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DOES INSTITUTIONAL CHANGE REALLY MATTER? INFLATION TARGETS, CENTRAL BANK REFORM AND INTEREST RATE POLICY IN THE OECD COUNTRIES

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We estimate forward-looking interest rate reaction functions for the G3 and some inflation targeters. Shifts in the conduct of monetary policy are detected for the USA and Japan. In contrast with the existing literature, we show that these countries only shifted to policies consistent with an implicit inflation-targeting regime in the 1990s. Inflation targets and central bank reforms in Sweden, the UK, Canada and New Zealand only led in some cases to changes in policy responses, and changes in policy *pre-date* the introduction of targets. We challenge the one-model-fits-all approach towards monetary policy that permeates much of the current literature.

1 Introduction

The importance of institutions has become a central tenet of modern macroeconomics. This is apparent in the literature on time inconsistency in monetary policy, which sees the appointment of an independent central bank and inflation targeting as key elements in achieving price stability (Persson and Tabellini, 1997). In this paper we report estimates of forward-looking interest rate reaction functions for a number of OECD economies. We demonstrate that significant changes in monetary policy behaviour have occurred in a number of OECD economies, but that such shifts in policy are not necessarily linked with institutional change.

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Our results contrast sharply with those obtained by others (Clarida *et al.*, 1998, 2000). Whilst confirming that central bank policies became more 'conservative' at the beginning of the 1980s (and inflation expectations followed), we also detect some significant shifts in the subsequent conduct of monetary policy in the USA and Japan. In addition, the introduction of inflation targeting and central bank reforms in countries like the UK, Sweden, Canada and New Zealand has not always led to major changes in the way in which central banks of those countries react to the objectives of economic policy. Our results suggest that institutional change either initiated a slow shift in policy or confirmed an earlier change in the policy stance. Our findings are consistent with the view that changes in society's attitude towards inflation—and the ensuing policies—are relatively more important than formal institutional arrangements (Posen, 1993; McCallum, 1996).

The paper is structured as follows. In Section 2 we outline the main contributions and results of the paper. In Section 3, we provide a link between estimated interest rate reaction functions and the theory of monetary policy design. This provides the background for our empirical models. Our empirical estimates are presented in Section 4, and Section 5 concludes.

2 The Existing Literature and Key Results

2.1 Context: The Existing Literature

There are several recent contributions on modelling interest rate reaction functions and we need to distinguish our contribution carefully from those of previous authors. In general three broadly different approaches have been used in modelling monetary policy behaviour. First, a number of researchers have used vector autoregressions (VARs) to estimate the way in which policy actions depend on a set of macroeconomic indicators, and how in turn policy actions are transmitted to key macro variables. Bernanke and Blinder (1992) use the US Federal Funds rate to analyse the transmission mechanism in the USA. Christiano *et al.* (1994) and Bernanke and Mihov (1995, 1997) *inter alia*[1] have refined this approach by analysing alternative measures of monetary policy and identification mechanisms for the estimated VARs. Second, some researchers have focused on estimating single-equation (structural) reaction functions for monetary policy instruments (see, for instance, Groeneveld *et al.*, 1996; Muscatelli and Tirelli, 1996; Clarida and Gertler, 1997; Clarida *et al.*, 1998, 2000). Third, Rudebusch (1995, 1996) uses data from forward-

[1]For an excellent survey, see Christiano *et al.* (1998) who analyse the advantages and pitfalls of the VAR approach to identifying monetary shocks.

looking financial markets to construct measures of unanticipated shocks to monetary policy.

In this paper we adopt the second of these approaches. The third approach, which uses financial market data, can be a useful method to test whether changes in monetary policy have an impact on inflation expectations. But it does not allow one to interpret policy changes in terms of policy reaction functions and the likely implications for output and inflation stabilization. The VAR approach has some advantages in that it allows one to jointly model both the endogenous policy response and the transmission mechanism by making only minimal assumptions about the causal links and the timing of the authorities' reactions to new macroeconomic data. However, the results from VAR models do seem to depend critically on the assumptions made about which variables to include in the VAR, and on the existence of a time-invariant transmission mechanism and reaction function (see Rudebusch, 1996). Given the number of variables one usually includes in a VAR and the limited number of observations, it becomes difficult to conduct any stability analysis by, say, using 'rolling VARs'. This is especially the case if there have been frequent changes in either the policy regime or the financial system which might affect the timing of the policy response and the nature of the transmission mechanism.[2]

Indeed, as noted by Christiano *et al.* (1998), VAR modellers usually prefer not to report or to interpret estimated policy rules, because if the actual policy rule is forward-looking the estimated coefficients of such VAR-estimated 'policy rules' will be difficult to interpret. Instead, VAR models are primarily designed to construct measures of monetary policy shocks for use in analysing the transmission of monetary shocks[3] (even though there are differing views of the robustness and usefulness of the monetary policy shock measures obtained from VARs—see Rudebusch, 1996; Bagliano and Favero, 1998; Christiano *et al.*, 1998). Overall, it does seem that VARs are less useful in undertaking an empirical analysis of regime changes in the conduct of monetary policy.

2.2 Key Results and Value Added

Our focus on single-equation (forward-looking) structural reaction functions is similar to that in Clarida and Gertler (1997) and Clarida *et al.* (1998, 2000), and allows us to analyse shifts in monetary policy regimes using recursive estimation techniques. We extend these earlier studies in

[2]Although Bernanke and Mihov (1995) do allow for a limited amount of time variation in their VAR model. For a recent application of Bayesian VAR analysis, see Muscatelli and Trecroci (2000).

[3]See for example Eichenbaum and Evans (1995).

the following ways. First, by presenting recursive estimates of these reaction functions, we can detect marked changes over the last two decades in the way monetary policy has been conducted. This allows us to find some new and surprising results. For instance, the announcement of explicit inflation targets and the move to more independent central banks in several OECD economies has not led to a major contemporaneous change in the way monetary policy reacts[4] to the final objectives of economic policy in the 1990s. This has important implications for the large theoretical literature that has emerged on central bank independence.

Second, unlike Clarida *et al.*, we use alternative methods to estimate our measures of expected inflation and potential output. This is based on the assumption that the private sector is imperfectly informed about the central bank's preferences, whereas the central bank is imperfectly informed about the permanent and cyclical components of output growth (see Blinder, 1998; Orphanides, 1999). Interestingly this also leads to new results. Whilst the interest rate reaction function for Germany is reasonably stable, there seem to be some signs that monetary policy rules in the USA and Japan are less stable than one might have imagined over the period 1985–99. Our estimates suggest that US policy has shifted to react to inflationary expectations more vigorously, and with a shorter lead. This fits well with most anecdotal accounts of US policy in the 1990s, but is in sharp contrast with the results of previous empirical studies. Moreover, we find that the output coefficient has the wrong sign in the mid-1980s, suggesting that real interest rates were too low at a time when the output gap was positive. Only in the 1990s does the output coefficient sign become positive once more. Our results stand in sharp contrast with those of Favero and Rovelli (1999), who argue that since 1982 the Fed acted as a strict inflation targeter, rejecting the hypothesis that output stabilization is an independent argument in the loss function of the Fed.[5] Japanese policy

[4]These results are consistent with those obtained in related work by Groeneveld *et al.* (1996), who reject the hypothesis of a structural break following the switch to inflation targeting in Canada, New Zealand and the UK. However, their models are backward-looking, use mainly domestic target variables, and focus solely on the overall stability of the fitted reaction functions during the early 1990s. Our modelling approach in this paper examines the stability of the model parameters over a longer sample and uses a measure of expected inflation and of potential output. There are also alternative approaches in the literature to assess the impact of inflation targets. For instance, Freeman and Willis (1995) examine credibility effects on the yield curve, and Almeida and Goodhart (1996) use a variety of different methods to assess the impact of inflation targeting on the behaviour of monetary authorities.

[5]They estimate a model of the US economy using a VAR specification for the output gap, inflation and a commodity price index over the period 1960–98. By doing this they estimate the parameters in the aggregate demand and supply functions. Then they use generalized method of moments (GMM) methods to estimate, over the period 1983–98, an interest rate rule which allows the central bank desired trade-off between output and inflation to be identified. Their approach requires full information, rational expectations and invariance of the structural model to changes of the monetary policy regime.

also seems to have exacerbated the cycle in the late 1980s, and only seems to have fallen in line with inflation targeting post-1990. This confirms the suspicion that Japanese policy was inappropriately geared to external objectives (the relationship with the USA) in the 1980s. A key contribution of this paper is that we move away from the full-sample rational expectations hypothesis adopted in Clarida *et al.* (1998, 2000), which sits rather uneasily with the 'learning under uncertainty' that characterizes central banks' decision-making (see Blinder, 1998; Woodward, 2000).

Third, our study provides some empirical evidence on the lead with which the expected inflation rate enters estimated reaction functions. This is of considerable interest to theoretical analyses of the trade-off between output and inflation variability in monetary policy design (see Haldane and Batini, 1998; Goodhart, 1999). There is no other empirical study that we are aware of which compares the forward-lookingness of monetary policy across OECD countries.[6]

3 Interest Rate Reaction Functions and the Theory of Monetary Policy Design

In this section we show how a forward-looking interest rate reaction function can emerge from a simple Barro–Gordon-type theoretical model of monetary policy design. Consider the following model for current inflation in the presence of costly price adjustment as in Calvo (1983) or Rotemberg (1983) (Rotemberg and Woodford (1998) propose a sticky-price model which has similar implications):

$$\pi_t = p_t - p_{t-1} = \beta\pi^{e}_{t+1} + \varphi(y_t - y^*) \tag{1}$$

where current inflation π depends on inflation expectations and the current output gap, and y^* is potential output. The output gap is given by

$$y_t - y^* = -s[(R_t - \pi^{e}_{t+1}) - r^*] + \varepsilon_t \tag{2}$$

Output deviations from the natural rate depend on a supply shock ε_t and the deviations of the expected real interest rate from its equilibrium value r^*. Following Svensson (1998), suppose that the monetary policymaker's loss function is given by

$$L = \chi(\pi_t - \pi^*) + (y_t - \tilde{y})^2 + \rho[R_t - E(R_t)]^2 + \rho_1(R_t - R_{t-1})^2 \tag{3}$$

where the authorities penalize not only deviations of output from an output target $\tilde{y}$, which exceeds the natural level y^*, and of inflation from a target π^* (as in Svensson, 1997a), but also penalize deviations and changes in the policy instrument.

[6]Although in a later paper Clarida *et al.* (2000) have considered reaction functions with different lead structures.

This formulation assumes that stabilization policy via interest rate changes is costly, and that for this reason shocks are never fully stabilized in the long run. Svensson's model highlights the risk of instability of an anti-inflationary policy by assuming that the policymaker penalizes deviations of R_t from zero. Instead the formulation in (3) assumes that the policymaker knows the level of inflationary expectations, and consequently chooses a sequence for R_t. However, in the event of shocks hitting the economy, the authority decides whether to deviate from the nominal interest rate implied by the state of inflationary expectations. Solving the model under discretion, so that the monetary authority minimizes (3) with respect to the nominal interest rate, taking expectations as given, yields an interest rate reaction function

$$R_t = wr^* - A^* + b^*\pi^{e}_{t+1} + c^*\varepsilon_t + d^*R_{t-1} \tag{4}$$

where the coefficients are

$$w = \frac{s^2\varphi^2\chi + s^2}{s^2\varphi^2\chi + s^2 + \rho_1} \qquad b^* = \frac{s^2\varphi^2\chi + s^2 + \chi\beta\varphi s}{s^2\varphi^2\chi + s^2 + \rho_1}$$

$$c^* = \frac{s\varphi^2\chi + s}{s^2\varphi^2\chi + s^2 + \rho + \rho_1} \qquad d^* = \frac{\rho_1}{s^2\varphi^2\chi + s^2 + \rho_1}$$

$$A^* = \frac{s(\tilde{y} - y^*) + \chi\varphi s\pi^*}{s^2\varphi^2\chi + s^2 + \rho_1}$$

Note that we need the interest rate adjustment costs ρ_1 to be small enough to avoid an unstable response following output shocks, as current inflation depends on expected future inflation.[7] Note also that in a reaction function such as (4) the interpretation of the constant term is different from that in Clarida *et al.* (1998, 2000). This is because Clarida *et al.* do not consider interest smoothing as part of the authorities' optimization process: instead this is added as an *ad hoc* adjustment. Basically, our model implies that the constant $wr^* - A^*$ is a function of the real interest rate, inflation target and inflationary bias, while in Clarida *et al.* it is referred to as simply the long-run component of the real interest rate. This demonstrates that one has to be careful in interpreting the estimated parameters of an interest rate reaction function, as these largely depend on the way the monetary authorities' loss function is specified.

[7]In general the system will be stable as long as $\chi\beta\varphi s > \rho_1$, which implies $b^* > 1$ and that the expected inflation response to the output gap is positive. Under rational expectations the reaction function in (5) yields the following equilibrium inflation rate:

$$\pi_t = \frac{\beta\rho_1}{\chi\beta\varphi s - \rho_1}r^* + \frac{\beta[s(\tilde{y} - y^*) + \chi\varphi s\pi^*]}{\chi\beta\varphi s - \rho_1} - \frac{\beta\rho_1}{\chi\beta\varphi s - \rho_1}R_{t-1} + \varphi(1 - sc^*)\varepsilon_t$$

If there is no uncertainty about the monetary authorities' policy objectives,[8] both inflation and interest rates will fluctuate stochastically around a given mean.[9] However, in practice, the authorities' policy goals may not be observable (see Faust and Svensson, 1998; Muscatelli, 1998, 1999) and may vary over time (see Cukierman, 1992). Suppose for instance that price and wage setters are uncertain about the policymaker's preferences over inflation (his/her credibility):

$$\chi_t = \chi_{t-1} + \omega_t \qquad \omega_t \sim (0, \sigma_\omega^2) \tag{5}$$

Suppose also that the policymaker cannot accurately predict the supply shock, but has to forecast it (this forecast being private information), and that wage and price setters cannot disentangle the uncertainty due to the supply shock ε and the preference shock ω.[10] The private sector will then perceive the interest rate reaction function as

$$R_t = r^* - \alpha_0 + \alpha_1 \pi_{t+1}^{\mathrm{e}} + \alpha_2 \varepsilon_t^{\mathrm{f}} + \alpha_3 R_{t-1} \tag{6}$$

where the α are functions of the same parameters (and $\alpha_1 > 1$ like b^*) as in (4) but with χ^{e} (the expected value of χ) and where ε^{f} is the forecast of the supply shock. The private sector will update their expectations of χ and ε^{f} each period on the basis of the variances of ε and ω in a standard signal extraction problem (see Cukierman, 1992; Muscatelli, 1999; Walsh, 1999).

Thus, following a regime change (e.g. the central bank being granted independence) where some parameter of the monetary authority's objective function shifts, if the regime change was not fully credible one would see a gradual adjustment of inflation and interest rates to a new average level.

In practice one can estimate a forward-looking reaction function for interest rates along the lines of (6) by constructing a series for expected inflation and the expected supply shock (or equivalently the expected output gap), using an optimal updating scheme for the expected variables (such as the Kalman filter). If one then observes the timing of significant shifts in the estimated reaction function parameters these should correspond to major shifts in the policymaker's preferences (institutional regime).[11]

[8] Muscatelli (1998, 1999) analyses a model of inflation targeting with uncertain central bank preferences.

[9] Given the nature of the supply shocks in the model, both inflation and interest rates will be stationary.

[10] In a monetary policy committee, the preference shock ω can capture fluctuations in votes between different 'wings' of the committee.

[11] They might also be due to shifts in the underlying structural model which change the way in which the authorities form their expectations about inflation and the output gap, but in this case we should observe changes in the models for expected inflation and the output gap.

It is worth noting that by estimating a simple forward-looking interest rate reaction function such as (6) one is not trying to capture the exact way in which the monetary authorities actually react to economic indicators which affect real economic activity and expected inflation. Instead estimated forward-looking reaction functions based on (6) capture the *implicit* way in which the central bank's operational rules/decisions translate into a reaction function in terms of expected inflation and output gaps. Thus, for example, one might find some instability in the estimated reaction function parameters which may not be due to a change in policy preferences, but which might be due to a shift in the intermediate targets used to achieve this outcome.[12] For instance, in the case of the UK, we know that in the early 1980s there was a move away from monetary targets once it became clear that monetary policy was becoming over-contractionary. But in general major and permanent shifts in the estimated parameters will reflect corresponding shifts in policy preferences.

Therefore, estimating reaction functions such as (6) does not allow one to directly estimate the authorities' preference parameters. It does allow one to judge whether the operational rules have been stable and whether the reliance on certain intermediate targets/indicators has taken place at the expense of meeting final output stabilization and inflation objectives. Where some variability is found in the estimated parameters this must be attributable either to changes in the authorities' preferences regarding output and inflation stabilization, or to some shift in the relationship between the final policy objectives and the macroeconomic indicators/intermediate objectives used by the authorities.

The theoretical literature on policy design has closely examined the performance of forward-looking (inflation expectations) policy rules (see Faust and Svensson, 1998; Haldane and Batini, 1998; Svensson, 1998). In part this is because of the emphasis given in some countries to the central bank's inflation forecast (cf. the Bank of England's regular inflation forecast based on current interest rate policies). In part it is because recent contributions to the inflation-targeting debate (Svensson, 1997b; Haldane and Batini, 1998; Rudebusch and Svensson, 1998) have shown the quasi-optimality of interest rate policy rules based on inflation forecasts. In general the form of the inflation-forecast-based rules considered by these authors is

$$r_t = \theta r_t^* + \varphi r_{t-k} + \gamma E_t \pi_{t+j} + \lambda (y_t - y_t^*) \tag{7}$$

where r_t is the short-term *ex ante* real interest rate, r_t^* represents the long-run equilibrium real interest rate, while $E_t \pi_{t+j}$ is the j-period-ahead

[12]This point is also stressed by Christiano *et al.* (1998) in the context of VAR models.

inflation rate expected at t. Past values of the interest rate to capture interest rate smoothing behaviour and the output gap are also included.[13] This can be rewritten in terms of the nominal interest rate:

$$R_t = \alpha + \varphi R_{t-k} + \omega E_t \pi_{t+j} + \lambda(y_t - y_t^*) \tag{8}$$

where $\omega = 1 + \gamma$, while α includes, as in equation (4), the equilibrium real interest rate and the inflation target.

Comparing (8) with (6) we see that, by generalizing the latter to include a longer lead for inflation and a longer lag for the interest rate smoothing term, and substituting the output gap for the supply shock forecast, (6) is identical to the forecast-based policy rule in (8).[14]

In what follows, we actually estimate reaction functions of the following type:

$$R_t = \alpha + \sum_{i=1}^{k} \varphi_i R_{t-i} + \omega E_t \pi_{t+j} + \lambda(y_t - y_t^*) \tag{9}$$

Typically we find that a maximum lag length of $k = 2$ is sufficient to capture the degree of interest rate smoothing. Having estimated the basic reaction function in (9), we then search for the appropriate lead (j) for the inflation forecast term $E_t \pi_{t+j}$ on the basis of goodness-of-fit.

As noted in Haldane and Batini (1998), the specification of reaction functions such as (9) allows one to analyse a number of issues. First, the parameters (ω, j), i.e. the weight the bank puts on expected inflation and the lead term on it, determine the responsiveness of the instrument to changes in the forecast and the forward-lookingness of the bank's horizon. In addition, the parameters (j, k, φ) capture the degree of inertia in the interest rate policy. Finally, a value of λ different from zero implies that the rule explicitly includes some reaction to deviations of output from potential.

One potential problem in estimating structural reaction functions is highlighted by Favero and Rovelli (1999), who argue that finding a significant output effect in the reaction function might simply mean that the central bank treats the current output gap as a leading indicator for expected inflation. In this case the output gap should be collinear with the proxy for expected inflation, or should predict inflation forecast errors. We were not able, as explained below, to find substantial collinearity

[13]Haldane and Batini (1998) note that the omission of an output gap term does not mean that the authorities do not stabilize output, since by adjusting the degree of interest rate smoothing and the lead in the inflation forecast one can trade off output stabilization against inflation stabilization.

[14]In practice the simple lag/lead specification for the interest rate reaction function derives from the simple assumptions made about the transmission mechanism in this model. This point is discussed in detail by Clarida *et al.* (2000). They note that their results are not too sensitive to the choice of lag/lead of the output gap or expected inflation term.

between our measures of inflation and the output gap, while the correlation with inflation forecast errors is small and often has the wrong sign.[15]

4 Empirical Estimates

4.1 The Monetary Policy Instrument Variables

As in other recent attempts to estimate monetary authorities' reaction functions (see Clarida *et al.*, 1998), we focus on short-term money market rates as the policy instrument.[16] Clearly there are difficulties in identifying a single interest rate measure as *the* monetary policy instrument for the whole of our sample period (see Bernanke and Mihov, 1995). One might want to use different interest rate measures as the policy instrument at different times (e.g. discount rates in the early part of the sample and repo or call money rates towards the end of the sample period). But such fine distinctions would inevitably be arbitrary, and in any case short-term money market rates will largely reflect the authorities' monetary policy stance under different operating procedures.

We do not, in this paper, adopt the procedure suggested by Orphanides (1999) of using unrevised data series (i.e. the data which were actually available to the authorities when the decisions were taken). The main reason for this is that we want our results to be comparable to those of other recent contributions (e.g. Clarida *et al.*, 1998, 2000), and this would not be possible if we used a completely different data set.

4.2 Measuring Inflation Expectations and the Output Gap

There are different methods to obtain measures of inflation expectations and the output gap. Clarida *et al.* (1998, 2000) use a quadratic trend to obtain a measure of potential output and hence deviations of actual output from this trend. In order to obtain a measure of inflation expectations, Clarida *et al.* use the errors-in-variables approach to modelling rational expectations of inflation whereby future actual values are used instead of the expected values, and GMM estimation is used as in standard dynamic rational expectations models.

Turning first to the output gap, one disadvantage of fitting non-linear trends to the data is that it involves using the full sample in the construction of the output trend. This involves making the assumption that the policymaker uses future information on the path of output in the

[15] These results are not shown here for brevity, but are available from the authors upon request.
[16] See the Data Appendix for details of the interest rate variables used.

evaluation of the potential output trend. Indeed, forecasting the future path of potential output is *the* key problem which central bankers face (see Blinder, 1998)—assuming it away by not modelling the learning process seems strange in our view.

Instead we proxy the learning process on the part of the monetary authorities by using the structural time series (STS) approach proposed by Harvey (1989) to generate series for the output gap and expected inflation. There are several advantages in using this approach. The first is that it provides a useful and intuitive way of decomposing a series into trend and cyclical components, which is particularly useful when one tries to estimate a series for an unobservable trend such as potential output. Second, the modelling approach lends itself readily to employing a Kalman filter estimation procedure, which allows one to proxy the learning process by policymakers and economic agents. Third, the STS models are parsimonious models which have reasonably rich autoregressive integrated moving-average processes as their reduced forms.

Essentially, we estimate models for real GDP and inflation for each country, seeking to disentangle the trend, cycle and irregular components.[17] In the case of GDP, a convenient decomposition of the series was made possible by applying the Kalman filter on the trend component. Subsequently, the latter was computed on the basis of one-step-ahead predictions of the state vector. This way, estimates of potential output are based only on past information, rather than on the full sample.

In the case of inflation, we simply computed one-step-ahead prediction errors from a univariate STS model to obtain a measure of expected and unanticipated inflation. Again, the models' parameters are updated only gradually, as new data are added. In both cases, the STS methodology makes the assumption that agents make the best use of all available knowledge in a regime of imperfect information. In contrast, use of a non-recursive estimation approach, such as errors-in-variables, has the defect of using information from the whole sample, thus ignoring policy regime shifts.

Figure 1 compares our measure of the output gap with that obtained from a Hodrick–Prescott filtering procedure for the USA.[18] This shows that our measure differs markedly from that used in previous studies, and indeed that quadratic or Hodrick–Prescott detrending procedures tend to exaggerate the cyclical component. The reason why our series seems more

[17]The STAMP 5.0 software was used to estimate the STS models. Output and inflation were found to be I(1), and to have a significant cyclical component. The estimated STS models are available on request from the authors. For a similar approach to forecasting inflation in the presence of potential structural breaks, see Stock and Watson (1999).

[18]Plots of the output gaps and inflation expectations measures are not reported for reasons of space, but are available on request. Fitting a quadratic trend, as in Clarida *et al.* (1998), produces a more marked cyclical pattern.

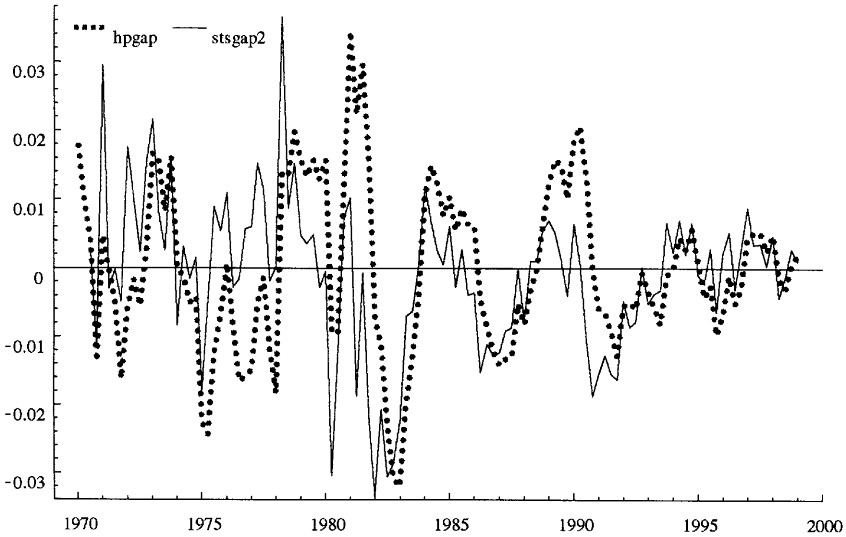

FIG. 1 USA—Hodrick–Prescott and Filtered STS Measures of the Output Gap

'noisy' is that it is what remains of the series once the trend component has been removed (i.e. the cycle plus the irregular component).[19] We also investigate whether the data-generation processes of output and inflation in the countries of our sample have undergone any significant break over the periods we study. If there are breaks in the underlying processes for potential output and inflation, one expects the policy authorities to learn gradually about such shifts. Hence our Kalman filter approach generating expected inflation and trend output measures would be a more appropriate one than the full-information assumption made in the existing literature.

There is now a wide range of technical contributions devoted to studying trend breaks in unit roots and the problems associated with the endogeneity of break points.[20] Here we limit ourselves to understanding whether the data-generation processes of output and inflation have undergone major shifts over our sample period, leaving aside the determination of the exact number and position of the break points. We employ the class of tests in Hansen (1992b) and Andrews and Ploberger (1994). They all belong to the broad category of Chow-type tests with unknown break point, and build upon the assessment of the significance of

[19]Whether the irregular component should be included is obviously a matter for debate. In practice it is unlikely to be a pure forecast error, and hence should probably be included. However, our estimates do not change markedly when only the cycle measure is included.

[20]For an extensive account of this debate, see Stock (1994).

the value of the likelihood ratio, Wald and Lagrange multiplier statistics derived from recursive switching regressions. We use the MeanF, SupF and the L_c variants advocated by Hansen (1992b). The first two tests have parameter constancy as their null against the alternative of sudden breaks, whereas the latter statistic is for the alternative of a smooth change. In our case, the testing strategy requires prior estimation of univariate models for output and inflation. We adopt simple autoregressive specifications including trends and constants when required, and the semiparametric, fully modified estimator of Phillips and Hansen (1990) and Hansen (1992a). The latter is a two-step methodology that first estimates the asymptotic covariance matrix of the system and then provides regression parameters.[21] Tests on the null of parameter stability are finally carried out.[22]

The test statistics are reported in Table 1 for all the countries we examine. In the case of inflation, the sample covers the years 1971Q3–1997Q4, whereas for GDP some data constraints for Sweden and New Zealand substantially shortened the sample we used. The first column of each section displays the estimated statistic for the three tests. For both output and inflation the results are clear-cut, with all countries displaying instability at standard significance levels. These estimates are robust to changes in the kernel and bandwidth parameter chosen to filter the residuals, as well as to alternative functional forms for the specified models. These results support our modelling approach.

4.3 Estimating Policy Rules

In estimating (9) the appropriate reaction lead to expected inflation (j) was usually found to be four quarters for most countries. This result broadly agrees with the findings of Batini and Haldane's dynamic simulations of a calibrated theoretical model, where the optimum lead length on the inflation forecast is found to lie between three and six quarters. However, in some key cases, as we shall see, different results emerge.

Lags of the dependent variable are always found to be significant. This is not surprising, as interest rate smoothing considerations appear to be a generally accepted feature of monetary policy (see Almeida and Goodhart, 1996; Bernanke and Mihov, 1997; Goodhart, 1999). We detect a substantial amount of policy inertia in all the countries examined.

[21]Additional details on the testing procedure, as well as on the results we summarize here, can be obtained from the authors upon request.

[22]Estimates were conducted by adapting a modified GAUSS code kindly provided by Bruce Hansen. Hansen (1992b) tabulates asymptotic critical values for each of the tests performed here.

Table 1
Tests for Parameter Instability

Test	*GDP*	*Inflation 1972–97*	*Inflation 1980–97*
USA			
L_c	0.548*	0.543*	1.762***
MeanF	7.946**	10.713***	9.495***
SupF	23.682***	68.269***	48.400***
Germany			
L_c	0.339	1.553***	0.651**
MeanF	4.931	14.422***	8.195***
SupF	47.889***	25.059***	58.105***
Japan			
L_c	0.298	0.743**	0.998***
MeanF	5.587*	5.913*	7.987***
SupF	15.877*	29.530***	28.035***
UK			
L_c	0.514*	0.624**	2.049***
MeanF	13.150***	5.915**	11.247***
SupF	43.787***	29.530***	48.371***
Canada			
L_c	0.254	1.406***	1.383***
MeanF	5.449*	12.803***	12.500***
SupF	24.662***	49.207***	55.908***
Sweden			
L_c	0.334	1.137***	0.959***
MeanF	7.846**	10.848***	24.192***
SupF	14.731*	31.085***	73.784***
New Zealand			
L_c	0.481*	1.097***	0.484*
MeanF	9.047***	21.772***	8.051***
SupF	26.849***	92.627***	27.016***

L_c, MeanF, SupF are defined as testing the null of stability against non-constancy on the parameters of univariate autoregressive models for inflation (four-quarter change in CPI) and real GDP. Constants and linear time trends were included when relevant.
*, **, *** Significance of the relevant F statistic at 10 per cent, 5 per cent and 1 per cent, respectively (for tabulated critical values, see Hansen,1992b).

One difference between our approach in this paper and that in other studies is that we do not take for granted, or assume, any structural break in the behaviour of the monetary authorities. Also, we have not imposed any particular structure for any shifts in monetary policy. This is because we want to test whether any change can be detected in correspondence to announced regime shifts.

For this reason, we first estimated the reaction function (9) for each country over the full sample period—extending in the G3's case back to the end of Bretton Woods—and conducted a recursive analysis on the magnitude and the significance of regressors. Using structural stability tests we were then able to detect major breaks in interest rate policy. As most major shifts in interest rate policies took place in the 1970s or early 1980s, we then re-estimated a reaction function for each country over the post-1980 period, and again performed recursive tests and stability analysis. This allowed us to detect any parameter shifts in the reaction functions since 1980, and to interpret these shifts and any structural breaks in the light of announced institutional changes or shifts in policy regime.

Finally, as in Clarida *et al.* (1998), we allow for the possibility that the monetary authorities might have responded to other intermediate objectives not included in our baseline specification in (9). The reason for doing this is twofold. First, if the baseline model does not perform well, we can check whether this is due to the targeting of some other intermediate objective. Second, institutional accounts of monetary policy suggest that these other variables might matter. Lagged values of money growth, changes in the exchange rates and influences from relevant foreign interest rates were included as additional regressors.

4.4 Interest Rate Reaction Functions: the G3 Countries

We now turn to our empirical results. Monetary institutions in the G3 (the USA, Germany and Japan) have been remarkably stable during the sample period; i.e. the relationship between the political system and monetary institutions has not changed in these countries.[23] In the USA and Germany the central bank enjoys/enjoyed a relatively high degree of independence (see Grilli *et al.*, 1991; Cukierman, 1992) and is best defined as a 'goal-independent' central bank,[24] i.e. a bank which is not held accountable for achieving a certain policy target. For instance German monetary policy has been defined as a regime of 'disciplined discretion' (Laubach and Posen, 1997), whereas US monetary policy during the Greenspan era has been defined as 'pre-emptive monetary policy without an explicit nominal anchor' (Mishkin and Posen, 1997). An interesting issue examined here is whether the success of the Fed in recent years has

[23]Since 1979, European Monetary System (EMS) membership might have constrained the Bundesbank's ability to retain control of monetary policy. Most discussions on the deutschmark's role in the EMS have concluded that the Bundesbank largely retained her independence (Von Hagen, 1995).

[24]For instance, both Neumann (1996) and Clarida and Gertler (1997) argue that the Bundesbank pursues multiple objectives and is flexible in attaining them, i.e. emphasis sometimes shifts from one policy target to another. For a similar view see Mishkin and Posen (1997). For a contrasting view, stressing continuity in the Bundesbank's use of monetary targets, see Issing (1997).

been achieved by changing the way in which interest rates respond to policy objectives.

Our estimated models are reported in Tables 2–8. For ease of exposition we report only the *long-run static solutions* of the model. This makes it more straightforward to interpret the effects in terms of the theory and with earlier studies which estimate the model long-run parameters directly using GMM. Each regression contains one or two lags of the dependent variable. These autoregressive terms are always highly significant and generally are large, with their sum equal to 0.7–0.9 in all cases. As we are only reporting long-run effects, the standard errors shown in Tables 2–8 are asymptotic standard errors for each estimated coefficient. Table 2 reports the estimated reaction function for Germany, respectively for the full sample period[25] and since 1980. The estimates for the whole sample show that interest rates reacted to inflation expectations (with a point estimate greater than one)[26] and output. Adding the US Federal Funds rates only marginally improves the fit of the interest rate reaction function. The variable addition tests show that neither money growth nor the exchange rate (measured as the index of real effective exchange rate) seemed to exert an independent significant effect on German interest rates. There seems to be a mildly significant effect of the Federal Funds rate, but this effect is not stable over the cycle, and probably picks up a correlation between the German and US business cycles. This confirms analogous results in Clarida and Gertler (1997) and Bernanke and Mihov (1997). Since 1971, the Bundesbank set target ranges for the growth of broad monetary aggregates, but over the last 15 years of its operation actual growth rates often exceeded (fell short of) the upper (lower) limit of the targeted band.[27] This confirms most modern accounts of the Bundesbank's monetary policy stance which suggest that monetary targets were not the Bank's primary objective but that discretionary undershoots and overshoots of the target bands were allowed where this did not impair the achievement of the inflationary objective.

The diagnostic tests for the estimated model in Table 2 show some signs of non-normality (and possibly autoregressive conditional heteroscedasticity (ARCH)) in the residuals, as is the case for all our estimated models, but this is due to the bunching of a small number of large residuals at the end of the 1970s, and this is apparent from the post-1980 estimates.

[25]Our sample period ends in 1999Q2, so the issue of the switch from the Bundesbank to the European Central Bank does not arise.

[26]As shown in Section 3, this is sufficient to ensure dynamic stability, providing the dynamic IS and aggregate supply equations (equations (1) and (2)) have the signs expected by economic theory. Estimating the dynamic IS and aggregate supply equations is beyond the scope of this paper.

[27]See Von Hagen (1995), Issing (1997).

Table 2
Germany

Regressor/ Sample	Constant	Expected inflation	Output gap	Federal Funds rate	Variable addition tests[c]		Summary statistics	
1970Q2–	−0.179	1.529	1.143		Money	0.456	R^2	0.886
1999Q2[a]	(2.016)	(0.4437)	(0.616)		growth	(0.634)	σ	0.875
							DW	1.79
					Exchange	0.560	AR 1–5 $F(5, 108)$	1.540 [0.1836]
					rate	(0.572)	ARCH 4 $F(4, 105)$	4.538 [0.0020]
							Normality $\chi^2(2)$	60.542 [0.0000]
							RESET $F(1, 112)$	1.259 [0.2643]
1970Q3–	−1.679	1.149	0.783	0.425	Money	0.194	R^2	0.892
1999Q2[b]	(1.823)	(0.3)	(0.401)	(0.164)	growth	(0.824)	σ	0.857
(adding Federal							DW	1.70
Funds rate)					Exchange	0.027	AR 1–5 $F(5, 106)$	1.842 [0.1109]
					rate	(0.974)	ARCH 4 $F(4, 103)$	4.185 [0.0035]
							Normality $\chi^2(2)$	65.493 [0.0000]
							RESET $F(1, 110)$	1.947 [0.1657]
1980Q1–	0.348	1.895	0.759		Money	0.561	R^2	0.958
1999Q2[a]	(1.327)	(0.376)	(0.372)		growth	(0.573)	σ	0.532
							DW	2.01
					Exchange	1.098	AR 1–5 $F(5, 69)$	0.734 [0.6003]
					rate	(0.339)	ARCH 4 $F(4, 66)$	2.419 [0.0574]
							Normality $\chi^2(2)$	20.209 [0.0000]
							RESET $F(1, 73)$	0.487 [0.4874]
1980Q1–	−0.296	1.514	0.531	0.237	Money	0.412	R^2	0.996
1999Q2[b]	(1.247)	(0.2968)	(0.280)	(0.116)	growth	(0.523)	σ	0.520
(adding Federal							DW	1.69
Funds rate)					Exchange	2.030	AR 1–5 $F(5, 68)$	1.138 [0.3487]
					rate	(0.158)	ARCH 4 $F(4, 65)$	2.370 [0.0615]
							Normality $\chi^2(2)$	13.733 [0.0010]
							RESET $F(1, 72)$	1.079 [0.3025]

[a] Derived from an RLS regression of the interest rate on a constant, four-quarter-ahead expected inflation, output gap and one lag of the dependent variable.
[b] As for the note above, but now with one lag of the Federal Funds rate on the right-hand side.
[c] We tested for the addition of other regressors. Zero restrictions on current and lagged money growth (M3, quarter and annual difference) and both the change and the lagged (shown in the table) value of the exchange rate *vis-à-vis* the US$ were tested by an *F* version of the Wald test. *p* values in parentheses.

Asymptotic standard errors of estimated parameters are in parentheses. AR is a Lagrange multiplier test for the hypothesis of no serial correlation; ARCH checks whether residuals have an ARCH structure, with no ARCH as the null; Normality tests the normality of residuals; RESET tests the null of no functional misspecification. *p* values in square brackets.

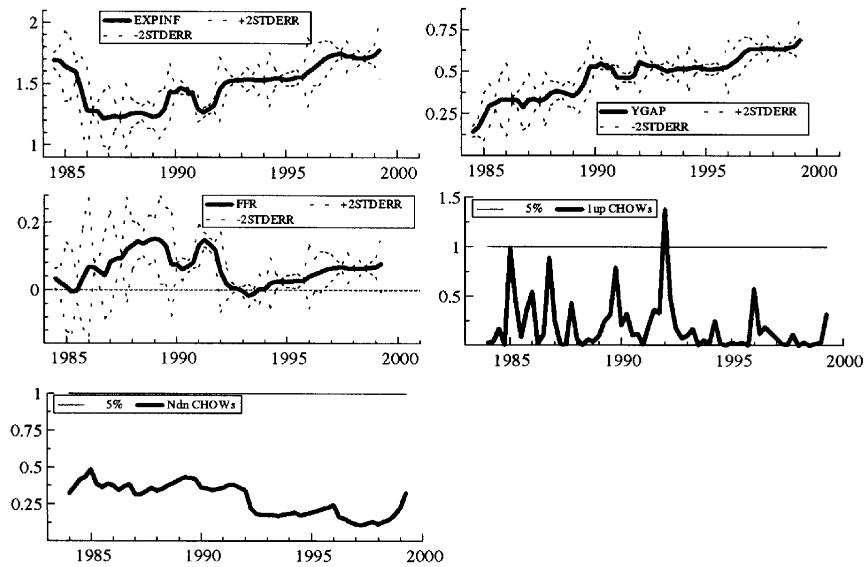

FIG. 2 Germany, 1980Q1–1999Q2: Recursive Coefficients and Standard Error Bands; One-step, N-step Up Chow tests (5 per cent)

The estimated reaction function for Germany does not display any major shifts, apart from slight variation around the time of reunification, with the estimated coefficients relatively constant across subsamples. We also found that a four-quarter lead for expected inflation works best for both the full sample and the post-1980 sample. Figure 2 shows one-step up and N-step down Chow tests, as well as the estimated coefficient and standard error bands and t values for the expected inflation, Federal Funds and output gap regressors for the post-1980 regression. The latter figure shows the long-run estimated coefficient as well as the asymptotic standard errors.[28] This confirms the stability of the Bundesbank's policy rule, but shows that the size of the estimated response to the output gap fell slightly after the unification shock in 1990–91. This shows, in line with recent work (Clarida *et al.*, 1998), that monetary policy in Germany reacted systematically to cyclical conditions, even though the Bundesbank's declared monetary strategy (see Issing, 1997) was expressed in terms of monetary targets. Note also that this result is not dependent on whether the money growth variable is added to the interest rate equation before or after the inflation and output gap term. Indeed, following the Exchange Rate Mechanism (ERM) crises

[28]Again note that these are the *asymptotic standard errors*, not the usual recursive least squares (RLS) standard errors. They were computed using the authors' own GAUSS routines and plotted using GiveWin. This explains why there is so much variation from period to period and they are not as smooth as RLS standard errors.

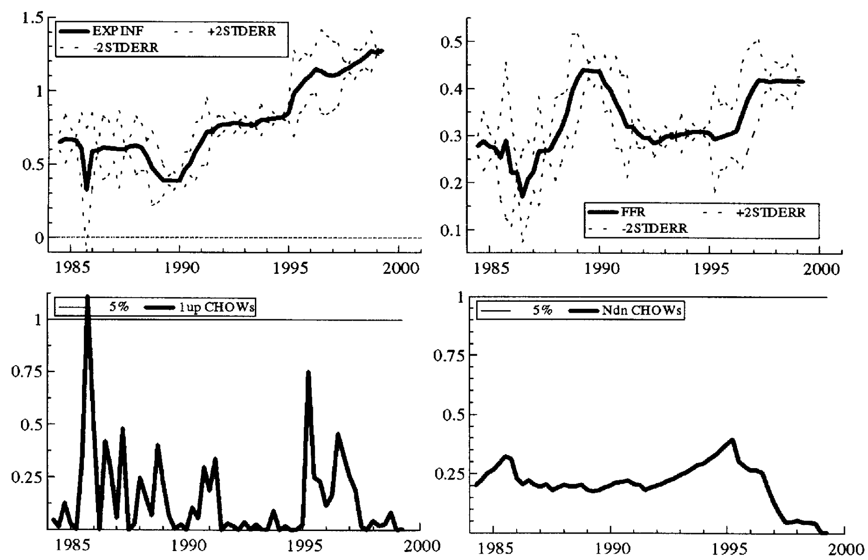

FIG. 3 Japan, 1980Q1–1999Q2: Recursive Coefficients and Standard Error Bands; One-step, N-step Up Chow tests (5 per cent)

in 1992–93, the Bundesbank's responsiveness to the output gap seemed to increase slightly. But the overall picture is one of overall stability, indicating stable underlying policy preferences and operational rules.[29]

Our estimates for the Japanese reaction function (Table 3, Fig. 3) over the whole sample show an insignificant coefficient on the output gap, whereas that on expected inflation is significant but well below one until post-1995. Furthermore, the equation displays some instability in the mid-1980s. We tried to improve on this by including additional regressors. It turns out that the US Federal Funds rate exerts a strong influence on Japanese policy. This confirms that in the 1980s Japanese monetary policy might have been hamstrung by agreements on managing the value of the yen:US$ rate. It also confirms the casual observation that Japanese policy might not have been sufficiently geared towards domestic targets (see *The Economist*, 17 July 1998) and that this might have contributed to the excessive deflation in Japan in the 1990s. Unfortunately what seems to have happened is that by the time that Japan's monetary policy became more consistent with the pursuit of output and inflation stabilization, the domestic economy was already in crisis.

[29]We would not wish to argue that the Bundesbank's policy stance has never shifted over time. Indeed, as Berger and Woitek (1998) have shown, the political composition of the Bundesbank's Council seems to have exerted an important influence on the way it formulated monetary policy since 1945.

Table 3
Japan

Regressor/ Sample	Constant	Expected inflation	Output gap	Federal Funds rate	Variable addition tests[c]		Summary statistics	
1970Q3–	2.527	0.7245	0.719		Money	0.97	R^2	0.964
1999Q2[a]	(1.011)	(0.1789)	(0.526)		growth	(0.38)	σ	0.601
							DW	2.21
					Exchange	3.49	AR 1–5 $F(5, 106)$	2.108 [0.0700]
					rate	(0.034)	ARCH 4 $F(4, 103)$	2.791 [0.0301]
							Normality $\chi^2(2)$	40.439 [0.0000]
							RESET $F(1, 110)$	6.670 [0.0111]
1970Q3–	1.99	0.5305	0.096	0.686	Money	0.307	R^2	0.971
1999Q2[b]	(1.23)	(0.098)	(0.311)	(0.137)	growth	(0.736)	σ	0.549
(adding Federal							DW	2.13
Funds rate)					Exchange	1.700	AR 1–5 $F(5, 105)$	1.590 [0.1692]
					rate	(0.19)	ARCH 4 $F(4, 102)$	3.432 [0.0113]
							Normality $\chi^2(2)$	20.266 [0.0000]
							RESET $F(1, 109)$	4.427 [0.0377]
1980Q1–	1.112	2.029	0.406		Money	1.39	R^2	0.968
1999Q2[a]	(1.053)	(0.479)	(0.475)		growth	(0.25)	σ	0.539
							DW	1.92
					Exchange	0.971	AR 1–5 $F(5, 67)$	1.018 [0.4143]
					rate	(0.383)	ARCH 4 $F(4, 64)$	1.605 [0.1837]
							Normality $\chi^2(2)$	27.08 [0.0000]
							RESET $F(1, 71)$	8.637 [0.0044]
1980Q1–	−1.000	1.277	Dropped	0.417	Money	0.302	R^2	0.970
1999Q2[b]	(1.152)	(0.4464)	because not	(0.167)	growth	(0.74)	σ	0.528
(adding Federal			significant				DW	1.84
Funds rate)					Exchange	0.640	AR 1–5 $F(5, 68)$	0.792 [0.5591]
					rate	(0.530)	ARCH 4 $F(4, 65)$	3.224 [0.0179]
							Normality $\chi^2(2)$	22.487 [0.0000]
							RESET $F(1, 72)$	10.385 [0.0019]

[a] Derived from an RLS regression of the interest rate on a constant, four-quarter-ahead expected inflation, output gap and two lags of the dependent variable.
[b] As for the note above, but now with one lag of the Federal Funds rate and with no output gap (because insignificant) on the right-hand side.
[c] We tested for the addition of other regressors. Zero restrictions on lagged money growth and changes in the current, once- and twice-lagged trade-weighted exchange rate were tested by an F version of the Wald test. p values in parentheses.

Asymptotic standard errors of estimated parameters are in parentheses. AR is a Lagrange multiplier test for the hypothesis of no serial correlation; ARCH checks whether residuals have an ARCH structure, with no ARCH as the null; Normality tests the normality of residuals; RESET tests the null of no functional misspecification. p values in square brackets.

The US reaction function estimated over the whole sample period (Table 4) is characterized by a coefficient on inflation which is not significantly larger than one and by an insignificant coefficient on the output gap. However, diagnostic tests and recursive graphics show a marked period of instability between 1979 and 1982, when the Fed switched from interest rate targeting to monetary base targeting, which implied greater instability in money market rates. Since then, the Fed has opted for the targeting of money market (Federal Funds) rates. Goodfriend (1995) argues that the 1979–82 parenthesis of monetary base targeting also marked the Fed decision to aggressively clamp down on inflation expectations, which was accomplished by 1985. Clarida *et al.* suggest a stable policy environment post-1982. However, our recursive estimates show that this is not the case. The output coefficient has the wrong sign in the late 1980s, suggesting that interest rates were too low at a time when the output gap was positive. Only in the 1990s does the output coefficient sign become positive once more.

Our estimates over the post-1980 sample in Table 4 confirm that some important changes seem to have taken place. US policy does seem to have been less constant over time than Germany's. Interest rates seem to react to inflation expectations on a shorter horizon (a two-quarter horizon is found to work best post-1985) and with a larger coefficient when the reaction function is re-estimated over the latter part of the sample. Our results reverse the conclusions of Clarida *et al.* (1998) using different estimation methods, as they find an estimated coefficient on inflation which is much greater than one.

The picture changes completely if we focus on the post-1985 sample (see Table 4). The equation is now stable, and includes a coefficient on expected inflation with a point estimate greater than unity (although it is not significantly larger than one). The recursive graphs also confirm that in the 1990s the Fed was adjusting interest rates to follow the output cycle much more closely.[30] Figure 4 shows a significant output gap effect post-1991.

Overall, the Fed looks very different from the Bundesbank until the 1990s. On this point our results differ sharply from those of Clarida *et al.* The usual accounts suggest that, having successfully restrained inflation expectations in 1979–82, the Fed exploited her reputation to implement

[30]One caveat emerges from the theoretical model discussed above: in a full-information context, i.e. when the private sector has learned about the bank's preferences, inflation expectations are highly collinear with the output cycle. This might bias the estimated coefficient on the inflation expectations regressor downwards. On the other hand, we find only a very small correlation between our measures of expected inflation and the output gap. The analysis of the coefficients' covariance matrix also confirms that the correlation between the coefficients is small and often has the wrong sign. These results are available on request.

TABLE 4
USA

Regressor/ Sample	*Constant*	*Expected inflation*	*Output gap*	*Variable addition tests*[c]		*Summary statistics*	
1970Q3–	1.276	1.19	0.508	Money	2.58	R^2	0.804
1999Q2[a]	(2.018)	(0.3585)	(0.511)	growth	(0.08)	σ	1.48229
						DW	2.09
				Exchange	0.675	AR 1–5 $F(5, 106)$	5.2527 [0.0002]
				rate	(0.513)	ARCH 4 $F(4, 103)$	20.662 [0.0000]
						Normality $\chi^2(2)$	104.05 [0.0000]
						RESET $F(1, 110)$	2.764 [0.0993]
1980Q1–	0.690	1.715	0.368	Money	5.69	R^2	0.816
1999Q2[a]	(1.875)	(0.433)	(0.418)	growth	(0.05)	σ	1.570
						DW	2.14
				Exchange	0.637	AR 1–5 $F(5, 68)$	3.862 [0.0039]
				rate	(0.531)	ARCH 4 $F(4, 65)$	42.388 [0.0000]
						Normality $\chi^2(2)$	94.183 [0.0000]
						RESET $F(1, 72)$	3.133 [0.0809]
1985Q1–	3.159	0.723	0.748	Money	0.483	R^2	0.936
1999Q2[b]	(0.767)	(0.237)	(0.237)	growth	(0.617)	σ	0.468
						DW	1.68
				Exchange	0.042	AR 1–5 $F(5, 49)$	1.621 [0.1721]
				rate	(0.958)	ARCH 4 $F(4, 46)$	0.718 [0.5837]
						Normality $\chi^2(2)$	0.031 [0.9847]
						RESET $F(1, 53)$	0.0235 [0.8787]

[a] Derived from an RLS regression of the interest rate on a constant, four-quarter ahead expected inflation, output gap and two lags of the dependent variable.
[b] Derived from an RLS regression of the interest rate on a constant, two-quarter ahead expected inflation, output gap and one lag of the dependent variable.
[c] We tested for the addition of other regressors. Zero restrictions on lagged money growth and changes in a lagged trade-weighted index of effective exchange rate were tested by an F version of the Wald test on the baseline model augmented with each new variable. p values in parentheses.

Asymptotic standard errors of estimated parameters are in parentheses. AR is a Lagrange multiplier test for the hypothesis of no serial correlation; ARCH checks whether residuals have an ARCH structure, with no ARCH as the null; Normality tests the normality of residuals; RESET tests the null of no functional misspecification. p values in square brackets.

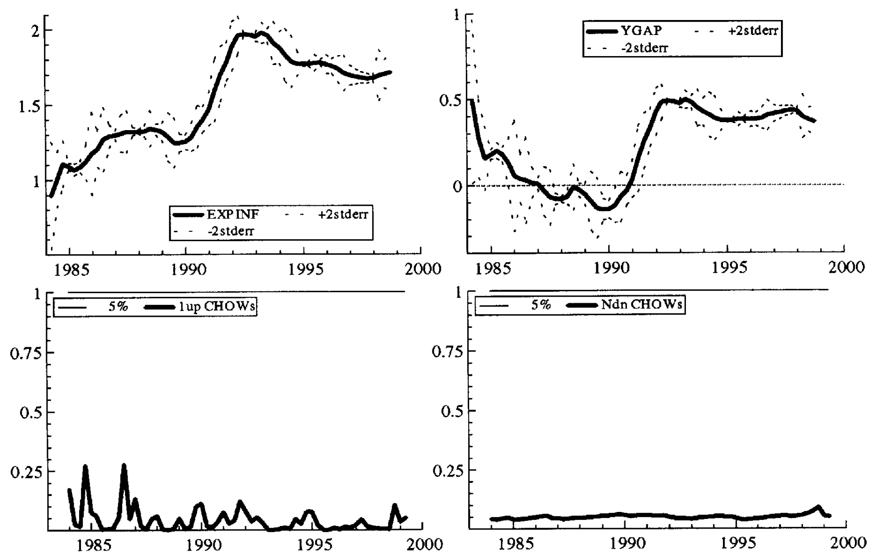

FIG. 4 USA, 1980Q1–1999Q2: Recursive Coefficients and Standard Error Bands; One-step, N-step Up Chow tests (5 per cent)

countercyclical policies. But the Fed's policy coefficients (particularly the output gap coefficient) suggest that a stable and correctly signed reaction function only operated since the early 1990s.

The US example is striking because it shows that a stable institutional set-up does not guarantee predictable policy outcomes. What we seem to observe in the US case is a progressive tightening of policy responses in the mid-1990s, probably linked to the uncertainty which existed in the Fed over what was happening to potential output (Woodward, 2000). Since 1995 the Fed's policy responses have become more predictable, but we await evidence from the current slowdown to see whether this stability will persist.

These findings also broadly illustrate a substantive difference between the Fed's and Bundesbank's monetary strategies. The Bundesbank appears to have responded more forcefully to movements in expected inflation than the Fed, judging from the inflation expectation coefficient. Some authors have suggested that this result is open to other interpretations. Mishkin and Posen (1997) label the Fed policy as 'just do it', or pre-emptive policy without a nominal anchor. Their argument is that monetary policy must act well in advance of a surge in inflation expectations since the full impact of monetary policy on inflation takes long lags. The main disadvantage of such a policy obviously lies in the difficulty of establishing a clear policy pattern with all the risks that this implies at times when the economy is being hit by major exogenous shocks. Our results suggest that such pragmatic and forward-looking policy should not be interpreted as if the Fed systematically reacted to longer-term expectations, as in the Bundesbank's case. In fact we found that shorter leads on the expected inflation variable (two instead of four quarters) seemed to work better in the case of the USA for the post-1985 sample. This confirms the casual observation that the Fed has chosen to signal its commitment to low inflation in recent years by reacting in advance to increases in inflationary expectations. However, in doing so, it has not always acted in a predictable way.

The other key results from this section are as follows. First, the G3 policy reaction functions look very different. One model does not fit all, in sharp contrast to the view expressed by Chinn and Dooley (1997). Second, despite having stable institutions, monetary policymaking in the G3 seems to have evolved gradually in different directions: in Germany it became more conservative post-unification. In Japan, it seems to have been led astray by inappropriate external objectives until recently. In the USA, the highly successful countercyclical monetary policy of the Fed seems to be purely a 1990s phenomenon. These discrepancies are not apparent in the existing empirical literature because of the tendency to only report full-sample estimates for the 1980s.

4.5 Interest Rate Reaction Functions: the Inflation Targeters

Turning to the other countries in our sample, we shall relate our results to major changes in the way in which monetary policy was conducted. As noted above, a variety of factors may cause shifts in estimated monetary policy reaction functions. Some of them, such as highly publicized institutional innovations and political changes, are easily identified from descriptive accounts of monetary policy and will be discussed here. Other shifts in the reaction functions may have occurred for 'technical' reasons. These include the instability of demand for money functions, which eventually caused the demise of monetary aggregates. Similarly, in other countries the authorities may have relied (formally or informally) on indicators or intermediate objectives that were subsequently abandoned. These too are important in understanding our results, and will be discussed as they show up in our estimates.

For most of the sample period, the central banks of the second group of countries (Canada, New Zealand, Sweden and the UK) have had limited independence in the conduct of monetary policy compared to the central banks of the G3 countries (see Grilli *et al.*, 1991; Cukierman, 1992). During the 1990s explicit inflation targets were announced in all countries, but there are important differences within the group in terms of institutional arrangements and the role the central bank plays in achieving the target. In fact only New Zealand's central bank and the UK's central bank (since 1997) have been given a legal mandate to achieve the inflation target.

In the UK, the Bank of England was only granted independence in 1997. However, there have been several changes in monetary strategy in the last two decades. The election of the Thatcher government in 1979 signalled a long-lasting shift in the collective attitude towards inflation. Instead of adopting an institutional approach the Conservative governments tried to build a reputation for their commitment to low inflation policies, experimenting first with monetary targets and then adopting a more eclectic approach to intermediate objectives from the mid-1980s. After a short spell of ERM membership in 1990–92, the government then opted for a new monetary policy framework involving the announcement of formal inflation targets. The Conservative government chose not to delegate the implementation of monetary policy to an independent and accountable central bank. Instead the government's own reputation was the ultimate guarantee of the policy commitment. However, the central bank played the key role of publicly assessing the overall consistency of the policy stance. The newly elected Labour government in 1997 then sought to further bolster the inflation-targeting framework by granting the Bank of England instrument independence. Monetary policy decisions are now taken by the Bank's Monetary Policy Committee.

Since the breakdown of M1 as an intermediate target in the early 1980s, until 1991 the Bank of Canada had not committed herself to any predetermined policy pattern, apart from the reiteration of the long-term goal of price stability. Neither intermediate target nor time frame was apparently cast in the attempt to pursue the long-run objective, while various monetary and credit aggregates (including the exchange rate with the US$) were used in turn as information variables. In 1991 the government and the bank set a sequence of year-to-year target bands for the inflation rate, so as to bring about a gradual reduction in inflation. However, the central bank was not granted a legislative mandate to achieve these inflation targets nor was a procedure established which would hold the central bank accountable for missing the targets. The 'doctrine of dual responsibility' traditionally attributes the ultimate responsibility for the results of monetary policy to the Minister of Finance. Thus, the Bank of Canada has enjoyed only a limited degree of formal independence (see Grilli *et al.*, 1991; Cukierman, 1992). Nonetheless, the central bank had been calling publicly for a stricter control on inflation since 1988, while from 1994 the degree of policy transparency has increased markedly (Mishkin and Posen, 1997).

Since 1977 Sweden had been pegging its currency unilaterally, first to a trade-weighted basket of currencies, then switching to the ECU in May 1991. However, the strength by which this commitment to the external anchor was pursued varied significantly, as numerous devaluations took place (Horngren and Lindberg, 1994). To some extent the Riksbank became less accommodating to inflation shocks after 1982. The marginal (overnight) rate was then extensively used to regulate large currency flows during the fixed exchange rate period. After the November 1992 crisis the Riksbank floated the krona and announced the unilateral adoption of an inflation target in January 1993.[31] However, the bank has never been granted an independent status, and political influences on the board are important (Svensson, 1995; McCallum, 1996).

Finally, we turn to the evolution of the monetary regime in New Zealand, which switched to inflation targeting in 1989. Historically, New Zealand's Reserve Bank had a degree of independence which ranked lowest amongst the OECD countries (see Grilli *et al.*, 1991; Cukierman, 1992). Correspondingly, New Zealand's inflation rate was well above the OECD average. Until the mid-1980s monetary policy relied mainly on regulation and administrative controls of capital markets. From 1985 the Bank turned to a more market-oriented approach to monetary control, and based policy decisions on a variety of indicators such as the exchange rate, the term structure of interest rates, monetary aggregates and output (see Leiderman and Svensson, 1995). The Reserve Bank Act, introduced in 1990 to establish a legislative commitment to price stability, gave the Government and the Central Bank Governor the mandate to agree on a policy target (it was decided that this should be an inflation target) and explicitly contemplates the possibility of the Governor's dismissal if the set target is not met.

Figure 5 plots the expected inflation series and the *ex ante* real interest rates computed using our expected inflation series for the group of inflation targeters in our study. It is interesting to note that in the case of Sweden, Canada and New Zealand *ex ante* real rates appear to have been pushed substantially higher and well above inflation expectations well before the announcement or the adoption of targets.[32] Also, inflation expectations, at least in the case of the UK, Sweden and New Zealand, seem to have been somewhat subdued prior to the announced regime changes. At first blush, the regime change seems to have simply consolidated the gains in terms of lower inflation.

Our estimates for the UK (Table 5) show that over the whole sample period the coefficient on inflation expectations is not significantly larger than one. Furthermore, the money market interest rate seems to have reacted to both the exchange rate and the money supply.

Given the instability in the estimated reaction function until the mid-1980s, we re-estimated the equation for the 1985–99 sample.[33] This shows that the policy horizon became substantially shorter after 1985, interest

[31]The term unilateral emphasizes the lack of a legislative mandate to achieve a specific inflation target. See Svensson (1995) for a detailed account of these events.

[32]UK in October 1992; Canada, January 1991; Sweden, January 1993; New Zealand, end 1989.

[33]Sterling came under pressure in early 1985 which might explain why the one-step Chow test is significant. However, this does seem to be a single outlier.

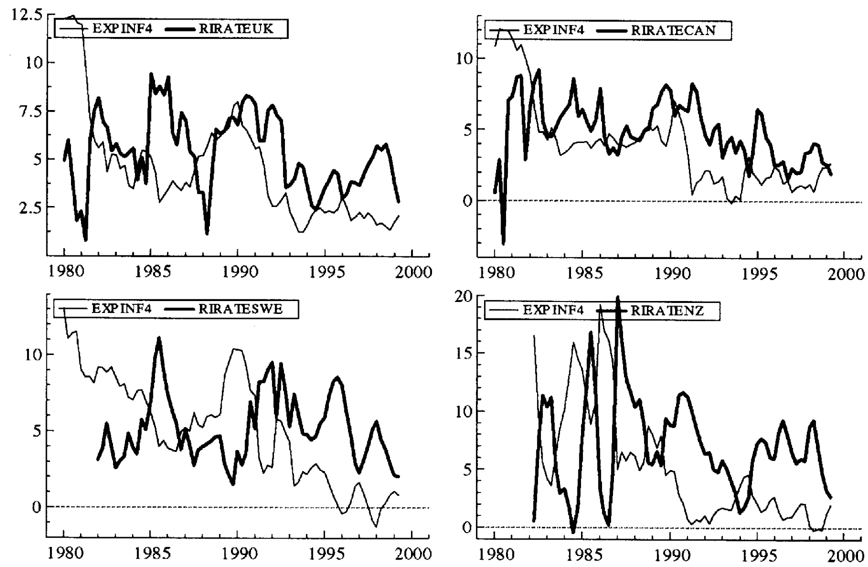

FIG. 5 UK, Canada, New Zealand, Sweden (clockwise): Real *ex ante* Interest Rates (bold line) and Four-quarter Expected Inflation

rates reacting to one-quarter-ahead expected inflation, and the coefficient on expected inflation becomes significantly larger than one. Within this, other minor shifts in policy regimes are also apparent (see Fig. 6), such as the 1985 sterling crisis and the exit from the ERM. The exit from the ERM (with the greater emphasis towards internal rather than external objectives) is apparent from the coefficient on expected inflation, which increases towards unity between 1992 and 1994, and the increased response to the output gap. Interestingly, at least with a sample up to 1999Q2, Bank of England independence in 1997 seems to have changed little. The watershed as far as UK monetary policy is concerned seems to have been 1992.

Our estimates for Canada over the full sample period (1975–99) yield somewhat puzzling results (see Table 6). When the US Federal Funds rate is added to the equation, the coefficients both on the output gap and on expected inflation are not significant. Clearly, as in the case of Germany and Japan, the Federal Funds rate absorbs part of the significance of the inflation variable. Even though M1 was the intermediate policy target in Canada between 1975 and 1982[34] (Freedman, 1995), we could not find a very significant role for the money supply in our estimated reaction function. Furthermore, there are clear signs of instability in the estimated

[34]In 1982 it was officially abandoned due to innovations in the financial sector.

Table 5
UK

Regressor/ Sample	*Constant*	*Expected inflation*	*Output gap*	*FIBOR*	*Variable addition tests*[d]		*Summary statistics*	
1975Q2–	4.479	0.810	0.991		Money	1.78	R^2	0.873
1999Q2[a]	(1.818)	(0.237)	(0.465)		growth	(0.17)	σ	1.149
							DW	1.93
					Exchange	3.434	AR 1–5 $F(5, 88)$	1.049 [0.3945]
					rate	(0.067)	ARCH 4 $F(4, 85)$	0.235 [0.9177]
							Normality $\chi^2(2)$	28.757 [0.0000]
							RESET $F(1, 92)$	0.140 [0.7092]
1975Q2–	1.555	0.662	0.953	0.660	Money	4.88	R^2	0.878
1999Q2[b]	(1.752)	(0.136)	(0.276)	(0.243)	growth	(0.03)	σ	1.132
(adding							DW	1.83
Germany's					Exchange	1.67	AR 1–5 $F(5, 87)$	0.923 [0.4703]
FIBOR)					rate	(0.2)	ARCH 4 $F(4, 84)$	0.260 [0.9028]
							Normality $\chi^2(2)$	22.947 [0.0000]
							RESET $F(1, 91)$	0.238 [0.6267]
1980Q1–	2.107	0.884	0.620	0.483	Money	0.668	R^2	0.916
1999Q2[b]	(1.643)	(0.208)	(0.257)	(0.276)	growth	(0.516)	σ	0.967
(adding							DW	1.89
Germany's					Exchange	0.003	AR 1–5 $F(5, 69)$	0.570 [0.7224]
FIBOR)					rate	(0.959)	ARCH 4 $F(4, 66)$	0.040 [0.9970]
							Normality $\chi^2(2)$	19.48 [0.0001]
							RESET $F(1, 73)$	2.253 [0.1376]
1985Q1–	2.875	1.402	0.571		Money	0.902	R^2	0.924
1999Q2[c]	(0.705)	(0.176)	(0.186)		growth	(0.412)	σ	0.898
							DW	1.40
					Exchange	4.29	AR 1–5 $F(5, 49)$	0.789 [0.5629]
					rate	(0.043)	ARCH 4 $F(4, 46)$	2.038 [0.1047]
							Normality $\chi^2(2)$	22.361 [0.0000]
							RESET $F(1, 53)$	0.767 [0.3850]

[a] Derived from an RLS regression of the interest rate on a constant, four-quarter ahead expected inflation, output gap and one lag of the dependent variable.
[b] Derived from an RLS regression of the interest rate on a constant, four-quarter ahead expected inflation, output gap and one lag of Germany's FIBOR (EURIBOR 1998Q4 onwards).
[c] Derived from an RLS regression of the interest rate on a constant, one-quarter ahead expected inflation, output gap and one lag of the dependent variable.
[d] We tested for the addition of other regressors. Zero restrictions on lagged money growth, changes in the current and lagged trade-weighted index of effective exchange rate were tested by an F version of the Wald test. p values in parentheses.

Asymptotic standard errors of estimated parameters are in parentheses. AR is a Lagrange multiplier test for the hypothesis of no serial correlation; ARCH checks whether residuals have an ARCH structure, with no ARCH as the null; Normality tests the normality of residuals; RESET tests the null of no functional misspecification. p values in square brackets.

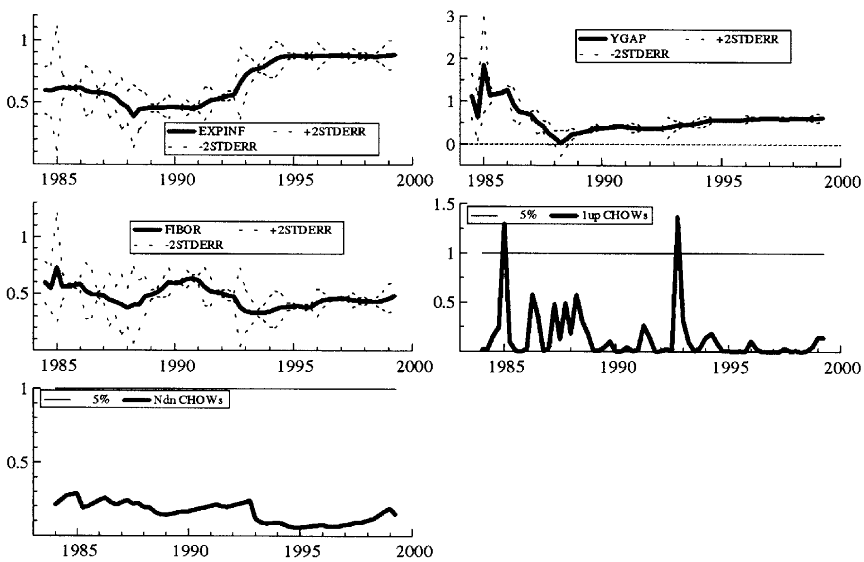

FIG. 6 UK, 1980Q1–1999Q2: Recursive Coefficients and Standard Error Bands; One-step, N-step Up Chow tests (5 per cent)

function in the late 1970s and early 1980s. Re-estimating the equation for the post-1980 sample, the coefficient on inflation expectations becomes significant but smaller than unity, whereas effective exchange rate variations now seem to be significant alongside the Federal Funds rate.

What about the impact of inflation targets? The introduction of targets does not seem to have caused a major change. At most there seems to have been a temporary impact on interest rate policy *just prior* to the introduction of inflation targets. Figure 7 shows a rise in the expected inflation coefficient around the period 1990–91, but the shift is small. Descriptive accounts of Canadian monetary policy in this period (Mishkin and Posen, 1997) point out that the inflation target was used as a guidance for expectations, but stress that on several occasions monetary policy was in fact constrained to react to external conditions, such as exchange rate developments and the behaviour of US monetary policy.

Our estimated reaction function seems to confirm this. Furthermore, the Bank did experiment for a short period with a short-run operational target, the index of monetary conditions (MCI). Although it has now been abandoned, MCI changes included variations in a short-term interest rate and in the trade-weighted exchange rate. Clearly, this highlights the importance of external constraints on the Bank of Canada's policy stance, and Canada does not seem to fit the model of an 'implicit inflation targeter'.

Table 6
Canada

Regressor/ Sample	*Constant*	*Expected inflation*	*Output gap*	*Federal Funds rate*	*Variable addition tests*[d]		*Summary statistics*	
1975Q3–1999Q2[a]	2.504	1.316	1.406		Money growth	0.674	R^2	0.804
	(2.013)	(0.3823)	(1.134)			(0.512)	σ	1.482
							DW	2.09
					Exchange rate	2.94	AR 1–5 $F(5, 106)$	5.253 [0.0002]
						(0.089)	ARCH 4 $F(4, 103)$	20.662 [0.0000]
							Normality $\chi^2(2)$	104.05 [0.0000]
							RESET $F(1, 110)$	2.764 [0.0993]
1975Q3–1999Q2[b]	0.817	0.3445	0.271	0.835	Money growth	2.268	R^2	0.827
	(0.946)	(0.1503)	(0.370)	(0.138)		(0.10)	σ	1.579
							DW	1.70
					Exchange rate	2.14	AR 1–5 $F(5, 85)$	4.939 [0.0005]
						(0.146)	ARCH 4 $F(4, 82)$	6.401 [0.0002]
							Normality $\chi^2(2)$	22.564 [0.0000]
							RESET $F(1, 89)$	2.174 [0.1439]
1980Q1–1999Q2[c]	0.809	0.603	Output gap not significant	0.695	Money growth	2.065	R^2	0.857
	(1.056)	(0.2358)		(0.197)		(0.134)	σ	1.569
			Exchange rate:				DW	1.60
			−0.536		Exchange rate	5.682	AR 1–5 $F(5, 67)$	7.689 [0.0000]
			(0.262)			(0.02)	ARCH 4 $F(4, 64)$	5.517 [0.0007]
							Normality $\chi^2(2)$	20.634 [0.0000]
							RESET $F(1, 71)$	2.646 [0.1083]

[a] Derived from an RLS regression of the interest rate on a constant, four-quarter ahead expected inflation, output gap and two lags of the dependent variable.
[b] Derived from an RLS regression of the interest rate on a constant, four-quarter ahead expected inflation, output gap, two lags of the dependent variable and one lag of the Federal Funds rate.
[c] Derived from an RLS regression of the interest rate on a constant, four-quarter ahead expected inflation, exchange rate, two lags of the dependent variable and one lag of the Federal Funds rate.
[d] Prior to estimating the third regression we tested for the addition of some regressors to the first regression. Zero restrictions on lagged money growth and the change in the trade-weighted index of effective exchange rate were tested by an F version of the Wald test. p values in parentheses.

Asymptotic standard errors of estimated parameters are in parentheses. AR is a Lagrange multiplier test for the hypothesis of no serial correlation; ARCH checks whether residuals have an ARCH structure, with no ARCH as the null; Normality tests the normality of residuals; RESET tests the null of no functional misspecification. p values in square brackets.

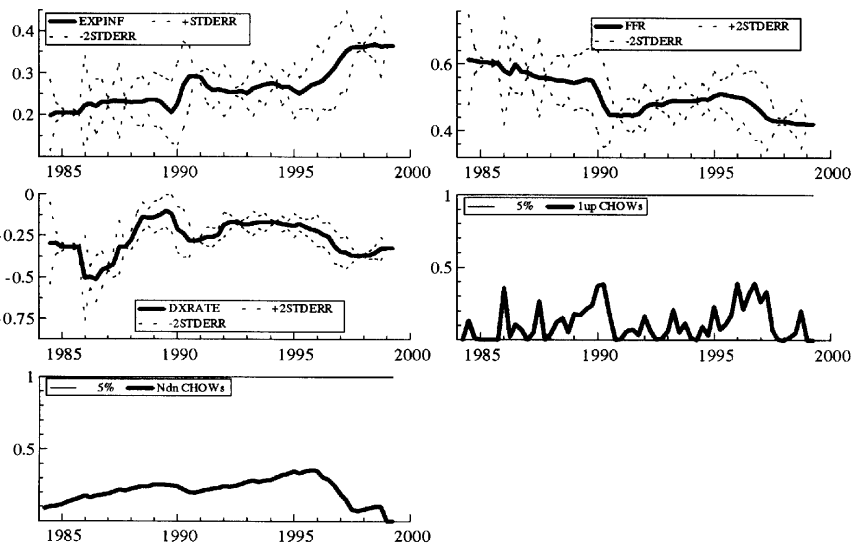

FIG. 7 Canada, 1980Q1–1999Q2: Recursive Coefficients and Standard Error Bands; One-step, *N*-step Up Chow tests (5 per cent)

The full-sample estimates (1982–99) for Sweden show a significant but relatively low coefficient on expected inflation, while the output gap is not significant at all (see Table 7). The main instability in the estimated reaction function corresponds to the time of the ERM crisis in 1992. Monetary policy in Sweden was externally tied to the ERM until 1992, when the krona was forced to devalue despite an unprecedented surge in domestic interest rates. Sweden has moved to inflation targeting since then. However, Svensson (1995) points out that the credibility of the new regime has been hampered by a number of factors, such as the deep political divisions over the conduct of monetary policy and the relatively large budget deficits.

Once a dummy is included for the ERM crisis in 1992, the coefficient on expected inflation rises and becomes more significant, but the point estimate remains below one, and the output gap variable remains insignificant at the 5 per cent level. The main story that emerges from Fig. 8 is the dominance of external factors (the German interest rate and the exchange rate) before 1992, and since 1992 a (slowly) growing importance of the domestic inflation variable. However, even by the end of our sample external variables seem to matter, and the coefficient on expected inflation remains below unity. The Swedish case is probably best described as 'credibility-building' since 1992, but without a major break due to the introduction of inflation targeting.

New Zealand has been the most often cited inflation-targeting

TABLE 7
SWEDEN

Regressor/ Sample	*Constant*	*Expected inflation*	*Output gap*	*FIBOR*	*Variable addition tests*[c]		*Summary statistics*	
1982Q2–	3.884	1.134	0.628		Money	1.162	R^2	0.902
1999Q2[a]	(1.514)	(0.2593)	(0.687)		growth	(0.319)	σ	1.066
							DW	1.90
					Exchange	3.57	AR 1–5 $F(5, 59)$	1.339 [0.2605]
					rate	(0.034)	ARCH 4 $F(4, 56)$	0.913 [0.4629]
							Normality $\chi^2(2)$	9.813 [0.0074]
							RESET $F(1, 63)$	4.156 [0.0456]
1982Q2–	2.02	0.773	Output gap not	0.669	Money	0.674	R^2	0.909
1999Q2[b]	(1.474)	(0.189)	significant	(0.268)	growth	(0.513)	σ	1.036
			Exchange rate:				DW	1.85
			0.495		Exchange	6.18	AR 1–5 $F(5, 59)$	1.391 [0.2408]
			(0.198)		rate	(0.015)	ARCH 4 $F(4, 56)$	3.040 [0.0244]
							Normality $\chi^2(2)$	2.415 [0.2989]
					FIBOR	4.092	RESET $F(1, 63)$	4.702 [0.0339]
						(0.021)		

[a] Derived from an RLS regression of the interest rate on a constant, four-quarter ahead expected inflation, output gap and one lag of the dependent variable.
[b] Derived from an RLS regression of the interest rate on a constant, four-quarter ahead expected inflation, the change in the trade-weighted index of effective exchange rate, the lagged three-month German FIBOR and one lag of the dependent variable.
[c] Before estimating the second regression we tested for the addition of some regressors to the first regression. Zero restrictions on lagged money growth, changes in the current trade-weighted index of effective exchange rate and the lagged three-month German FIBOR were tested by an *F* version of the Wald test. *p* values in parentheses.

Asymptotic standard errors of estimated parameters are in parentheses. AR is a Lagrange multiplier test for the hypothesis of no serial correlation; ARCH checks whether residuals have an ARCH structure, with no ARCH as the null; Normality tests the normality of residuals; RESET tests the null of no functional misspecification. *p* values in square brackets.

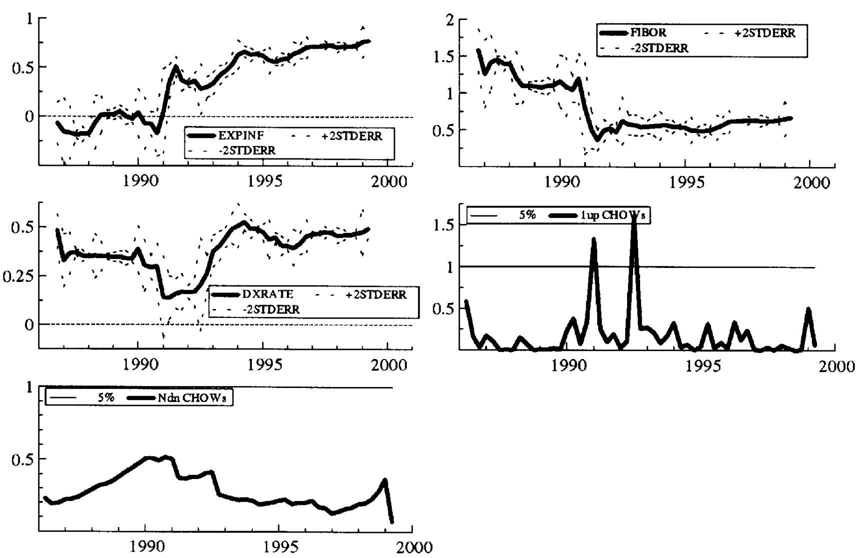

FIG. 8 Sweden, 1982Q2–1999Q2: Recursive Coefficients and Standard Error Bands; One-step, N-step Up Chow tests (5 per cent)

experiment, not least because in this case the legal arrangements designed to regulate the bank activity follow the prescriptions of monetary policy design theory more closely than elsewhere (see Walsh, 1995). The estimated equation for the full sample (see Table 8, Fig. 9) shows that interest rates seem to have reacted only to expected inflation from 1994 (the estimated coefficient is significantly larger than one) whereas the output gap does not seem to matter much.[35] The Bank contract cites the exchange rate as a possible justification for deviating from the announced policy, and interestingly we find a significant exchange rate effect over the 1983–99 sample. The other main point to note from Fig. 9 is the gradual rise in the inflation expectations coefficient after 1990 from a value below unity. The New Zealand case does seem to be one where the introduction of targets made a difference to interest rate policy, but the change was gradual and external objectives also remain important. Since 1995 policy seems to be essentially stable with little variation in the estimated coefficients. The other main point to note is that inflation targeting does not seem to have allowed the authority a greater leeway to stabilize output fluctuations, in contrast to the usual propositions in the theoretical literature.

[35]Hutchison and Walsh (1998) suggested that the Reserve Bank looked at output stabilization as an additional objective, but the output gap term is not significant in our estimates. Nevertheless, as pointed out previously, the absence of an output gap term in the reaction function does not preclude some degree of output stabilization.

TABLE 8
NEW ZEALAND

Regressor/ Sample	*Constant*	*Expected inflation*	*Output gap*	*Variable addition tests* [c]		*Summary statistics*	
1983Q2–	5.776	1.144	−0.392	Money	4.504	R^2	0.942
1999Q2 [a]	(0.724)	(0.103)	(0.466)	growth	(0.038)	σ	1.381
						DW	2.14
				Exchange	2.015	AR 1–5 $F(5, 56)$	0.741 [0.5964]
				rate	(0.161)	ARCH 4 $F(4, 53)$	3.233 [0.0192]
						Normality $\chi^2(2)$	4.257 [0.1190]
						RESET $F(1, 60)$	3.029 [0.0870]
1983Q2–	5.839	1.172	Output gap not	Money	2.554	R^2	0.950
1999Q2 [b]	(0.83)	(0.119)	significant	growth	(0.078)	σ	1.26
			Exchange rate:			DW	2.10
			0.239	Exchange	9.959	AR 1–5 $F(5, 55)$	1.586 [0.1794]
			(0.129)	rate	(0.002)	ARCH 4 $F(4, 52)$	5.289 [0.0012]
						Normality $\chi^2(2)$	15.315 [0.0005]
						RESET $F(1, 59)$	4.077 [0.0480]

[a] Derived from an RLS regression of the interest rate on a constant, one-quarter ahead expected inflation, output gap and one lag of the dependent variable.
[b] Derived from an RLS regression of the interest rate on a constant, two-quarter ahead expected inflation, output gap and two lags of the dependent variable.
[c] Before estimating the second regression we tested for the addition of other regressors. Zero restrictions on lagged money growth, lagged changes in the trade-weighted index of effective exchange rate were tested by an F version of the Wald test. p values in parentheses.

Asymptotic standard errors of estimated parameters are in parentheses. AR is a Lagrange multiplier test for the hypothesis of no serial correlation; ARCH checks whether residuals have an ARCH structure, with no ARCH as the null; Normality tests the normality of residuals; RESET tests the null of no functional misspecification. p values in square brackets.

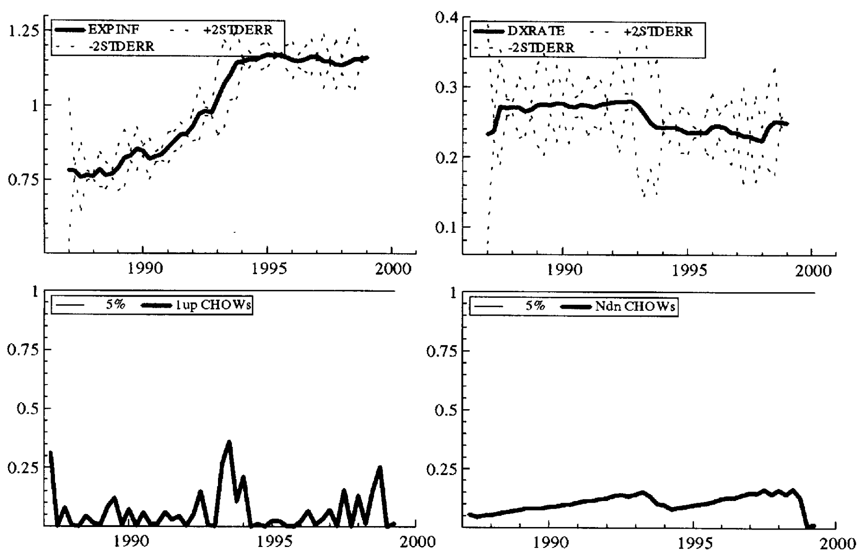

FIG. 9 New Zealand, 1982Q4–1999Q2: Recursive Coefficients and Standard Error Bands; One-step, *N*-step Up Chow tests (5 per cent)

5 Conclusions

In this paper we estimate forward-looking interest rate reaction functions for the G3 economies and for a group of countries which recently adopted explicit inflation targets and central bank reforms as the centrepiece of their monetary strategies. In addition to the detailed results for each country set out above, a number of general conclusions emerge from our empirical results.

First, with the exception of the UK and New Zealand, the adoption of inflation targets does not by itself seem to have caused a major shift in policy. In the case of New Zealand the shift was gradual post-1990. In Canada it seems that any major changes in the responsiveness of interest rates to expected inflation took place well before the adoption of inflation targets. In most cases (New Zealand, Sweden, Canada) there is little evidence that domestic interest rates react significantly to the output gap. Although this still implies some degree of output stabilization in response to aggregate demand shocks, it does reflect a lower priority on output stabilization following aggregate supply shocks. Only time will tell if, after a period of reputation-building, monetary policy will respond more vigorously to both inflationary shocks and output shocks. In the case of the UK the main policy shift seems to date from 1992, with little change since 1997 in the behaviour of the Bank of England.

Second, in countries where there were explicit intermediate targets (such as monetary aggregates in Germany) these were usually used as an anchor for expectations, but this did not seem to strictly constrain policy in practice. Monetary policy often followed a broader set of macroeconomic objectives. Our results confirm those of previous researchers who find that in practice the Bundesbank targeted inflation and output and reacted to external conditions.

Third, where the policymaker is subject to some implicit constraint due to external conditions (as in the case of Canada, Sweden, Japan and New Zealand) this can sometimes lead to a less clear picture regarding the monetary authorities' response to expected inflation and to the cycle. Interestingly even where the authorities abandon an exchange rate peg (Sweden), external considerations still tend to matter. More generally, the adoption of inflation targets and the move to greater central bank independence appear to have taken place without sacrificing these external objectives.

Fourth, even in G3 countries where there have been no central bank or other institutional reforms (e.g. the USA, Japan), we find that policies did evolve to a considerable degree in the 1980s and 1990s. Only since the 1990s do these countries begin to resemble the Bundesbank in terms of their estimated interest rate reaction functions. The variability of policy in the USA until 1995 seems to have been linked to uncertainty about the trend in potential output.

Fifth, with the exception of Germany and the UK (since 1992), most of the monetary authorities in our sample do not seem to follow stable simple forward-looking policy reaction functions based on output gaps and expected inflation (and, *a fortiori*, Taylor rules). This suggests that caution has to be exercised in using an inflation-targeting framework to model the preferences of the monetary authorities (see Clarida *et al.*, 1998; Favero and Rovelli, 1999).

Finally, we should focus on some important differences in the behaviour of central banks regarding output stabilization. On the one hand in the USA we seem to have the apparent 'just do it' attitude of the Fed, who since 1990 exploits her reputation to focus on the cycle, bolstered to some extent by a shorter horizon on expected inflation in the estimated reaction function. At the other extreme there are those monetary authorities who feel that they have yet to build up a reputation, e.g. the apparently exclusive focus of the Bank of New Zealand on domestic inflation and exchange rate considerations. Whether this 'reputation-building' phase will also apply to central banks that have only recently acquired their independence, such as the European Central Bank, remains an open question. Interestingly, in the case of the Bank of England, output stabilization does not appear to have been sacrificed with the adoption of the inflation target.

Data Appendix

The data we used were quarterly series up to 1999Q2, extracted from OECD *Main Economic Indicators*, apart from a few cases in which the source is equivalently quoted. In most cases we were able to employ seasonally adjusted data.

For each country we measured output using the GDP at constant prices series. For Sweden and New Zealand the available constant price series for GDP do not date back further than 1980 and 1982Q2, respectively. The inflation series were defined as simple four-quarter log-differences in the all-items consumer price index (CPI), except for Britain, where it was the equivalent change in the index of retail prices excluding mortgage interest payments (not available before 1975).

The index of effective exchange rates (trade weighted) was the measure for the exchange rates. Also, spot exchange rates *vis-à-vis* the US dollar were tried for Japan, Germany, Canada, New Zealand and the UK; *vis-à-vis* the German mark for the UK and Sweden.

The rate on US Federal Funds was used as the foreign interest rate for Japan, Germany, Canada and New Zealand. The three-month FIBOR German rate was the foreign rate for the UK and Sweden.

Below we briefly outline the short-term interest rates we chose as policy indicators, along with the monetary aggregates we applied in the generation of regressors. The rates are generally converted from monthly series.

Country	*Modelled interest rate variable*	*Money*
USA	*Federal Funds Rate*. As noted in the main text, during the early to mid-1980s the Federal Funds rate provides an accurate measure of the Fed's policy stance. The only exception is the Volcker experiment in the 1979–82 period, when the Fed's operating procedures could be better summarized by a different instrument choice (*inter alia*, Bernanke and Mihov, 1995; Goodfriend, 1995)	M1
Japan	The *Call Money Rate* (rate between financial institutions; *source:* Bank of Japan) is directly affected by the Bank of Japan reserve management policy, through discount window and open market operations	M2 plus CD
Germany	The Bundesbank's intentions are mainly reflected by the rate in the market for interbank reserves, the *Call Money Rate*. In fact, the discount window lending to commercial banks exclusively affected the behaviour of this rate until 1985, when the banks started to be supplied with reserves by repurchase operations. Since then the call rate shadows the rate on these loans (repo rate) (see Bernanke and Mihov, 1997; Clarida and Gertler, 1997)	M3[a]
UK	We use an *Overnight Interbank Rate* series post-1983. This is not available pre-1983, and we use the *Rate on 90-day Treasury Bills*, which displays a very close correlation with the interbank lending rate, for those observations (*source:* IMF, IFS)	M4

Country	*Modelled interest rate variable*	*Money*
Canada	The Bank of Canada introduced in 1996 the concept of monetary conditions index (MCI) as its short-run operational target. The changes in the index are defined as a weighted average of the changes in the 90-day commercial paper rate and the changes in a trade-weighted Can$ exchange rate. Although the MCI was computed backward and onward from 1987, the *Overnight Money Market Rate* (available from 1975) is clearly a superior indicator of the Bank's policy stance	M1, M2+[b]
Sweden	During the fixed exchange rate regime the overnight rate in the interbank market represented the Riksbank's favourite instrument to keep the desired krona's parity. Then, after the switch to the inflation targeting regime, the repo rate has become the Bank's operational instrument. For homogeneity and continuity we used the *Rate on 3-month Treasury Discount Notes* (not available before 1982), which roughly shadows the behaviour of both marginal and repo rates	M3
New Zealand	The *Rate on 90-day Bank Bills* (not available before 1974) was our choice. Until March 1985 New Zealand has pursued a policy of adjustable pegged exchange rate. 'The instrument since 1985 has been the quantity target for settlement balances held at the Reserve Bank. Settlement cash is used by commercial banks for end-of-day settlements with each other and the government. Should the banks run out of cash during the settlement period, further cash is available from the Reserve Bank by discounting Reserve Bank bills of short maturity at a penalty rate of 1.5 per cent above market rates.... Such an approach allows interest rates to move quickly, particularly when the change involves a politically unpopular increase in interest rates' (Leiderman and Svensson, 1995, p. 35). It is then understandable why banks prefer to act in the bank bills market, whose short-term interest rate tends to react rapidly to changes in policy intentions.	M1

[a] The Bundesbank announced targets for the growth of Central Bank Money until 1987, when it switched to M3, which we chose. The two move very closely together, apart from two episodes of divergence in 1988 and 1990–91. Notwithstanding that the official target is announced in terms of base-money growth, the evidence points to Germany as an 'atypical' inflation targeter, who influences the money markets through changes in a day-to-day rate (Von Hagen, 1995; Bernanke and Mihov, 1997; Mishkin and Posen, 1997).

[b] Until 1982 the Bank of Canada was committed to targeting M1. It now also follows closely the behaviour of M2+ (Freedman, 1995).

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