

**Differences in Monetary Policy Transmission?  
A Case not Closed**

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## 1 Introduction

The approach of Economic and Monetary Union in Europe has inspired a large number of studies seeking to uncover possible differences in the impact of monetary policy changes on output and inflation among European countries. Important differences in the impact of the single monetary policy would be a potential source of cyclical divergence and could impose significant adjustment demands on other economic policies.

Hitherto, academic studies concerned with the pooling of monetary sovereignty in EMU have tended to focus more on the likely incidence of asymmetric shocks and adjustment mechanisms to such shocks. In comparison, differences in the transmission of the common monetary policy seemed a relatively minor issue. It is only recently, and to some extent in the aftermath of the comprehensive study on financial structures and the monetary transmission mechanism in BIS (1995), that the potential significance of differences in monetary transmission has become a popular topic of analysis, cf. e.g. Barran, Coudert and Mojon (1996), Britton and Whitley (1997), Ramaswamy and Sloek (1998), and Dornbusch, Favero and Giavazzi (1998).

Heuristically, the case for asymmetries in monetary policy transmission is easy to make. Such asymmetries are usually seen to stem from cross-country differences in financial structures, for example in terms of the development of credit markets and the shares of fixed and variable-rate borrowing in mortgage, company and government financing. The claim one encounters most frequently is that a given change in short-term interest rates may have larger effects in an economy like the UK's with a heavily indebted household sector and variable mortgage rates than in some continental European economies where fixed-rate financing is widespread and the response of bank lending rates tend to be sluggish. Another widespread assertion is that monetary policy would have a relatively smaller impact on spending and output in Italy, where the household sector holds large variable-rate claims on the government, since changes in interest rates might be accompanied by an off-setting fiscal impact on disposable income.

However, knowledge of the differences in financial structure does not translate easily into robust conclusions about the likely impact of monetary policy. This is partly because, at the theoretical level, there is typically no clear-cut relationship between such features and the power of monetary policy. For instance, sluggish adjustment of bank lending rates may help shield companies from interest rate movements, but banks may apply non-price rationing of loans, thus amplifying the effects of monetary policy through the so-called "credit channel"<sup>1</sup>. The distinction between fixed rate or variable-rate financing may not matter for the substitution between current and future consumption except in the presence of liquidity constraints, and the effects of interest rate changes will depend on the financial positions as well as the marginal propensities to spend for both borrowers and lenders. At an empirical level, the

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<sup>1</sup> On the credit channel of monetary policy, see e.g. Bernanke and Gertler (1995). For an empirical perspective, see Davis (1995)

different particularities of a given country in terms of financial structure may have off-setting effects on the power of the monetary transmission.

The analysis of the potential role of differences in monetary policy transmission in EMU is further complicated by the fact that transmission mechanisms are likely to change in the new regime as countries become subject to a common policy and regulatory environment. In this paper, we will concentrate on the issue of correctly identifying monetary policy actions and their causal effect on the economy in the current (or more precisely, historical) set-up, although we believe that the question of gauging the extent to which these differences will disappear with EMU is of at least equal importance.

We start in the following section by reviewing the existing empirical evidence on differences in monetary transmission between Member States. In our view, the econometric evidence does not provide a coherent picture of such differences, much less of a ranking of how monetary policy has impacted historically in the four major EU economies. We then go on to focus on a particular aspect which may underlie the differences in results but which typically receives little attention in this part of the literature, namely the problem in identifying the causal effect of monetary policy. Movements of the economy following a change in the stance of monetary policy may be due to the policy action itself or they may be related to the underlying causes that spurred that action. In practical terms, the question is when to interpret a co-movement of a monetary policy variable and another variable as causality, and when to interpret it as a response of two endogenous variables to common causes. In examining the scope of the identification problem, we use the structural VAR framework, which allows a very general approach while remaining computationally manageable. Our results show that the way structural identification of monetary policy is handled can have very substantial effects on the results. According to our estimations, once the uncertainty involved in the structural identification is accounted for, no statistically significant differences in monetary transmission can be found for a group of three large EU countries (Germany, France and the United Kingdom).

## **2 Evidence on the strength of the monetary transmission**

The existing empirical evidence on differences in the impact of monetary policy on output and prices in European countries has been reviewed in Britton and Whitley (1997) and in Dornbusch, Favero and Giavazzi (1998). The summary provided here focuses on the four major EU economies. We review the main results in the perspective of a brief recap of the principal methodological differences and potential shortcomings of various methods.

Since the evidence refers to the strength of the monetary transmission in the past, i.e. before the inception of EMU, the results are influenced by the differing monetary and exchange rate regimes under which countries have hitherto operated. Whether and to what extent any past differences in monetary transmission will remain in EMU is another matter. We focus here on the historical evidence.

In the spirit of Britton and Whitley (1997), we divide the empirical studies into the following categories:

- (i) large-scale macroeconometric single-country models (MEM1s);
- (ii) large-scale multi-country models (MEM2s);
- (iii) small-scale structural models (SSMs);
- (iv) single equation models (SEMs); and
- (v) structural VAR models (SVARs).

The results are reported in Table 1. In the interest of brevity, we focus wherever possible on the estimated impact on output of a given monetary policy action at a particular horizon, namely in the year “t+1” following the policy action (quarters 5-8 after the initiation of the action)<sup>2</sup>. This horizon is chosen because in almost all the studies surveyed, the maximum impact of the monetary policy action on output falls within that year in most countries. For the study by Dornbusch, Favero and Giavazzi (1998), the impact over this horizon is not available to us. In this case, the impacts after 8-12 months and after 24 months are reported. Focusing on a single representative number for the output effect allows us to make a simple ranking of the power of monetary policy in each of the four major EU economies according to each study. Ignoring the dynamics of the output responses tends, if anything, to underestimate the variation in results for a given country between different studies (to the extent they are comparable). In general, the rankings obtained in this way correspond well with a casual assessment of the full dynamic impact.

In interpreting the results reported in the table, it is important to keep in mind that they are often not comparable across studies, and sometimes not strictly comparable across countries within a given study. This holds in particular for two reasons:

*(i) Different exchange rate assumptions for model simulations:* For the simulations of macroeconomic models and the small structural models, the exchange rate assumptions are crucial. For instance, in the study reported in the first line of Table 1 (BIS (1995)), the simulations for France, Germany and Italy are all carried out under the assumption of fixed exchange rates within the ERM<sup>3</sup>, whereas the simulation for the UK assumes a flexible exchange rate. Also the impulse response functions in the structural VARs implicitly assume different exchange rate responses across countries.

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<sup>2</sup> In some cases where the data were not available to us, the impact has been gauged from the graphical representations in the original sources. We do not consider the impact on prices, not because it is not important, but because we wish to keep the discussion as short as possible and because most of the studies reviewed focus on output. Several of them do not report the price impact. For those that do, the results seem to corroborate our general conclusion concerning the output effects, namely that there is no consistent pattern across studies, neither of the size nor of the ranking of the impact in various countries.

<sup>3</sup> They do not take into account that this would require a simultaneous increase in interest rates in the ERM area and thus have repercussions on each country via reduced foreign demand.

(ii) *Different monetary shocks*: Different studies report the response to different monetary shocks. These have been standardised to the extent possible. The table contains four basic types of monetary shock. The first is an increase in the short-term interest rate sustained for two years (used in the MEM1 results, some MEM2 results and the SSM results). The second is a sustained 1% decrease in the money target (used in several MEM2 results)<sup>4</sup>. The third is a simultaneous 1% permanent shock to interest rates in all EU countries keeping intra-EU exchange rates fixed (used in SEM results). The fourth and last type is a one standard deviation shock to the interest rate equation in VAR models; in these cases both the initial size of the shock and the subsequent movement in interest rates may vary between countries. The importance of these differences is illustrated in the last two lines of Table 1 which show the results of a particular VAR (Gerlach and Smets (1995)) using both one standard deviation shocks, and a standardised shock amounting to a temporary one percent increase in interest rates sustained for two years.

It follows from the above considerations that none of the studies reported, with the exception of Dornbusch, Favero and Giavazzi (1998), attempts to estimate the impact of the type of common monetary shock which will occur in EMU, namely a simultaneous and equal change in policy interest rates with fixed exchange rates among participating countries, a simultaneous and equal move in the common euro exchange rate against third countries, and a similar response along the yield curve. The data provided in Table 1 should therefore not uncritically be interpreted from the perspective of assessing possible differences in the EMU regime.

Moreover, in the majority of the studies under examination, differences in monetary transmission are not established with statistical rigour. We follow the practice of loose language here. Hence, when stating that an impact in one country is “stronger” or “weaker” than in another we merely mean that the two point estimates differ, with no implication of the possible statistical significance of the difference. In some cases, as in the comparison of macroeconomic models, the context does not allow any kind of rigorous statistical testing. However, even in the cases in which the approach would have allowed statistical evaluation of the differences, the possibility has often not been exploited.

Inevitably, all of the approaches face a number of methodological problems, some of which are specific to the approach, and some of which are common (even if they manifest themselves differently in the approaches considered). A crucial issue of the latter variety is the correct identification of the causal effect of monetary policy on the economy. The problems involved are of the ordinary textbook variety, mundane in almost all econometrics. Essentially, when trying to pinpoint the effects of monetary policy on, say, output and price level, one must recognise not only that monetary policy is an endogenous variable which responds to the same set of underlying causes

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<sup>4</sup> In some cases, the original sources show the impact of a 10% increase in the money target (IMF Multimod) or a 3% increase in the money target phased in over 1 year (Taylor (1993)). These shocks have been standardised to a 1% contraction in the money target assuming linearity.

as output and prices, but also that there is a direct reverse causality from the two variables to monetary policy. The difficulty is to decide when to interpret correlation as causality (or, equivalently, the lack of correlation as the lack of causality)<sup>5</sup>. Econometrically, the question is about omitted variables and simultaneity.

In large-scale models, simultaneity among a wide range of macroeconomic variables is almost routinely ignored when, as is typically the case, equations are estimated one-by-one. Small-scale structural models may be estimated jointly but, as with large-scale models, the exclusion restrictions imposed to achieve identification are typically not well accounted for<sup>6</sup>. The latter generally also applies to single-equation models. Structural VARs have the advantage that they force the researcher to deal explicitly with the identification problem but, as with the other approaches, there is a high degree of arbitrariness involved in choosing a basis for identification. The problem of omitted variables looms large in all approaches, including the large structural models since they are too large to be estimated as a system and individual equations are typically estimated separately.

#### *Large-scale macroeconomic single-country models (MEMIs)*

The most commonly quoted evidence based on national econometric models is a set of independent simulations carried out by the G10 national central banks as part of the BIS study of financial structures and the monetary transmission mechanisms; these results are reported in Smets (1995). According to these results, monetary policy has essentially identical effects on output at the one-year horizon in Germany, France and Italy (in the simulations assuming fixed intra-ERM exchange rates)<sup>7</sup>. Among the four large economies, only the UK stands out. According to the simulations, the effect of

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<sup>5</sup> This problem was already known to Kareken and Solow (1963): “Imagine an economy buffeted by all kinds of cyclical forces, ... Suppose by heroic...variation in the money supply...the Federal Reserve manages deftly to counter all disturbing impulses and to stabilize the level of economic activity absolutely. Then, an observer...would see peaks and troughs in monetary change accompanied by a steady level of aggregate activity. He would presumably conclude that monetary policy has no effects at all, which would be precisely the opposite of the truth.”.

<sup>6</sup> As an example, a structural model may contain a money demand equation relating  $M$  to  $Y$ ,  $P$  and  $i$  (in standard notation), and a money supply reaction function which may depend e.g. on  $Y$  and  $P$ . In this case, identification has been achieved by excluding the interest rate from the money supply equation. But is it reasonable to assume that a central bank which targets the money supply does not look at the interest rate at all? Or, if central bank conducts monetary policy by setting a short-term interest rate, is it reasonable to assume that it does not take into account developments in the money supply at all? If the answer to these questions is “no” then  $Y$ ,  $P$  and  $i$  all belong in *both* the money supply and the money demand equations. Therefore, it is not clear whether regression estimates will deliver one or the other, or a mixture of the two.

<sup>7</sup> In table 1, the reported impact of national monetary policy in the Netherlands is considerably smaller. This is a consequence of the relative openness of that economy. In the scenario of a common increase in interest rates in the ERM (or EMU) countries, demand would fall also among the Netherlands’ trading partners, and the overall effect would be closer to that reported for France, Germany and Italy.

monetary policy on output is about twice as large in that country as in Germany, France and Italy. This is partly due to the fact that the UK model, as stated above, is simulated under a different assumption about exchange rates.

While national macroeconomic models may reflect a detailed knowledge of the respective economies, a potential problem with this approach is that the specification of individual country models may, in practice, vary depending on the priors, views and assumptions of their respective model-builders. This pertains to a wide range of areas, e.g. different methodologies on whether and how to incorporate long-run constraints on the economy or whether to model expectations as adaptive or forward-looking and model-consistent (cf. Smets (1995)). Clearly, assumptions in these areas are highly relevant for an assessment of the effects of monetary policy. Such differences in model specification are likely to obscure, and could well dominate, any genuine differences in real economic structures and behaviour.

#### *Large-scale macroeconomic multi-country models (MEM2s)*

Other macroeconomic models tend to produce results which are often qualitatively different from those obtained using the national central bank models. Table 1 summarises results from the US Federal Reserve's MCM model (reported in BIS (1995)), the IMF's Multimod standard simulations (Masson *et.al.* (1990)), the Commission services' Quest II model (Roeger and In't Veld (1997)), and the model results reported in Taylor (1993). For the first model, results are reported for the same shock as in the national central bank models but with fully endogenous exchange rates. The output responses are similar between Germany and France, but considerably smaller in Italy. As before, they are clearly stronger in the UK. The deviating results for Italy and the UK can to a large extent be accounted for by differences in the net interest-bearing asset position of the household sector.

For the other three multi-country models, results are reported for a shock of a 1% decrease in the money target with endogenous exchange rates. In all of the three models, there are no significant differences in the output response between different countries. The only exception is the Taylor model for the UK where the output response is, in contrast with the previous findings, considerably smaller than in Germany, France and Italy. This is mainly due to a much smaller estimated response of domestic investment and consumption in the UK in that model.

Multi-country models tend to impose a similar structure across countries. While this may limit the potential biases underlying the use of national models, it could mean that multi-country models are less able to capture specific features of individual economies. In part, this may account for the tendency for such models to give broadly similar results for different countries.



**Table 1. Empirical assessments of the impact of monetary policy on output in various European countries**

(Impact on real GDP in year  $t+1$  (5 to 8 quarters after original shock, percentage deviation from baseline))

<i>Study</i>	<i>Shock</i>	<i>D</i>	<i>F</i>	<i>I</i>	<i>UK</i>	<i>E</i>	<i>NL</i>	<i>Ranking</i> ( <i>D, F, I, UK</i> )	<i>Comments</i>
<b><u>Single country macro models</u></b>									
National Central Bank models (BIS (1995))	Type 1	-0.4	-0.4	-0.4	-0.9	0.0	-0.2	D = F = I (< UK)	Fixed ERM rates for D, F and I; endogenous exchange rate for UK
<b><u>Multi-country macro models</u></b>									
Fed MCM model (BIS (1995))	Type 1	-0.7	-0.7	-0.3	-1.2	-	-	I < D = F < UK	Endogenous exchange rates
IMF Multimod standard multiplier	Type 2	-0.5	-	-	-0.5	-	-	D = UK	Endogenous exchange rates
Quest II (Commission Services)	Type 2	-0.4	-0.4	-0.3	-0.4	-0.4	-0.3	I < D = F = UK	Endogenous exchange rates
Taylor (1993)	Type 2	-0.4	-0.4	-0.4	-0.1	-	-	UK < I = D = F	Endogenous exchange rates
<b><u>Small structural models</u></b>									
Britton and Whitley (1997)	Type 1	-0.5	-0.5	-	-0.3	-	-	UK < D = F	Each country estimated separately
Britton and Whitley (1997)	Type 1	-0.4	-0.4	-	-0.4	-	-	D = F = UK	All countries estimated jointly
<b><u>Reduced form equation</u></b>									
Dornbusch, Favero and Giavazzi (1998)	Type 3	-0.5	-0.5	-1.1	-0.5	-0.4	-	UK = D = F < I	Effect after 8-12 months
Dornbusch, Favero and Giavazzi (1998)	Type 3	-1.4	-1.5	-2.1	-0.9	-1.5	-	UK < D = F < I	Effect after 2 years
<b><u>Structural VARs</u></b>									
Ramaswamy and Sløk (1997)	Type 4	-0.6	-0.4	-0.5	-0.5	-0.3	-0.6	F < I = UK < D	Baseline model
Barran, Coudert and Mojon (1996)	Type 4	-0.6	-0.4	-0.2	-0.4	-0.4	-0.3	I < F = UK < D	Baseline model (model 1)
Gerlach and Smets (1995)	Type 4	-0.3	-0.2	-0.2	-0.6	-	-	F = I < D < UK	1 standard deviation shock
Gerlach and Smets (1995)	Type 1	-1.2	-0.6	-0.6	-0.8	-	-	F = I < UK < D	1% interest rate hike for 2 years

*Types of monetary shock:* Type 1: 1% point rise in short-term interest rates sustained for at least two years; Type 2: 1% permanent decrease in money target;

Type 3: 1% simultaneous permanent increase in short-term interest rates; Type 4: 1 standard deviation interest rate shock

### *Small structural models (SSMs)*

The small structural model estimated by Britton and Whitley (1997) is a variant of the Mundell-Fleming model incorporating the Dornbusch overshooting mechanism. The parameter estimates suggest that the sensitivity of output to the real interest rate may be lower in the United Kingdom than in Germany or France. However, the differences are generally not large enough to be statistically significant. When the models are estimated separately, the simulated output effects are noticeably smaller in the UK than in the other two countries, but when the three country models are estimated jointly, the output response in the UK is virtually identical to Germany and France.

One specific criticism against this type of model is that it may be too parsimonious and too highly aggregated to capture cross-country differences in economic structure.

### *Single equation models (SEMs)*

Dornbusch, Favero and Giavazzi (1998) estimate a single equation for output growth (industrial production) in each of six countries based on (1) past output growth, (2) past output growth in the other countries, (3) present and past values of “expected” and “unexpected” interest rates<sup>8</sup>, and (4) present and past values of each country’s bilateral exchange rate against the DM and the US dollar. The estimation method allows for simultaneity in the determination of output across countries, but not for simultaneity in the determination of output, prices and monetary policy in each country.

An advantage of this specification is that it allows one to control for the intra-European exchange rate channel, and thus to simulate the impact of a common interest rate change with fixed exchange rates among the European countries. The results are clearly different from the previously reported findings<sup>9</sup>. The impact on output, as reported in table 1, is similar in Germany, France and Spain but much stronger in Italy (and Sweden, not reported in table 1). At the other end of the spectrum, the UK stands out with a significantly smaller long-term impact.

The estimation of single equation models of this sort suffers from the same methodological problems of identification as the other approaches<sup>10</sup>. Moreover, it is

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<sup>8</sup> The authors decompose interest rate changes into “expected” and “unexpected” components using fitted values and residuals of an estimated monetary policy reaction function for each of the six countries (Germany, France, Italy, the United Kingdom, Spain and Sweden). For countries other than Germany, the authors find large residuals during periods of exchange rate turbulence that may be associated with shocks to risk premia. In the preferred specification of the output growth equations, only expected interest rate changes are included (the unexpected ones are statistically insignificant).

<sup>9</sup> The long-run output response which is obtained after two years is clearly stronger than in the other models reported. This is partly due to the simultaneous interest rate shock which allows for spill-overs between countries.

<sup>10</sup> The analysis is akin to the so-called “St. Louis Fed” approach due to Anderson and Jordan (1968). See Cochrane (1994b) for a critique of this approach.

by no means clear what economic theory would lead to such an equation being a good approximation to the true reduced form equation for real output growth.

### *Structural VARs*

Table 1 reports the results of three recent studies which have applied the VAR methodology to assess the possible differences in monetary transmission mechanisms in Europe. The way in which we report the results, at the one-year horizon and for output only, does not capture fully the richness of these results. However, it does appear that while each individual study may have identified differences in the shape and timing of the responses between countries, the pattern is not consistent across studies. In our view, the simplification adopted here by disregarding the time profile of output responses does not seriously underestimate the indications of differences in the monetary transmission mechanism.

In interpreting the results, it should be kept in mind that in some of these studies the type of monetary action considered -- namely, a one standard deviation shock to the interest rate equation -- may deviate substantially between countries.

Ramaswamy and Sloek (1998) report that the full effects on output in one group of EU countries (including Germany and the UK) take longer to occur, but are almost twice as large as in another group (including France and Italy). The results in Barran, Coudert and Mojon (1996) indicate, for their basic model including five variables, a similar output response in France and the United Kingdom, a somewhat stronger impact in Germany, and a somewhat smaller impact in Italy. Gerlach and Smets (1995), using long-run identifying restrictions, report similar output responses across countries<sup>11</sup>, except in the case of the United Kingdom, where the effects are somewhat larger<sup>12</sup>. When the monetary policy shocks are standardised by assuming that the central bank raises the nominal interest rate by 100 basis points for eight quarters, after which the interest rate returns to baseline, the results are somewhat different. Now, monetary policy is most powerful in Germany with an impact about twice as large as in France and Italy and with the UK in between.

Although the VAR models have the advantage of dealing explicitly with the identification of monetary policy<sup>13</sup>, they suffer from a number of disadvantages. First, their “black-box” nature makes it difficult to relate estimated parameters and impulse responses to structural differences in the economies. For instance, if monetary policy is found to be more powerful in one country than in another, the

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<sup>11</sup> This is confirmed for Germany, France and Italy in Smets (1997).

<sup>12</sup> The latter is in contrast with Smets (1996) where the output effect is smaller in the UK.

<sup>13</sup> Ramaswamy and Sloek (1997) assumes that there is no contemporaneous (within quarter) effect of monetary policy on output and prices (this type of identification scheme in a three-variable VAR often leads to the so-called “price puzzle”, which suggests that the scheme may not identify monetary policy shocks correctly, cf. e.g. Leeper, Sims and Zha (1996)); Barran, Coudert and Mojon (1996) also assume a recursive Cholesky decomposition (using five variables, including the world export price index, the “price puzzle” is largely gone, but the assumption that the exchange rate does not respond contemporaneously to the interest rate seems implausible); and Gerlach and Smets (1995) use long-run identifying restrictions (which, as Faust and Leeper (1997) have shown, can be misleading).

model does not indicate what economic structures cause this result. Second, since the VAR approach focuses on the instances when monetary policy-makers deviated from their “normal” reaction function, it says relatively little about the economic consequences of the way in which the central bank normally conducts monetary policy<sup>14</sup>.

*Summing up the evidence*

The results reviewed here are obtained under different assumptions and with different methods; they are often not comparable between studies, and sometimes not even within studies. However, the general conclusion emerges that very different results can be obtained for the same country using different models. These differences are often larger than the differences which appear across countries using a given model or method. Moreover, different studies tend to provide a different ranking of the four major EU economies with respect to the power of monetary transmission. Although it may be perfectly reasonable to have more faith in the results of one study or method over the others, the variation in the results indicates that econometric analysis has not provided consistent evidence about either the extent nor the ranking of possible differences in monetary transmission across EU countries.

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<sup>14</sup> Although the VAR approach focuses on monetary “shocks” or “innovations”, it does not rule out the possibility that systematic and predictable monetary policy affect the course of the economy in an important way (Bernanke, Gertler and Watson (1997)).

### 3 The indeterminacy of monetary shocks — structural identification of VARs

In what follows, we will analyse the identification of a causal structure between monetary policy and the economy in the framework of the vector autoregressive model. It is worth emphasising that the problems discussed are in no way specific to the particular modelling approach — it is just as easy to misrepresent co-movement as causality using any of the approaches outlined above. The reason for concentrating on the VAR framework is that it forces one to tackle the identification problem in a particularly explicit form. At the same time, it offers a straightforward tool for examining the extent of the problem with a minimal set of prior restrictions. And, as a final bonus, it is perhaps the approach most widely used in cross-country comparisons, and therefore a good candidate for presenting the methodological point.

The generic VAR model can be written as

$$(1) \quad [I - A(L)]X_t = u_t, \quad E(u_t u_t') = \Sigma.$$

Here  $X_t$  is the vector of endogenous variables,  $u_t$  is the residual vector, and  $A(L) = A_1 L + \dots + A_k L^k$  is a lag polynomial. Given that  $X_t$  is a stationary process, it has a Granger representation

$$(2) \quad X_t = B(L)u_t,$$

where  $B(L) = [I - A(L)]^{-1}$ . In general, the error term  $u$  has a non-diagonal covariance matrix, indicating that the data contains some regularities left unexplained by the model. Whether this is the case or not, the model (2) does not yet provide a basis for an analysis of causal relations. Instead, it is to be seen as the reduced form of the underlying *structural* model

$$(3) \quad X_t = B(L)G e_t, \quad E(e e') = I, \quad GG' = \Sigma.$$

Here  $e_t = G^{-1}u_t$  are the normalised *structural* shocks. The structural shocks are supposed to be economically meaningful, genuine primary shocks, exogenous to everything else in the model, including each other — hence the diagonality of their covariance matrix. This exogeneity property allows one to analyse their effects on the model variables without having to bother with simultaneity considerations.

The problem is that the identification of the correct  $G$  is far from straightforward. The requirement  $GG' = \Sigma$  falls well short of identifying a unique structural form. For any candidate  $G$  there exists an infinity of other possible structural identification schemes  $ZG$ , where  $Z$  is any orthonormal matrix, i.e.  $ZZ' = I$ . Each  $Z$  defines a different structural identification scheme, which fits the data equally well as the original scheme. Since the data cannot differentiate between the different identification schemes, this has to be done by the econometrician.<sup>15</sup>

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<sup>15</sup> Whether the residual covariance matrix  $\Sigma$  is diagonal or not is unimportant in this context. This property used to play a role when the structural identification was understood as merely finding the appropriate recursive structure. For a diagonal  $\Sigma$ , all recursive identification schemes are equivalent. In

This problem is well understood in the econometric profession. Generally, the solution has been to impose a sufficient number of additional restrictions on  $G$  (or on a function of  $G$ ) to identify a unique  $G$ . Traditionally, the usual (indeed, almost mechanical) procedure was to assume a recursive contemporaneous structure, i.e. assume  $G$  to be triangular. Recent studies have devoted more careful attention to the issue, imposing a variety of exclusion restrictions, either on the short-run or on the long-run effects of a shock. Typical restrictions include requiring that monetary policy shocks do not have immediate effects on some variables with high inertia, or that they do not have long-run effects on variables which are thought to be ultimately determined by supply-side factors.

Both short-run and long-run restrictions suffer from certain pitfalls. The short-run restrictions are typically based less on theoretical considerations than on heuristic arguments. Sometimes the heuristic arguments used are in manifest contradiction with some theoretical results. For example, the popular assumption that monetary policy does not have short-run (i.e. within-period) effects on output is at odds with the neoclassical view that monetary policy *only* has short-run effects on output.

Often there is a better theoretical justification for relying on the long-run properties of the model in the structural identification. When applied to the real world, many theoretical arguments derived from models of a frictionless economy are best thought as applying to the long-run equilibrium of the economy. However, even for the long run it is often difficult to interpret the model implications as holding *exactly*. For example, although the long-run growth of the economy is determined by the supply side which probably is relatively unaffected by monetary policy, claiming that a monetary episode causing a prolonged and significant slowdown of the economy cannot have *any* effect on the supply side would be stretching the argument too far. For this reason, it would appear fairly dubious to conclude that a particular shock which is estimated to have had real effects in the long run (the actual length of which is determined by the estimation period) cannot be of monetary origin. Hence, in our view, the case for using exact restrictions on the long-run effects is not much stronger than with the short-run restrictions.

Furthermore, as Faust and Leeper (1995) have shown, relying on long-run restrictions to derive short and medium-run implications can be highly misleading. In a VAR model, the relationship between long-run and short-run properties is a direct result of the model's autoregressive structure. However, this structure is meant to be a reasonable approximation, not a faithful representation, of an arbitrary true data generating process. A VAR with a reasonable lag length can be expected to track the short-run properties of the true process well, but for long horizons, this is by no means clear. In short, basing the identification scheme on long-run restrictions places an unjustified emphasis on a particular property of the VAR framework which should be viewed as a convenient simplification rather than as a true feature of the data generating process.

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the more general non-recursive setting, however, the space of possible identification schemes (possible orthonormal matrices  $Z$ ) is just as large for a diagonal  $\Sigma$  as it is for a non-diagonal one.

All in all, there is no simple way to solve the identification problem. A variety of approaches has been tried, producing, as showed in the previous section, rather different outcomes. In this light, it is somewhat curious to see that in the literature, the identifying assumptions used to define monetary policy are, in most cases, devoted merely a single sentence, as the were technical assumptions with few consequences. There is seldom any discussion about possible alternative identification schemes, or about the sensitivity of the results with respect to the particular identification restrictions (King and Watson, 1997, and Uhlig, 1997, are rare exceptions<sup>16</sup>).

As a final methodological point, it is notable that the studies searching for cross-country differences in monetary policy transmission invariably impose the same identification scheme on all countries. In other words, any observed differences in the responses are conditional on the assumption that the same qualitative restrictions regarding the effects of monetary policy hold for each country. Given that the purpose of the exercise is to identify differences in the monetary policy transmission, conditioning the results on an identical set of assumptions seems rather counterintuitive. In our view, a rigorous cross-country comparison should allow for different identification schemes. In what follows, we propose an approach that allows this.

#### **4 The identification problem in practice — an example**

The lack of exact knowledge regarding the correct contemporaneous structure does not necessarily mean that no meaningful econometric conclusions are possible. If, after applying some reasonable restrictions to exclude those identification schemes that are clearly implausible, the remaining schemes produce closely similar predictions, then the exact choice of identification scheme is largely irrelevant and robust conclusions can be drawn. But is this really the case in the bulk of econometric work? In what follows, we try to show that this is not the case.

To illustrate the extent and consequences of the identification problem, and to shed light on the methodology we use later in the cross-country comparison we have estimated a 3-variable VAR model for the United Kingdom. The variables included are the logarithmic levels of the real GDP and the consumer price index, and the level of nominal 3-month money market interest rate. The data are quarterly, seasonally non-adjusted, and span from the beginning of the 1970s to 1997 Q3. We estimated the VAR in levels, which somewhat simplified the calculations later in the section.

We created an identified baseline model by using the Cholesky decomposition with the ordering output - price level - interest rate.<sup>17</sup> This baseline structural model determines a base of three orthogonal shocks, the third of which we, following the

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<sup>16</sup> King and Watson (1997) test the long-run neutrality of money in a two-variable VAR for a range of identification schemes. Uhlig (1997) selects the identification scheme by maximising an ad hoc objective function over a range of identification schemes.

<sup>17</sup> Hence, our baseline model corresponds to a familiar recursive assumption that monetary policy reacts to the contemporaneous values of all variables but does not have within-period effects on any of them.

usual tradition, label as our baseline monetary shock. To this orthogonal base we apply a series of orthogonal rotation matrices  $Z_i$ . Any three-dimensional rotation matrix  $Z_i$  can be represented as a function of three parameters  $A_i$ ,  $B_i$  and  $C_i$ :

$$Z(A_i, B_i, C_i) = \begin{bmatrix} 1 & 0 & 0 \\ 0 & \cos(A_i) & \sin(A_i) \\ 0 & -\sin(A_i) & \cos(A_i) \end{bmatrix} \begin{bmatrix} \cos(B_i) & 0 & -\sin(B_i) \\ 0 & 1 & 0 \\ \sin(B_i) & 0 & \cos(B_i) \end{bmatrix} \begin{bmatrix} \cos(C_i) & -\sin(C_i) & 0 \\ \sin(C_i) & \cos(C_i) & 0 \\ 0 & 0 & 1 \end{bmatrix}$$

Loosely speaking, each of the three parameters determine the amount the matrix  $Z$  rotates the orthogonal base around a certain dimension. The last parameter  $C_i$ , for example, rotates the base around the third dimension, leaving the last structural shock unchanged but "blending" together the effects of the first two shocks. By varying the three parameters in the range  $[0, 2\pi]$  one can produce all possible transformation matrices.

As in the present context we are only interested in the effects of monetary policy on output and prices, we do not have to examine all possible triplets of the parameters. First, the parameter  $C_i$  determines the extent to which the transformation matrix rotates the base around the third axis, chosen to represent a genuine monetary policy shock. This means that the parameter  $C_i$  only affects the identification of the first two shocks, but leaves the third one, the monetary policy shock, unaltered. Hence, when searching for viable identification schemes for the monetary policy shock, we only need to concentrate on the first two parameters  $A_i$  and  $B_i$ . Moreover, we can reduce our search to those combinations of  $A_i$  and  $B_i$  in which both parameters take values from the range  $[0, \pi]$ <sup>18</sup>. Any parameter combination outside this range can be remapped into it by a suitable transformation.<sup>19</sup>

To map the whole space of identification schemes, we divided this range into a grid of  $100 \times 100$ , thus producing a set of 10000 schemes to evaluate. For the identification scheme determined by each point in the grid, we generated the corresponding impulse responses.

Next we needed a methodology that allows us to compare different identification schemes on a sensible basis. In particular, the nature of a monetary shock can vary considerably between identification schemes. A monetary shock can show up as a very short-lived or as a sustained change in the interest rate. A sustained increase in interest rates is likely to have larger effects elsewhere in the economy than a transitory peak. However, such differences should be interpreted as reflecting differences in estimated monetary policy reaction function rather than in the monetary policy transmission. Whether this distinction is relevant or not depends on the intended use of the model. If the purpose is to identify likely differences in the responses to the

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<sup>18</sup> Any other range of length  $\pi$  obviously does equally well. For presentational purposes, we will in some occasions choose the range  $[-\pi/2, \pi/2]$ .

<sup>19</sup> More precisely, any parameter combination  $(A, B)$  can be remapped into this range by choosing suitable integers  $k_1$  and  $k_2$  in the formula  $(A + k_1, B \cos(k_1\pi) + k_2\pi)$  without affecting the results.



common monetary policy, then such differences due to monetary policy reaction functions should be filtered out.

To account for such differences, we decided, instead of looking at the effects of monetary policy shock of uneven size and duration, to examine the response of the economy to a standardised monetary tightening.<sup>20</sup> That is, for every identification scheme, we simulated a series of monetary policy shocks such that they produced a 1 percentage point increase in the short-term interest rate, sustained over a period of four years. In this way we obtained a set of output and price responses which can be compared in a meaningful way.

Obviously, not all identification schemes in our set of 10000 make good candidates for monetary policy. In most cases, their predictions of the effects of monetary policy are highly implausible. To provide an initial idea of how sharp the predictions get when one begins to require intuitively "plausible" behaviour, we reduced the set of all identification schemes criteria in several steps.

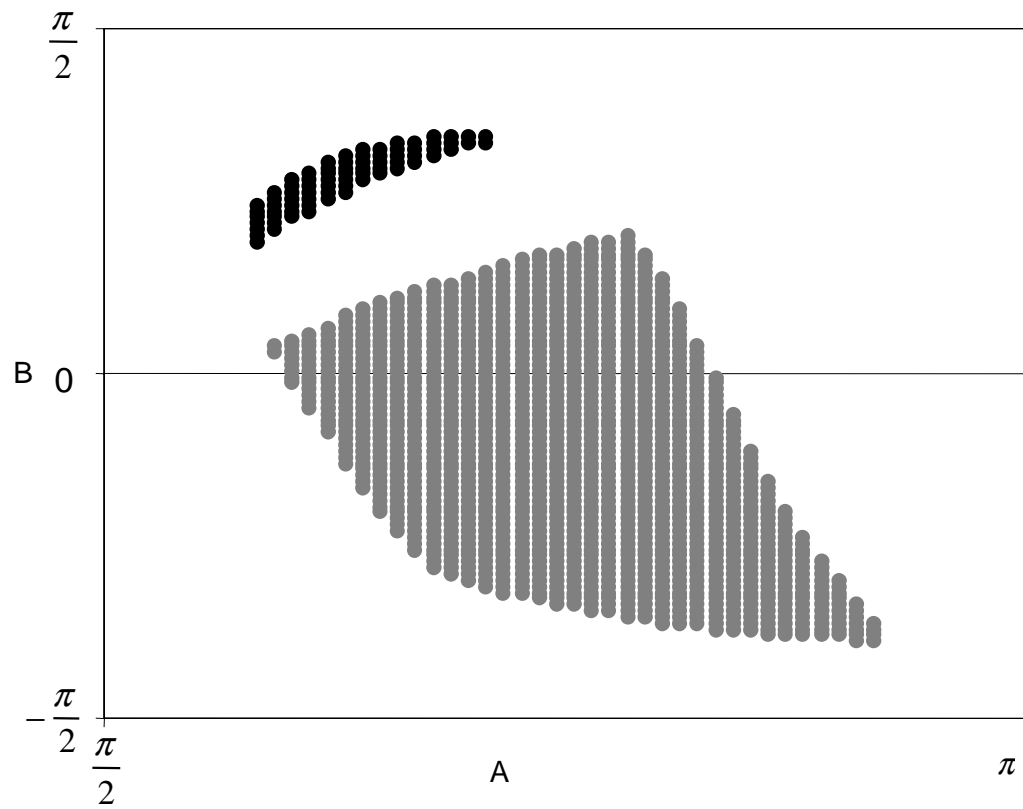
First, we filtered out those pairs of responses which failed to fulfil a minimal medium-term "plausibility" criterion. We required that the accumulated response of the level of both real output and the CPI in the UK to the sustained 1 percentage point monetary tightening is, after two years, between zero and -10%. There were 1372 identification schemes which fulfilled this condition. Chart 1 illustrates the situation. There are two distinct regions of parameter combinations which fulfil this criterion: one large area with 1299 observations, and a separate small slice with 73 observations. As we show below, the responses corresponding the two separate areas differ qualitatively, so we differentiate the two areas by marking the larger one with black markers and the smaller by grey markers.

Chart 2 plots the corresponding responses of output (horizontal axis) and prices (vertical axis) after 2 years for these 1372 identification schemes. As in chart 1, the two separate areas are again distinguished by grey markers for the larger block and black markers for the small block. The chart shows that the identification schemes in the larger block span the whole area of responses from zero to 10% both for output and for prices. The smaller block of identification schemes covers likewise a sizeable part of the upper right hand corner of the chart.

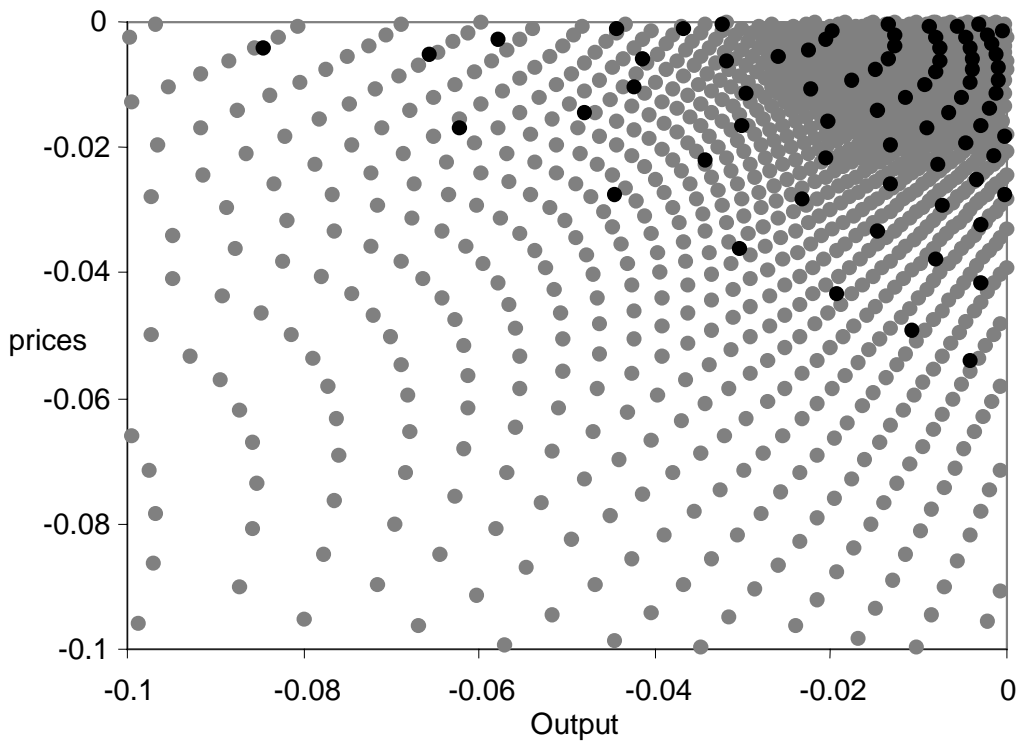
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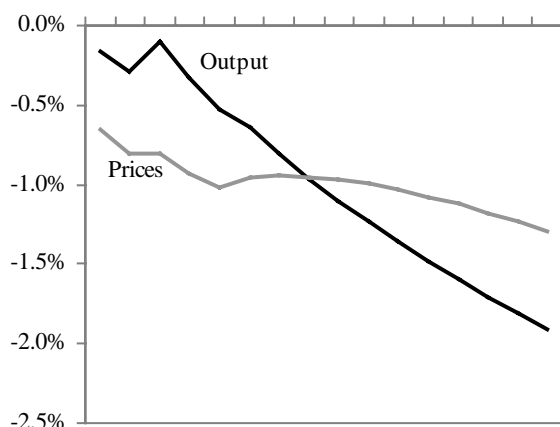
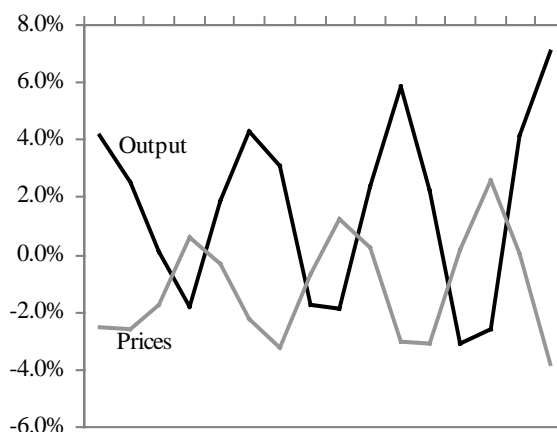
<sup>20</sup> A similar approach has been used by others, see for example Gerlach and Smets (1995).

**Chart 1.** The "plausible" set



**Chart 2.** Responses after 2 years



**Chart 3a****Chart 3b**

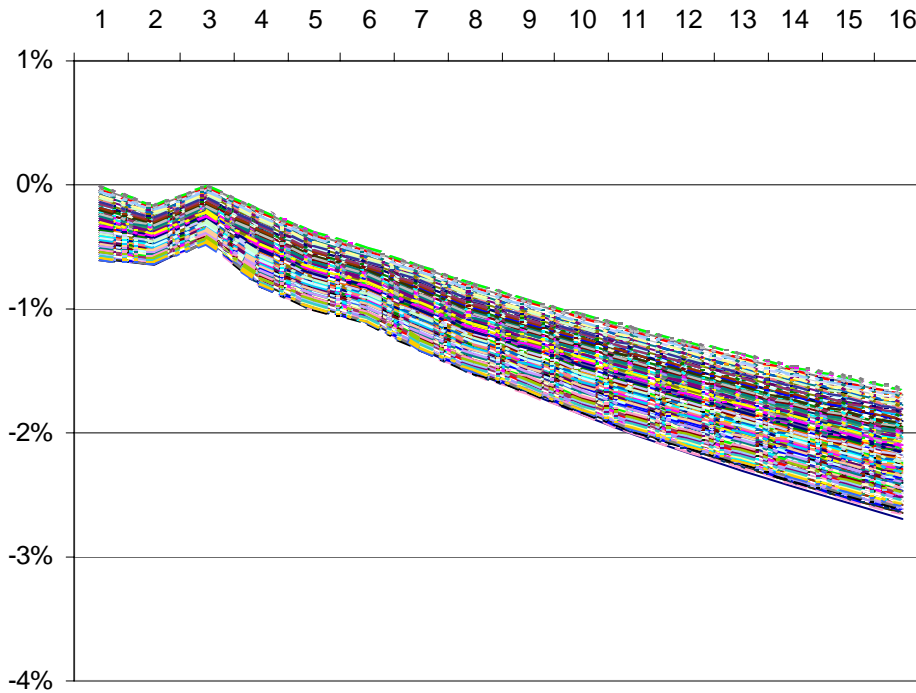
The conclusion is that for any arbitrary combination of accumulated output and price responses at the two-year horizon, an identification scheme can be found that provides it. Of course, basing our selection criteria only at the two-year point leaves a lot open. A closer inspection would reveal that a great majority of the 1372 sets of responses are implausible in one way or another (there are many more ways to fail than to succeed). Charts 3a and 3b plot two response functions — the first one from the larger (grey) set (corresponding to parameter values of  $A_i = 2.387$  and  $B_i = -.11$ ) and the second one from the smaller (black) set ( $A_i = 2.01$  and  $B_i = 0.974$ ) — both of which have the two-year accumulated responses close to 1% for both prices and output. The first graph looks generally plausible, except for the somewhat uncomfortably rapid onset of the price response. The responses in chart 3b, on the other hand, exhibit clear explosive oscillating behaviour. This property is shared by all identification schemes in the smaller (black) set, so in what follows, we will exclude these schemes.

Since the first plausibility criterion proved to be excessively loose, we proceeded to reduce the plausible set further. We did that by imposing additional restrictions in the short and medium term. Specifically, we required that the (negative) within-quarter impact of the monetary tightening must be smaller than 0.6% (in absolute value) both on the price level and on output. We also required the accumulated responses of the two variables to be negative after two and four years of tight money.<sup>21</sup>

We found 240 identification schemes which fulfilled each of these conditions. The full impulse response functions for these are presented in charts 4 and 5. Although the two charts resemble the familiar "point-estimate-and-an-error-range" charts, the correct interpretation is actually quite different. All the responses are observationally equivalent — the data support each response equally well. From a purely statistical point of view, there is no presupposition that a response in the middle of the range is

<sup>21</sup> These limits are chosen to facilitate the presentation of the methodological point — they do not intend to faithfully represent our priors. The cross-country comparison in the next section involves a more careful handling of the plausibility conditions.

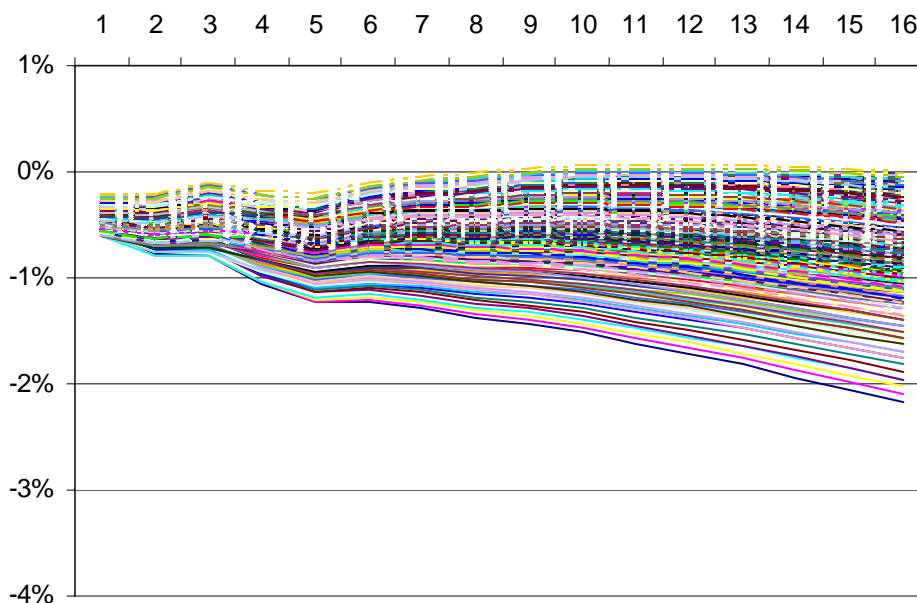
**Chart 4** Output responses



more plausible than one near the edge. All such conclusions must be based on considerations independent of the data.

An interesting observation is that the initial price responses are quite tightly packed at around -0.4%. There is no identification scheme with the immediate price response smaller than -0.2% which at the same time fulfils our medium-term restrictions. The implication is that with our data set, attempting to identify monetary shocks by restricting the within-quarter response of price level to zero (as in Ramaswami & Sloek, 1998, Barran, *et. al.* 1997, and many others) would necessarily lead to response

**Chart 5** Price responses



functions which fail our medium-term plausibility test. With regard to output responses we see that the within-quarter responses are more widely spread and cover the full range from zero to -0.6%. We consider this to be of plausible range of magnitude, given that we are dealing with a sizeable tightening of a full percentage point.

Equipped with this particular set of criteria, we would have to settle for a relatively wide range of viable estimates. For example, it can be calculated that the negative effect of a tightening at the beginning of the year on aggregate output of the UK economy during the full year could be somewhere between less than 0.1% and more than 0.6%. Similarly, in the third year of the tightening, the effect would be roughly between 1% and 2%. Choosing a single ("best") identification scheme involves many trade-offs and a careful weighting of one's priors.

Finally, it is worthwhile to bear in mind that these are point estimates which ignore the considerable degree of statistical uncertainty always involved with VAR impulse responses. Hence, each estimate of the impulse response should be viewed as coming with a (normally fairly wide) confidence interval attached.

We think that this example manages to show that although the "plausible" identification schemes may only occupy a small part of the space of all possible schemes, it is far from simple to pick out the single "most plausible" identification scheme. Any plausible parametrisation is surrounded by a continuum of equally plausible and observationally equivalent parametrisations. As one moves further away from the original parametrisation, the parametrisation becomes less "plausible" and, at some point, becomes "implausible". However, before reaching that point, the impulse responses may have evolved considerably, leaving one not with a single estimate, but with a range of impulse responses. In our opinion, a rigorous cross-country comparison has to take this into account and base the comparison on the plausible ranges instead of on single points in those ranges, selected by a mechanical criterion. Such an approach will certainly yield less clear-cut results in terms of identifying cross-country differences but would avoid the risk of spurious accuracy, which is inherent in results based on single points within the ranges.

## **5 Differences in monetary transmission — a cross-country comparison**

### **5.1 Methodology**

The previous two sections have sought to show that the present mainstream in the analysis of monetary policy transmission using VARs suffers from serious shortcomings. It is not obvious how these shortcomings could be overcome. Choosing between a large number of observationally equivalent parametrisations is an inherently fuzzy business which ultimately depends on the researchers' subjective priors. A high degree of subjective input leaves the econometrician exposed to charges of manipulation of the results.

We do not pretend to have a comprehensive solution to these problems. However, we would like to suggest an approach which would fulfil some minimum standards of objectivity and allow statistical inference. Having estimated VAR models for two countries, it would seem natural to consider the evidence of differing monetary transmission strong if, in the space of all possible identification schemes, there do not exist two schemes, one for each country, such that i) they produce plausible response functions for their respective countries and ii) the differences in the resulting responses to standardised shocks can be reasonably explained by random variation.

This criterion may seem weak, and indeed, it is. Even if a vast majority of the viable identification schemes in one country showed, say, a stronger response to monetary shocks than those in the other country, as long as there is a single pair of schemes which produce responses statistically similar enough, then the criterion would fail to find a difference. However, it is difficult to think of any statistically justified test with a higher power.<sup>22</sup>

To apply this approach in practice, one has to first decide what criteria to apply when judging the plausibility of the response functions. To minimise the amount of arbitrariness involved, we suggest that two general principles should be adhered to. First, the definition of plausibility should be the same for each country. Clearly, the use of differing "plausibility windows" for different countries would bias the results. Second, we think that it is important to use a window which reflects the true degree of uncertainty in the priors, which usually means a fairly wide window *at every maturity*. In particular, we think that on the basis of our prior knowledge it would be unjustifiable to restrict the impulse responses (or some function thereof) to pass through a single point — as would normally be the case if either the short-run or the long-run effects were constrained to zero.<sup>23</sup>

In this section we demonstrate the approach by applying it in the comparison of monetary policy transmission in three countries: France, Germany, and the UK.

## **5.2 Data and the models**

Our VAR models consist of three endogenous variables: real GDP, the CPI index, and the three-month money market interest rate. For Germany, we included as an exogenous variable the nominal world commodity price index, which was measured in US dollars to guarantee a reasonable amount of true exogeneity. Our quarterly data come from the OECD and the IMF (IFS) databases, are seasonally non-adjusted, and cover the period from the beginning of the 1970s to the third quarter of 1997.

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<sup>22</sup> In the spirit of Bayesian econometrics, one could envisage an approach in which each identification scheme would be assigned a prior probability. In cross-country comparison, one would then take into account both the parameter uncertainty and the (prior) probability distribution over the identification schemes. We have not pursued this avenue.

<sup>23</sup> Restricting the within-period effect of monetary policy on the price level and GDP to zero might be justifiable with very high-frequency data. Given that the available data is usually quarterly we do not think small within-period effects can be excluded.

The output and price variables entered our VAR models in logarithmic differences, interest rates in levels. Four lags were found to be sufficient for UK and France while five lags was used for Germany. We did not search for co-integration, partly because theory would not lead us to expect any long-run relationships, and partly because the thrust of this paper is elsewhere.

An obvious complication in the cross-country comparison is that the estimation of a separate model for each country would enable rigorous statistical evaluation of the observed differences only in case the models were independent, i.e. if there is no cross-model correlation of residuals. Given the interdependencies of the European economies, this would likely be an incorrect assumption. Indeed, with our set of VAR models, the residuals of the different systems turned out to be highly correlated, in particular between the residuals of the GDP and interest rate equations. Hence, to account for such dependencies and to enable cross-country comparison, we estimated the three VAR models as a single system of seemingly unrelated equations.<sup>24</sup>

Allowing cross-country interdependencies adds another dimension to the identification problem. Instead of the structural decomposition of, in our case, three 3\*3 residual covariance matrices one is faced with a single 9\*9 covariance matrix with many non-zero off-diagonal elements. One could imagine decomposing the effects to, say, international monetary shocks, country-specific monetary shocks, etc. We chose the simpler approach of concentrating on the familiar single-country identification problem - i.e. for each country, we performed the structural decomposition of the 3\*3 covariance (sub-)matrix. The resulting "primary" shocks were orthogonal within the same sub-system but not necessarily between different systems. For example, the "primary monetary shocks" remained intercorrelated between countries. We think that structural identification at the multi-country level might be an interesting path for future work.

The estimation of the VARs as a system changed the point estimates of the parameters only marginally, but the standard errors of the parameters narrowed notably. Also, some effect was observed when simulating the confidence intervals for the cross-country differences in the impulse responses. Due to the positive cross-country correlation of the estimated impulse responses (mainly in the GDP responses between France and Germany), some of the confidence intervals of the differences were narrower than what would have been the case if the simulations had been based on random simulations on independent models.

### **5.3 Estimation results**

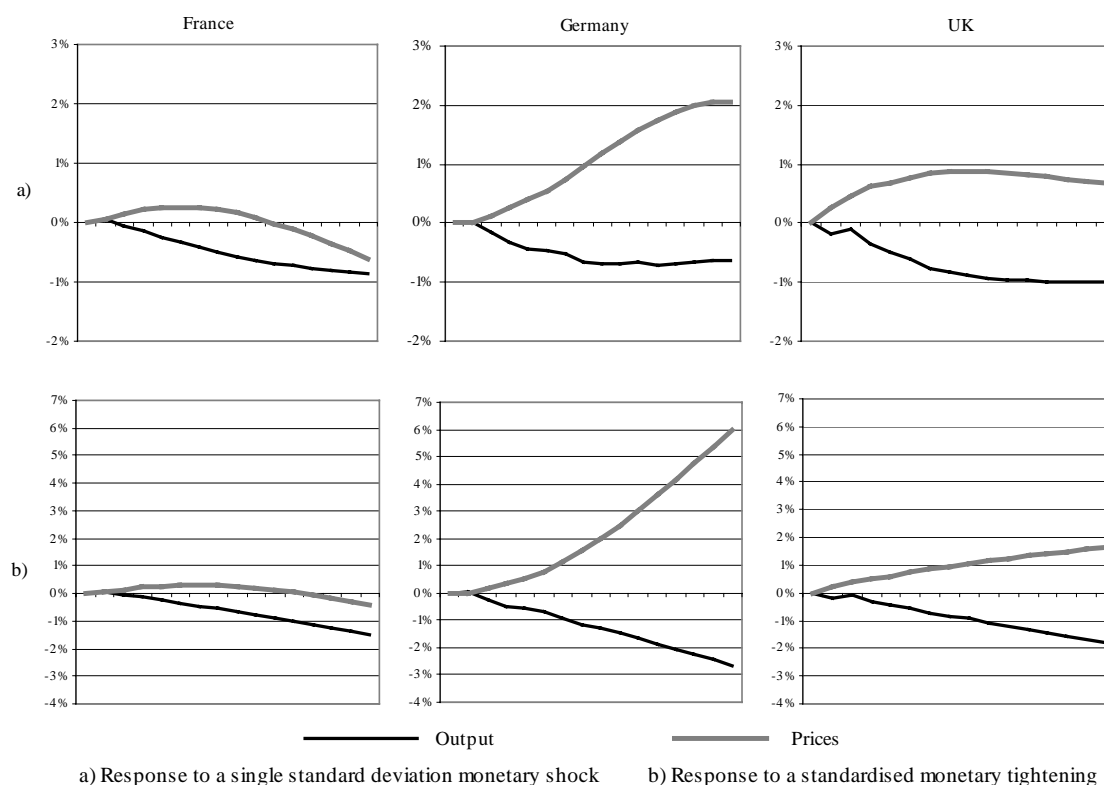
#### **Baseline model**

To create a baseline model, we again applied the most commonly used identification scheme, i.e. the triangular (Cholesky) decomposition with the interest rate variable ordered as the last. In other words, we defined a monetary policy shock as an

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<sup>24</sup> As an exception, the analysis of the long -run properties of the models was done within the framework of independent country-specific VARs.

**Chart 6** Impulse responses for the baseline model



unanticipated change in interest rate with no within-period effect on prices or output. The resulting responses of prices and output are plotted in chart 6. The upper row presents the responses to a monetary shock of one standard deviation, the size and duration of which varies from one country to another. The second row presents the impulse responses to a standardised monetary tightening. The standardised monetary tightening is defined as in the previous section; i.e. it consists of a series of monetary shocks resulting in an interest rate increase of one percentage point which is sustained at that level for the whole period of four years.

The estimated responses of real output to a monetary shock appear, at first sight, rather plausible. They are reasonably well in line with the results obtained by Ramaswami and Sloek (1998) using the same identification scheme. It appears that the response of output to a monetary shock is the largest either in Germany or in the UK, depending on the definition of the shock, while the response in France is the smallest in both cases.

Unfortunately, a glance at the price responses reveals that this "plausible" outcome is likely to be an illusion. For each of the three countries, we observe the "price puzzle", i.e. the response of the price level to a monetary tightening is positive over the whole simulation horizon, except for France, for which the effect turns negative after two years. To us, this is clear evidence of mis-identification of monetary policy shocks. Our interpretation is that the model picks out as monetary policy shocks changes in interest rates which were actually taken as a response to emerging inflationary pressures. As a result, the estimated responses are not responses to monetary shocks, but responses to some mixture of true monetary shocks and other shocks, for example, supply shocks. Hence, a more careful structural identification seems required.



## Long-run properties

Notwithstanding the problems of using long-run (i.e. infinite horizon) restrictions to identify VARs (see the discussion in section 3), the fact that they play an important role in the literature warrants an examination of the long-run properties of our VAR models.

Essentially, we would prefer that the models deemed plausible exhibit an approximate long-run neutrality of money. It is not obvious in which form such a requirement should be introduced. For various reasons, we do not believe that imposing restrictions directly on the point estimates of the long-run effects is justified. First, as discussed above, neither theoretical considerations nor the econometric structure are robust enough to justify requiring exact long-run neutrality. Furthermore, examination of the results shows that the long-run effects exhibit fairly serious instability, in the sense that small perturbations in the identifying restrictions can cause large changes in the estimated long-run effects.

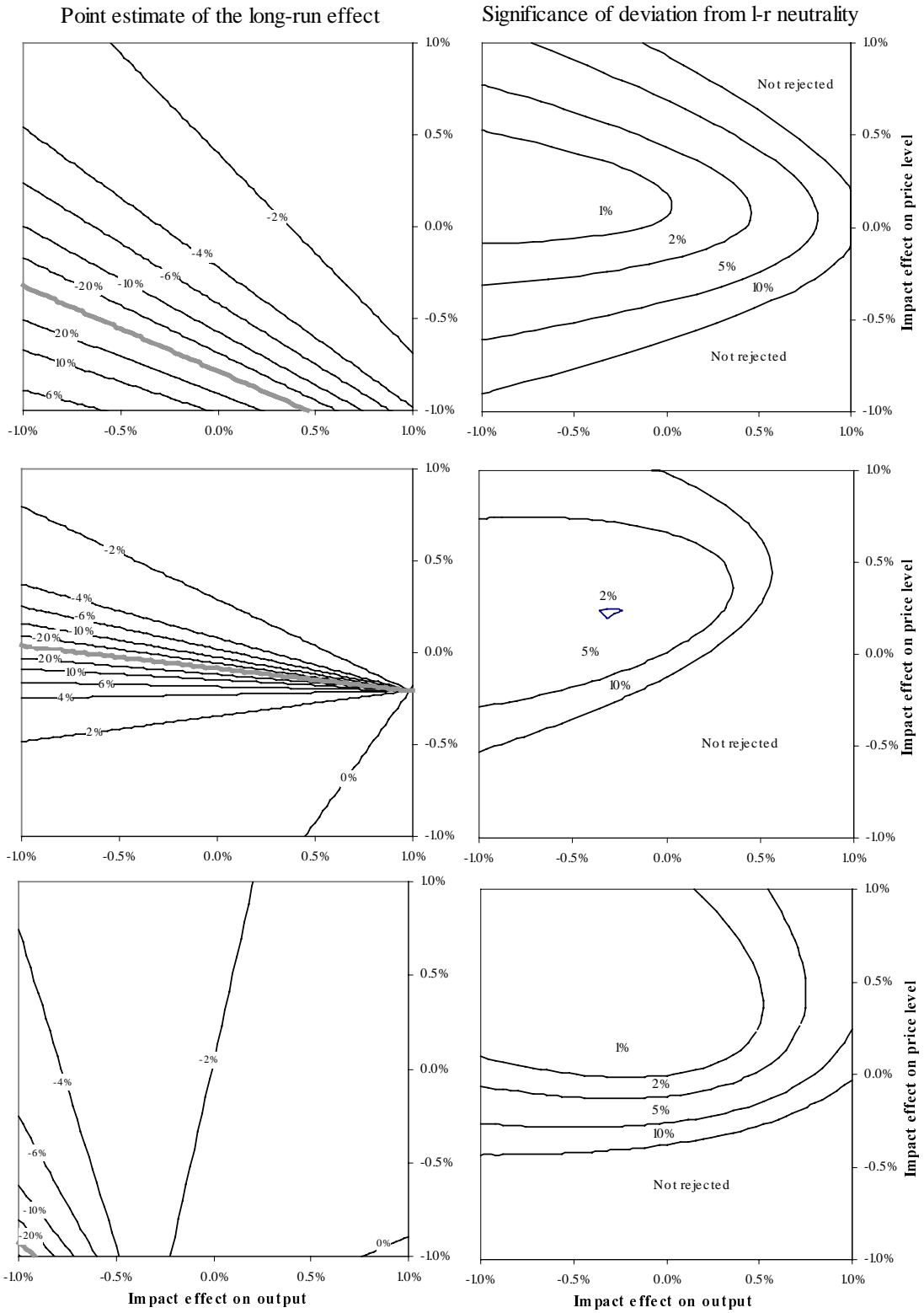
Finally, there is the question of the precision with which the long-run effects are estimated. Intuitively, the fact that over finite horizons, the standard errors of the estimated impulse responses tend to increase drastically with the length of the horizon, would suggest that for an infinite horizon the precision of the estimates should be poor. This is not straightforward to verify rigorously since although the point estimates of the long-run effects are easy to derive, their standard errors — or preferably their full probability distributions, which are not likely to be symmetric — are not. King and Watson (1997) present a method for calculating approximate standard errors in a 2-variable VAR, but that approach is not viable in our larger model.

We developed a somewhat simpler approach to verify the consistency of different identification schemes with long-run neutrality. Long-run neutrality implies a particular, highly non-linear, condition on the VAR coefficients. By linearising this condition, one can derive an asymptotic Wald-type test for the null hypothesis that the long-run effect of a particular shock on output is zero. (See the Appendix for details.) Hence, although we cannot calculate the confidence intervals for the long-run effects, we can test whether the deviation from long-run neutrality is statistically significant for each identification scheme.

Our findings for each country are summarised in Chart 7, which condenses a substantial amount of information on the long-run properties of the data. Basically, the chart links the short-run properties of each identification scheme to the corresponding long-run implications. In each panel of the graph, the horizontal axis defines the impact (within-quarter) effect of a 1 %-point monetary tightening on output, and the vertical axis its impact effect on the price level. Together, these two define a unique identification scheme.

For each identification scheme determined by its short-run effects, the left-hand panels define the point-estimate of the long-run effect on output. The monetary tightening is, once again, scaled to yield a one percentage point increase in the short interest rate, sustained for four years. The thick grey line marks the singularity frontier, on which the long-run effect switches from plus to minus infinity, i.e. is

**Chart 7** Long-run properties



undetermined.<sup>25</sup> The right-hand panels provide the marginal significance level of the deviation of the estimate from zero, as given by the asymptotic Wald test.

For example, the traditional "no-within-period-effects" identification scheme would be found at the origin of each panel. The upper left-hand panel shows that such identifying assumptions would imply that in Germany, the monetary tightening would have a negative long-term effect of nearly 4% on output. From the corresponding right hand panel it can be read that this point estimate is significantly different from zero at the 2% significance level. For France, the estimated long-term effect would be about -6%, but just outside the 5% significance frontier. Finally, for the UK, the point estimate is just -2%, but nevertheless significantly (at 2% level) below zero.

In order to obtain a point estimate close to long-run neutrality, one would have to accept ludicrously high impact effects. For Germany, if one retains zero impact effect on output, the implied negative impact effect on price level would be approximately 15%. Conversely, maintaining a zero within-period effect on price level would have to be counterbalanced by a more than 2% positive impact effect on output in order to yield long-run monetary neutrality. For France and the UK, the numbers are somewhat smaller, but nevertheless nonsensically high.

If long-run restrictions so manifestly fail to yield sensible results, why is it then that many studies have seemingly successfully exploited them, producing intuitively sensible outcomes? While it may be that better data and/or modelling techniques (e.g. the inclusion of a larger number of variables) may have a role to play, we are inclined to look for other explanations.

In many studies, the set of results reported is not complete, in the sense that only part of the responses are shown. Typically strong restrictions on monetary policy effects on one variable (output) get reflected in an unrealistic response of another (for example, prices). Without a full set of results it is difficult to judge the overall plausibility of the model.

It is also possible that the focus on non-standardised monetary shocks can mask implausible outcomes. We suspect that particularly in high-dimensional models, the model can create a shock with long-run monetary neutrality by finding an interest rate path on which the high initial rates are counterbalanced by lower interest rates later, so that there is no effect on the *cumulative* interest rate. In other words, each monetary shock cancels its own initial effects. In our framework this is not possible, as the path of the interest rate is fixed.

To summarise, it does not seem possible to formulate for the three countries a uniform criterion for long-run effects of monetary policy while still allowing sensible contemporaneous effects. Our interpretation is that the long-run estimates which our VAR-models yield are simply too fragile to be helpful in identifying the structural

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<sup>25</sup> Intuitively, this happens when an identification scheme produces a zero *cumulative* impulse response of interest rate; i.e. when initially higher interest rates are exactly matched by lower interest rates later. In such a case the standardisation procedure, to get a comparable four-year tightening, creates an infinite series of monetary shocks which do not converge to zero, thereby effectively multiplying the long-run effect of the original, non-standardised shock by infinity.

shocks. Hence, in what follows, we will only use short and medium-term plausibility conditions.

### Cross-country comparison

We proceeded to apply in practice the methodology outlined at the beginning of this section by defining a window of plausible responses for prices and output at the short and medium run. All plausibility criteria apply to the accumulated responses to a standardised monetary tightening of one percentage point, as defined above. The window was defined for three different horizons: within-period, 2 years, and 4 years. To avoid ruling out differences by definition, we required a plausible candidate for monetary shock to fulfil the following relatively loose criteria:

		Within period	After 2 years	After 4 years
Output response	Upper	<0%	<0%	<-1.0%
	Lower	>-0.6%	>-3.0%	>-7.0%
Price response	Upper	<0.5%	<0%	<-1.0%
	Lower	>-0.6%	>-3.0%	>-7.0%

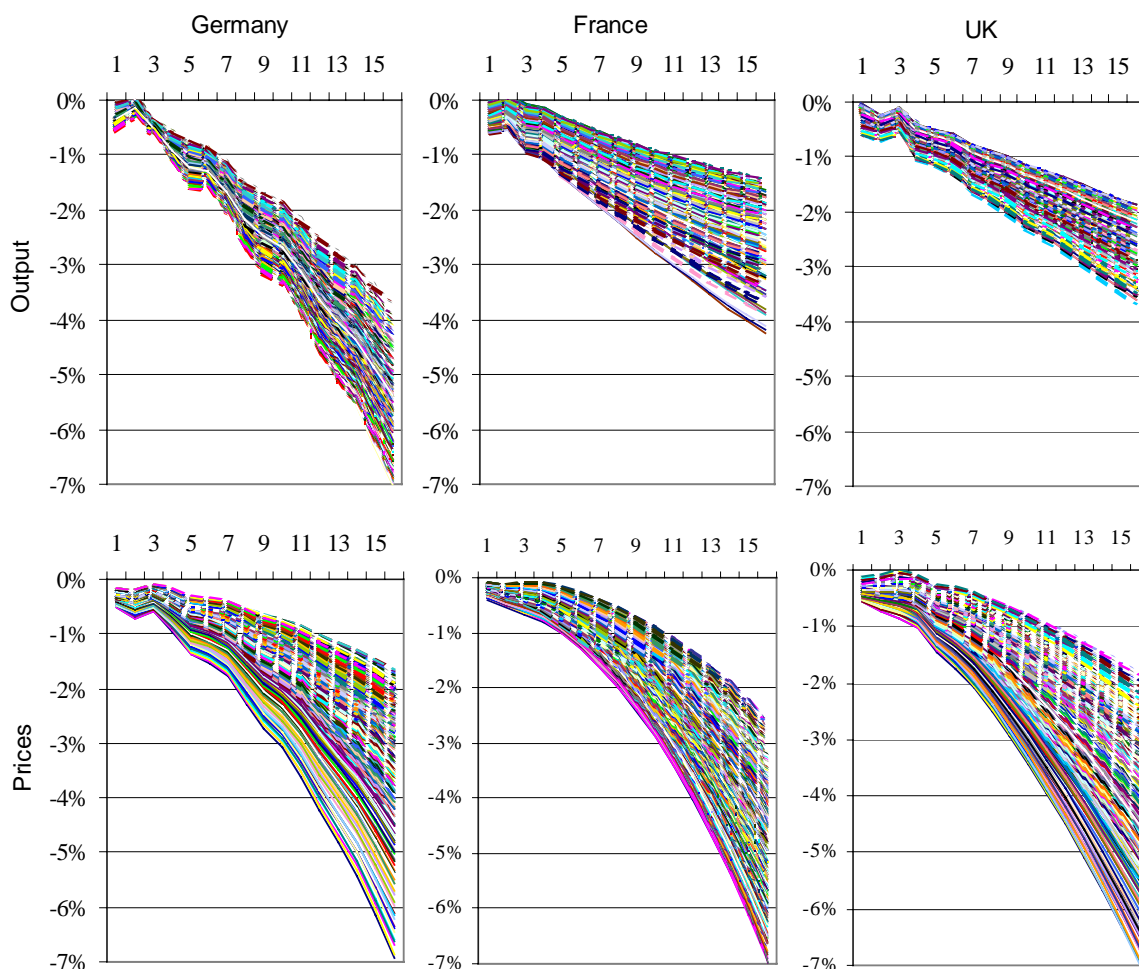
To identify the viable identification schemes, we searched the space of all identification schemes by rotating the orthogonal base defining the three shocks in steps of  $\pi/100$  (i.e.  $1.8^\circ$ ). Of the 10 000 possible identification schemes, the number of schemes which fulfilled our plausibility criteria was 136 for Germany, 445 for France, and 184 for the UK. The impulse responses corresponding to these schemes are plotted in chart 8. In general, we observe that the plausible price responses of Germany and the UK span roughly similar area while the responses for France are more widely distributed. With regard to the responses of output, the responses are similar for France and the UK, while for Germany, the viable responses are more pronounced.<sup>26</sup>

One feature shared by each country is that the immediate response of the price level is always strictly, and mostly considerably, negative. We are somewhat uncomfortable with the abruptness of the effects. Although we do not think that it is appropriate to exclude the possibility of within-period effects of monetary policy on prices, our subjective priors would have been better consistent with a more gradual onset of the effect.<sup>27</sup>

<sup>26</sup> As can be observed by comparing chart 8 to charts 4 and 5, the impulse responses for UK have changed somewhat. This is because the impulse responses in chart 7 are based on the coefficients from system estimation in differences, while charts 4 and 5 come from OLS estimation in levels.

<sup>27</sup> One explanation is that our three-variable model is too small to allow the identification of "pure monetary shocks", in which case we would actually measure a mixture of responses to a number of shocks. Perhaps our monetary shocks get mixed with policy responses to "financial shocks", i.e. turbulence in the financial markets with direct effects on prices and output. Including a narrow

**Chart 8.** Feasible responses to a standardised monetary tightening



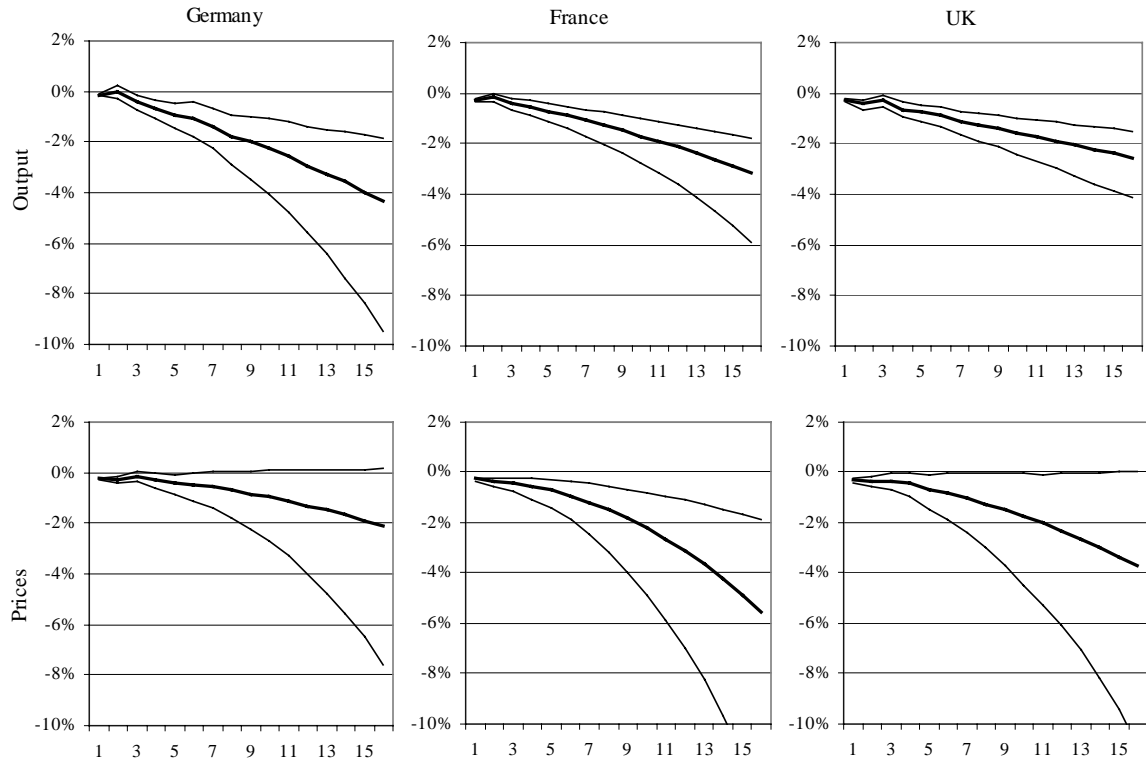
Equipped with a rough knowledge about the plausible area for each country, we proceeded to the statistical comparison of a set of three identification schemes. As explained above, we wanted to pick three identification schemes as similar as possible. As we found no computationally feasible formal method to identify the statistically most similar triple, we used visual inspection to select three identification schemes which were closely similar and, at the same time, corresponded as closely as possible with our priors. The impulse responses corresponding to the selected identification schemes are plotted in chart 9.<sup>28</sup> Each sub-plot includes the median response together with the 5% and 95% bounds based on 1000 random draws from the joint probability distribution of the SUR system.

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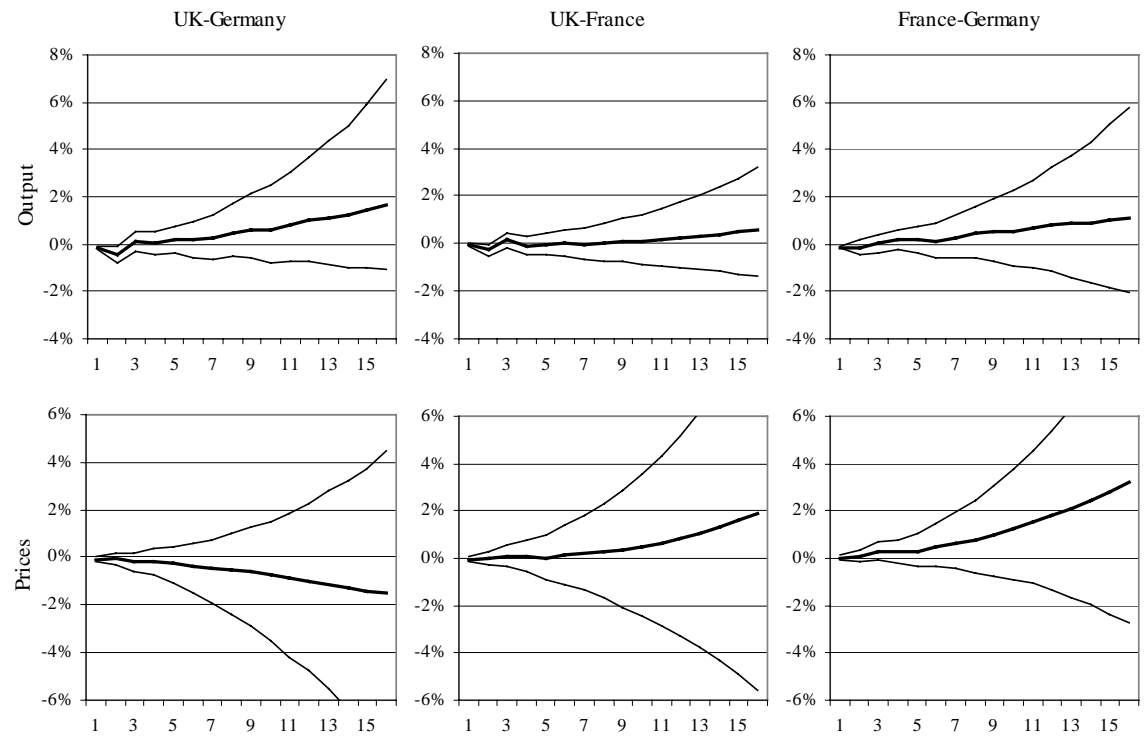
monetary aggregate or the central bank's foreign reserve position in the model might help to better identification of monetary shocks.

<sup>28</sup> We ended up with a triple selected from broadly the same part of the identification space. The rotation matrices were  $Z(-0.63, -0.09, 0)$  for Germany,  $Z(-0.48, -0.42, 0)$  for France, and  $Z(-0.65, -0.22, 0)$  for the UK. The resulting impulse responses are linear combinations of the baseline impulse responses based on triangular (Cholesky) decomposition. For example, with the parameter combination chosen for Germany, the impulse response of each variable to what is defined as a monetary shock is a combination of its responses to the three baseline shocks (i.e. output, prices, and interest rate) with weights  $(-0.09, -0.59, 0.80)$ .

**Chart 9** Selected impulse responses to a standardised monetary tightening



**Chart 10** Differences in monetary transmission



The results are qualitatively in line with what is usually obtained with VARs. The confidence intervals become, in some instances, quite wide for longer horizons. In particular, the models seem to have little or no predictive content on the price response in France beyond a horizon of a year or two. The output responses are generally more tightly estimated than price responses, except for Germany, for which both responses have a similar degree of uncertainty.

In relation to the degree of statistical uncertainty in the estimates, the cross-country differences appear to be small. This is confirmed by chart 10 which plots the cross-country differences in the responses, together with the 90% confidence intervals. Taking into account the parameter interdependencies between the individual VAR models (i.e. performing the random simulation at the system level) resulted in slightly tighter confidence intervals for the differences than would have been the case with random draws from independent distributions. Generally, zero is firmly within the confidence intervals, except for very short horizons, for which some small, but just marginally significant differences can be measured. However, the significance of these differences is exaggerated by the fact that they are mainly determined by the rotation matrices, which were kept unchanged across the simulations.

In conclusion, our empirical analysis fails to produce evidence of the existence of significant differences in monetary policy transmission among our sample of countries. The sets of identification schemes which fit into our definition of plausible differ somewhat in their general shape. However, a large enough overlap remains to allow the possibility of statistically indistinguishable responses.

Obviously, there may be room for improvement in our VAR models. Our three-variable model is quite possibly too narrow to allow the identification of the pure monetary shocks. If the behaviour of our variables is driven by four or more other shocks (with linearly independent effects), then the effects of pure monetary shocks cannot be identified within our model.<sup>29</sup> The appropriate response would be to include additional endogenous variables in the model. Unfortunately, the modelling approach applied here gets increasingly difficult to apply in practice as the dimensionality of the model increases.

A similarly potentially useful but computationally difficult extension of our approach would involve a closer examination of implied monetary policy response functions. It may be that a particular identification scheme, which produces plausible responses of money and output to monetary policy, actually involves implausible responses of monetary policy to the shocks driving the economy. By defining a "plausibility window" also for monetary policy responses could therefore, in principle, further

policy shocks falls short of identifying unique responses to the other shocks. In order to identify the latter, we need to choose also the third parameter  $C$ . In other words, for each pair of responses of output and prices to monetary policy shocks, there is a continuum of possible monetary policy response functions, arranged by the parameter

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<sup>29</sup> If the true process generating our three variables is  $X_t = B(L)e_t$ , where the shock vector  $e_t$  is a  $k \times 1$  random vector,  $k > 3$ , then the process does not have an autoregressive representation in our three variable framework and the responses to the primary shocks cannot be identified.

C. If none of the monetary policy responses on the continuum is acceptable, then that identification scheme can be rejected. Needless to say, verifying this for a large number of identification schemes would be computationally very heavy.

## 6 Conclusions

In recent years, a large number of econometric studies have sought to identify cross-country differences in monetary policy transmission in the pre-EMU regime. Among the most common methodologies in these studies has been the one of a vector autoregressive (VAR) model. The results have generally varied a great deal and, as a result, no consensus regarding the likely extent or nature of these differences has arisen. In this paper we explore one particular source of differences in the results of the VAR models, namely the structural identification of the econometric model.

Our results show that the choice of the structural identification scheme has a substantial effect on the results. The set of plausible identification schemes is typically still very large after taking into account theoretical (and ad hoc) information, and consequently the resulting set of feasible impulse response functions exhibits wide variation. In our view, this finding casts doubt on the studies that, based on the comparison of results produced by a single identification scheme, claim to find cross-country differences in the monetary policy transmission mechanism.

We propose an alternative approach which explicitly recognises the degree of uncertainty involved in choosing the identification scheme. Our approach consists of defining a "plausibility window" for the impulse responses and accepting all impulse responses that fall within this window as viable representations of the monetary policy transmission. The analysis of cross-country differences in monetary transmission would then be based on these plausible sets of impulse responses. We apply this approach to the analysis of monetary policy transmission in three European countries, namely France, Germany, and the UK. Our results show that it is possible to choose a set of identification schemes such that the resulting impulse responses are broadly plausible and, at the same time, do not differ to a statistically significant extent between the three countries. Hence, the conclusion is that our data does not enable us to confirm the existence of significant cross-country differences in monetary transmission for those countries.

We do not claim to have shown that no cross-country differences exist in the monetary transmission of the EU countries. On the contrary, in light of the existing differences in financial market structures and price and wage mechanisms among the EU countries, we consider it quite plausible that monetary transmission differs as well. Rather, our point is that in order to prove the existence of such differences, a much more careful handling of certain econometric issues is required. So far, the existence of the differences has not been established. Given the complexity of the issue and the extent of identification uncertainty, we have serious doubts about whether macro-level econometrics will ever be able to resolve the issue.



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## Appendix: Testing long-run neutrality of monetary policy

The estimated VAR model can be written as in equation (1)

$$(A1) \quad [I - A(L)]X_t = u_t, \quad E(u_t u_t') = \Sigma.$$

The purpose is to test whether identifying the model by a contemporaneous structure given by the matrix  $Z$ , where  $ZZ' = \Sigma$ , is consistent with zero long-run effect of monetary shock on output. We write the model:

$$(A2) \quad [I - A(L)]X_t = Ze_t, \quad E(e_t e_t') = I.$$

The long-run effects of the shock vector  $e_t$  are given by the matrix  $[I - A(1)]^{-1}Z \equiv L_t$ . The long-run matrix  $L$  is highly non-linear in the original parameters and hence the distribution of its elements cannot be directly solved. However, denoting the inverse of  $L$  by  $K \equiv Z^{-1}[I - A(1)]$ , it is easy to verify that the upper right-hand element  $l_{13}$  of the matrix  $L$ , giving — with our ordering — the long-run effect of a monetary shock on output, is of the form

$$(A4) \quad l_{13} = \frac{(k_{12}k_{23} - k_{13}k_{22})}{X},$$

where  $k_{ij}$  denote the elements of  $K$ , and  $X$  is a finite-valued nonlinear function of the elements of  $K$ . Hence, testing the long-run neutrality of money boils down to testing the null hypothesis:

$$(A5) \quad h_0 : H(\mathbf{k}) \equiv k_{12}k_{23} - k_{13}k_{22} = 0.$$

where  $\mathbf{k} = \text{vec}(K)$ . Linearising around the OLS estimate of  $\mathbf{k}$  yields

$$(A6) \quad h_0 : H(\tilde{\mathbf{k}}) + \left. \frac{\partial H(\mathbf{k})}{\partial \mathbf{k}} \right|_{\mathbf{k}=\tilde{\mathbf{k}}} (\mathbf{k} - \tilde{\mathbf{k}}) = 0,$$

To derive a test for this restriction, notice first that  $A(1)$  is a simple summation matrix of the original parameters, so its probability distribution is asymptotically normal and easy to calculate. Using the vector notation, we write the distribution of  $A(1)$  as

$$(A7) \quad \text{vec}[\tilde{A}(1) - A(1)] \underset{\alpha}{\sim} N(0, \tilde{\Omega}_L)$$

where the tilde denotes the OLS estimate and  $\tilde{\Omega}_L$  is the appropriate parameter covariance matrix. The distribution of  $\tilde{\mathbf{k}}$  can be derived as follows:

$$(A8) \quad \begin{aligned} \tilde{\mathbf{k}} - \mathbf{k} &= \text{vec}[Z^{-1}(A(1) - \tilde{A}(1))] \\ &= (I \otimes Z^{-1}) \text{vec}[A(1) - \tilde{A}(1)] \underset{\alpha}{\sim} N[0, (I \otimes Z^{-1})\tilde{\Omega}_L(I \otimes Z^{-1})']. \end{aligned}$$

Using this, the asymptotic distribution of the constraint is

$$(A9) \quad H(\tilde{\mathbf{k}}) - H(\mathbf{k}) \underset{\alpha}{\sim} N[0, H'(\tilde{\mathbf{k}})(I \otimes Z^{-1})\tilde{\Omega}_L(I \otimes Z^{-1})'H'(\tilde{\mathbf{k}})'],$$

where  $H'(\tilde{\mathbf{k}}) \equiv \left. \frac{\partial H(\mathbf{k})}{\partial \mathbf{k}} \right|_{\mathbf{k}=\tilde{\mathbf{k}}}$ . This yields the Wald test statistic

$$(A10) \quad H(\tilde{\mathbf{k}})' [H'(\tilde{\mathbf{k}})(I \otimes Z^{-1})\tilde{\Omega}_L(I \otimes Z^{-1})'H'(\tilde{\mathbf{k}})]^{-1} H(\tilde{\mathbf{k}})$$

which is asymptotically  $\chi^2(1)$  distributed under the null.