Monetary Transmission in Diverse Economies

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Monetary Transmission in Diverse Economies

The transmission mechanism of monetary policy explains how monetary policy works – which variables respond to interest rate changes, when, why, how, how much and how predictably. It is vital that central banks and their observers, worldwide, understand the transmission mechanism so that they know what monetary policy can do and what it should do to stabilise inflation and output. The volume sets out different aspects of the transmission mechanism. Some chapters scrutinise the relevance of practical issues such as asymmetries, recent structural changes and estimation errors using data on the USA, the euro area and developing countries. Other chapters focus on modelling crucial aspects such as productivity, the exchange rate and the monetary sector. These issues are counterpointed by contributions that analyse contemporary monetary policy in Japan and the UK.

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V. Anton Muscatelli and Carmine Trecroci

5.1 Introduction

A considerable empirical literature has emerged on the estimation of policy reaction functions and the identification of the underlying preferences of monetary authorities (see Groeneveld, Koedijk and Kool, 1996; Muscatelli and Tirelli, 1996; Clarida and Gertler, 1997; Clarida, Galí and Gertner, 1998; Favero and Rovelli, 1999; and Muscatelli, Tirelli and Trecroci, 1999). Some of these contributions examine whether recent changes in institutional structure, such as the shift to inflation targeting, have had an impact on the conduct of monetary policy.[1] The evidence is mixed. For instance, Muscatelli, Tirelli and Trecroci (1999) show that there is only slight evidence that the introduction of inflation targeting affected forward-looking policy reaction functions in the United Kingdom, New Zealand, Sweden and Canada. In contrast they find some evidence of policy instability in Japan and the United States in the 1980s and 1990s, even in the absence of institutional change.

Of course one would also expect significant shifts in monetary policy that bring about a reduction in inflation expectations to affect the transmission mechanism of monetary policy. The standard New Keynesian model of aggregate demand and supply,[2] which has been used extensively for policy analysis (see Svensson, 1997; Rudebusch and Svensson, 1999; McCallum and Nelson, 1999a,b; and Rudebusch, 2000), suggests that forward-looking expectations are important on both the demand and the supply side. A typical formulation of a New Keynesian model is

$$y_t = E_t y_{t+1} - \alpha_1(i_t - E_t \pi_{t+1}) + \varepsilon_{1t} \qquad (5.1)$$

$$\pi_t = \beta_1 E_t \pi_{t+1} + (1 - \beta_1)\pi_{t-1} + \beta_2 y_t + \varepsilon_{2t}, \qquad (5.2)$$

where y is the output gap, i is the nominal interest rate, π is the inflation rate and E is the expectations operator. In empirical applications, lags of output

could be added to equation (5.1) to capture costly adjustment of habit persistence, and a more complex lagged adjustment of inflation could be considered in equation (5.2). The presence of forward-looking expectations in these models means that any changes in monetary policy should lead to structural breaks in the aggregate supply relationship (5.2) or in the intertemporal aggregate demand equation (5.1). This is of course the standard Lucas (1976) critique. Hutchison and Walsh (1998) find that, in the case of New Zealand, the short-run output–inflation trade-off in a relationship like that in equation (5.2) increased in the 1990s after the monetary reform. For the United States, Rudebusch and Svensson (1999) find little evidence of structural breaks[3] in their (backward-looking) models of output and inflation estimated over a sample period 1961(Q1)–1996(Q4).

This paper makes two contributions. First, for the United States and the United Kingdom we check for any evidence that shifts in the transmission mechanism, as represented by the aggregate demand and supply relationships, have resulted from shifts in agents' expectations. We use a class of invariance tests developed by Engle and Hendry (1993) to test whether the Lucas critique has any force in this case. We find evidence of shifts in the output and inflation models in the past two decades. Second, because the invariance tests are not designed to examine the timing of shifts in the transmission mechanism, we use an alternative method to detect contemporaneous shifts in interest rate policy reactions and in the output and inflation equations. We estimate a Bayesian vector autoregression (VAR) of the monetary transmission mechanism for the United States and the United Kingdom, using a simple trivariate specification with the interest rate, output gap and inflation rate. In contrast to standard full-sample VAR estimates, we find the policy rules that emerge from our Bayesian VAR estimates to be more plausible and interpretable in terms of policy reactions. Our results show that policy in the United Kingdom and the United States has evolved since the late 1980s. We also detect some minor shifts in the way in which output and inflation have responded to policy, which may indicate expectations effects. Evidence, therefore, does exist that the Lucas critique is important in understanding the transmission mechanism. And, although it may have been less important in the relatively benign macroeconomic environment of the 1990s, it may become more significant in the future. The Lucas critique may be especially relevant at a time when the spectre of oil shocks has returned.

The rest of this paper is structured as follows. In section 5.2 we report our invariance tests of conditional models for the aggregate demand and supply relationships. In section 5.3 we present our Bayesian VAR results. Section 5.4 concludes the paper.

5.2 Invariance tests

For our purposes, we wished to test whether forward-looking expectations enter into equations (5.1) and (5.2), the output and inflation relationships.[4] The approach followed here entails estimating backward-looking conditional models for output and inflation and testing whether variations in the moments of the regressors in each model influence the parameters of the backward-looking conditional model.[5] Thus, we first estimate the output and inflation equations as backward-looking models, using autoregressive distributed-lag formulations:[6]

$$y_t = c_1 + \sum_{i=1}^{n} \alpha_{yi} y_{t-i} + \sum_{i=0}^{n} \alpha_{ri} \bar{r}_{t-i} + \varepsilon_{1t} \tag{5.3}$$

$$\pi_t = c_2 + \sum_{i=1}^{n} \beta_{\pi i} \pi_{t-i} + \sum_{i=0}^{n} \beta_{yi} y_{t-i} + \varepsilon_{2t}, \tag{5.4}$$

where c_i are constants and $\bar{r}$ is a measure of the real interest rate, calculated using the current inflation rate, $\bar{r}_t = (i - \pi)_t$. If forward-looking expectations of the interest rate and inflation are important in determining current output and inflation, then these equations will not be invariant to changes in policy regime because they will be convolutions of the 'deep' parameters of the forward-looking models and the forecasting equations. The impact of forward-looking expectations can be tested by checking if shifts in the first and second moments of the real interest rate or output affect the regressions in equations (5.3) and (5.4) (see appendix 1). To obtain measures of these moments, we fit marginal models for output and the real interest rate using simple autoregressive models.[7]

Table 5.1 shows the most parsimonious version of estimated equations (5.3) and (5.4). Our conditional models generally follow the specifications reported in Rudebusch and Svensson (1999) for the United States and in Hall et al. (1999) for the United Kingdom. Definitions of variables used can be found in the data appendix (appendix 2). In computing the output gap, we follow two alternative approaches. The first is to follow these earlier contributions in using the Bureau of Economic Analysis (BEA) estimate (y^{bea}, for the United States) and the OECD estimate (y^{oecd}, for the United Kingdom), of potential output, and computing the output gap as the deviation of actual from potential output as a percentage of potential output. The second approach is to fit a univariate structural time-series (STS) model (see Harvey, 1989) for output. This decomposes output into stochastic trend, cycle and irregular components, and we define potential output as the stochastic trend element. A convenient decomposition can be obtained by applying the Kalman filter on the trend component and using the one step ahead predictions of the state vector. The output gap measure obtained is labelled y^{kal}. The difference between the two output gap series in the case of the two countries is shown in figures 5.1 and 5.2. As can be readily seen, the

Table 5.1 *Backward-looking conditional models of output and inflation*

(a) United States

$$\begin{array}{lllll} y_t^{bea} = 0.031 & + 1.187 y_{t-1}^{bea} & - 0.279 y_{t-2}^{bea} & - 0.010 \bar{r}_{t-1} & \\ \quad (2.44) & (15.87) & (-3.81) & (-2.68) & \end{array}$$
$R^2 = .90 \quad \sigma = 0.0997$

$$\begin{array}{lllll} y_t^{kal} = 0.021 & + 0.512 y_{t-1}^{kal} & - 0.290 y_{t-2}^{kal} & - 0.007 \bar{r}_{t-1} & \\ \quad (1.467) & (6.81) & (-3.88) & (-1.56) & \end{array}$$
$R^2 = .59 \quad \sigma = 0.1187$

$$\begin{array}{lllllll} \pi_t = 0.119 & + 0.619 \pi_{t-1} & + 0.024 \pi_{t-2} & + 0.188 \pi_{t-3} & + 0.132 \pi_{t-4} & + 1.20 y_{t-1}^{bea} \\ \quad (0.82) & (7.98) & (0.266) & (2.09) & (1.706) & (4.46) \end{array}$$
$R^2 = .85 \quad \sigma = 0.972$

$$\begin{array}{llllll} \pi_t = 0.117 & + 0.711 \pi_{t-1} & + 0.027 \pi_{t-2} & + 0.228 \pi_{t-3} & + 1.06 y_{t-1}^{kal} \\ \quad (0.74) & (9.32) & (0.094) & (2.91) & (2.24) \end{array}$$
$R^2 = .84 \quad \sigma = 1.012$

Sample: 1958(Q4) to 2000(Q1).

(b) United Kingdom

$$\begin{array}{llllll} y_t^{oecd} = 0.001 & + 1.711 y_{t-1}^{oecd} & - 0.766 y_{t-2}^{oecd} & + 0.023 \bar{r}_{t-2} & - 0.893 q_{t-4} \\ \quad (-) & (24.95) & (-11.713) & (1.74) & (-1.90) \end{array}$$
$R^2 = .98 \quad \sigma = 0.325$

$$\begin{array}{lllllll} y_t^{kal} = 0.038 & + 0.454 y_{t-1}^{kal} & + 0.497 y_{t-2}^{kal} & + 0.199 y_{t-3}^{kal} & - 0.278 y_{t-4}^{kal} & + 0.077 \bar{r}_{t-2} \\ \quad (0.246) & (4.47) & (4.55) & (1.88) & (-2.91) & (2.33) \\ \quad - 2.195 q_{t-4} & & & & & \\ \quad (-1.96) & & & & & \end{array}$$
$R^2 = .88 \quad \sigma = 0.881$

$$\begin{array}{lllllll} \pi_t = 0.255 & + 1.081 \pi_{t-1} & - 0.165 \pi_{t-2} & - 0.132 \pi_{t-3} & + 0.513 y_{t-2}^{oecd} & - 0.381 y_{t-3}^{oecd} \\ \quad (1.60) & (11.28) & (-1.27) & (-1.65) & (3.14) & (-2.35) \\ \quad + 0.091 pim_{t-1} & - 0.146 ulc_{t-1} & + 0.255 ulc_{t-2} & & & \\ \quad (4.40) & (-3.30) & (5.56) & & & \end{array}$$
$R^2 = .97 \quad \sigma = 0.777$

Sample: 1977(Q1) to 1994(Q4).

BEA and OECD measures lead to measures of the output gap that are much smoother, but that show a greater variance over the sample period.

The real interest rate effect $\bar{r}$ is measured as an average effect:[8] the deviation of the four-quarter average nominal interest rate, i, from the four-quarter inflation rate in the relevant price index, that is, $\bar{r}_t = (1/4) \sum_{i=0}^{3} i_t - (1/4) \sum_{i=0}^{3} \pi_t$.

In the case of the United Kingdom we also add some additional regressors to the models as specified in equations (5.3) and (5.4), following Hall et al.

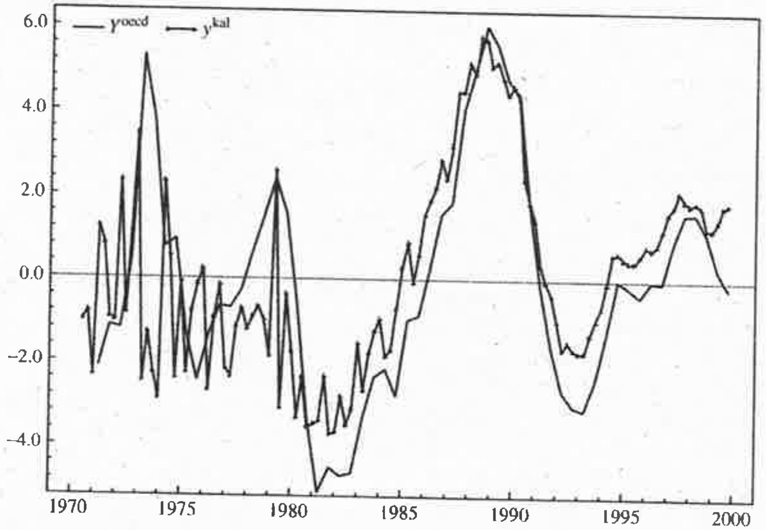

Figure 5.1 Output gaps: UK.

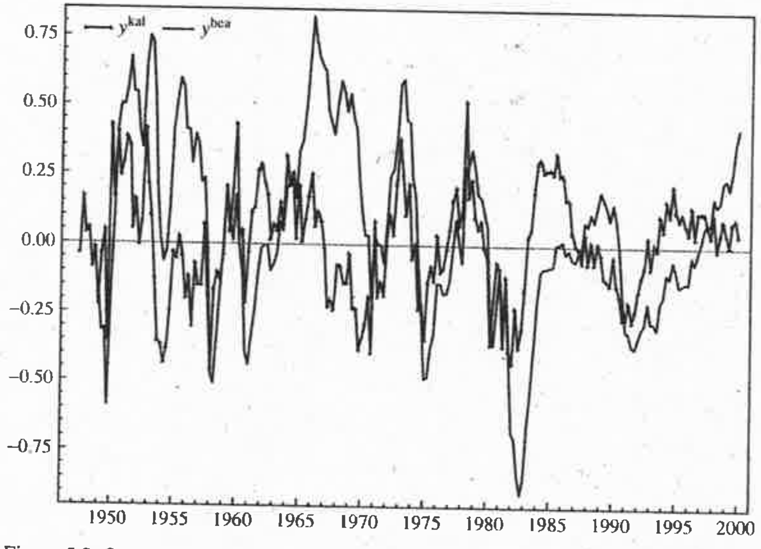

Figure 5.2 Output gaps: USA.

(1999).[9] In the case of the aggregate demand equation, we add a measure of the real effective exchange rate, *q*. In the case of the inflation equation, we add as regressors a series for the domestic currency value of import prices, *pim*, and a series on nominal unit labour costs, *ulc*. The former reflects the fact that the United Kingdom has a more open economy.

The results obtained for the United States are very similar to those reported in Rudebusch and Svensson (1999), which is not surprising since the sample we used is quite similar. The use of y^{kal} instead of y^{bea} does not yield very different results, although the fit using y^{bea} is slightly better; this is unsurprising given that it is a smoother series and has a higher sample variance than y^{kal}. The United Kingdom results are similar to those of the United States, except that the real interest variable has the wrong sign (but is insignificant if the y^{oecd} definition of the output gap is used). The additional regressors all have the right signs, and again the fit is generally better if the y^{oecd} measure is used. Next we conduct invariance tests using the y^{oecd} and y^{bea} measures of the output gap.

The marginal models fitted for the real interest rate and output are reported in table 5.2. It should be noted that these regression models are not designed to have a structural interpretation as outlined in appendix 1. They merely allow us to construct some series for the conditional mean of the interest rate and output, which can be used in the invariance tests. In order to obtain measures of the second moments of these variables, we fit fourth-order ARCH models for the fitted residuals ($\hat{\eta}_i^2$) from the regressions in table 5.2. In table 5.2 we also report the estimates of single-equation, forward-looking policy reaction functions[10] as a potential alternative to a naïve autoregressive model for the real interest rate. In practice, however, the best results from the invariance tests are obtained using the simple autoregressive model. As stressed in Engle and Hendry (1993), the marginal models can include dummy variables to capture important policy shifts. For the interest rate models we find that a dummy used to capture the 1992 Exchange Rate Mechanism (ERM) crisis improves the fit of the UK model, and a dummy used to capture the change in the Federal Reserve's operations in 1979–82 improves the fit of the US model.

Our invariance tests are reported in table 5.3. They are computed for two different sample periods for each country: the full sample available to us (1959(Q4)–1999(Q4) for the United States, and 1978(Q1)–1999(Q4) for the United Kingdom); and a shorter sample from 1984 through 1999. The latter period is usually seen as one of greater monetary stability, after the disinflation following the second oil shock.

The invariance tests show some evidence in favour of the Lucas critique for both the full sample and the post-1983 subsample. In the case of the United States, the null of invariance is rejected in both the output and inflation equations for the full sample period. If we focus only on the post-1983 period, however, we find evidence of the failure of invariance only in the case of the inflation

Table 5.2 *Marginal models of the interest rate, output and inflation*

(a) United States

Simple autoregressive model for interest rate:

$$r_t = 0.130 + 1.478 r_{t-1} - 0.532 r_{t-2} + 0.383 D79/82$$
$$\quad (2.73) \quad (22.12) \quad (-8.19) \quad (2.944)$$
$$R^2 = .97 \quad \sigma = 0.392$$

Forward-looking interest rate reaction function (1958(Q4)–1999(Q1)):

$$i_t = 0.091 + 1.062 i_{t-1} - 0.150 i_{t-2} + 0.124 \pi^e_{t+4} + 1.21 y^{\text{kal}}_t$$
$$\quad (0.49) \quad (13.75) \quad (-1.98) \quad (3.493) \quad (2.66)$$
$$R^2 = .92 \quad \sigma = 0.949$$

Forward-looking interest rate reaction function (1983(Q1)–1999(Q1)):

$$i_t = 0.062 + 1.264 i_{t-1} - 0.390 i_{t-2} + 0.910 \pi^e_{t+4} + 0.277 y^{\text{kal}}_t$$
$$\quad (0.33) \quad (12.19) \quad (-3.99) \quad (2.317) \quad (3.45)$$
$$R^2 = .96 \quad \sigma = 0.443$$

Simple autoregressive model for output gap:

$$y^{\text{bea}}_t = 0.003 + 1.195 y^{\text{bea}}_{t-1} - 0.072 y^{\text{bea}}_{t-2} - 0.224 y^{\text{bea}}_{t-3}$$
$$\quad (0.40) \quad (10.09) \quad (-0.38) \quad (-2.01)$$
$$R^2 = .94 \quad \sigma = 0.055$$

Sample: 1958(Q4) to 2000(Q1), unless otherwise stated.
Note: $D79/82$ is a dummy variable used to capture the impact on federal funds rates of the Federal Reserve's change in operating procedures between 1979 and 1982.

(b) United Kingdom

Simple autoregressive models for real interest rate:

$$r_t = 0.420 + 1.160 r_{t-1} - 0.234 r_{t-2} - 0.178 r_{t-3} + 0.167 r_{t-4}$$
$$\quad (2.65) \quad (11.24) \quad (-1.57) \quad (-1.22) \quad (1.92)$$
$$R^2 = .91 \quad \sigma = 0.893$$

Simple autoregressive models for real interest rate (1983(Q1)–1999(Q4)):

$$r_t = 0.51 + 1.27 r_{t-1} - 0.36 r_{t-2} - 0.52 D92/93$$
$$\quad (2.24) \quad (11.26) \quad (-3.17) \quad (-2.43)$$
$$R^2 = .90 \quad \sigma = 0.45$$

Forward-looking interest rate reaction function (1977(Q1)–1999(Q1)):

$$i_t = 0.773 + 0.825 i_{t-1} + 0.163 \pi^e_{t+4} + 0.148 y^{\text{kal}}_t$$
$$\quad (1.91) \quad (17.45) \quad (3.83) \quad (3.03)$$
$$R^2 = .88 \quad \sigma = 1.12$$

Forward-looking interest rate reaction function (1983(Q1)–1999(Q1)):

$$i_t = 0.99 + 0.629 i_{t-1} + 0.57 \pi^e_{t+1} + 0.220 y^{\text{kal}}_t$$
$$\quad (2.53) \quad (7.61) \quad (3.9) \quad (3.99)$$
$$R^2 = .91 \quad \sigma = 0.891$$

Simple autoregressive model for output gap:

$$y^{\text{bea}}_t = -0.011 + 1.80 y^{\text{oecd}}_{t-1} - 1.034 y^{\text{bea}}_{t-2} + 0.389 y^{\text{bea}}_{t-3} - 0.2 y^{\text{oecd}}_{t-4}$$
$$\quad (-0.034) \quad (17.15) \quad (-4.79) \quad (-1.80) \quad (-1.90)$$
$$R^2 = .98 \quad \sigma = 0.326$$

Sample: 1977(Q1) to 1999(Q4), unless otherwise stated.
Note: $D92/93$ is a dummy variable used to capture the impact on UK interest rates of the exit from the ERM, which allowed interest rates to fall rapidly between 1992(Q4) and 1993(Q4).

Table 5.3 *Invariance/superexogeneity tests*

Variable	USA: 1959(Q4)–1999(Q4)	USA: 1984(Q1)–1999(Q4)	UK: 1978(Q1)–1999(Q4)	UK: 1984(Q1)–1999(Q4)
Output equation				
$\hat{\eta}_r$	1.87*	0.56	1.50	2.60**
$\hat{x}^2_r$	−1.62*	0.11	0.07	−0.58
$\hat{\sigma}^2_{rr}$	0.57	−1.12	1.71*	−0.75
$(\hat{\eta}_r, \hat{\sigma}^2_{rr})$	−0.84	−0.14	−1.37	−2.40**
$(\hat{x}^2_r \hat{\sigma}^2_{rr})$	−0.09	−0.83	−1.24	0.27
Inflation equation				
$\hat{\eta}_r$	−5.18**	−3.19**	2.20**	−3.24**
$\hat{x}^2_r$	−0.38	−0.70	2.27**	0.37
$\hat{\sigma}^2_{rr}$	−1.54	−0.24	−0.69	0.49
$(\hat{\eta}_r, \hat{\sigma}^2_{rr})$	−0.03	0.16	−0.83	−0.66
$(\hat{x}^2_r \hat{\sigma}^2_{rr})$	−1.79*	−0.32	1.86*	0.68
$\hat{\eta}_y$	−1.54	1.10	0.34	0.7
$\hat{x}^2_y$	−0.46	−0.84	0.89	−0.63
$\hat{\sigma}^2_{yy}$	−0.55	0.48	−0.96	0.29
$(\hat{\eta}_y, \hat{\sigma}^2_{yy})$	2.02**	−1.15	−0.19	−0.06
$(\hat{x}^2_y \hat{\sigma}^2_{yy})$	0.07	0.53	0.36	2.01**

Note: Numbers reported in table are t-values for addition of variable to conditional models for output and inflation. The notation for the variables follows the description of the tests in appendix 1.
*test statistic significant at the 10% level; **test statistic significant at the 5% level.

equation. In the case of the United Kingdom the results are very similar, but stronger. There is evidence of non-invariance in both the full sample and the post-1983 sample, in both the output and inflation equations.

Of course one has to be cautious in concluding from invariance/superexogeneity tests alone that Lucas critique type effects are present. One could argue that the empirical size of these tests will depend critically on the correct specification of the marginal models.[11] It should also be noted that, even when the null hypothesis of invariance is rejected, we cannot date the shift in policy regime using these tests.

As stated in the introduction, we have strong evidence from previous studies (for example, Muscatelli, Tirelli and Trecroci 1999) that some policy regime shifts have occurred over the past two decades, even in the United States. We now turn to an alternative method of checking for shifts in the transmission mechanism that correspond to interest rate policy regime shifts; that is, ascertaining whether shifts in the transmission mechanism are attributable to the Lucas critique.

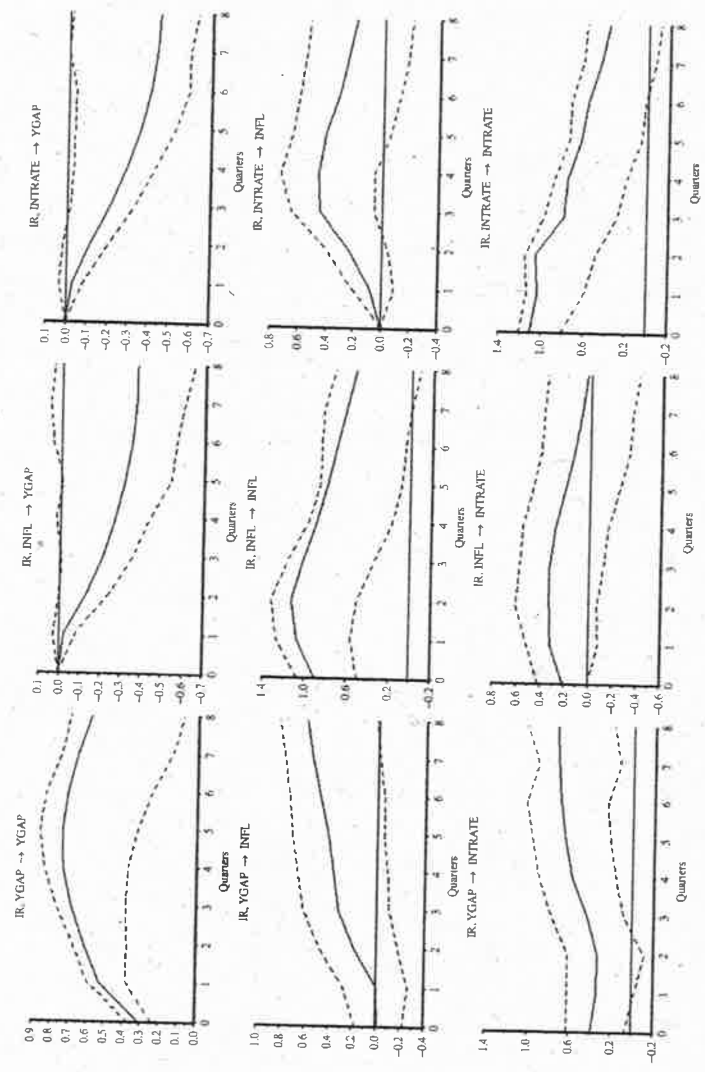

Figure 5.3 Standard trivariate VAR model: UK.

Notes: YGAP = output gap; INFL = quarter-on-quarter inflation rate; INTRATE = nominal policy interest rate. The upper and lower lines are 95% confidence interval bands.

5.3 The monetary transmission mechanism: Bayesian VAR estimates

5.3.1 *Standard VAR analysis*

Before turning to our Bayesian VAR estimates, we estimate a standard traditional VAR of the transmission mechanism as a benchmark. Figures 5.3 and 5.4 show the impulse responses and 95% confidence bands for the UK and USA (respectively) from a trivariate VAR that includes the nominal policy interest rate, the quarter-on-quarter inflation rate and the output gap (using the y^{oecd} and y^{bea} definitions).[12] Four lags of the variables are included. For reasons of space we show impulse responses only for up to eight quarters in figures 5.3 and 5.4. Although some of the impulse responses seem not to converge towards zero (especially for the United Kingdom), in fact the VAR is stationary, and convergence generally occurs for all the impulse responses after 8 to 16 quarters. The impulse responses have been derived using a Cholesky factorisation with a causal order as follows: the output gap, inflation and the interest rate. The VAR estimates[13] have been conducted for nearly the same sample as our invariance tests, namely 1961(Q3)–2000(Q1) for the United States and 1977(Q4)–1999(Q4) for the United Kingdom.

The United States results are, not surprisingly, very similar to those reported in Rudebusch and Svensson (1999) for an equivalent trivariate VAR for the United States during a similar sample period. The impulse responses in figure 5.4 show a positive response of the Federal Funds Rate to shocks to the output gap and the inflation rate. There is a slight price-puzzle effect following an interest rate shock, but generally the other impulse responses are as one would have expected: an output gap shock leading to a significant inflation response after three quarters and an inflation shock leading to a fall in output, significant after eight quarters; interest rate shocks lead to output falls after two to three quarters.

In the case of the United Kingdom the price puzzle is also present, but the impulse response is insignificant (except in quarters 3 and 4). The only other feature of note is that the response of the interest rate to an inflation rate shock is insignificant, reflecting the inclusion of the 1970s data in our full sample.

These impulse responses reveal the typical weaknesses of VAR analysis. One problem with unrestricted VARs is that they are overfitted, which explains why some of the responses are insignificant. Another problem is the interpretation of the implicit policy rule of the monetary authorities. As noted by Rudebusch (1998), the implicit policy rule that emerges from a just-identified VAR model is implausible, suggesting a response of the funds rate to a unit shock in the inflation rate of less than one. But the problem with the policy rule is also that it covers periods during which the Federal Reserve's behaviour changed dramatically. The same applies to the United Kingdom where, as we have already noted, the response of the interest rate to an inflation shock has the right

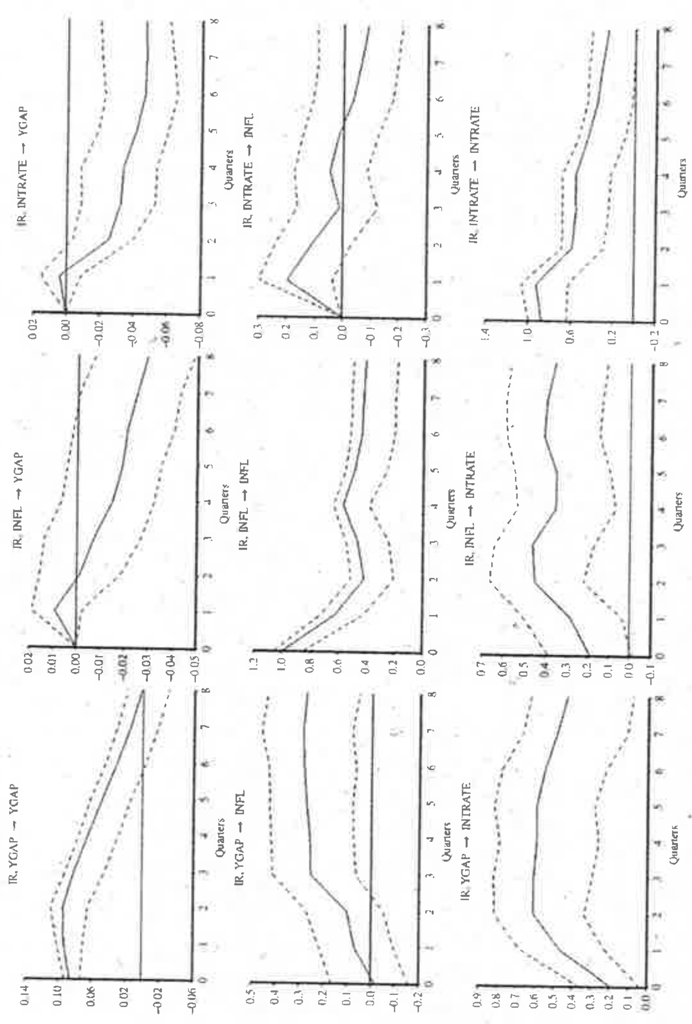

Figure 5.4 Standard trivariate VAR model: USA.

Notes: YGAP = output gap; INFL = quarter-on-quarter inflation rate; INTRATE = nominal policy interest rate. The upper and lower lines are 95% confidence interval bands.

sign but is insignificant, reflecting the very different nature of monetary policy in the early part of the sample.

Despite their limitations, VARs can be useful as descriptors of the dynamic correlations of jointly endogenous variables. In our case, we are particularly interested not only in examining policy responses, but also in examining whether any changes in policy regime elicited a change in transmission mechanism responses.

The alternative to VARs has been to estimate single-equation interest rate reaction functions (as in Clarida, Galí and Gertner, 1998; and Muscatelli, Tirelli and Trecroci, 1999). These single-equation interest rate reaction functions can be estimated using rolling or recursive estimation. However, single-equation approaches typically ignore the fact that we are dealing with jointly endogenous variables. It is therefore important to check that the results obtained from single-equation models are not due to the arbitrary imposition of a theoretical structure on the reaction function. VARs remain useful in this context because they impose the minimum amount of theoretical structure. However, we have to allow explicitly for changes or evolution in monetary policy regimes. In order to do this, we have to allow for some time variation in the parameters of the VAR.

Before turning to our models we should mention that there are some caveats to interpreting shifts in the response of interest rates to the output gap and inflation as shifts in policy preferences. These caveats apply to both single-equation reaction functions and VARs. Setting interest rates usually involves a complex decision-making process, in which a policy committee or council bases its decisions on a host of different macroeconomic indicators and models (see Bank of England, 1999b; Vickers, 1999). Where interest rates are set by committee, each member of the decision-making body might rely on a different subset of indicators and a different preferred 'model' of the transmission mechanism. Estimated interest rate reaction functions clearly cannot, and do not set out to, model every aspect of this complex decision-making process. Instead they focus on simple policy rules governing interest rate responses to expected inflation and the output gap (and to short-term inertia). If the estimated reaction functions are found to be stable, this stability can be seen as indicating that the actual information set used by the policy authorities displays a stable relationship with the final policy objectives (output and inflation) of the estimated rule *and* that the policy authorities' preferences did not change over the estimation sample. On the other hand, instability in a reaction function indicates either a shift in policy preferences *or* a shift in the relationship between the final objectives of policy and the wider information set (including the many indicators) used by the policy committee.

Some observers see VAR reaction functions as particularly prone to problems of interpretation (see Rudebusch, 1998). Part of the reason for this drawback is that most monetary policy VARs include many indicator and intermediate objective variables. Because our approach focuses only on a trivariate VAR, it has more in common with those contributions that estimate single-equation reaction functions (Clarida, Galí and Gertner, 1998; Muscatelli, Tirelli and Trecroci, 1999; Nelson, 2000a). However, we are able to consider not only potential shifts in interest rate policy, but also the consequences of these policy shifts for the transmission mechanism.

5.3.2 *Bayesian VAR analysis*

The estimation of VAR models with time-varying coefficients was pioneered by Doan, Litterman, and Sims (1984). We will now employ this Bayesian approach to VAR estimation, which allows the parameters of the VAR to evolve as more observations are added.[14] This technique has intuitive appeal for modelling a situation in which monetary policy changes have occurred, because policy regime changes are likely to be followed by a gradual evolution of responses. This is even the case where shifts in policy regime seem sudden. In the United Kingdom, after exit from the ERM and the implementation of inflation targets, it arguably took some time before the new system of inflation targeting became fully functional. In 1997, after the Bank of England acquired independence, it took time for the new Monetary Policy Committee to learn how to react optimally to new information. Even in the case of the United States, where the underlying institutional structure has not changed, the 1990s was a period in which the authorities had to learn gradually about the changing nature of the underlying macroeconomic relationships. For instance, in the mid-1990s the Federal Reserve (the Fed) tightened monetary policy at a time when it was uncertain about the 'productivity miracle' and the inflationary consequences of a tightening labour market. Similarly, if the monetary authorities' policy stance changes over time, we would expect private sector expectations to evolve gradually as they learn about the effects of changes in policy response. The fact that the Bayesian VAR takes this into account makes it a particularly attractive approach.

The estimation procedure can be outlined as follows. We begin with the definition of a standard VAR(p):

$$X_t = c + \sum_{j=1}^{p} A_j X_{t-j} + \varepsilon_t, \tag{5.5}$$

where X_t is an $n \times 1$ vector of endogenous variables, A_js are the $n \times n$ matrices of parameter coefficients and ε_t is an $n \times 1$ vector of disturbances for which

$$\begin{aligned} &E\{\varepsilon_t\} = 0 \\ &E\{\varepsilon_t \varepsilon_t'\} = \Sigma \\ &E\{\varepsilon_t \varepsilon_s'\} = 0, \quad \forall t \neq s. \end{aligned} \tag{5.6}$$

In our case, $X_t' = (y, \pi, i)_t$ and $n = 3$. Following Lütkepohl (1991) and Hamilton (1994), we can write the model in the following way:[15]

$$\begin{aligned} X &= AZ + U, \\ X &= (X_{p+1} \quad X_{p+2} \quad \dots \quad X_T); \\ A &= (c \quad A_1 \quad \dots \quad A_p); Z = (Z_p \quad Z_{p+1} \quad \dots \quad Z_{T-1}) \end{aligned} \tag{5.7}$$

$$Z_t = \begin{pmatrix} 1 \\ X_{t-1} \\ X_{t-2} \\ \vdots \\ X_{t-p} \end{pmatrix}; U = (\varepsilon_{p+1} \quad \varepsilon_{p+2} \quad \dots \quad \varepsilon_T),$$

where now we have only $T^* = T - p$ observations available in each equation. Assuming that the model is stationary, the VAR has the following finite MA representation:

$$X_t = \sum_{j=0}^{\infty} B_j \varepsilon_{t-j}, \tag{5.8}$$

where the B_js are the $n \times n$ MA parameter matrices. Given the information set Ω, if the residual variance–covariance matrix Σ is diagonal, the impulse response function will be defined as

$$IR_X(h, \delta, \Omega_{t-1}) = E\{X_{t+h} | \varepsilon_t = \delta, \Omega_{t-1}\} - E\{X_{t+h} | \Omega_{t-1}\}, \tag{5.9}$$

that is, the difference between the expected value of X_t at horizon h, given that a shock δ hits the system in time t, and the expected value of X_t in the absence of shocks. The MA parameters φ can then be intepreted as responses of X_{t+h} to a shock in t on variable j:

$$\varphi_{j,h} = B_h e_j, \tag{5.10}$$

where e_j is a vector of zeros with one as the jth element. If, on the other hand, Σ is not diagonal, contemporaneous interactions amongst the variables prevent any interpretation of the VAR residuals as fundamental disturbances, and the system is not identified. The method of identification employed here orthogonalises the shocks according to a Choleski decomposition of the residual variance–covariance matrix: $PP' = \Sigma$, where P is a lower triangular matrix. This way, the orthogonalised responses are recovered as

$$\varphi_{j,h}^{O} = B_h P e_j. \tag{5.11}$$

This defines the standard VAR model.

Next we consider the possibility of time-varying parameters. If we assume the VAR coefficients are time dependent, j from equation (5.5) can be written as

$$x_{t,j} = \boldsymbol{Z}' \begin{pmatrix} c_j \\ \beta_{j1}^{1} \\ \vdots \\ \beta_{jn}^{1} \\ \vdots \\ \beta_{j1}^{p} \\ \vdots \\ \beta_{jn}^{p} \end{pmatrix} + \varepsilon_{t,j} = Z'\beta_t + \varepsilon_{t,j}, \tag{5.12}$$

where the βs represent the elements of the VAR parameter matrices. A state–space representation of such a model would have equation (5.12) as the measurement equation. The Doan et al. (1984) procedure assumes that the VAR coefficients follow an AR(1) process, and the transition equation of the system is therefore

$$\boldsymbol{\beta}_t = c + \boldsymbol{T}\boldsymbol{\beta}_{t-1} + \upsilon_t. \tag{5.13}$$

Doan et al. suggest using a Bayesian prior distribution for the initial value of the coefficient vector, $\boldsymbol{\beta} \sim N(\overline{\boldsymbol{\beta}}, \boldsymbol{P}_{1|0})$, and then allowing the parameters to be updated according to some law of motion. In fact, we assume that the VAR parameters behave as follows:[16]

$$\boldsymbol{\beta}_t = (1-\pi_1)\overline{\boldsymbol{\beta}} + \pi_1 I_{np+1}\boldsymbol{\beta}_{t-1} + \upsilon_t. \tag{5.14}$$

In equation (5.14), the parameter vector follows an autoregressive process, in which the weighting parameter π_1 determines the importance of the steady-state value for the coefficient vector $\overline{\boldsymbol{\beta}}$. The disturbance term is uncorrelated with the disturbances in the original VAR: $Cov[\varepsilon_t, \upsilon_t] = 0$. The expected value $\overline{\boldsymbol{\beta}}$ consists of a vector with a one as the element corresponding to the own variable $x_{t-1,j}$ at lag 1 for each equation, and zeros elsewhere. This prior distribution holds that changes in the endogenous variable modelled are so difficult to forecast that the coefficient on its lagged value is likely to be near unity, while all other coefficients are assumed to be near zero. This prior distribution is independent across coefficients, so that the mean square error of the state vector is a diagonal matrix.

The matrix $\boldsymbol{P}_{1|0}$ is given by

$$\boldsymbol{P}_{1|0} = \begin{pmatrix} g\hat{\tau}_1^2 & 0' \\ 0 & (\boldsymbol{B}\otimes\boldsymbol{C}) \end{pmatrix}, \tag{5.15}$$

where

$$\boldsymbol{B} = \begin{pmatrix} \gamma^2 & 0 & 0 & \dots & 0 \\ 0 & \gamma^2/2 & 0 & \dots & 0 \\ 0 & 0 & \gamma^2/3 & \dots & 0 \\ \vdots & \vdots & \vdots & & \vdots \\ 0 & 0 & 0 & \dots & \gamma^2/p \end{pmatrix},$$

$$\boldsymbol{C} = \begin{pmatrix} 1 & 0 & 0 & \dots & 0 \\ 0 & \omega^2\hat{\tau}_1^2/\hat{\tau}_2^2 & 0 & \dots & 0 \\ 0 & 0 & \omega^2\hat{\tau}_1^2/\hat{\tau}_3^2 & \dots & 0 \\ \vdots & \vdots & \vdots & & \vdots \\ 0 & 0 & 0 & \dots & \omega^2\hat{\tau}_1^2/\hat{\tau}_n^2 \end{pmatrix}$$

Also, Q, the variance of υ_t, is equal to

$$Q = \pi_2 \boldsymbol{P}_{1|0}. \tag{5.16}$$

Doan, Litterman and Sims (1984) suggest the use of predefined values for the parameters in equations (5.15) and (5.16). The following assumptions are made:

$$\omega^2 = \frac{1}{74}, \qquad g = 360, \qquad \pi_1 = 0.999 \quad \text{and} \quad \pi_2 = 10^{-7}.$$

Note that this assumes that the coefficient vector $\boldsymbol{\beta}$ converges only very slowly towards the mean. Finally, if γ defines the analyst's confidence that the first-order autoregressive coefficient $\phi_{ii,1}^{(1)}$ relating y_{it} to $y_{i,t-1}$ attached to each series, $i = 1, \dots, n$ is near unity, Doan et al. recommend $\gamma^2 = 0.07$.

This general time-varying formulation turns the estimation problem into one of forecasting in each period the optimal state vector based on information available up to the previous period. Using the normality and independence assumptions about the disturbances, the computation of the state vector is obtained simply by applying the Kalman filter (Harvey, 1989; Hamilton, 1994). Doing so allows us to obtain filtered estimates of the VAR parameters and the residual variance–covariance matrix for each observation in the sample. Orthogonalised impulse responses are finally computed according to the standard Choleski decomposition, producing a set of different impulse responses (5.10) for each period of our sample.

Let us first examine the impulse responses for the United Kingdom. Figures 5.5 and 5.6 show the impulse responses[17] of the interest rate at the first-, second-, third-, fifth- and eighth-quarter horizon following unit shocks to the inflation rate and the output gap, respectively, over the sample period 1988–99.

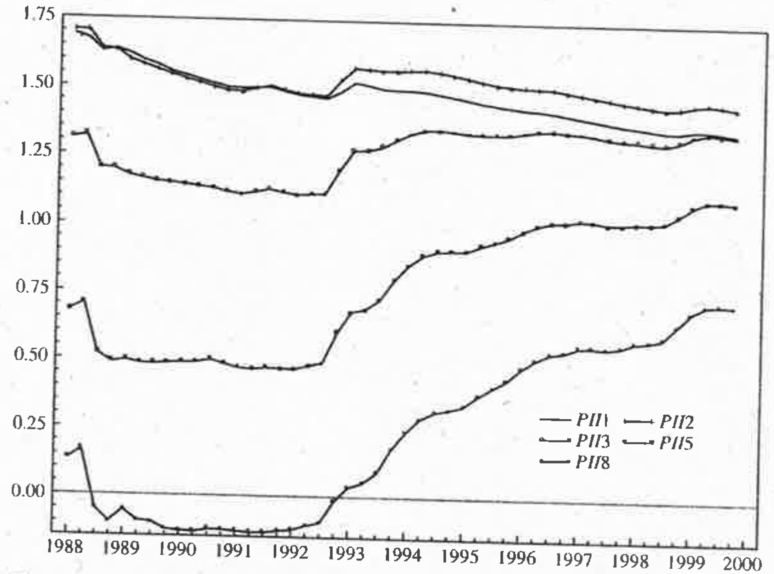

Figure 5.5 UK Bayesian VAR: effect on interest rate of an inflation shock.
Note: *PIIx* denotes the impulse response of the interest rate (*I*) following an inflation shock (*PI*) after *x* periods.

Note that, in figure 5.5, *PIIx* denotes the impulse response function for the interest rate (*I*) following an inflation shock (*PI*) after *x* periods. In figure 5.6, *YIx* denotes the impulse response of the interest rate (*I*) following a shock to the output gap (*Y*) after *x* periods. Four points can be noted.

First, from figure 5.5 we can clearly see that the interest rate response to the inflation rate is much greater than unity, at least for the first three quarters following an inflation shock. This confirms that the some of the usual critique of VAR policy rules (Rudebusch, 1998; Rudebusch and Svensson, 1999) is directly attributable to the use of constant coefficient VARs.

The second point to note, from both figures 5.5 and 5.6, is that the implicit policy rule seemed to change in the early 1990s. In this our analysis confirms evidence from some earlier contributions (Muscatelli, Tirelli and Trecroci, 1999; Nelson, 2000a). There is a sharp shift in 1992. From figure 5.5 it is apparent that the interest rate response to an inflation shock becomes more sustained, even after the third quarter, suggesting a more concerted and persistent response of monetary policy to the inflation shock. Figure 5.6 shows that the response to the output gap has also changed: prior to 1992 there was a perverse response to the output gap, whereas a more consistent response over time is apparent later in the 1990s.

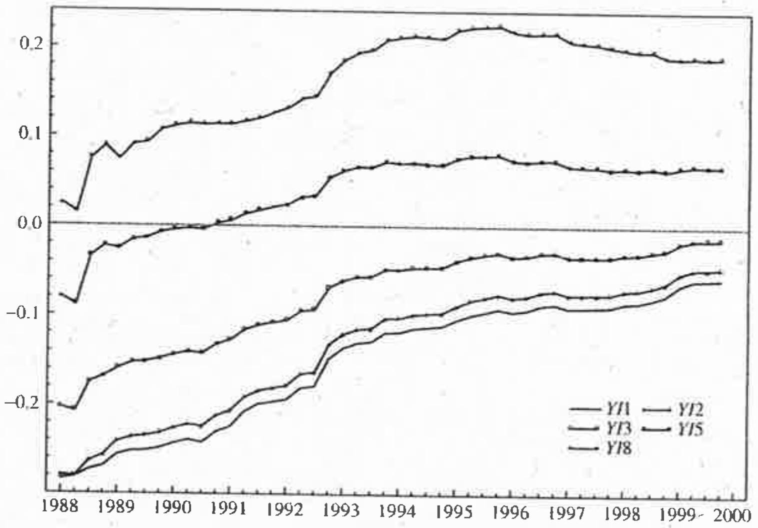

Figure 5.6 UK Bayesian VAR: effect on interest rate of an output gap shock.
Note: *YIx* denotes the impulse response of the interest rate (*I*) following a shock to the output gap (*Y*) after *x* periods.

Thirdly, Bank of England independence does seem to have had some effect on policy responses, with the interest rate responding more decisively to inflation shock from late 1998. However, it is probably too early to say whether this change represented a shift in the policy rule, or whether it was due to the need to tighten policy more dramatically after the 1997 election, following the failure of the Conservative administration to tighten policy sufficiently early in 1997.

Fourthly, the evolution of the policy rule after 1992 explains the fact that the invariance tests pick up shifts in the output and inflation equations. It appears that the shift in policy stance has affected the transmission mechanism (the response of output and inflation to interest rates). This is shown in figures 5.7 and 5.8.[18] The response of inflation to shocks in the output gap (*YPIx* in figure 5.7) seems to have changed only little post-1992. Most of the evolution (essentially a more inflationary response to an output shock after eight quarters) seemed to take place during the ERM membership phase (1990–92). The UK result seems to contradict the Hutchinson–Walsh (1998) result for New Zealand, namely that the acquisition of credibility should increase the output–inflation trade-off.[19] Figure 5.8 shows that there has also been a marked decrease in the responsiveness of output to the interest rate in the United Kingdom (*IYx*) after 1992. This result has implications for the degree of policy activism to be used. This result can be interpreted in three ways. The first is that in the early part of the

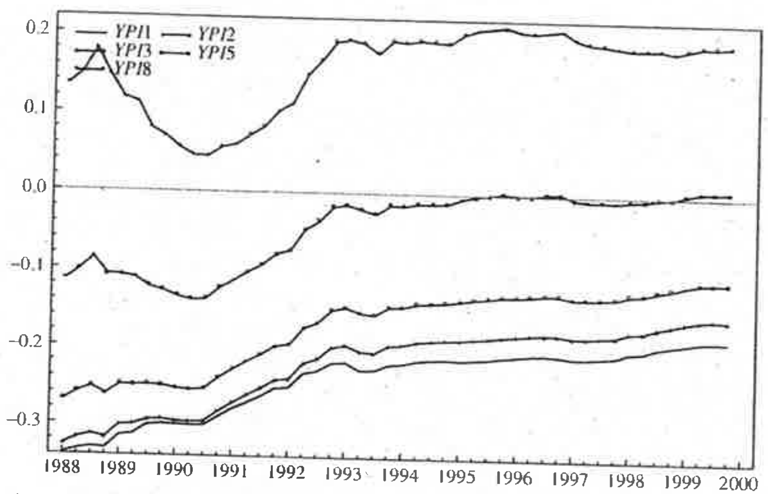

Figure 5.7 UK Bayesian VAR: effect on inflation rate of an output gap shock.
Note: *YPIx* denotes the impulse response of the inflation rate (*PI*) following a shock to the output gap (*Y*) after *x* periods.

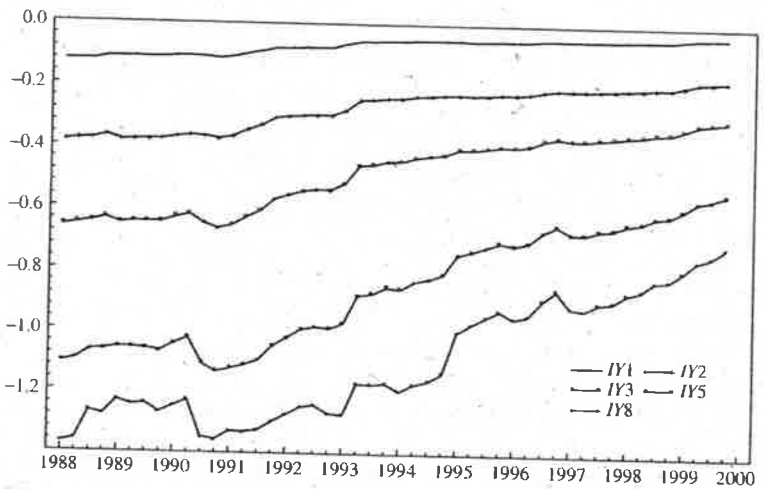

Figure 5.8 UK Bayesian VAR: effect on output gap of an interest rate shock.
Note: *IYx* denotes the impulse response of the output gap (*Y*) following a shock to the interest rate (*I*) after *x* periods.

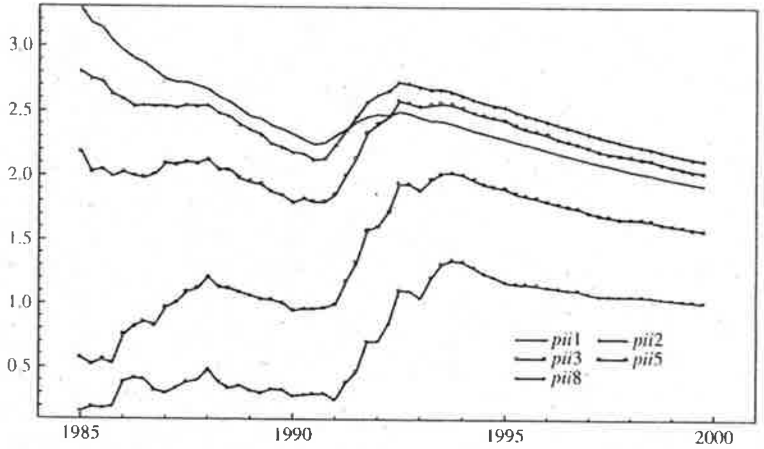

Figure 5.9 US Bayesian VAR: effect on interest rate of an inflation shock.
Note: *piix* denotes the impulse response of the interest rate (*i*) following an inflation shock (*pi*) after *x* periods.

1988–98 period nominal interest rate shifts had a more dramatic impact on real interest rates as inflationary expectations adjusted downward. Once inflation expectations adjusted downward in the mid-1990s, nominal interest rate shocks had a smaller impact on real interest rates, and hence on aggregate demand. The second possibility is that a non-linear relationship exists between interest rates and aggregate demand, so that the responsiveness of the output gap is different at different interest rate levels. The third interpretation is that the UK economy became less sensitive to interest rate increases in the late 1990s as the problems with mortgage indebtedness of the early 1990s were gradually reduced.

Turning next to the United States, we see from figures 5.9 and 5.10 that, despite the absence of formal institutional change, the Fed's reaction to inflation and output gap shocks in the 1980s and 1990s evolved markedly. The change, however, has not been monotonic, as was the case in the United Kingdom. In figure 5.9, we see that the interest rate response to inflation (*piix*) increased sharply between 1991 and 1994, but fell back sharply afterward. This pattern fits with general commentaries on Fed policy, which suggest that in 1994–95 the Fed reconsidered its policy stance in the light of evidence on productivity growth. Again, in contrast to the points made in Rudebusch (1998) and Rudebusch and Svensson (1999), the impulse responses show a greater-than-unit response of the Federal Funds Rate to a unit shock in the inflation rate, so that real interest rates are raised by the Fed following a shock. The pattern of impulse responses is similar to that for the United Kingdom in terms of timing (the peak is reached

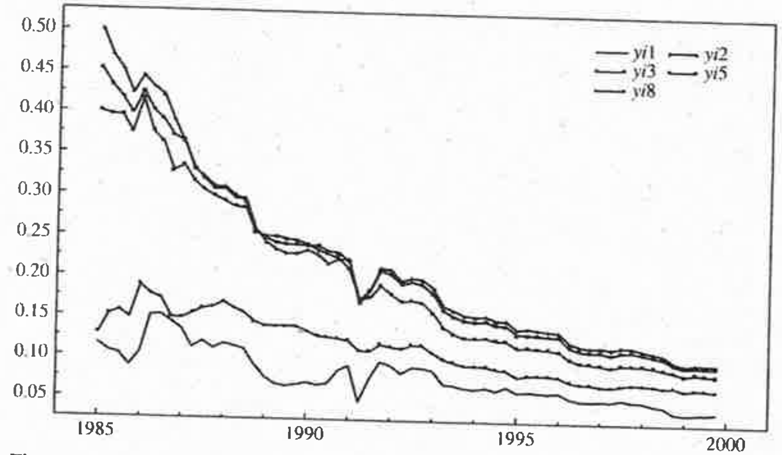

Figure 5.10 US Bayesian VAR: effect on interest rate of an output gap shock.
Note: *yix* denotes the impulse response of the interest rate (*i*) following a shock to the output gap (*y*) after *x* periods.

two quarters after the shock), but the US responses are larger. In figure 5.10 we see that, in contrast, the response of the Fed to the output gap (*yix*) tended to decrease from 1992 onwards, especially over the medium horizon ($x = 3,5$).[20] The final observations from our sample are of particular interest because, using the BEA potential output data, the output gap shows a major surge in 1999–2000. This is because the implicit assumption is that the rate of growth in potential output has not increased, that is there has been no 'productivity miracle'. The interpretation that emerges from the VAR is then that monetary policy became less countercyclical in the late 1990s. Using our alternative series for the output gap, based on a Kalman filter procedure (see figure 5.2), would yield slightly different results for the late 1990s, because this technique interprets the recent increase in output growth as partly due to an increase in potential output.

Have recent changes in policy responses affected the nature of the transmission mechanism in the United States? Figures 5.11 and 5.12 show the response of output and inflation to an interest rate shock and an output gap shock, respectively.[21] Figure 5.11 shows that, as opposed to the case in the United Kingdom, there is very little evidence of the United States experiencing a shift in the response of output to interest rate shocks, despite the presence of apparent policy shifts. This finding is consistent with the evidence from the application of invariance tests to our estimated aggregate demand equation (see table 5.3). The absence of expectations effects in the output equation might be attributable to the absence of formal institutional change in the United States, which means that the transmission mechanism has remained more stable. Figure 5.12 shows

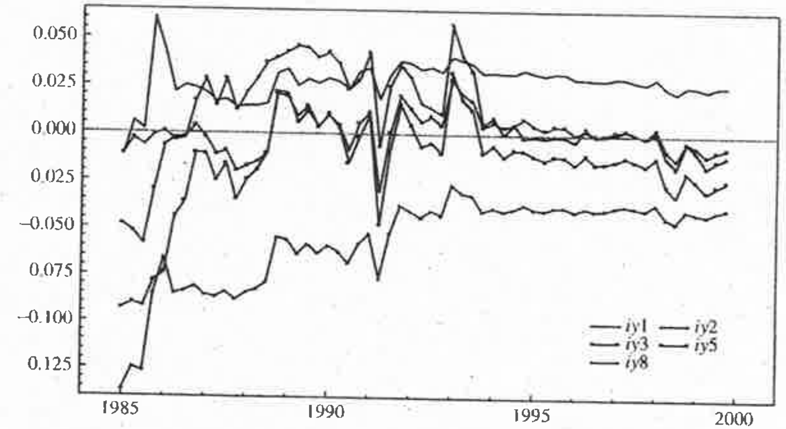

Figure 5.11 US Bayesian VAR: effect on output gap of an interest rate shock.
Note: *iyx* denotes the impulse response of the output gap (*y*) following a shock to the interest rate (*i*) after *x* periods.

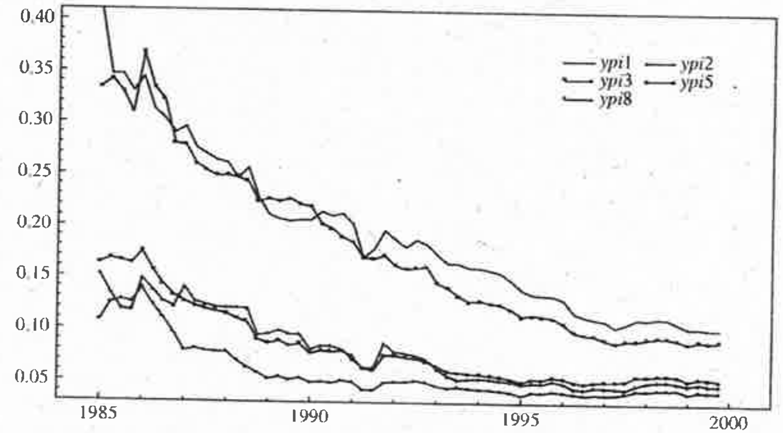

Figure 5.12 US Bayesian VAR: effect on inflation rate of an output gap shock.
Note: *ypix* denotes the impulse response of the inflation rate (*pi*) following a shock to the output gap (*y*) after *x* periods.

again that, in contrast to the United Kingdom situation, the short-run output–inflation correlation has shifted. An output gap shock resulted less readily in an inflationary surge in the 1990s. This result is in line with the results of Hutchison and Walsh (1998) for New Zealand, but whether it reflects an expectations effect or the impact of more rapid productivity remains an open issue. It might also explain why the Fed's interest rate response to the output gap declined over the same period. Although the impulse responses in figure 5.12 seem to change, it is hard to ascribe the shifts to an expectations effect, because they do not match variations in the implicit VAR interest rate policy response to inflation.

5.4 Conclusions

In this paper we have examined how interest rate policy responses to both output and inflation shocks in the United Kingdom and the United States have evolved during the past two decades, and how the transmission mechanism has been affected by apparent shifts in the policy rules followed by the Fed and the UK monetary authorities. Tests for invariance/superexogeneity seem to show some evidence in support of the Lucas critique, confirming the importance of forward-looking expectations in aggregate demand and supply. The evidence relates both to full sample estimates, which cover known policy shifts (the switch to much tighter inflation control beginning in the 1980s), and post-1983 findings reflecting formal institutional change limited to the United Kingdom. In the UK analysis, there is evidence that the Lucas critique has some weight in both the output and inflation equations, and its role might be linked to the presence of formal institutional change (the adoption of inflation targeting and the granting of instrument independence to the Bank of England). However, even for the United States we find evidence of expectations effects in the post-1983 sample estimates.

Our Bayesian VAR analysis shows a much richer picture. It confirms the finding that interest rate policy rules shifted in both the United States and the United Kingdom in the 1990s. In the United Kingdom this was linked to institutional change, and in particular to the adoption of inflation targets, but also, in a minor way, to the granting of independence to the central bank in 1997. US policy changes have been less systematic, with a more activist policy pursued until the mid-1990s, and a less responsive interest rate policy since 1995. This trajectory is probably attributable to the Fed's changing attitude toward the US 'productivity miracle'. The shift in policy rule seems to coincide, at least in the United Kingdom, with shifts in the transmission mechanism, lending some weight to the notion that the Lucas critique may need to be taken into account in setting interest rates if policy-makers significantly change their policy stance. Although independent central banks may be more immune to problems of credibility, it would appear that forward-looking expectations are still an important dimension of monetary policy design.[22] This fact has important implications for the vast theoretical literature that has emerged on the optimal design of feedback rules for monetary policy (see Haldane and Batini, 1998; Rudebusch and Svensson, 1999; Rudebusch, 2000; and Onatski and Stock, 2000). This literature has far too often ignored forward-looking expectations in the specification of underlying structural models.

Appendix 1: invariance/superexogeneity tests

The following tests are developed fully in Engle and Hendry (1993). This appendix briefly summarises the approach taken to the statistical tests of the Lucas critique in section 5.2.

Consider the following joint distribution of the variables x_t and y_t, which are assumed to be conditional normal, with conditional means

$$\begin{aligned} E(x_t|\Omega_t) &= \mu_t^x \\ E(y_t|\Omega_t) &= \mu_t^y \end{aligned} \qquad (5A.1)$$

and covariance matrix

$$\Sigma_t = \begin{bmatrix} \sigma^{yy} & \sigma^{yx} \\ \sigma^{xy} & \sigma^{xx} \end{bmatrix}, \qquad (5A.2)$$

where the means and covariances need not be constant but may depend on the information set, Ω. This information set contains past values of the two variables and also current and past values of other valid conditioning variables, z_t.

Consider a model that relates the two conditional means of x_t and y_t:

$$\mu_t^y = \beta\mu_t^x + z_t'\gamma. \qquad (5A.3)$$

Given the expectation of y_t conditional on x_t, and its conditional variance, we can write the following conditional model for y_t:

$$y_t = \beta x_t + z_t'\gamma + (\delta_t - \beta)(x_t - \mu_t^x) + \omega_t, \qquad (5A.4)$$

where $\delta_t = (\sigma_t^{yx}/\sigma_t^{xx})$, $\omega_t = \sigma_t^{yy} - ((\sigma_t^{yx})^2/\sigma_t^{xx})$.

The efficient estimation of (5A.4) requires: (i) the weak exogeneity of x_t, which is satisfied if $\delta_t = \beta$; (ii) constancy of the regression coefficients, which requires the constancy of δ_t; and (iii) invariance, which requires the parameter β to be invariant to changes in the process generating x_t. Weak exogeneity and invariance imply that the variable x_t is superexogenous for the parameters of interest in our model.

To test for superexogeneity, Engle and Hendry examine the impact of changes in the moments of x_t on β. Allowing β to be a function of the moments of x_t and using a linear expansion for $\beta(.)$ yields:

$$\beta = \beta_0 + \beta_1\mu_t^x + \beta_2\sigma_t^{xx} + \beta_3(\sigma_t^{xx}/\mu_t^x) \qquad (5A.5)$$

(where higher-order terms of the expansion can be considered). The test regression then becomes

$$\begin{aligned} y_t = {} & \beta_0 x_t + z_t'\gamma + (\delta_t - \beta_0)(x_t - \mu_t^x) + \beta_1(\mu_t^x)^2 + \beta_2\mu_t^x\sigma_t^{xx} \\ & + \beta_3\sigma_t^{xx} + \varepsilon_t. \end{aligned} \qquad (5A.6)$$

Table 5A.1 *Data sources*

Time series	United Kingdom	United States
Price index	CPIX up to 1987(Q1) (OECD), RPIX thereafter (ONS)	GDP chain-weighted price index (BEA)
Output	GDP	GDP
Output gaps	Interpolated OECD estimates	Bureau of Economic Analysis estimates
Short-term interest rate	London Clearing Banks Overnight Rate (IFS)	Federal Funds Rate (IFS)
Real exchange rate	Real effective exchange rate, index 1995 = 100 (OECD)	
Unit labour cost	Unit Labour Cost, in Manufacturing Industries, Sterling Pounds, index, 1995 = 100 (OECD)	
Import prices	Import Price Index, 1995 = 100 (IFS)	

Table 5A.2 *Definition of variables*

Symbols adopted	Variable definition
y^{bea}	Output gap, BEA series. Gap between actual and potential, as a percentage of potential output
y^{kal}	Output gap, Kalman-filtered series. Gap between actual and potential, as a percentage of potential output
y^{oecd}	Output gap, OECD series. Gap between actual and potential, as a percentage of potential output
π	Quarterly inflation in percentage points at an annual rate
i	Short-term interest rate, in percentage points at an annual rate
$\bar{r}$	Real-interest rate, difference between four-quarter average i and π
q	Real effective exchange rate, in logs
pim	First difference in the (log of) Import Price Index
ulc	First difference in the (log of) Unit Labour Cost Index

To implement this test regression, one has to fit marginal models for the variable(s) x_t and use the residuals for this model to obtain a measure of $(x_t - \mu_t^x)$, and the fitted values to obtain a measure of μ_t^x; and one can model the variance σ_t^{xx} through, say, an ARCH model fitted to the residuals of the marginal model for x_t.

Appendix 2: data

Data sources and definitions are provided in tables 5A.1 and 5A.2. Data were obtained from IMF International Financial Statistics (IFS), OECD Main Economic Indicators, the United Kingdom Office of National Statistics (ONS) and the US Bureau of Economic Analysis (BEA). Where available, seasonally adjusted series were employed.

Acknowledgements

We are extremely grateful to our colleague Ulrich Woitek for providing us with the basic GAUSS code used to perform the Bayesian VAR estimates in section 5.3. We are grateful for helpful comments from participants at the Bank of England Centre of Central Banking Studies Conference on the transmission mechanism held at the CCBS in June 2000, and at the Royal Economic Society Conference 2000, held in St. Andrews. In particular, we should mention (without implicating them) Larry Ball, Jon Faust, Lavan Mahadeva and especially Ed Nelson, who acted as discussant on the paper.

Notes

1 In addition to making contributions that estimate policy reaction functions, a number of authors have used a variety of methods to test whether the introduction of inflation targeting had an impact on inflation expectations (see Freeman and Willis, 1995; King, 1995) or on a range of monetary policy indicators (see Almeida and Goodhart, 1996).

2 See, *inter alia*, Goodfriend and King (1997), Walsh (1998), Clarida, Galí and Gertner (1999), McCallum and Nelson (1999a,b), Rotemberg and Woodford (1999).

3 They use Andrews (1993)-type stability tests.

4 A full account of the relevant concepts of weak exogeneity, constancy, invariance and superexogeneity is provided elsewhere in the literature and will not be repeated here (see Favero and Hendry, 1990; Engle and Hendry, 1993).

5 More details of the test procedure used here are provided in appendix 1.

6 In the case of the United Kingdom some additional regressors are added to take account of the greater openness compared with the United States economy.

7 As an alternative, one could fit marginal models for the interest rate that have a structural interpretation. We experimented with the use of estimated policy reaction functions such as those estimated in Muscatelli, Tirelli and Trecroci (1999). The results obtained, however, were not very different from those that used an autoregressive model for the interest rate.

8 This provided a better fit than the use of a distributed lag of individual interest rates.

9 However, we do not follow Hall et al. in adding terms for foreign demand, fiscal policy and real money balances in the aggregate demand equation. In any case, our models are not strictly comparable because our dependent variable is the output gap rather than growth in real output.

10 These have the same structure as the forward-looking policy reaction functions reported in Muscatelli, Tirelli and Trecroci (1999) and Clarida, Galí and Gertner (1998).

11 These tests can also suffer from the problem of low power.

12 For the United Kingdom we decided not to include additional variables such as the real exchange rate. Although such variables might be helpful in describing the transmission mechanism in an open economy, restricting our attention to the output gap and inflation allows us more readily to interpret the policy rules that emerge from the VAR. Including the real exchange rate would also raise issues regarding the appropriate ordering of the variables in the Cholesky factorisation, given that the United Kingdom has experienced different exchange rate regimes over the sample period.

13 All estimates in this paper are produced using GAUSS routines.

14 For a recent application of Bayesian VAR analysis to the US economy during the Great Depression, see Ritschl and Woitek (2000).

15 Henceforth we assume that p is known and fixed.

16 The description here closely follows the notation used in Hamilton (1994), ch. 13.

17 We do not report confidence bands, because for each period and each impulse response we would need to report a confidence interval, making the graph difficult to read. However, having checked the impulse responses and confidence intervals (constructed using bootstrap methods) for critical periods discussed in the text, when the policy regimes were subject to change, we can confirm that the impulse responses discussed are significantly different from zero at the 95% confidence interval. Detailed results are available on request from the authors.

18 The response of inflation to an interest rate shock still shows a minor price puzzle (one that tends to be statistically insignificant), and is not shown here. It appears from the VAR impulse responses that the main channel of transmission from interest rates to inflation is through the output gap.

19 See also Lucas (1973) on the slope of the short-run aggregate supply relationship.

20 This is partly in line with the findings of Favero and Rovelli (1999), although their single-equation policy reaction function estimates seem to suggest that the output gap did not matter to the Fed throughout their sample period (1983–98).

21 Once again the impulse responses of inflation to the interest rate show a slight price puzzle, which disappears after three quarters, and eventually a negative response of inflation to an interest rate increase. But there is little evidence of a change in the transmission lag or in the size of interest rate shocks to inflation.

22 It is perhaps also worthwhile reminding ourselves that time-inconsistency does not only arise in static models where there is a conflict between the objectives of policy-makers and the monetary authorities. It is also present in models where the authorities target the natural rate of unemployment (so no Barro–Gordon type time-inconsistency problem exists), but forward-looking variables enter into the structural model.

6 The transmission mechanism of monetary policy near zero interest rates: the Japanese experience, 1998–2000

Kazuo Ueda

The Bank of Japan (BOJ) has gone through a unique experience in the past few years. When I joined the Bank's newly formed policy board in April 1998, the overnight call market rate, the key policy instrument of the BOJ, was already below 0.5%. The economy was in the midst of the most serious recession in the postwar period, although it took us a little while to realise this. We guided the call rate down to virtually zero in the first quarter of 1999 and followed up by promising to keep it there until deflationary concerns had been dispelled. Finally, in August 2000, we brought the rate up to 25 basis points after having kept the zero rate for one and a half years.

In this short paper, I would like to discuss some of the key aspects of the evolution of our thinking on monetary policy over the period 1998–2000. In so doing, I would like to focus specifically on the characteristics of the 1997–98 Japanese recession, the transmission process of monetary policy in the neighbourhood of a zero rate and the background thinking behind the rate hike in August 2000.

6.1 The nature of the 1997–98 recession

It is appropriate to begin with a brief discussion of the nature of the recession that started in 1997(Q4), which is what the BOJ was trying to respond to in 1998–2000. Clearly, the most important characteristic of the recession was the credit crunch caused by the slow and inappropriate handling of the bad loan problem. In addition, the Asian economic crisis, a premature tightening of fiscal policy in 1997 and the Russian crisis in 1998 added to the severity of the recession.

More specifically, the onset of the Asian economic crisis and the absence of fundamental measures to address the bad loan problem finally resulted in the failure of three medium- to large-sized financial institutions in the fall of 1997.