable extent from its earlier behaviour, and I base my identification strategy on that fact.[13]

I assumed therefore that between 1995 and 2003 the largest unexpected monetary loosening occurred in Q1:95, while during the same period the band widening in 2001 was the largest contractionary monetary policy shock.[14] While specifying the date of the latter step is straightforward, it is not clear in which quarter it was effective. Although the change of the exchange rate regime took place in May, the appreciation continued in the third quarter as well. It is therefore more reasonable to formulate the restriction as ‘the bigger shock of the two relevant quarters should be at the same time the biggest between 1995 and 2003’. One can argue that the tightening shock itself was the widening of the fluctuation band of the Forint. It is important, however, to emphasise that we are trying to locate monetary policy shocks using exchange rate and interest rate data, and the figures show even more substantial appreciation from Q2:2001 to Q3:2001 than between the first two quarters, with the central bank not trying to dampen it by lowering short-term interest rates.

The method of estimation is quite similar to that of the previous strategy. For each joint random draw from VAR-posterior and the unit sphere of possible shock vectors I calculated the historical shock series, and the draws not meeting the restrictions described above were rejected. The posterior distributions were constructed from the successful draws. The main results are summarised in Figure 3.

The most striking feature of these charts is their similarity to previous ones. This observation indicates that the two sets of identification restrictions, namely those imposed on impulse responses and those imposed on shock history, are nearly equivalent. In other words, the implied history of monetary shocks identified by impulse responses of interest rate and exchange rate typically correspond to our prior belief about when the biggest contractionary and expansionary monetary policy surprises took place during the past nine years. On the other hand, fixing the extreme points of implied history produces impulse responses that are typically in accordance with our intuition regarding the behaviour of the nominal interest and exchange rate in the aftermath of a monetary policy shock.

There are, however, differences as well. While the response of output is almost the same in both cases, identification based on historical evidence seems to dampen the price puzzle. The response of the price level to monetary contraction has an appealing sign during the first seven quarters, although later

[13] One can argue that during those episodes the policy rule changed itself. Since my model does not deal with structural breaks that affect the reduced form or the identification scheme, it seems to be a good approximation to look at those policy actions as extreme shocks.

[14] Note that those shocks are not expected to be the largest on the full sample, i.e. from 1992.

it rises above zero. This may be related to the bigger appreciation of nominal exchange rate after the monetary shock.

Considering the intuition behind identification through shock history, it is not surprising that we have more chance of eliminating such shocks as 'future inflation shocks' suspected to be responsible for the price puzzle. Using historical evidence we force impulse responses (or other statistics we are interested in) to be close to the effect of disturbances in certain periods. In particular, we located the two extreme points of implied monetary policy shock history (the biggest tightening and easing), thus our impulse responses will be similar to the effect of the Bokros-depreciation and (with opposite sign) of the band widening. Since we identify on time domain rather than on the space of impulse responses, we can avoid mixing up shocks that have similar effects. If we are really convinced that these two periods were dominated by monetary policy shocks and we can rule out that at the same time a pair of another type of shocks occurred in both periods with both signs, we can get rid of all pseudo monetary policy disturbances that trigger plausible responses but which do not originate in monetary policy.

On the other hand, the impulse response of nominal interest rate is a bit annoying in the second identification approach. The quick correction after the shock and then the second increase are difficult to interpret. This is due to a special type of impulse response that occurred quite often during random search for plausible monetary policy shock vectors. According to this 'alternative' rate scenario, the initial response is a decline in the short interest rate followed by a gradual increase above zero. The high probability of generating such response from random draws influences the posterior distribution, especially in the first two years.

In Figure 4 only the median impulse responses are plotted, allowing for a more convenient way of comparison of different identification approaches. However, it should be noted that all the differences can be considered as insignificant in the following sense: in each case the middle two-thirds of impulse response posterior distribution (thin lines on Figures 2-3) contains the median impulse response obtained from the alternative strategy.

4 Robustness check: alternative estimates

As a check of robustness I also estimated a 4-variable VAR model on monthly data. Since on our sample the model is likely to contain structural breaks, I re-estimated the monthly model on a shorter sample beginning in 1995. As is demonstrated below, switching to the monthly model does not change the picture significantly. On the other hand, impulse responses estimated on the shorter sample are quite different from full sample results, and resemble those obtained for developed countries in the literature.

4. 1 Estimation on monthly data

The observations of the monthly model range from M1:1992 to M3:2004. CPI, nominal interest rate and exchange rate series are from the same sources as in the quarterly model. Real GDP was replaced by constant price industrial production, which is available at monthly frequency. I used the seasonally adjusted series corrected for calendar effects produced in MNB.

Lag length of 2 was suggested by most information criteria. Two lags eliminated the bulk of autocorrelation of residuals. The LM-test still detected significant autocorrelation at lag 5, but inclusion of more lags did not help with this problem. I therefore used 2 lags.

Because of the higher frequency, I assumed in both identification strategies that monetary policy influences the output only with lags. While sign restrictions could be imposed on impulse responses in an analogous way to the quarterly case (I chose the length of constrained period to be 12 months, which corresponds to the 4 quarters of our previous exercise), locating the most contractionary and most easing monetary policy shocks in time may require some justification. The Bokros loosening in 1995 is likely to have had its maximum magnitude in March, reflected in a roughly 6% depreciation of the Forint. The contractionary effect of exchange rate band widening in 2001 appeared most sharply during May and June based on exchange rate data. The monthly appreciation rates were roughly 3% and 4% respectively. This seems to contradict the quarterly identification strategy, since we expected the maximum tightening to appear in either the 2nd or the 3rd quarter. This contradiction, however, is of purely technical nature; it is a consequence of taking period averages.

The results are quite similar to those of the quarterly model, as Figure 6 demonstrates (results from restricted impulse response approach are not reported). The most important differences are the faster (but still moderate) response of output and the smoother path of interest rate in the monthly model, but all differences are small compared to the sampling and identification uncertainty.

4. 2 Estimation on subsample

Finally, I estimated the monthly model on a shorter sample. The 12 years of previous estimation are supposed to be full of regime changes. These structural breaks may have blurred the picture we obtained from full sample estimation. Shortening the period under investigation may produce sharper results.

Among the most important changes were the announcements of two systematic monetary policy regimes. In the beginning of 1995 a crawling narrow band exchange rate system was introduced. The central bank announced the changing devaluation rate of the exchange rate band in advance. In 2001 the fluctuation band was widened and an inflation targeting framework replaced the previous regime. Both dates can be considered as significant turning points in preferences of monetary policy and in its behaviour .

The results seem to confirm that there was indeed an important structural break during the first half of the nineties, and it might have been the regime change of monetary policy. Despite the smaller sample, the posterior distribution became more concentrated around the median (compare Figure 7 with Figure 5), especially in the case of price level, nominal (and real) interest rate and nominal exchange rate. As mentioned earlier, the uncertainties in price level and nominal exchange rate behaviour were correlated, and could be attributed to identification uncertainty. If we restrict our dataset to contain only observations from 1995 on, identification of monetary policy shock became much easier in the sense that only a small set of possible shock vectors met the identifying restrictions. This finding is reinforced by the rather technical experience that random search produced more rarely plausible 'monetary policy vectors' than in the full sample case.

Moreover, there are spectacular differences regarding the point estimates, too.[15] From the point of view of monetary transmission mechanism, the most important change is perhaps the reaction of price level. The immediate response to monetary tightening of typical magnitude is virtually zero, and it starts to decline at the end of the first year. The pace of the decrease is very slow, the greatest effect (0.1-0.15%, depending on identification strategy) can be observed during the fourth-fifth years after the shock. This is in sharp contrast with full sample estimates, where an initial drop in prices was followed by a rise above zero, even if 'history restrictions' were imposed. Due to the fact that the latter phenomenon occurred irrespective of identification scheme and data frequency, we can attribute the bulk of price puzzle to the data prior to 1995.

The behaviour of the nominal interest rate and exchange rate is of great importance, too. While on full sample one standard deviation monetary policy contraction resulted in a permanently (for 2-3 years) 30-40 basis points higher short rate, since 1995 a typical monetary tightening appears in the form of 20-30 basis points higher short interest rate that quickly declines. One year after the shock the distance from baseline path is only less than 10 basis points. On the other hand, this more moderate interest rate policy has virtually the same immediate effect on the nominal exchange rate: a 1% appreciation within a few months, just like in the full sample case. In contrast to the full sample case, the return to the baseline is more gradual and the nominal exchange rate never becomes weaker than in the baseline. We can interpret this result as monetary policy became more effective after 1995 in influencing the nominal exchange rate. This is probably due to the nature of the monetary regimes after 1995. In the crawling peg regime, the pre-announced devaluation rate of the narrow fluctuation band was generally credible. In the inflation targeting regime the inflation forecast was conditioned on the nominal exchange rate as a policy variable, therefore market participants had a quite clear picture about the 'desired' future development of the HUF/EUR. The improvement in efficiency, therefore, can be attributed to the more efficient orientation of exchange rate expectations.

[15] For a convenient comparison see Figure 8.

The response of the output seems to be the most robust result across identification and sample choices. Although the short sample with history restriction produced the less smooth decline (it drops immediately to the minimum value of -0.25%), the size of the recession and the beginning of the recovery is roughly the same in all cases: the level of output decreases by roughly 0.3% within the first two years after the shock and starts to increase at the end of the third year.

Together with the price level response, this behaviour exhibits the main characteristics of sticky-price models. Because of the slow adjustment of prices, it is the output that reacts first to the contraction. The price adjustment is coupled with the gradual return of output to its natural level. This pattern is in accordance with survey results, too. Based on a survey among Hungarian companies conducted in 2001, Tóth (2004) concludes that before changing their prices, Hungarian firms typically try to meet shifts in demand by first changing their output.

It is worth noting that the difference between estimates on different time span is much bigger than the difference caused by switching to the alternative identification strategy. On data starting in 1995, both restriction sets produced almost the same picture that fits the typical findings in the literature. We can conclude, therefore, that our identification strategies are a good characterisation of monetary policy shocks, and this becomes obvious when they are applied on a relatively homogenous sample.

5 Conclusions

The purpose of this paper was to estimate the dynamic effect of monetary policy on several variables, in particular on output and consumer prices using Hungarian data. Due to possible data problems and the supposed existence of structural changes, two variable sets were used, one of them on two different, but nested samples. Due to doubts regarding the applicability of widely used identification approaches, in particular zero restrictions, sign restrictions were imposed on impulse responses. In order to obtain more credibility, an alternative identification scheme was also proposed. The latter tried to capture the main features of a monetary policy shock by using historical evidence of some periods when monetary policy is known to have surprised market participants.

Although the results are weak in the sense that even the middle two-thirds of the distributions of possible impulse responses contain zero in most cases, the robustness of the point estimates to the identification strategy on the one hand, and the coincidence of the shorter sample estimates with the results of the literature on the other, allow a few firm conclusions to be drawn.

All of our estimates produced the result that one standard deviation unanticipated monetary contraction results in 1% quick nominal appreciation and 0.3% reduction in output. The latter starts to recover after three years. Although the real exchange rate appreciates quite significantly in the first 1-2 years, it returns to its equilibrium after 3-4 years.

Comparing results across different estimates, we can conclude that it is more feasible to estimate the effect of Hungarian monetary policy on data starting in 1995, as long as we do not believe that monetary contraction can cause rising prices one year later.[16] Excluding observations prior to 1995 also has the advantage of obtaining sharper results. The shape of the impulse responses obtained on short sample are quite similar to those which can be found in the literature. They can also be reconciled with the predictions of sticky-price models. Based on these estimates, a typical unanticipated monetary policy contraction amounted to a roughly 25 basis points rate hike and resulted in a quick 1 per cent appreciation of nominal exchange rate during the past 9-10 years. This was followed by 0.3% lower output and 0.1-0.15% lower consumer prices. The impact on prices was slower than on output, it typically reached its minimum only 4-6 years after the shock.

As far as our identification strategies are concerned, the difference between the two was minor. Imposing restrictions on history may help to exclude some puzzles stemming from a too loose identification of other strategies, but in our case the sampling error suppressed possible improvements. In my view, however, it may add to the credibility of the other identification scheme and to the reliability of the results.

As far as possible improvement of the estimates is concerned, the sampling uncertainty seems to be a binding constraint. The data is given, the sample cannot be extended backwards. Short sample estimates revealed that even the observations prior to 1995 provide very noisy information about the underlying relationships. Including more variables in the VAR would lessen the degrees of freedom considerably.

[16] Nonetheless, this way of choosing the best specification is still subject to the criticism outlined in subsection 2.1.

On the other hand, reducing the uncertainty stemming from my cautious approach to identification is possible, at least in theory. Identifying more periods when something is known about the direction of monetary policy surprises may produce narrower error bands. In practice, however, after having identified the biggest historical surprises, there remained not much dispersion in implied monetary shock history; therefore exclusion of substantial amount of shock vectors based on history may not be carried out with high credibility.

In the case of restrictions on impulse responses, much improvement could not be achieved, unless we are willing to sacrifice some part of the convincing power of our assumptions. Lengthening the number of periods throughout which sign restrictions are imposed would inevitably arouse the suspicion of arbitrariness. Imposing additional restrictions on variables' reaction we are particularly interested in (price, output) would make the interpretation of the results difficult. In the case of the nominal exchange rate and short interest rate this problem is not so serious, since their reactions are at the very beginning of the monetary transmission's causality chain. Therefore we can have firmer prior belief about their behaviour, especially regarding the first few periods.

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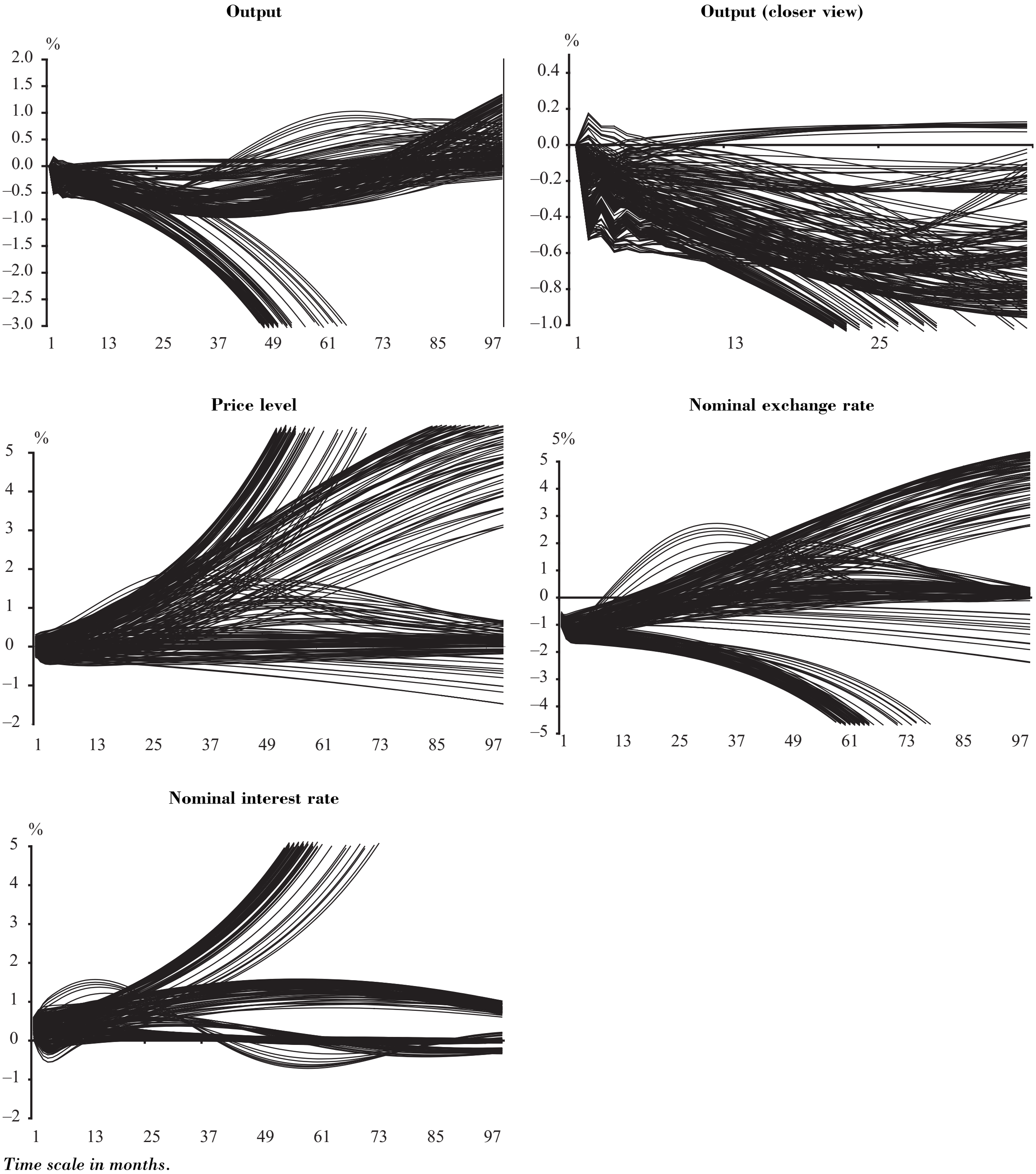
Appendix: Figures

Figure 1

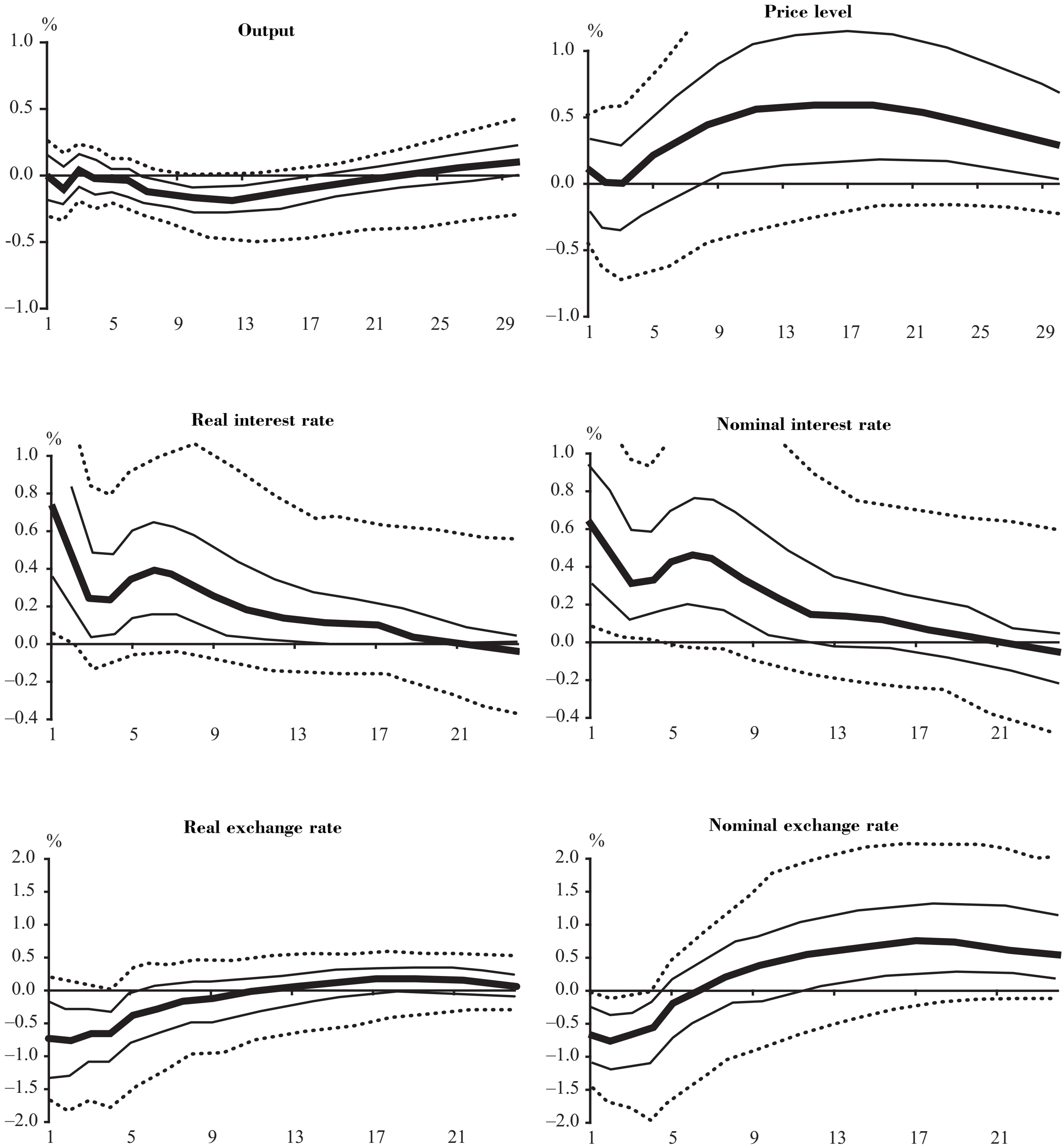
Examples of the effect of sampling and identification uncertainty: impulse responses to plausible monetary policy shocks (estimates on monthly data from 1992 allowing for explosive roots)



Time scale in months.

Figure 2

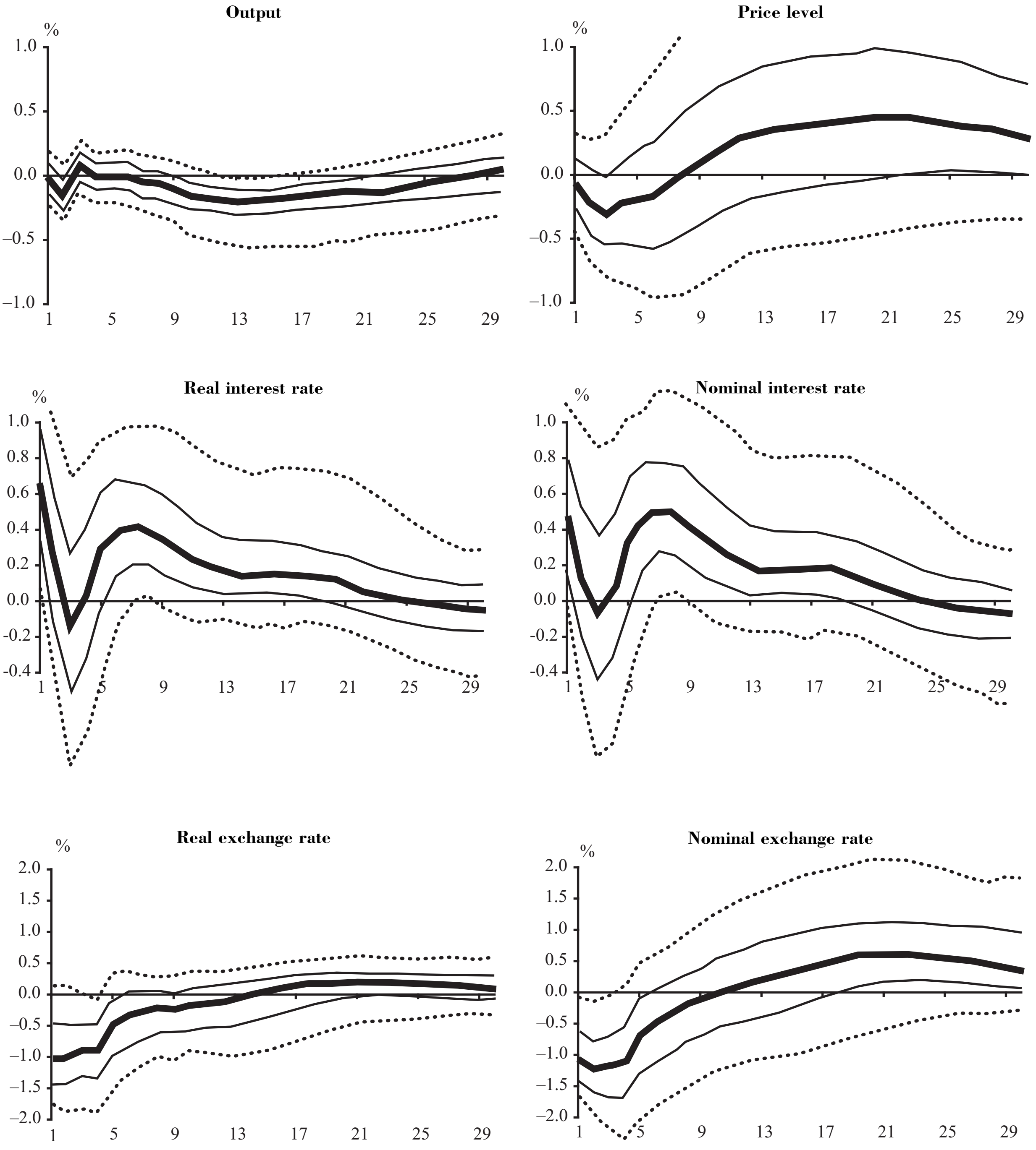
Impulse responses to a one standard deviation monetary policy shock; posterior distributions from the sign restriction approach (full sample)



Time scale in quarters. The middle 95.4% (± 2 st. dev. for normal distribution) of the distribution ranges between the dotted lines, the 68% (± 1 st. dev. for normal distribution) between solid lines. The thick line connects median values for each period.

Figure 3

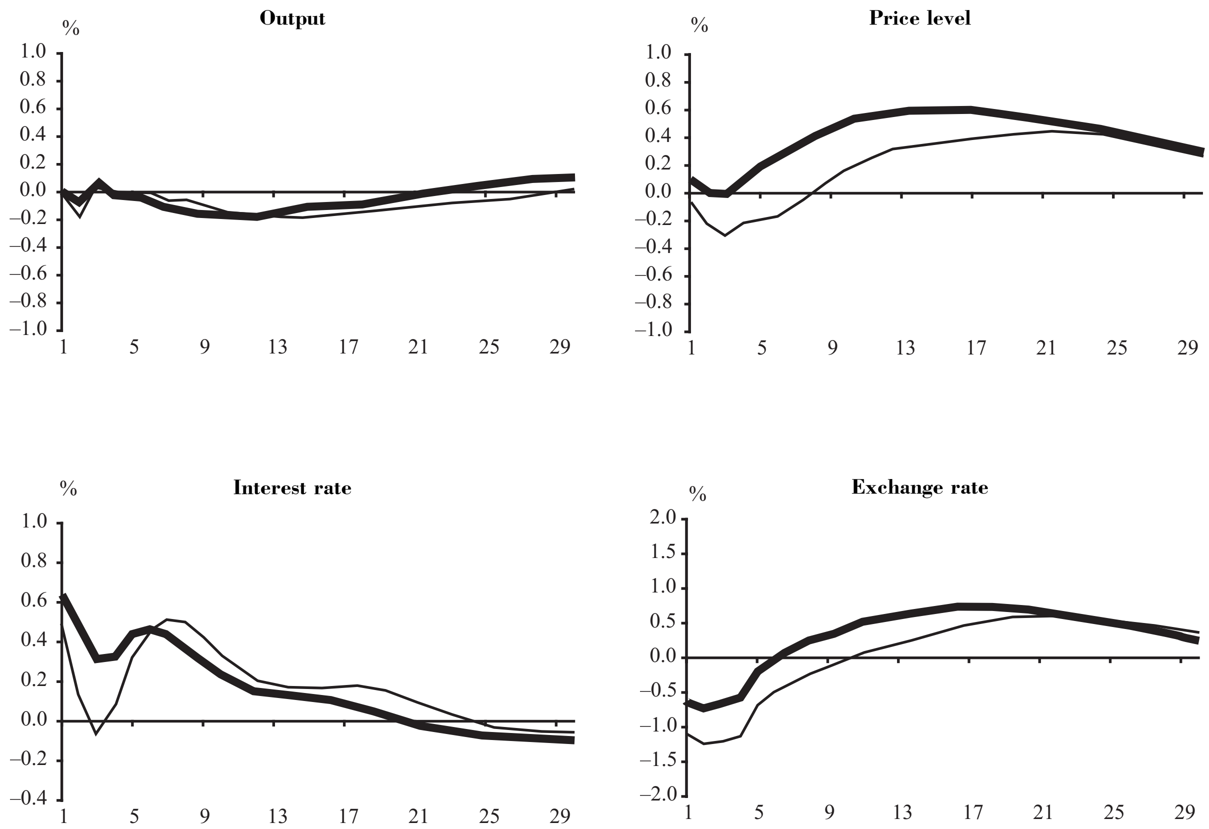
Impulse responses to a one standard deviation monetary policy shock; posterior distributions from the 'history restriction' approach (full sample)



Time scale in quarters. The middle 95.4% (± 2 st. dev. for normal distribution) of the distribution ranges between the dotted lines, the 68% (± 1 st. dev. for normal distribution) between solid lines. The thick line connects median values for each period.

Figure 4

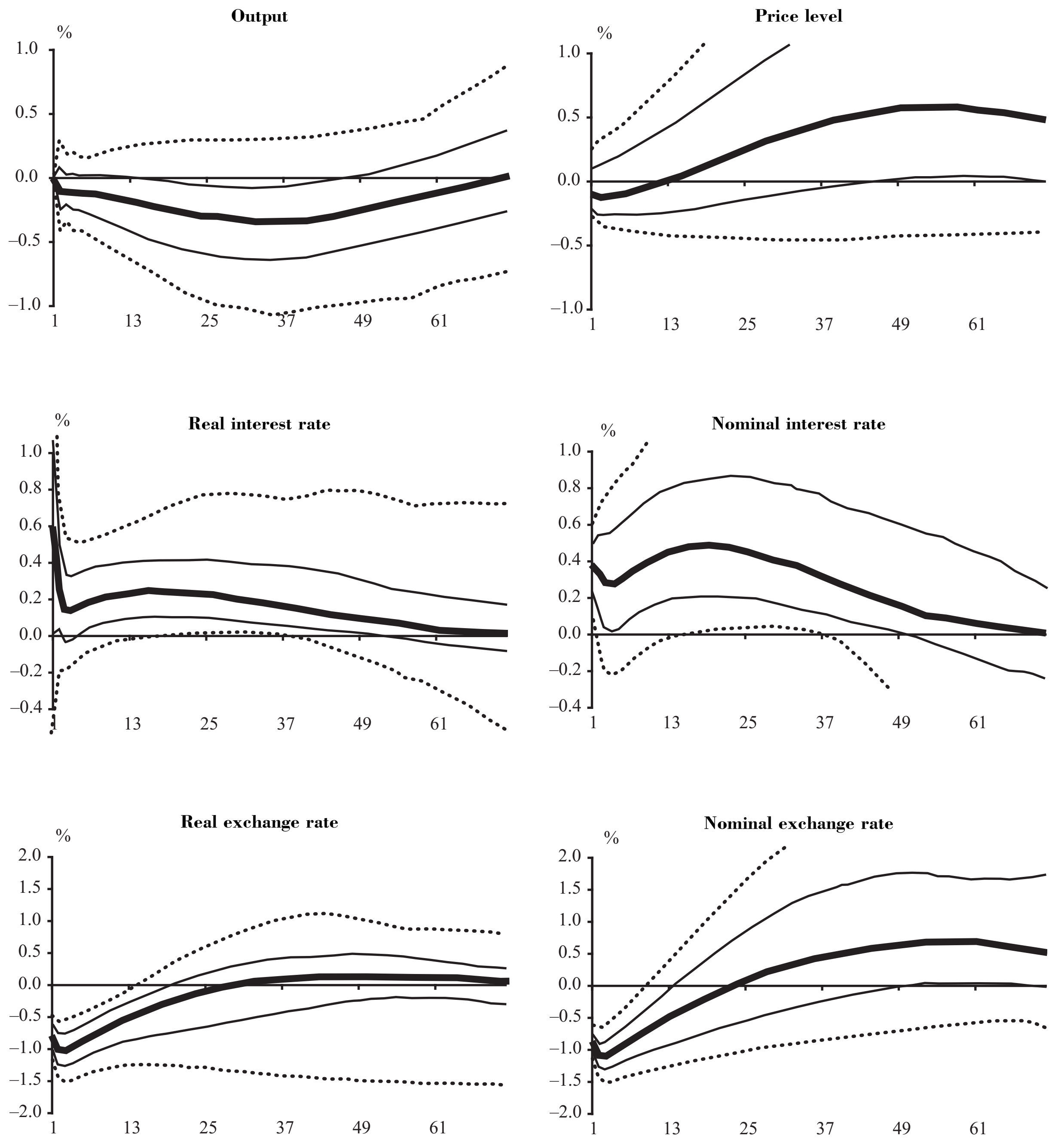
Comparison of impulse responses from competing identification approaches (estimates on quarterly data, full sample)



Time scale in quarters. Thick line: restrictions imposed on impulse response functions. Thin line: restrictions imposed on implied shock series.

Figure 5

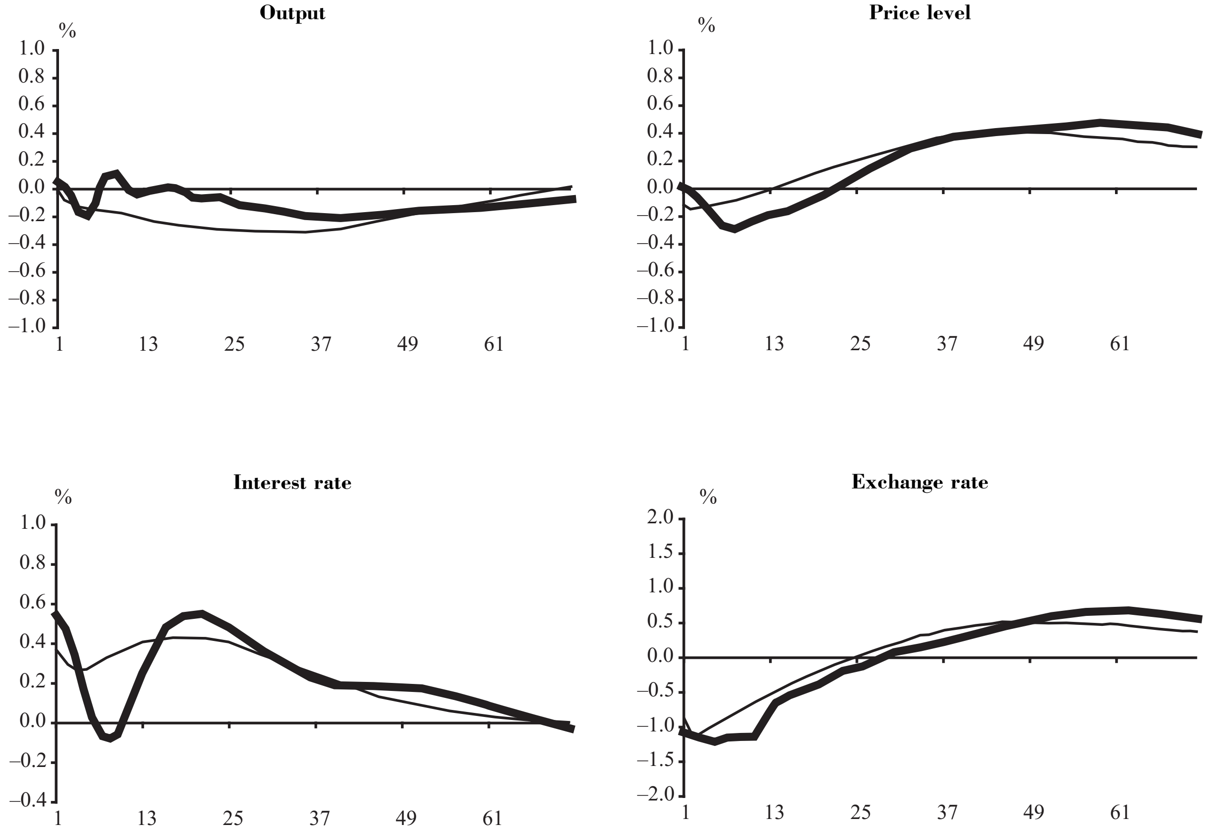
Impulse responses estimated on monthly data; posterior distributions from the 'history restriction' approach (full sample)



Time scale in months. The middle 95.4 % (± 2 st. dev. for normal distribution) of the distribution ranges between the dotted lines, the 68 % (± 1 st. dev. for normal distribution) between solid lines. The thick line connects median values for each period.

Figure 6

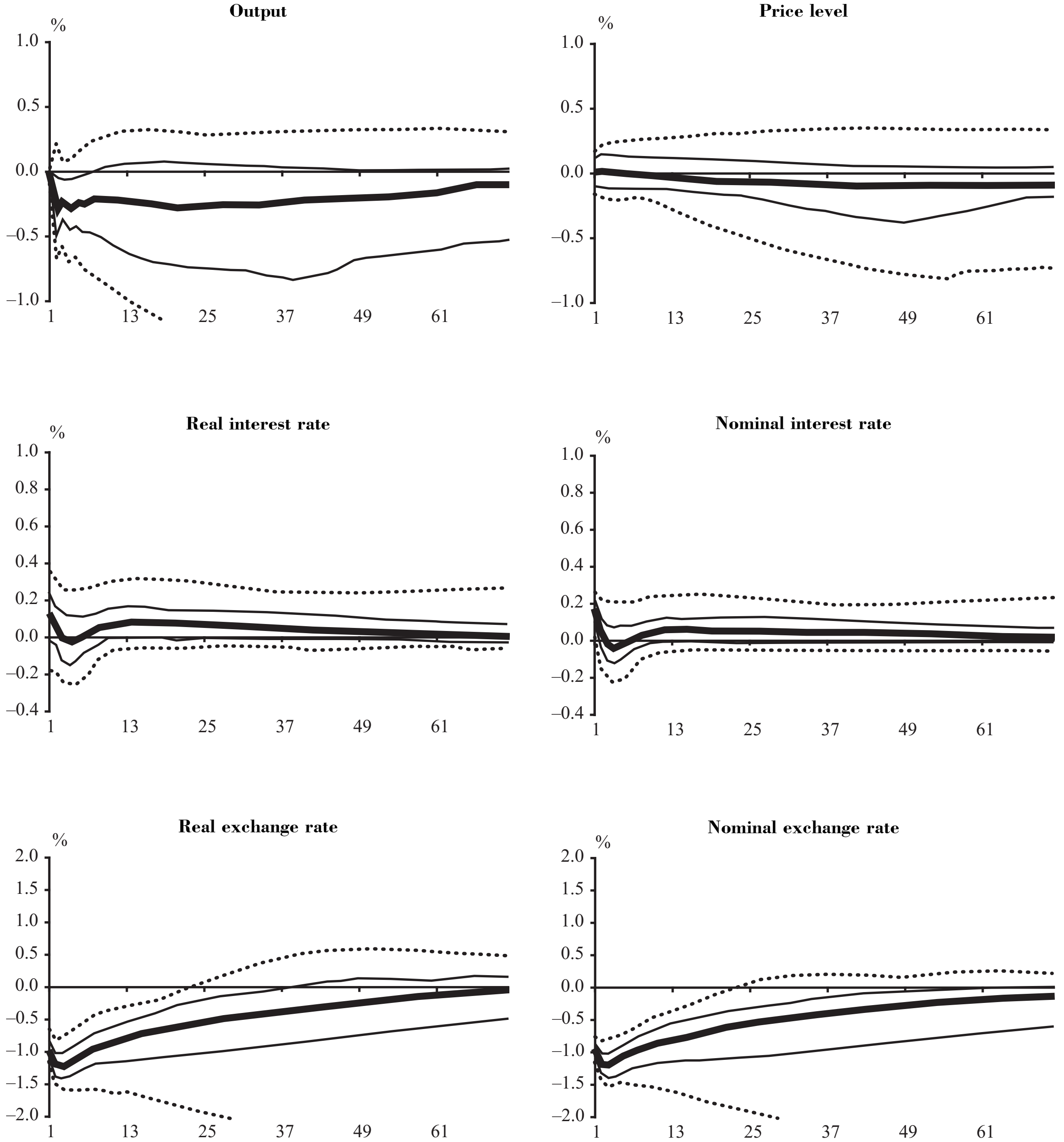
Comparison of impulse responses from quarterly and monthly models; identifying restrictions on shock history (full sample)



Time scale in months. Thick line: quarterly model. Thin line: monthly model. Impulse responses of the quarterly model were converted to monthly frequency by interpolation preserving quarterly averages and achieving maximum smoothness.

Figure 7

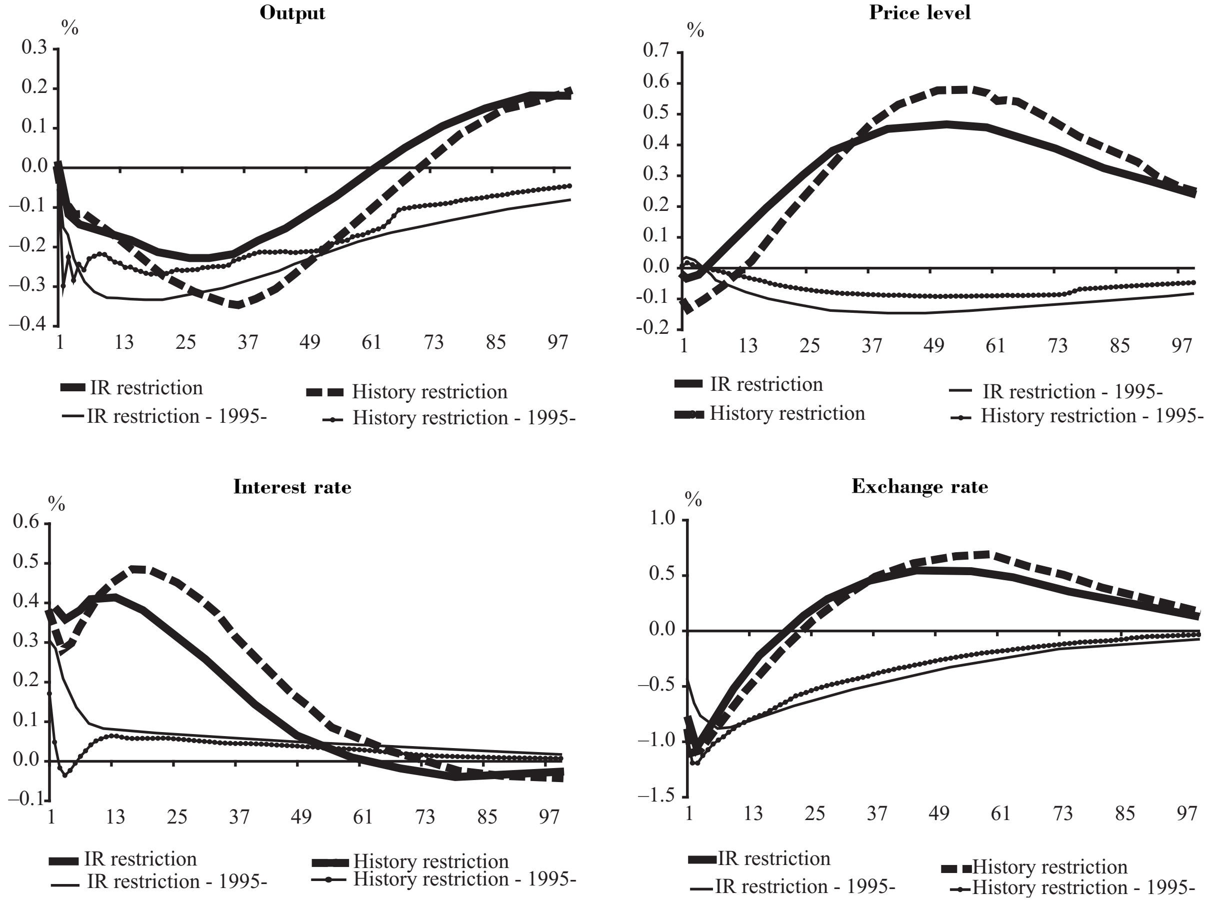
Impulse responses estimated on shorter sample using monthly data; posterior distributions from the 'history restriction' approach



Time scale in months. The middle 95.4% (± 2 st. dev. for normal distribution) of the distribution ranges between the dotted lines, the 68% (± 1 st. dev. for normal distribution) between solid lines. The thick line connects median values for each period.

Figure 8

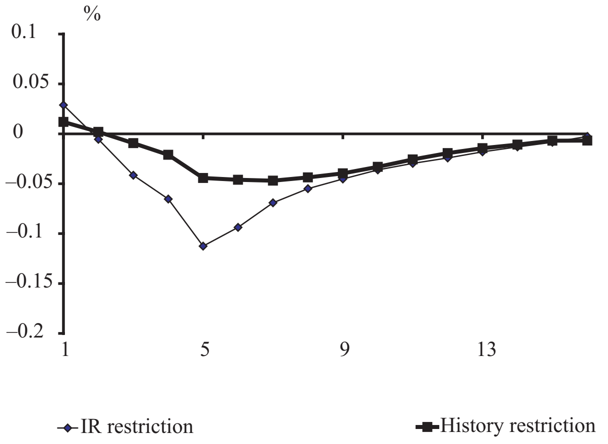
Comparison of impulse responses from different samples; monthly data, identifying restrictions on impulse responses (IR) and on shock history



Time scale in months.

Figure 9

Impulse response of 12-month consumer price inflation to a monetary shock, which corresponds to a roughly 25 basis points rate hike coupled with a 1 per cent appreciation of the nominal exchange rate (derived from the response of price level, short sample estimates with monthly data)



Time scale in quarters. Percentage point deviation from baseline scenario.

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