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Monetary Frameworks and Institutional Constraints: UK Monetary Policy Reaction Functions, 1985–2003

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Abstract

Monetary policy reaction functions are estimated for the UK over three periods – 1985–90, 1992–97 and 1997–2003 – in order to disentangle two effects: the switch from an emphasis on exchange rate stabilization to inflation targeting, and the introduction of instrument-independence in 1997. The external factors considered include US as well as German interest rates, and this leads to the identification of 'domestic' and 'international' models of the reaction function. The results suggest that it is the changes in the institutional arrangements rather than those in the targeting regime which have been decisive in the development of policy in this period.

I. Introduction

Casual observation of macroeconomic outcomes suggests that UK monetary policymaking has greatly improved over the last two decades. The two principal changes to which this improvement might be attributed are in the role of monetary policy targets and in the institutional arrangements. On the former, the monetary authorities moved decisively away from exchange rate stabilization after the exit from the Exchange Rate Mechanism (ERM) of the

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European Monetary System in 1992 and adopted inflation targeting. On the latter, the Monetary Policy Committee (MPC) of the Bank of England was given control of interest rates in 1997. In this paper, we estimate reaction functions for the UK interest rate over three periods, in order to disentangle these two influences empirically and to draw out the implications for the conduct of monetary policy.[1]

Modern work on reaction functions derives largely from Taylor's (1993) proposal of a simple policy rule for the Federal Reserve in which the interest rate is varied in response to inflation and the output gap (only), and his demonstration that US monetary policy from 1986 could be characterized as following such a rule. While the context of Taylor's and much other work is a closed economy such as the US, authors including Ball (1999b) and Svensson (2000) have investigated the operation of similar rules in open economy models. Empirical work by Clarida, Galí and Gertler (1998), Nelson (2000) and others has considered some open economy aspects in estimating reaction functions for countries such as the UK.

Without additional information, the policy preferences of the authorities and the structure of policy rules in circumstances where the rule is not explicit cannot, in general, be identified from the parameters of estimated Taylor rules. Dennis (2003) discusses the conditions under which the authorities' policy preferences (and the policy rule) can be identified for a class of linear policy rules. Muscatelli, Tirelli and Trecoci (2000) use recursive estimation as a way of pinning down the effect of changes in institutional constraints. In this paper, we approach the identification issue from a different direction. Drawing on a detailed narrative of key events in recent UK monetary history, we estimate separate (reduced-form) reaction functions over three specific periods. Using within- and out-of-sample tests of model stability we argue that changes in the reduced-form parameters of the reaction functions across these periods represent discrete changes in the role played by the external constraints and policy preferences in shaping monetary policy in the UK.

Our empirical investigation differs from those of other researchers in the field in two respects. First, we consider influences from US interest rates and the dollar/sterling exchange rate as well as those from German interest rates and the DM/pound exchange rate which have been considered elsewhere. This leads us to identify and contrast 'domestic' and 'international' models of the UK reaction function for different periods. Second, by estimating reaction functions for periods before and after the adoption of inflation targeting and then before and after the introduction of instrument independence, we are able to distinguish the effect of the change in the institutional constraints on policy

[1]The authors are grateful for comments to Larry Ball, participants in a seminar at the Bank of England and at the 2001 MMF annual conference, and three anonymous referees, but none of these carry any responsibility for the errors that remain.

from the effect of the adoption of inflation targeting. Our results indicate that, while the adoption of inflation targeting in 1992–93 led to some change in the reaction function as compared with the second half of the 1980s, that of instrument-independence in 1997 was associated with a much more profound change in monetary policy.

Section II identifies the periods used in the econometric investigation. Sections III and IV set out the theory and methodology underlying the econometric investigation presented in section V. Section VI concludes.

II. Monetary policy episodes[2]

In the *pre-ERM* period between April 1985 and September 1990 there were no firm targets for monetary growth or inflation, but external factors were important in UK interest rate decisions, notably but not only during the informal DM-shadowing phase from March 1987 to March 1988. In October 1990, the UK entered the ERM, but it left in September 1992; this period is too short for estimation. Between October 1992 and April 1997, the *post-ERM* period, there was a clear monetary framework in the form of inflation targets, but interest rate decisions continued to be under the control of the Chancellor of the Exchequer (Minister of Finance). From May 1997 in the *MPC* period the Bank of England became instrument-independent with a continuing inflation target set (but not varied) by the government.

Figure 1 shows the UK policy rate together with the 3-month interbank interest rate, which is the dependent variable in the regressions below,[3] and the corresponding German and US rates. The UK policy rate was initially on a downward trend but with large and frequent oscillations, and policy was then tightened rapidly and substantially from mid-1988, to decline only on entry to the ERM. The exit from the ERM was accompanied by further large interest rate cuts, followed later by a small upswing in 1994–95 and some cuts (plus one rise) in 1996. The final period covers two upswings and two downswings, with successively lower peaks and troughs. The UK 3-month interbank rate moves closely with the policy rate: the correlations between the series are 0.995 for the pre-ERM, 0.879 for the post-ERM and 0.989 for the MPC periods. The German and US rates follow only broadly similar trajectories, with the German rate the lowest except for the post-unification years of 1991–94 and the US rate

[2]See Cobham (2002a, b) for a detailed examination of the monetary frameworks and (2002b, Chapter 8) for an analysis of the concerns mentioned by the authorities themselves in their explanations of interest rate changes.

[3]We use a market rate partly because that has become conventional in this sort of exercise and partly because doing so avoids possible econometric problems associated with estimating models of the policy rate when the latter is constant over extended periods and is changed only in discrete multiples of (in this case) 0.25% (see Galí *et al*., 2004, pp. 60, 61).

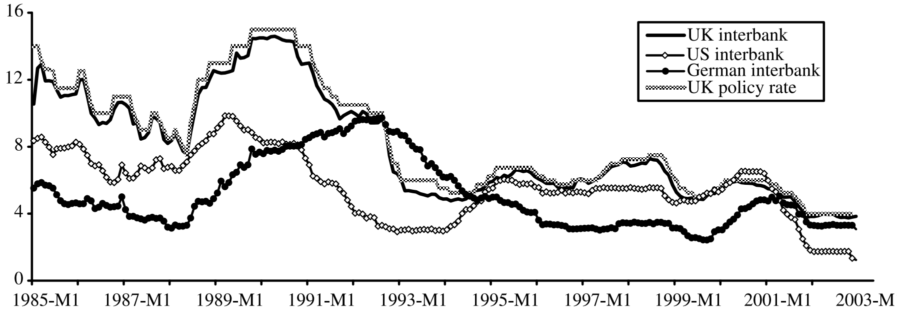

Figure 1. UK, US and German interbank rates and UK policy rate

mostly above the German but below the UK rate; the three rates are closer together in both levels and turning points in the final period.

III. The existing literature and methodological issues

Taylor's original (1993) rule, under which the policy interest rate is varied in response to inflation and the output gap [as in equation (1)], was both a stylization of the reaction function which the Federal Reserve had apparently been following for several years and a proposal as to the sort of reaction function which central banks should follow. Various authors have since demonstrated that the 'optimal policy rule' can be expressed in this form. Ball (1999a), for example, shows that in a backward-looking model of a closed economy the 'optimal policy rule' is of the form

$$r_t = \gamma_0 + \gamma_1 y_{t-1} + \gamma_2 \pi_{t-1}, \tag{1}$$

where y is the output gap, r is the difference between the real interest rate and its equilibrium level, and π is the difference between inflation and its average level. More recently, attention has focused on forward-looking models in which inflation depends on *expected* inflation and the output gap and the *ex ante* real interest rate is the main explanatory variable in the output gap equation.[4] These models lead to an 'optimal policy rule' such as

$$r_t = \delta_0 + \delta_1 y_t + \delta_2 E_t \pi_{t+k}, \tag{2}$$

where $E_t\pi_{t+k}$ is the inflation rate expected at some relevant period in the future. These models use *ad hoc* formulations of the loss function, but Woodford (2003) has shown that similar formulations can be derived from a full micro-founded model of consumer welfare.

[4]See McCallum and Nelson (1999) on the micro-foundations of the underlying IS-LM specification.

Svensson (1997) has argued that in this case, the inflation forecast performs the role of an intermediate target. Empirical forward-looking models were first estimated by Clarida *et al*. (1998), while forward-looking rules have been explored analytically by Batini and Haldane (1999), who conclude that inflation-forecast-based rules are superior in welfare terms to backward-looking specifications (which are nested within them).

Authors have also considered such rules in an open economy context. Ball (1999b) extends his (1999a) backward-looking model by introducing the real exchange rate and adding an equation for the exchange rate as a function of the interest rate and a white noise shock. The optimal rule then involves a monetary conditions index, i.e. a weighted average of the interest rate and the exchange rate, responding to the output gap, inflation and the lagged exchange rate. Svensson (2000) considers open economy issues in a forward-looking framework, which includes the foreign interest rate (modelled as determined by a Taylor-type rule) and the foreign exchange risk premium, as well as the real exchange rate, deriving optimal rules for the domestic interest rate which mostly include one or more of these external variables. Kirsanova, Leith and Wren-Lewis (2004) have argued that in a model with UIP shocks there is a case for including an 'exchange rate gap' term in the reaction function. On the other hand, Clarida *et al*. (2001) argue that optimal monetary policy should have the same form for an open as for a closed economy, and should not respond to the exchange rate or to foreign interest rates, while Taylor (2001) argues that the standard rule already includes an 'indirect' effect from the exchange rate, via its implications for inflation and output, and does not need to be supplemented by a direct effect.

The theoretical and empirical analysis of this class of open economy policy reaction functions has raised a number of issues. First, while Taylor's original rule specified a weight of 0.5 on the output gap, writers such as Ball (1999a) and Rudebusch and Svensson (1999) have argued for a much higher, and Rotemberg and Woodford (1999) for a much smaller, response to output. In addition, a number of empirical studies have found that the weight on the output gap for the US was significantly higher than 0.5 in the post-1979 period (e.g. Clarida *et al*., 2000, report a weight of 0.93). Second, most empirical researchers have found that they need to include a lagged dependent variable in Taylor-rule regressions to allow for interest rate persistence or smoothing (e.g. Clarida *et al*., 2000, report a coefficient of 0.79).[5]

A third issue concerns the expected coefficient on inflation in open economy Taylor rules and estimates of it. In the closed economy literature much importance has been attached to the issue of whether the coefficient on inflation is greater or less than 1 (e.g. Taylor, 1999; Clarida *et al*., 1999), and

[5]Sack and Wieland (1999). See also Rudebusch (2002) and English, Nelson and Sack (2003).

some empirical work has produced estimates well above unity (e.g. Clarida *et al.*, 2000 report an estimate of 2.15). While Ball (1999b) found that the coefficient on inflation in his open economy model should not differ much from that in his (1999a) closed economy model, Svensson (2000) found that in the flexible CPI-targeting case the weight on inflation should be much smaller than in the standard Taylor rule. In empirical terms Clarida *et al.* (1998) found that when they added the German interest rate to their baseline equations for France, Italy and the UK the coefficient on inflation fell from around 1 to around 0.5.

A fourth issue concerns the modelling of external relationships in open economy models. Different authors have characterized the external factors in different ways. Ball's (1999b) model includes the real exchange rate (only), while Svensson's (2000) model also includes the foreign interest rate and the foreign exchange risk premium. Most empirical work – including Clarida *et al.* (1998) – has found it necessary to include external factors for open economies, typically in the form of foreign interest rates rather than exchange rates. It seems obvious that for policymakers exchange rates must be proximately more important than foreign interest rates, but the appropriate response to exchange rate movements will depend on the nature and source of the change (temporary or permanent, domestic or external, portfolio vs. aggregate demand shocks, etc.),[6] and the noise to signal ratio in exchange rates is high. Moreover, in the words of the MPC, 'it would not be sensible for policy to react to high frequency movements in the exchange rate, as this could lead to a volatile path for interest rates from month to month, and might make it more difficult for others to understand the motives for interest rate changes'.[7] If responding continuously to exchange rates is uncongenial, policymakers might see pegging the domestic interest rate to some foreign rate as an effective way both of acquiring credibility and (given uncovered interest parity) of stabilizing the exchange rate over the medium term.[8]

A fifth issue, first raised by Orphanides (2003), is that output data are typically revised significantly at later dates, but policymakers have to base their decisions on the data available to them at the time. While most empirical work on reaction functions uses the (revised) data most recently published at the time of the study, authors such as Orphanides for the US and Nelson and Nikolov (2003) for the UK have estimated reaction functions on data available at the time. However, Orphanides and van Norden (2002) have shown for the US that the difference between the standard output gap series, calculated from the revised data for the full sample, and the 'real-time' output gap series,

[6]See Cecchetti *et al.* (2000, Chapter 2).

[7]MPC, minutes of its March 1999 meeting, paragraph 28. See also the minutes of June 1999 meeting, paragraph 7.

[8]We are indebted to Larry Ball for discussion on this issue.

computed under the assumption that at each point in time the authorities re-estimate trend output (and hence the output gap) on the basis of the most recently released data (only), is mainly the result of the rolling (real-time) estimation of trend output, rather than of later revisions to GDP data. This is due to what Orphanides and van Norden (2002, p. 582) refer to as 'the pervasive unreliability of end-of-sample estimates of the trend in output'. Thus, while a real-time measure of the output gap is clearly more appealing at a theoretical level, its mechanical application plays down the extent to which policymakers are able to recognize when the economy is booming and when it is in recession, and for that reason the full-sample output gap series may give a more accurate impression of what the policymakers believed at the time.

The empirical literature concerned with estimating monetary policy reaction functions has been strongly influenced by the work of Clarida *et al.* (1998), to which we have already referred. This provided a new framework for estimating forward-looking reaction functions (using an errors-in-variables approach and GMM estimation), which was then applied to the US, Germany, Japan, France, Italy and the UK. The most detailed study on the UK is that of Nelson (2000), which reports results for five separate periods. Table 1 provides a summary of the main results found in the existing literature on the UK for periods that correspond to or overlap with ours; the numbers in the inflation and output gap columns are (where relevant) the sum of the long run coefficients on different lags and leads.

All the studies report significant coefficients on the lagged dependent variable, mostly between 0.7 and 0.9, with Nelson's estimates rather lower and Dornbusch, Favero and Giavazzi (1998) finding a severe unit root problem (which prevented them from reporting long-run coefficients). Results for the inflation rate differ widely. Those studies that included the German interest rate mostly found that the coefficient on inflation was below unity. Nelson finds evidence of higher coefficients in later periods (though his inflation coefficient is restricted to zero for his middle period), and Muscatelli *et al.* (2000) report a relatively high value of 1.4. Kuttner and Posen (1999), however, obtained a coefficient of 1.64 for the 1980s but only 0.52 for the 1990s.[9] Results on the output gap differ less widely, though there is some tendency for coefficients to be higher in the later periods (Angeloni and Dedola, 1999; Nelson, and – for their unemployment variable – Kuttner and Posen).

Four of the studies included the German interest rate.[10] Clarida *et al.* report a value of 0.60 for their period (1979–90), Nelson and Angeloni and Dedola report coefficients greater than unity for parts of the 1980s, but the latter also

[9]Goodhart (1999) reports that Stephen Wright (2002), who had found in his paper an inflation coefficient of 0.8 for 1961–94, obtained a value of 1.6 when he reran his test for the 1990s.

[10]Muscatelli *et al.* (2000) also report additional variable tests that are significant for the German interest rate and the exchange rate.

TABLE 1

Summary of previous best results on UK reaction function

Authors	*Period*	*Lagged dependent variable*	*Inflation*	*Output gap*	*German interest rate*	*Other variables included*
Clarida *et al*. (1998)	1979M6–1990M10	0.87*	0.48*	0.28*	0.60*	—
Dornbusch *et al*. (1998)	1986M4–1995M4	0.98*	—	—	—	—
Angeloni and Dedola (1999)	1980M1–1987M12	0.87*	0.32*	0.60*	1.32*	M3, real exch rate
	1988M1–1997M4	0.86*	0.93*	0.73*	0.45*	$/DM exch rate
Kuttner and Posen (1999)	1984M1–1989M12	0.86*	1.64*	[−0.21]	—	—
	1992M10–1999M4	0.79*	0.52	[−0.29*]	—	—
Muscatelli *et al*. (2000)	1985Q1–1996Q3	Not reported	1.40*	0.64*	—	—
Nelson (2000)	1979Q2–1987Q1	0.37*	0.38*	0.15*	—	—
	1987M3–1990M9	0.52*	0.00	0.45*	1.11*	—
	1992Q4–1997Q1	0.29*	1.27*	0.47*	—	—

Notes:
*Indicates significant at 95% level.
Clarida *et al*. GMM estimation of single equation; results refer to baseline with German interest rate variant.
Dornbusch *et al*. FIML simultaneous estimation of reaction functions for Germany, France, Italy, Spain, Sweden and UK; unit root problem so no long run coefficients given (short run coefficients: current inflation 0.20, lagged inflation −0.26, output gap 0.29, current German interest rate 0.64, lagged German interest rate −0.63).
Angeloni and Dedola Bivariate GMM estimation of reaction functions for Germany and the UK; backward and forward inflation variants in both cases.
Kuttner and Posen OLS; unemployment used instead of output gap; long run coefficients calculated from short-run results reported.
Muscatelli *et al*. RLS (recursive least squares); variable addition tests also carried out for money growth, exchange rate, German interest rate.
Nelson IV first and third sub-periods forward-looking, in the second subperiod the inflation coefficients are restricted to zero.

find a significantly lower coefficient on the German interest rate for their second period.

IV. Estimation

Our estimated reaction functions are based on the 'augmented' Taylor rule tradition reviewed in section III, and are derived as restrictions on the following equation (see also Clarida *et al.*, 1998):

$$r_t = (1-\rho)\alpha + (1-\rho)\beta E_t\pi_{t+j} + (1-\rho)\gamma E_t(y-\bar{y})_{t+k} \\ + (1-\rho)\delta E_t\mathbf{q}_{t+l} + \rho r_{t-1} + \varepsilon_t, \quad (3)$$

where r denotes the interest rate (the three month interbank rate), π denotes the (year-on-year) inflation rate, $(y-\bar{y})$ is a measure of the output gap,[11] and $\mathbf{q}$ is the vector of external factors. E_t denotes the expectations operator and the time subscripts $\{j, k, l\}$ reflect the authorities' forecast horizons for inflation, output and external factors, respectively (the data are described in detail in the appendix). The parameter ρ captures interest rate smoothing behaviour on the part of the authorities, although, given the reduced-form nature of (3), it also captures other dynamic misspecification errors in the equation. We estimate variants of equation (3) over the three periods identified in section I: the pre-ERM period (April 1985–September 1990); the post-ERM period (October 1992–April 1997) and the MPC period (May 1997–July 2002).[12]

Our specification differs from those found in the existing literature insofar as our vector of international factors consists of both US and German policy interest rates. The UK was a member of the ERM for only two years (a period over which we obviously do not estimate the reaction function) and prior to this shadowed the DM closely for only 13 months from March 1987 to March 1988. However, as Cobham's (2002b) analysis of the making of monetary policy in the UK suggests, the authorities' concerns about international developments were broader than simply a focus on Germany, a point confirmed by our empirical analysis. Conditional on interest rates, however, we found that the corresponding bilateral exchange rates play no statistically significant role in the UK authorities' reaction function.[13]

[11]In contrast to Clarida *et al*. (1998) we measure output directly in terms of real GDP rather than industrial production, but interpolate the gaps to obtain monthly data. Industrial production covers a decreasing share of total GDP, and over the sample the correlation between real GDP and industrial production changes sharply, especially towards the end of the sample. Industrial production is therefore not a sufficient statistic for the evolution of real economic activity.

[12]The estimation sample is truncated at July 2002 to allow up to a 12-month lead on inflation.

[13]This is true regardless of whether the bilateral exchange rates are deemed to be elements of $\mathbf{q}$ or instruments for the endogenous variables in the model. For the pre-ERM period only, we found it is possible to define a model in which UK interest rate movements can be explained in terms of *both* foreign interest rates and exchange rates. However, conditioning on exchange rates adds little to the predictive power of the reaction function since the weights on the exchange rates are extremely small, and the extended model is dominated by the international model reported in Table 2.

Since equation (3) represents a forward-looking reaction function it can be estimated using an errors-in-variables estimator such as Hansen's (1982) GMM estimator in which expected values of regressors are replaced with their actual future values and instrumented accordingly. The vector of instruments, **Z**, consists of lagged values of the interest rate, inflation, and the output gap for each country. Hence we assume that whereas interest rate determination in the UK may respond to German and US factors, the opposite is not the case: only German and US inflation and output serve as instruments for their respective own interest rates.

Our estimating equation takes the form

$$r_t = (1-\rho)\alpha + (1-\rho)\beta_t\pi_{t+j} + (1-\rho)\gamma(y-\bar{y})_{t+k} + (1-\rho)\delta\mathbf{q}_{t+l} + \rho r_{t-1} + v_t, \tag{4}$$

where v_t is now a non-Gaussian error term consisting of the pure stochastic error, ε_t, plus the authorities' forecast errors on inflation, output and foreign variables, conditional on their information at time t. If forecasts are made over more than a single period ahead the error term and hence the covariance between it and the instrument set, $(\mathbf{Z}'\mathbf{vv}'\mathbf{Z})$, will have a moving average representation of order $(n-1)$, where n is the forecast horizon. Under these conditions the Generalized Method of Moments (GMM) estimator suggested by Hansen (1982) can be used to generate consistent estimates of $(\mathbf{Z}'\mathbf{vv}'\mathbf{Z})$.

Preliminary estimation of equation (4), especially for the pre-ERM period, suggested that when US and German interest rates are included in **q**, the domestic variables became insignificant. This led us to estimate rival 'domestic' and 'international' reaction functions, where $\delta = 0$ and $\beta = \gamma = 0$, respectively, alongside the general, nesting equation (4) across all three periods. We then discriminate between the rival specifications using a combination of encompassing and over-identification tests.

We employ two encompassing tests of the rival models. The first is Ericsson's encompassing test of the null that each model variance-dominates its rival (see Ericsson, 1983), and the second a 'joint' test against the null that each model encompasses the linear combination of the two.[14] We then use the Davidson and Mackinnon (1993) over-identifying restriction test against the null that the international factors satisfy the valid instrument condition (i.e. $\mathrm{Cov}(\mathbf{Z}'\mathbf{v}) = 0$) but do not enter independently into the reaction function. The intuition is that if the authorities react directly to external factors this implies that the parameter vector δ in (4) must be non-zero. Hence these external

[14]Smith (1992) describes how encompassing tests for models estimated by GMM may be constructed. Since these procedures have not yet found their way into standard estimation packages, we report encompassing tests based on the IV rather than GMM estimation of the parameters of (4). The full IV estimation results, which are available on request, are very similar to those derived using the GMM estimator.

factors cannot be employed as instruments and some variant of the 'international' or 'nesting' model is a more appropriate representation of the policy rule. However, if external factors are part of the information set used by the UK authorities to forecast (current) output and (future) inflation but do not enter the reaction function itself, so that they can be restricted to lie in the vector **Z** only, we accept the domestic model specification in which the role of external factors is indirect. This test has a Chi-squared distribution under the null.

V. Results

Table 2 summarizes the long-run parameter estimates from our three variants of (4) estimated across the different sub-samples following the strategy outlined above. The results reported here represent the outcome of a more extensive specification search in which we examined alternative specifications of the elements of the vector of international effects, the forecast horizons, and the vector of instruments. From this we find in general that in the post-ERM and MPC periods the authorities react to contemporaneous values of the output gap and of the international factors (so that $k = l = 0$), but to a 9-month-ahead forecast horizon for inflation. For the pre-ERM period 12-month-ahead horizons for inflation and the output gap appear to be optimal, although given that domestic factors play such a minor role in determining interest rates in this period, the results do not change significantly if we assume different horizons, including those used for the later periods.[15] Finally, again given the limited sample sizes in each of the periods, we select our vector of instruments parsimoniously. The results reported in Table 3 employ a maximum of four (monthly) lags on each variable, although experimentation with alternative specifications suggested that this limitation was readily accepted by the data.[16]

Pre-ERM (April 1985–October 1990)

Our results suggest that over the pre-ERM period interest rate setting in the UK was overwhelmingly constrained by external factors. Not only does the international model unambiguously dominate the domestic model, it also dominates the joint nesting model: conditioning on foreign interest rates

[15]These horizons were derived from a systematic search over the grid defined by the range $j, k = 0$, 3, 6, 9, 12, where we selected the optimal horizon as that which minimized the equation standard error.

[16]The principal alternative involved using the first, third, sixth and twelfth lags. However, given the fact that all the series used are highly autoregressive, the dominant lag was invariably the first one and hence the results were relatively insensitive to the precise specification of longer lags.

TABLE 2

Long-run reaction function parameters

	Pre-ERM 1985(4)–90(9)			*Post-ERM 1992(10)–97(4)*			*MPC 1997(4)–2002(7)*		
	Nesting	*Domestic*	*International*	*Nesting*	*Domestic*	*International*	*Nesting*	*Domestic*	*International*
Dependent variable: 3 month UK interbank interest rate, GMM estimation (monthly data), [t-statistics in brackets]*,**,†,‡,§									
Constant α	0.5%	6.2%	1.3%	1.1%	4.0%	1.6%	−1.4%	0.6%	5.8%
	[0.35]	[2.35]	[1.36]	[1.81]	[4.21]	[2.69]	[0.31]	[0.35]	[1.96]
Inflation β	−0.16 [0.96]	1.19 [3.07]	—	0.20 [0.88]	0.69 [1.96]	—	2.24 [1.79]	1.89 [2.66]	—
Gap γ	0.13 [0.84]	−0.45 [1.41]	—	−0.23 [2.09]	0.32 [2.60]	—	0.12 [0.08]	1.30 [5.15]	—
Iger $\delta 1$	1.30 [5.03]	—	1.09 [10.18]	0.11 [1.82]	—	0.18 [3.41]	−0.76 [0.82]	—	−1.78 [1.41]
Ius $\delta 2$	0.60 [3.89]	—	0.56 [3.44]	0.73 [7.35]	—	0.67 [8.56]	0.86 [0.83]	—	1.13 [2.28]
Lagged r ρ	0.50**	0.87**	0.57**	0.57**	0.64**	0.60**	0.90**	0.85*	0.95*
	[4.14]	[11.01]	[6.72]	[14.99]	[13.96]	[18.78]	[13.13]	[13.51]	[30.86]
SD¶	2.087%			0.595%			1.047%		
Eq SE††	0.437%	0.486%	0.429%	0.142%	0.185%	0.146%	0.146%	0.146%	0.169%
Within-period stability tests‡‡									
	0.383	0.327	0.422	0.160	0.157	0.165	0.303	0.292	0.471
Between-period stability tests§§									
1985–90	—	—	—	10.802**	6.599**	10.130**	11.894**	7.407**	5.451
1992–97	0.565	0.280	0.491	—	—	—	2.202**	2.403**	1.715
1997–02	0.323	0.813	0.717	2.200**	1.100	2.461**	—	—	—
Non-nested encompassing and over-identification tests¶¶,††† (M1 = domestic model; M2 = international model)									
M1 E M2 (Ericsson IV)		4.910			6.614			1.710	
M1 E nesting model		[0.000]			[0.000]			[0.2371]	

continued overleaf

TABLE 2

(continued)

	Pre-ERM 1985(4)–90(9)			*Post-ERM 1992(10)–97(4)*			*MPC 1997(4)–2002(7)*		
	Nesting	*Domestic*	*International*	*Nesting*	*Domestic*	*International*	*Nesting*	*Domestic*	*International*
M2 E M1 (Ericsson IV)			−0.015			−1.042			6.775
M2 E nesting model			[0.8840]			[0.1479]			[0.000]
Davidson–Mackinnon OID test		6.54			18.35			6.98	

Notes:
* and ** denotes autoregressive parameter significantly different from 1 at 5% and 1% significance level.
†See equation (4) and data appendix for variable definitions.
‡For the domestic model additional instruments are: 1–4 lags of inflation and the output gap plus lags 2–4 of the interest rate; for the international model additional instruments are 1–4 lags of US and German interest rates and output gaps (full-sample only) plus lags 2–4 of the UK interest rate. The instrument set for the nested models is the union of domestic and international factor models.
§GMM estimation embodies a correction for the MA($n-1$) error term. For the domestic and nested models this correction is of order MA(11) for 1985–90 and MA(9) for 1992–97 and 1997–2002. For the international model the error term is assumed to be non-autocorrelated [i.e. MA(0)].
¶,††Sub-sample standard deviation of the policy interest rate (SD) and equation standard error (eq. SE), respectively.
‡‡Hansen (1992) LM test of null of within-sample parameter constancy (5% critical value = 0.470).
§§Chow test of null that estimated model displays parameter stability over alternative samples.
¶¶For each model we report the test of the null that the domestic model encompasses the international model (denoted M1 E M2) and *vice versa*. The Ericsson IV test has a normal distribution under the null, the nesting model an *F*-distribution. For the latter, we report the probability of the Type I error.
†††Test of the null that external factors can be restricted to lie in the vector of instruments only. The test has a Chi-square distribution with $r = 8$ (5% critical value = 15.51).

TABLE 3

Long-run reaction function parameters (policy rate)

	Pre-ERM 1985(4)–90(9)		*Post-ERM 1992(10)–97(4)*		*MPC 1997(4)–2002(7)*	
	Domestic	*International*	*Domestic*	*International*	*Domestic*	*International*
Dependent variable: Bank of England policy interest rate, GMM estimation (monthly data), [t-statistics in brackets]						
Constant α	6.8%	1.1%	4.3%	1.3%	1.0%	4.2%
	[1.27]	[1.05]	[3.59]	[1.66]	[0.89]	[1.94]
Inflation β	0.93	—	0.65		1.77	—
	[0.91]		[1.53]		[3.25]	
Gap γ	−0.38	—	0.19		1.38	—
	[0.70]		[1.17]		[6.65]	
Iger $\delta 1$	—	1.12	—	0.30	—	−1.30
		[9.65]		[4.66]		[1.56]
Ius $\delta 2$	—	0.63	—	0.69	—	1.20
		[3.54]		[6.89]		[2.80]
Lagged r ρ	0.91*	0.57**	0.71**	0.63**	0.83*	0.94
	[17.99]	[6.66]	[12.33]	[13.94]	[15.31]	[31.33]
SD	2.190%		0.559%		1.085%	
Eq SE	0.536%	0.462%	0.209%	0.176%	0.147%	0.175%

implies that neither domestic inflation nor the output gap enter the estimated reaction function significantly and, in fact, the inflation term is 'incorrectly' signed. By contrast, the international model fits the data well. The equation standard error is approximately 43 basis points per month against a mean interest rate for that period of 11.3% and an unconditional monthly standard error of 209 basis points. Other things equal, UK interest rates reacted more or less point-for-point to movements in German rates but much less to movements in US interest rates.[17]

Post-ERM (October 1992–April 1997)

The UK's dramatic exit from the ERM on 'Black Wednesday' in October 1992 and the adoption of inflation targeting with floating exchange rates saw domestic factors playing a much greater role in determining the stance of UK monetary policy, although external factors continued to significantly influence the authorities' interest rate setting. This is broadly confirmed by the results in the central panel of Table 2, which indicate that while exit from the ERM coincided with a weakening of the link to German interest rates, it did not,

[17]In an earlier version of this paper (Adam, Cobham and Girardin, 2001) we showed that the UK rate was more responsive to the German rate during the formal DM-shadowing episode (from March 1987 to March 1988).

in fact, fully 'liberate' UK monetary policy from the constraints imposed by external factors. Hence, although the domestic model is more recognizable as a conventional Taylor-style monetary policy rule – both inflation and the output gap are now positive and statistically significant – the estimated inflation coefficient is less than unity, and, as the encompassing tests indicate, this model is still statistically dominated by the international model, albeit less decisively than in the earlier period. The effect of the external factors does, however, change noticeably compared to the pre-ERM period. Post-1992, it would appear that the US interest rate exerted a much stronger impact on the UK authorities' rate-setting behaviour than the German rate.

Monetary Policy Committee (May 1997–July 2002)

Since the creation of the MPC in May 1997, there is strong evidence in support of a conventional Taylor rule representation of UK interest rate setting. The domestic model is now well-defined, both statistically and in terms of the conventional wisdom on the Taylor rule. But more importantly, the encompassing tests suggest that it dominates both the international and nesting models while the over-identification tests indicate that the international factors can be restricted to enter only the vector of instruments, implying that external factors impact on interest rate setting but do so indirectly via their impact on the authorities' inflation and output expectations.

The estimated coefficients on the lagged interest rate, expected inflation and the output gap are systematically higher than those in Taylor's original formulation and are towards the high end of the distribution of estimates from other studies of UK monetary policy reaction functions (Table 1), although close to those (of 0.79, 2.15 and 0.93) reported by Clarida *et al*. (2001) for the US in the Volcker–Greenspan period.

Robustness and identification

The validity of our approach rests on two maintained hypotheses. The first is that our estimates genuinely reflect the (reduced form) parameters of the authorities' reaction function rather than other factors driving UK interbank interest rates, for example, interest rate arbitrage effects; and the second is that these parameters are not constant across the three regimes. On the former our general strategy has been to estimate reaction functions for periods defined on the basis of turning points in the narrative history. However, we can also check directly for the importance of interest rate arbitrage effects by re-estimating the reaction functions for the UK policy rate rather than the interbank rate, on the argument that arbitrage effects might be expected to drive the latter but not the former. Table 3 reports the results of re-estimating the reaction functions where the Bank of England's policy rate is the dependent variable. The reaction functions reported in this form inevitably fit the data less well,[18] but there are no significant changes in the parameter estimates (for example, the inflation and output gap coefficients for the MPC period are now 1.77 and 1.38 as opposed to 1.89 and 1.30), and the encompassing or over-identifying tests lead us to exactly the same inferences as before. We are confident, therefore, that our results do in fact reflect the parameters of the authorities' rule rather than private sector interest-rate arbitrage effects.

We test the second hypothesis, that the sample partition is genuinely reflected in the data, using standard parameter-stability tests. We test for *within-period* parameter stability using Hansen's (1992) LM test, and for *between-period* stability using conventional split-sample Chow tests. Both sets of test statistics are reported below each set of reaction functions in Table 2. The LM tests suggest that only in the case of the 'international' model in the MPC period is there any evidence of parameter instability, and even then the null is only very marginally rejected. The interpretation of the between-period Chow tests is a little less straightforward, for two reasons. First, it only makes sense to consider tests based on models that are statistically well-defined within their own sub-period.[19] Second, the discriminatory power of the Chow test will depend on the underlying variance of the dependent variable across the sample: if the 'base-period' equation standard error is large in absolute terms it will be correspondingly harder (easier) to reject the null of parameter constancy across an extended sample where the equation standard error is lower (higher). This is a feature of the data here since the equation standard error falls from around 0.50% in the pre-ERM period to less than 0.20% in the post-ERM period, reflecting, in large measure, the decline in the mean interest rate rather than an increase in the explanatory power of the (best) model. Nonetheless, and bearing these caveats in mind, a clear pattern emerges. Starting with the domestic model we note that for the post-ERM era, the stability tests decisively reject the null that this model exhibits constant parameters over the pre-ERM period. We find the same result comparing the MPC era domestic model with the post-ERM period.[20] A similar pattern emerges if we look at the international

[18]Given the econometric issues referred to above, this is unsurprising.

[19]To see this, notice that the 1985–90 domestic model appears to exhibit parameter stability over all sub-samples. However this is a pure artefact: the model is so poor within its own sub-sample, that it is trivially easy to accept any arbitrary restriction that the parameter estimates derived from a different sample are consistent with (i.e. lie within the confidence interval of) the original model estimates.

[20]Notice that this result is not symmetrical since the Chow test for the 1992–97 model extended to 2002 has a value of 1.10 against a critical value of 1.56. However, we already know that this original model is relatively poorly specified over its 'own sample' and hence the null will be correspondingly harder to reject.

model, which displays symmetric non-constancy between the post-ERM and MPC periods. Taken together, these results confirm the validity of our sample-splitting strategy.

We have also checked that our main results are robust to an alternative specification of the output gap. The regressions reported use the full-sample output gap, defined as the deviation of actual real GDP from a deterministic trend where the latter is estimated using the latest vintage of UK real GDP data. In line with Orphanides and van Norden's (2002) findings for the US, we have shown elsewhere (Adam and Cobham, 2004) that the bulk of the difference between this series and the 'real-time' output gap is due to the rolling (real-time) estimation of trend output, rather than of the revisions to GDP data that occur subsequent to its first release, and the full-sample series may therefore be preferable.[21]

In this case, however, the two series yield remarkably similar results. For the pre-ERM and post-ERM periods our encompassing tests indicate that, with the real-time as well as the full-sample data, the international model dominates the domestic and nesting models. In the MPC period, the tests suggest that neither model dominates the other (although the domestic model is 'closer' to encompassing the joint and international models), but this reflects the fact that across the periods the weight on the real-time output gap is consistently lower than that on the full-sample gap measure, particularly during the 1992–97 period, when the output gap is large.[22]

VI. Summary and conclusions

The institutional context in which monetary policy has been formulated in the UK has changed significantly since the mid-1980s. To investigate how these changes have influenced monetary policy decision-making, we estimate and compare the characteristics of monetary policy reaction functions over three major periods: the pre-ERM era from 1985 to 1990, the post-ERM period from 1992 to 1997, and the era of the MPC, which began in May 1997. Our analysis delivers two key results. The first is that US as well as German influences should clearly be included in the UK monetary policy reaction function. When that is done, however, it turns out that domestic variables have no contribution to make in the pre-ERM period and only a weak contribution at best in the post-ERM period: US and German interest rates on their own provide the best explanation of UK interest rates in the former period and also, though less unequivocally, in the latter. In the MPC period, on the other hand, our results

[21]The real-time gaps were calculated from the data base provided by Eggington, Pick and Vahey (2002), updated from *Economic Trends* and *Economic Trends Annual Supplement*.

[22]Regression results using real-time output gap data are available from the authors on request.

suggest that interest rates are explained by domestic factors, with the international influences entering only as instruments for output and inflation.

The second, and most important, result of the paper is that the major change in the conduct of monetary policy was not the introduction of inflation targeting in 1992, but the granting of instrument-independence to the Bank of England in 1997. After 1997, the reaction function is significantly different, with interest rates set on the basis of domestic variables and the international influences contributing only as instruments for UK output and inflation. The reaction function is also much closer to those found for the US and other G3 countries by Clarida *et al*. (1998, 2000). The implication is that institutional constraints really make a difference.

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Appendix A: Data and estimation

Data

With the exception of UK GDP, the data used in the paper are taken from Datastream Advance Version 3.5. Datastream mnemonics are reported in parentheses.

Interest rates: The 3-month interbank interest rates for UK (LDNIB3M), Germany (FIBOR3M) and the US (USCOD3M). **Inflation rates** are computed from the UK retail price index excluding mortgage interest payments (UKRPAXMIF) and consumer price indices for the US and Germany (USOCPCONF and BDOCPCONF, respectively). **Exchange rates** are nominal spot rates between sterling and the US dollar (UKXRPD) and sterling and the Deutschemark/Euro (BDWU5005). The **output gap** for Germany and the US is defined as the monthly interpolation of the deviation from a linear quadratic trend of real quarterly GDP (BDRGDP, and USRGDP, respectively). Trend output is estimated over the period 1984–2000. UK real GDP is taken from Economic Trends and Economic Trends Annual Supplement (various issues) (code CGCE from 1998 onwards).

Estimation

The GMM estimates were generated using the non-linear estimation command in TSP (Version 4.5). The $(\mathbf{Z}'\mathbf{v}\mathbf{v}'\mathbf{Z})$ covariance matrix was estimated using the NMA option with lag length $n - 1$, where n is the forecast horizon. We use the Bartlett Kernel to ensure $(\mathbf{Z}'\mathbf{v}\mathbf{v}'\mathbf{Z})$ is positive semi-definite.