Department of Economics
Working Paper Series

Measuring Monetary Policy Interdependence

Oscar Jorda
University of California, Davis

Paul Bergin
University of California, Davis

January 01, 2001

Paper # 06-9

This paper measures the degree of monetary policy interdependence between major industrialized countries from a new perspective. The analysis uses a special data set on central bank issued policy rate targets for 14 OECD countries. Methodologically, our approach is novel in that we separately examine monetary interdependence due to (1) the coincidence in time of when policy actions are executed from (2) the nature and magnitude of the policy adjustments made. The first of these elements requires that the timing of events be modeled with a dynamic discrete duration design. The discrete nature of the policy rate adjustment process that characterizes the second element is captured with an ordered response model. The results indicate there is significant policy interdependence among these 14 countries during the 1980-1998 sample period. This is especially true for a number of European countries which appeared to respond to German policy during our sample period. A number of other countries appeared to respond to U.S. policy, though this number is smaller than that suggested in preceding studies. Moreover, the policy harmonization we find appears to work through channels other than formal coordination agreements.

UCDAVIS

Department of Economics
One Shields Avenue
Davis, CA 95616
(530)752-0741

http://www.econ.ucdavis.edu/working_search.cfm

March 2002

Measuring Monetary Policy Interdependence*

Abstract

This paper measures the degree of monetary policy interdependence between major industrialized countries from a new perspective. The analysis uses a special data set on central bank issued policy rate targets for 14 OECD countries. Methodologically, our approach is novel in that we separately examine monetary interdependence due to (1) the coincidence in time of when policy actions are executed from (2) the nature and magnitude of the policy adjustments made. The first of these elements requires that the timing of events be modeled with a dynamic discrete duration design. The discrete nature of the policy rate adjustment process that characterizes the second element is captured with an ordered response model. The results indicate there is significant policy interdependence among these 14 countries during the 1980-1998 sample period. This is especially true for a number of European countries which appeared to respond to German policy during our sample period. A number of other countries appeared to respond to U.S. policy, though this number is smaller than that suggested in preceding studies. Moreover, the policy harmonization we find appears to work through channels other than formal coordination agreements.

Paul R. Bergin

and

Òscar Jordà

Dept. of Economics
U.C. Davis
One Shields Ave
Davis, CA 95616-8578
e-mail:
prbergin@ucdavis.edu
ojorda@ucdavis.edu

*The authors wish to thank Rob Feenstra, Greg Sutton, Alan Taylor and participants of the Macro/International Economics workshop at U.C. Davis for useful comments and suggestions. Selva Demiralp and Kristin Van Gaasbeck provided extraordinary research assistance. This research was funded by a grant from the Institute of Governmental Affairs at U.C. Davis.

1 Introduction

This paper investigates the empirical regularities of monetary policy setting among a sample of 14 industrialized countries, with special focus on the international interdependence among these policies. A new empirical approach is proposed, which avoids the pitfalls implied by the VAR approach that traditionally has been used to address this issue. This approach allows us to look in a new way at the nature of policy coordination and the relative leadership roles played by the U.S. Federal Reserve, the Bank of Japan, and the German Bundesbank.

International monetary policy coordination has become a subject of renewed interest of late. In particular, Obstfeld and Rogoff (2001) initiated a re-exploration of international policy coordination using models with microeconomic foundations, representing a significant methodological departure from past literature. They conclude that in this context the welfare gains of coordination are likely to be quantitatively small in comparison to the gains from domestic stabilization policy. In rebuttal, subsequent theoretical work has suggested a variety of economic features that could potentially generate greater motivation for nations to coordinate their monetary policies. See for example, Benigno and Benigno (2001a,b), Canzoneri, Cumbi and Diba (2001), and Clarida, Gali and Gertler (2001). Given this theoretical controversy, it is natural to ask the empirical question of how much coordination we observe in actuality.

There is a history of empirical research on this question. Studies focusing on the major industrial countries generally find evidence that the U.S. acts as a leader for policy makers in certain countries, but the mechanisms through which this coordination takes place are often unclear. (See Dominguez (1997), Furman and Leahy (1996), Chung (1993), Burdekin and

Burkett (1992), Burdekin (1989), and Batten and Ott (1985).) Another branch has focused on coordination among European countries, generally finding that Germany had a limited leadership role among European countries prior to monetary union. (See Garcia-Herrero and Thorton, 1996; Katsimbris and Miller, 1993; Biltoft and Boersch, 1992; Karfakis and Moschos, 1990; and von Hagen and Fratianni, 1990a,b).

Our paper represents a significant methodological departure from this preceding empirical literature. We build upon the recently developed methodology of Hamilton and Jordá (2002), which has been successfully employed to study monetary policy in the U.S. We extend this methodology to a set of countries, and we explicitly allow for interdependence among these national policies. Our approach is based on a novel data set on overnight, interbank interest rate targets that central banks in our sample use to communicate and operationalize monetary policy.

Methodologically, we first argue against traditional time series techniques based on dynamical conditional correlations in semi-structural vector processes, which suffer from the common identification assumptions that mar the monetary vector autoregression (VAR) literature. Perhaps more importantly however, we show via Monte Carlo experimentation that the peculiar statistical properties that policy rate targeting imbue on market interest rates tend to severely distort these measures of association. In particular, these experiments demonstrate that, for example, Granger causality tests often will misrepresent the true nature of existing monetary policy linkages that are measured with market interest rate data.

These methodological pitfalls prompt us to pursue a more modest approach but which is immune to the deficiencies described above. In particular, our goal is to determine whether

G-3 policy moves help predict the timing of domestic policy adjustments and the direction in which interest rates are modified, all conditional on domestic macroeconomic conditions. Exploring these issues poses special econometric challenges that arise from the irregular nature in which policy rate targets are adjusted over time as well as the discrete nature in which these adjustments are made. Hence, we will use the autoregressive conditional hazard (ACH) model proposed in Hamilton and Jordá (2002) to tackle the predictability in the timing whereas a discrete ordered response model specification will permit us to deal with the discreteness of policy rate adjustments.

The empirical results are elucidating along several dimensions. First, our study can be seen as a broad investigation of individual central bank behavior, and in this sense, it provides important lessons to the literature that explores monetary policy rules. The special nature of the data and the particulars of our empirical design suggest that Taylor type rules do not universally describe monetary policy well, once the natural inertia of interest rates and macroeconomic data has been filtered. In particular, while U.S. policy responds to inflation in a statically significant manner, we were unable to find similar evidence among any of the other countries, including Germany, which traditionally enjoyed a reputation for an aggressive stance against inflation. Further, while it is a current debate among monetary theorists whether it is optimal for policy rules to include the exchange rate, we find evidence that this variable has in practice played a very prominent role in the policy rules of many countries.

With regard to monetary policy interdependence, Germany emerges as a significant factor for a block of European countries, in agreement with most past work. This indicates that policy interdependence existed among this European core long before the creation of the explicit monetary union. Further this European interdependence entailed more than simply responding to exchange rate movements, as required by the European Exchange Rate Mechanism (ERM), but appears to have involved direct responses by policy makers to changes in German policy.

We also find evidence that a limited number of countries respond to U.S. policy changes. This list includes Germany and the UK, but in contrast to some past work, Japanese policy does not appear to respond in a significant way to the U.S.

2 Pitfalls of Interest-Rate Dynamic Correlations

It would be natural to explore the degree of monetary policy interdependence with vector time series techniques using interest rate data on the countries in our sample, as is done in Nogues and Grandes (2001), Bensidoun, Coudert, and Nayman (1997), and DeGennaro and Kunkel (1994), for example. However, a large number of central banks in our sample operationalize monetary policy by targeting an overnight, interbank, interest rate. A consequence of this modus operandus is that interest rates in these countries have unusual statistical properties which make them unsuitable for conventional econometric analysis. This section briefly explains the nature of these problems with Monte Carlo experimentation and serves to motivate the empirical design in the remainder of the paper.

Rudebusch (1995) demonstrated in a clever simulation that the manner the U.S. Federal Reserve adjusts the federal funds rate target affects the statistical properties of traditional term structure regressions in a manner consistent with the poor results reported in that literature. We wish to extend Rudebusch's (1995) arguments to show that central bank interest-rate targeting behavior distorts conventional measures of dynamic covariation between interest rates of different countries.

Let r_t^A and r_t^B denote the overnight rate in countries A and B, respectively. Although central banks choose a target level for the overnight rate, they have only imperfect control over it. Therefore, the overnight rate tends to fluctuate around the target level. In the U.S., these deviations typically amount to only a few basis points and tend to be short lived (see Hamilton (1996)). Accordingly, we model the overnight rate as,

$$r_t^i = \overline{r}_t^i + \varepsilon_{it} \qquad \varepsilon_{it} \sim N(0, \sigma_i^2) \qquad i = A, B \tag{1}$$

where $\overline{r}_t^i$ denotes the target overnight rate. In practice, the residuals, ε_i, may be a general stationary sequence but to keep the exercise simple, we will simulate them as *i.i.d.* normal random variates. The variance of these residuals, σ_i^2, is a natural measure of the central bank's control over the interbank market. The process by which the target is adjusted is described by,

$$\overline{r}_t^i = x_t^i \overline{r}_{t-1}^i + (1 - x_t^i)(\overline{r}_{t-1}^i + y_t^i) \tag{2}$$

where $x_t^i = 1$ if the target is changed in period t, 0 otherwise, and $y_t^i \in \{k_1, k_2, ..., k_m\}$, reflecting the fact that changes in the target are done in a small finite set of discrete increments. Each period, there is some probability that the target will be changed, $h_t^i = P(x_t^i = 1|\Omega_{t-1})$, which can be interpreted as a discrete time hazard. As usual, Ω_{t-1} represents the information set up to time $t - 1$. Similarly, conditional on changing the target, the magnitude of the change, k_j, can be described by,

$$P(y_t^i = k_j | x_t^i = 1, \Omega_{t-1}) = P(c_{j-1} < y_t^{i*} < c_j | x_t^i = 1, \Omega_{t-1}) \qquad j = 1, 2, ..., m \tag{3}$$

with $c_0 = -\infty$, $c_m = \infty$, and y_t^{i*} an auxiliary latent index that reflects the central bank's ideal for the overnight rate but which may be costly to attain in every single period. To complete the simulation, we specify a simple bivariate model for h_t^i that will capture the coordination in the timing between countries A and B, and a bivariate model for Δr_t^{i*} that will capture coordination in the direction of the change instead. First, note that h_t^i is a probability, hence a convenient way of ensuring $h_t^i \in [0, 1]$ is to take the following logistic transform

$$h_t^i = \frac{1}{1 + e^{\lambda_t^i}} \tag{4}$$

$$\begin{cases} \lambda_t^A = & \omega_A + & \rho_A \lambda_{t-1}^A + & \beta \lambda_{t-1}^B + & e_t^A \\ \lambda_t^B = & \omega_B + & \rho_B \lambda_{t-1}^B + & & e_t^B \end{cases} \qquad e_t^i \sim N(0, 0.1)$$

so that $x_t^i = 1$ if $h_t^i > u_t^i$, 0 otherwise, where u_t^i is a uniform [0,1] random variable. The bivariate specification in (4) makes explicit two important features: (1) persistence in target changes, which as we shall see in later sections, is a common feature in our data set, and (2) policy setting in country A depends, via the parameter β, on country B's policy setting. If $\beta > 0$, the likelihood of a target change in A increases with the likelihood of a target change in country B, and vice versa. If $\beta = 0$, target setting occurs independently.

Similarly, to capture any coordination in the direction of changes in the target that may occur independently from coordinating the timing, consider the following bivariate model for the latent indices,

$$\begin{cases} y_t^{A*} = & \gamma_A y_{t-1}^{A*} + \quad \psi y_{t-1}^{B*} \quad + \quad v_t^A \\ y_t^{B*} = & \gamma_B y_{t-1}^{B*} + \qquad\qquad\quad + \quad v_t^B \end{cases} \qquad v_t^i \sim N(0,1) \tag{5}$$

The parameters γ_i capture persistence in the direction of target changes, which is also a commonly observed feature of the data, and the parameter ψ makes explicit the correlation in the direction in which the targets in countries A and B are changed. If $\psi > 0$, countries A and B tend to change their targets in the same direction and vice versa, if $\psi = 0$, there is no relation.

The basic setup described in equations (1)-(5) establishes a well-defined hierarchy: while decisions on country B's target are based completely on domestic information, country A's target timing and magnitude of changes depend on previous actions by country B whenever β and ψ are non-zero. In other words, *B Granger-causes A*. Table 1 contains the results of Monte-Carlo experiments in which simulated overnight rates are generated for different parameter values, using equations (1)-(5). For each combination of parameters, we generate 100 series of length 400, where the first 100 observations are deleted to avoid initialization problems. Then, on each of the series, we perform a standard Granger-causality test, collect the p-value of all these tests and report the Monte-Carlo average. Figure 1 displays an example of what these simulated interest rate series look like to show they closely resemble the data in our study.

Several results deserve comment. The most immediately apparent is the extraordinary sensitivity of the Granger-causality test. Even in the quasi-ideal scenario in which there is substantial coordination in the timing ($\beta = 0.75$) as well as coordination in the direction ($\psi = 0.75$), Granger-causality tests will routinely fail to pick up these features whenever the frequency with which the target is changed is relatively low (below 30% of the time) and/or the central bank's control of the overnight rate around the target is poor ($\sigma_\varepsilon > 0.25$). Other features of the model appear to be less important in determining this sensitivity except perhaps the degree of persistence in the direction of the target changes (γ_A and γ_B small). It is important to note that this failure of the Granger-causality test can occur under the most favorable of situations, with high values of the parameters β, and ψ, with a relatively precisely controlled overnight rate and with realistic parameters values. The Granger-causality test metric appears to be well suited to capture directional coordination since it will detect this type of coordination in the absence of coordination in the timing as long as the target is changed sufficiently often (40% of the time and above). However, even when the timing of target changes is strongly related ($\beta = 0.99$), as the value of ψ decreases, even slightly (from 0.75 to 0.5 and below), the Granger-causality test will fail to pick the relationship.

These results are strongly suggestive that conventional econometric techniques are unsuitable for the purposes of our analysis. Consequently, the remainder of the paper devises an alternative empirical strategy which is based on the availability of an uncommon data set, whose properties we discuss in the next section.

3 A Novel Data Set on Policy Rate Targets

A significant aspect of our investigation on monetary policy linkages requires that we identify the policy stance of each country in our sample. As we discussed in the previous section, it is often common to do this by isolating the exogenous component of a monetary policy indicator based on a semi-structural VAR. However, in addition to the pitfalls discussed in the previous section, these measures of the policy stance are problematic since they depend on a number of untestable identifying assumptions and specification choices that can produce dramatically different outcomes (for a discussion on these issues see Rudebusch, 1998; Sims, 1998; and Evans and Kuttner, 1998).

We prefer to avoid this controversy altogether by pursuing a radically different approach. In an exhaustive survey from the Bank of International Settlements (BIS), entitled "Monetary Policy Operating Procedures in Industrial Countries," Claudio Borio[1] (1997) provides an in-depth discussion on central bank operations for 14 major industrialized OECD countries, namely: Austria (AT), Australia (AU), Belgium (BE), Canada (CA), Switzerland (CH), Germany (DE), Spain (ES), France (FR), Italy (IT), Japan (JP), the Netherlands (NL), Sweden (SE), the United Kingdom (UK), and the United States (US).

Table 2 reports the specific variables and samples available for each country while Figure 2 displays the policy target data along with the corresponding overnight interest rates. The countries in our sample fall into two broad groups: (1) Australia, Canada, Japan and the U.S. where the most representative policy variable is the overnight interbank rate; and (2) Austria, Belgium, Switzerland, Germany, Spain, France, Italy, the Netherlands, Sweden, and the U.K. in which the policy variable is usually a tender rate applicable mainly to repurchase agreements (repos) and whose maturity varies from one day to one month. Along with these data, we collected domestic macroeconomic data on inflation, output and exchange rates

[1] We thank Greg Sutton and the Statistical Assistance Section of the Bank of International Settlements for providing us the data.

from the International Finance Statistics database produced by the I.M.F. The specifics for these data are also reported in Table 2.

It is instructive at this point to take a first look at our data. In particular, we ask whether the probability that domestic target rates are changed, depends on whether similar changes transpire in any of the G-3, regardless of the direction of the change. Thus, let $x_t = 1$ if in period t, the target was changed in a given country, 0 otherwise. We denote with $\mathbf{w}_t = (x_t^{DE}, x_t^{JP}, x_t^{US})'$ the vector of explanatory variables that consists on the similarly defined x variables for each of the G-3 countries. The per-period probability conditional on the actions of the G-3 can be therefore be expressed as a discrete time hazard, $h_t = P(x_t = 1|\mathbf{w}_t)$. A simple way to compute this hazard is with the logistic transform,

$$\begin{aligned} h_t &= \frac{1}{1+e^{\lambda_t}} \\ \\ \lambda_t &= \delta + \boldsymbol{\gamma}'\mathbf{w}_t \end{aligned} \tag{6}$$

with obvious log-likelihood

$$L(\theta) = \sum_{t=1}^{T} x_t^i \ln(h_t^i) + (1 - x_t^i)\ln(1 - h_t^i) \tag{7}$$

Table 3 reports for each country the baseline hazard and the hazard whenever there is a policy action in each of the corresponding G-3 targets. The data is organized at a monthly frequency with the double purpose of matching, the timing and length of the maintenance periods common in the countries that we consider, as well as the frequency at which our explanatory variables are available in the analysis reported below. The latter allows direct comparability with the results in Table 5.

A cursory look at Table 3 reveals perhaps familiar results. For example, Austria changes its target, on average, once every 10 months. However, if there is a change in the German Lombard rate, the probability of a target change increases to a 1 in 2 chance for that month. The effect is more dramatic in Belgium (the baseline hazard is 17% but when Germany adjusts, the hazard increases to 70%), France and the Netherlands. Notice that although the effect of German policy on the U.S. is statistically significant, it only represents a 1% increase in the probability that the U.S. will adjust the federal funds rate target that month. By contrast, although the effect in Spain is not significant, the hazard increases by more than 10% from a baseline level of 38%. With respect to Japan, it appears it is only influential with respect to German policy (the baseline is 48% but increases to 69% with a Japanese move). The effects of U.S. policy appear to affect Austrian, Australian, Japanese and Dutch policy but in every case the effects are not particularly strong.

Several results deserve comment. First is the fact that these are estimates of probabilities and therefore, highly nonlinear. Consequently, we have situations in which relations are statistically significant but economically weak (for example, the effect of U.S. policy in Japan causes the baseline to decline from 6% to 2%, hardly an economically meaningful effect), and vice versa, effects that are economically noticeable but statistically weak (a move in Japanese policy causes the Australian baseline to almost double from 18% to 32%). Two more caveats to interpreting these results include (1) the effects that obviously omitted variables may have – domestic macroeconomic variables may explain away some of these relations or they make others become significant; and (2) the effect that omitted dynamics may have on the standard errors. Both of these issues are explored below.

Summarizing, it is not surprising that Germany's influence is felt in a list of countries that has often been characterized as the "European core." Italy's absence from this list may reflect the fact that our data sample for this country (September 1992 - November 1996) is dominated by Italy's absence from the EMS following the 1992 crisis. The U.S. influence does not appear very strong, which is at this stage, a surprising result. Other specific countries worthy of mention include Switzerland, whose measure of monetary policy is the flexible Lombard rate whereas in practice, the primary focus of the central bank is the volume of giro deposits. Borio (1997) thus suggests that "[...] interest rates are of limited significance in conveying policy intentions." (p. 17). Similarly, Japan is the only country that is still using a quantitative signal as a key mechanism for steering an interest rate operating target. Sweden uses a combination of a target for interest rate tenders with variable interest rate auctions when markets fluctuate around desirable levels (this mechanism is similar to that used by the Bundesbank).

Although these results are suggestive, it is difficult to discern at this point what motivates these interconnections. Our approach in the following sections will be to try to describe each central bank's policy rule and choice of response timing. If, after controlling for responses to a range of domestic macroeconomic factors, we are unable to control for the timing component, we will conclude we have uncovered evidence of direct policy linkages. The next sections discuss the econometric aspects of this strategy.

4 Measuring Monetary Policy Interdependence

The Monte Carlo experiments in Section 2 and the preliminary evidence in Section 3 already provide a rough outline of the empirical strategy of the remainder of the paper. In particular, we will measure monetary policy interdependence along two dimensions. On one hand, we will investigate to what extent the timing of policy rate adjustments is affected by actions of the G-3. On the other hand, we will examine whether domestic policy rate adjustments are made in the same direction as adjustments in the G-3, all conditional on domestic macroeconomic conditions. We begin by presenting the model specification of the timing model.

4.1 The Timing of Policy Rate Adjustments

Predicting when a central bank will adjust its policy rate next can be conceived as a discrete-time duration process, as we saw in the previous section in expression (6). However, unlike that specification, we wish to incorporate two new elements in the analysis: (1) adding explanatory variables on domestic macroeconomic conditions, and (2) allowing for a general dynamic specification. The first element is designed to rule out interdependence that is mediated by synchronized business cycles and exchange rate considerations while the second element is designed to account for serial dependence in target setting and to provide more accurate standard errors.

There is substantial agreement in the literature that the manner central banks determine the optimal level of their interest rate objectives can be well modelled by a Taylor rule (see Taylor, 1999 for an extensive survey). This monetary policy rule establishes the dependence of the policy instrument's optimal level on deviations of inflation from a target (usually 2% annual rate) and some measure of the output gap. In addition, small open economies also include a term related to the exchange rate variability. Therefore, it seems natural to assume that these variables would also determine the likelihood with which a central bank will decide to adjust its policy rate in time.

Based on this disquisition, we construct the following variables for each of the countries in our sample: the inflation variable π_t measured as the deviation of the annualized percentage change in the consumer price index from a 2% norm; the output variable g_t measured as the deviation of the annualized percentage change in the industrial production index from a 2.5% norm; and the exchange rate variable fx_t measured as the monthly percentage change in the nominal exchange rate weighted by the composition of each country's trading partners. All of these variables are added to the vector of explanatory variables $\mathbf{w}_t$ defined in Section 3 so that $\mathbf{w}_t = (\pi_{t-1}, g_{t-1}, fx_{t-1}, x_t^{DE}, x_t^{JP}, x_t^{US})'$.

Several features of these variables deserve comment. First, to maintain the analysis at a monthly frequency rather than quarterly (for which we would not have had sufficient degrees of freedom), we had to compromise on the measures on the inflation and output variables commonly used elsewhere in the literature (typically consisting on the GDP deflator and the gap between GDP and some measure of potential GDP). Second, we preferred to avoid the well-known and controversial issue of measuring the output gap by using instead deviations from a 2.5% annual growth norm. Thirdly, these domestic macroeconomic variables are lagged to reflect the information available to the central bank at the time of decision. The final tally of the specific variables and sources can be found in Table 2.

The final element of the model's specification consists of endowing the process with dynamics that account for possible serial dependence in the timing of domestic policy rate adjustments. Hamilton and Jordá (2002) introduced the autoregressive conditional hazard (ACH) model for this precise purpose in the context of U.S. monetary policy setting. Thus, collecting all these elements together, the conditional probability that a country will adjust its policy rate at time t, presented in expression (6), can be expanded into a new version of the ACH model, namely

$$h_t = \frac{1}{1 + e^{\lambda_t}} \tag{8}$$

$$\lambda_t = \delta + \alpha u_{s(t-1)} + \beta \lambda_{t-1} + \boldsymbol{\gamma}' |\mathbf{w}_t|$$

and where $u_{s(t-1)}$ denotes the amount of time elapsed between the two most recent policy rate changes as of date $t-1$, and where λ_t now becomes a dynamic latent index whose structure is similar to that in an ARCH model. Notice that the explanatory variables $\mathbf{w}_t$ enter in absolute value since we are interested in the triggering mechanisms that prompt the central bank to adjust the policy rate and hence, this is likely to occur whenever these variables depart from above or from below their targeted values. We will call the version in expression (8) ACH(1,1) to refer to the fact that one lag of each the duration and the latent index enter the specification. The expression for this ACH(1,1) can be easily generalized to include longer lag lengths p and q and has the advantage of leaving the parameter space unrestricted.

The sum of the coefficients α and β provide a summary measure of the degree of serial dependence in addition to allowing the effect of the explanatory variables to be completely dynamic but parsimonious. To see this, notice that the difference equation of the latent index λ_t in expression (8) implies that the effect of the terms in the vector $\mathbf{w}_t$ is exponentially decaying over time at a rate β^t. The likelihood for the ACH model is the same as that in expression (7), which can be maximized by conventional numerical techniques.

4.2 The Direction of Policy Adjustments

Conditional on the decision to adjust the policy rate, the central bank then needs to decide on the nature and magnitude of this adjustment. This is typically done in discrete increments (of 25 basis points in the U.S., for example) rather than continuously and therefore, it is natural to model this feature with a conventional ordered response model. Conditioning on $x_t = 1$ implies that the specification of the model is in *event time* rather than in the usual calendar time scale. This attribute has the considerable advantage of distilling demand fluctuations and other extraneous sources of variation from interest rate adjustments and into the actual nature of the intended policy moves.

Denote the magnitude of the policy rate adjustment when it happens by y_s. Since policy rates are adjusted in discrete increments, we think of y_s as taking a finite collection of values, $y_s \in \{k_1, ..., k_m\}$ and we make note that the index s refers to those periods in which $x_t = 1$ only. Therefore, the probabilistic model that determines the observation of the variable y_s can be written as

$$P(y_s = k_j | \Omega_{s(t-1)}) = P(c_{j-1} < y_s^* < c_j | \Omega_{s(t-1)}) \tag{9}$$

$$y_s^* = \theta' \mathbf{w}_s + \varepsilon_s \qquad \varepsilon_s \sim N(0,1)$$

where y_s^* is a latent index, the thresholds c_j, $j = 0, 1, ..., m$ are such that $c_0 = -\infty, c_{j-1} < c_j$, and $c_m = \infty$. $\mathbf{w}_s$ is the vector of explanatory variables defined in the previous section but where we replace the $x_t^i, i = DE, JP, US$ with $y_t^i, i = DE, JP, US$ so that $\mathbf{w}_s = (\pi_s, g_s, fx_s, y_s^{DE}, y_s^{JP}, y_s^{US})'$. $\Omega_{s(t-1)}$ denotes the information set available just before the s^{th} event.

The parameter estimates of θ' are identified up to scale only, however, they are easily interpretable. For example, a positive value of $\widehat{\theta}_{DE}$ would suggest that domestic rates are more likely to be increased following a German increase and vice versa. The model in expression (9) has a well known likelihood function (see Maddala, 1983) which can be maximized by conventional numerical techniques. Finally, we restrict the number of categories, k_j, to four to homogenize across the countries in our sample and to allow for better comparability. Hence, the four categories can be loosely interpreted as: k_1 = strong decrease, k_2 = decrease, k_3 = increase, and k_4 = strong increase.[2] Table 4 reports for each country the correspondence between this classification system and the specific discrete changes of each country. As an illustration, changes in the U.S. federal funds rate target correspond to $k_1 = -0.50\%, k_2 = -0.25\%, k_3 = 0.25\%$, and $k_4 = 0.50\%$.

5 Results

5.1 The Timing Model

Table 5 displays the results of the ACH estimation of the timing model. For each country, the label *Constant* refers to the baseline hazard assuming that inflation and output are at their targeted levels (2% and 2.5%, respectively), that the exchange rate remained constant during the previous month, and that none of the G-3 central banks modified their policy rates during the month. The label *Inflation* refers to the hazard prevailing when inflation deviates by 1% from its targeted value. A similar experiment describes the entries associated with the labels *Growth* and *Ex. Rate*. The labels *Germany, Japan, U.S.* correspond to

[2] The only exception to this reclassification is Switzerland. Due to the large number of observed values, it made more sense to proceed with a continuous model instead.

the estimated hazards when these countries modify their policy rates that month. Finally, because the parameters α, and β measure the degree of serial dependence in the timing data, we preferred to report the parameter estimates and their standard errors directly.

A first observation is that the standard Taylor rule variables, inflation and output growth, do not apply equally well to all countries. While inflation is an important variable for U.S. policy, it does not appear as significant for any of the other countries. This is perhaps most surprising for Germany, given its traditional reputation for an aggressive stance toward inflation. Output growth is important for five of the countries, including Germany. Notably, the exchange rate also is important for five countries. So while monetary theorists have cast doubt on whether it is optimal to include the exchange rate in a country's monetary policy rule, our evidence suggests that quite a number of countries have done this in practice.

The next observation is that as in Table 3 previously, Table 5 shows a clear group of countries in the European core which respond to German policy. Note that this is in addition to the fact that these countries also respond in a significant manner to exchange rate movements. This appears to indicate a long-standing degree of policy dependence in Europe, one which is distinct from the demands of the ERM to respond to deviations in the exchange rate. In contrast, for countries usually thought to be in the European periphery, such as Spain and Italy, we do not discern a significant response to German policy. Also, Sweden and Switzerland, true to their reputations for independence, also appear to be independent of foreign influence in their monetary policies. The United Kingdom is an interesting case, in that the coefficient on German policy is statistically significant, but it is smaller than the baseline value. This indicates that once we control for British responses aimed at stabilizing its exchange rate, British policy makers appear to be actually less likely to change domestic monetary policy when the Germans have recently taken action.

Consider now the role of the U.S. While the U.S. does not respond to Germany, German policy does respond to the U.S. One surprise, compared to previous literature, is that there is not clear evidence that Japan responds to U.S. policy. The United Kingdom does respond to the U.S., which is interesting, given that we found above that the U.K. does not respond in a favorable way to German policy changes. It is worth noting that each of the relationships with the U.S. listed above, inferred from the results in Table 5, is the opposite of that which one would infer from Table 3. Apparently it is very important for judging the relationship with the U.S. to account for the dynamics and Taylor rule variables, as is done for the results reported in Table 5. This was not the case for the relationships with Germany reported above; the core group of countries relating to Germany appears nearly the same in both Tables 3 and 5.

Considering lastly the influence of Japanese policy, this plays a role only for Germany and for Australia. It is interesting that Australian policy appears to have closer links to its neighbor in Asia, than its culturally more similar but geographically more remote trade partners in Europe and North America.

A final word on these results pertains the dynamics captured by the ACH formulation. These are particularly important for Austria, Switzerland, Germany, France, the Netherlands and the U.K., where the sum $\alpha + \beta$ is, respectively: 0.81, 0.63, 0.84, 0.33, 0.45, and 0.67. They are somewhat less relevant but significant for Australia, Belgium, Japan, and the U.S. where only α is significant and attains the values 0.15, 0.09, 0.03, and 0.12 respectively. The only countries for which the dynamics are not relevant are Canada, Spain, Italy and Sweden, although we have already remarked the lack of degrees of freedom for Canada and Sweden. Overall, these numbers confirm that accounting for serial correlation patterns in the data is beneficial for the majority of countries in our sample.

5.2 The Directional Model

Table 6 reports the ordered probit results, which directly display the coefficient estimates of the relevant parameters (along with the standard errors) since they are easily interpretable. We begin by remarking that, although the sample sizes are generally quite small (this is due to the fact that we conduct the analysis in event time and thus depend on the number of actual policy adjustments made by each central bank), the coefficients are rather precisely estimated and the overall fit of the models is quite good.

The estimates on the parameters of the domestic macroeconomic variables are generally consistent with the typical behavior one would expect of central banks and embodied in a typical Taylor rule. Inflation coefficients are positive and significant for Austria, Belgium, Switzerland, Germany, Spain, Japan and the U.K., indicating that a rise in inflation above the 2% target is tempered by an increase in the policy rate. Similarly, output deviations from the 2.5% norm cause interest rates to increase although this effect is measured significantly for Canada, Switzerland and the U.S. only. The exchange variable, which triggered central bank responses in many of the countries in the timing model, is significant for Canada, Germany and the U.S. only.

What about the pattern of responses to the G-3? The timing model suggested that Austria, Belgium, France and the Netherlands responded significantly to the Bundesbank.

The responses of the directional model are largely consistent with this view, with significant responses for Austria, Belgium, and France but in all cases of the correct sign: these countries adjust interest rates at the same time and in the same direction as Germany.

Perhaps the only other pattern worth remarking is that of Australia. The timing model suggests that Australia responded strongly to Japan (the baseline hazard goes from 86% to 97%). The directional model also shows a strong response to Japan but in this case, it is in the direction of moving interest rates against the Bank of Japan.

5.3 Robustness – The Role of Formal Agreements

The empirical experiments in the previous section demonstrate that there is a substantial amount of monetary policy interdependence. The relations existing between Germany and Austria, Belgium, France and the Netherlands are likely to be a direct result of the formal ties that the European Monetary System (EMS) represented, now the European Monetary Union. By the same token, Italy's apparent lack of relation may be explained by its absence from the EMS from September 1992 to November 1996. Incidentally, the 1992 crisis in the EMS significantly affected Sweden although the sample available for this country only starts after 1994, well past this juncture. One may suspect that this particular episode may also adversely affect the results reported in the previous section. However, our methods are unaffected by this crisis since the logical consistency of our approach relies instead on the responsiveness of central banks in terms of timing and direction of adjustments made – an issue that would be clearly problematic had we taken the approach of using some measure of dynamic covariation. Nevertheless, to highlight that our results are not sensitive to the 1992 crisis, we experimented with a sample truncated in 1994. These experiments were broadly consistent with our findings in the previous section and therefore are not reported here to save space.

There is yet another dimension in which we experiment with the robustness of our findings. Interesting work by Dominguez (1997) explores the influence of formal international agreements. Dominguez (1997) studied the communiques produced from meetings of the G7, G-5 and G3 dating from 1975 to 1993, and identified when there was a call for a coordinated effort to reduce inflation or lower interest rates. Commitments to fight inflation coincide with periods where actual inflation is rather high, beginning with the first summit in 1975 through the London Summit in 1984, and again from mid 1988 to April 1989. The focus shifted to economic growth and commitments to lower interest rates in 1986, 1987, 1991 and 1992. Overall, Dominguez identifies in the sample fifteen cases of commitments to lower inflation and nineteen commitments to lower interest rates.

We experimented with the dates reported in Dominguez's (1997) study for the G-5 countries, namely Germany, France, Japan, the U.K., and the U.S. This choice is primarily motivated by the length of the available samples for these countries and the overlap with the Dominguez (1997) sample. We specified complete timing models that included the domestic macroeconomic variables, the G-3 timing variables, a G5/G7 meeting dummy if that month corresponded to a meeting reported in Dominguez (1997), and the ACH dynamics. Rather than reporting a new table of estimates, we computed the baseline hazard and the hazard computed with the dummy variable for the G-5/G-7 meetings. The baselines for Germany, France, Japan, the U.K. and the U.S. are 62%, 4%, 62%, 43%, and 70%, respectively. The G5/G-7 meeting dummy was not significant for any country except for Japan. The corresponding values for these hazards are 60%, 2%, 25%, 50%, and 67% – virtually identical to the baselines except for Japan, where a month in which a G-5/G-7 meeting takes place is usually a month in which the Bank of Japan is half as likely to adjust its policy rate.

6 Implications

The contributions of this paper span in several dimensions. At a methodological level, we demonstrated that interest rate targeting modifies the time series properties of market interest rate data in a manner that renders traditional vector time series techniques inapplicable. Our Monte Carlo evidence complements the results reported in Rudebusch (1995) regarding the effects of interest rate targeting on term structure regressions.

The solution we propose to this problem is based on an uncommon data set on policy rate target data and consists on decomposing central bank policy into the decision to adjust versus the actual type of adjustment. This strategy is based on previous work by Hamilton and Jordà (2002) for the U.S. and is modified and used here for the first time on other countries to detect monetary policy interdependence.

This empirical design has several advantages that go beyond the econometric pitfalls discussed early on. In particular, recent research by Rudebusch (2002) and Van Gaasbeck (2002) suggest that Taylor rules may appear to accurately describe central bank behavior due to the inherent persistence in macroeconomic data. We completely dispense with this artifact by concentrating instead on the snap shots that target rate adjustments provide of actual central bank policy. Hence, our results are elucidating, not only with regard to the issue of monetary policy linkages, but also as an alternative analysis of domestic monetary policy in industrialized OECD countries.

We find evidence of a clear European core, that responded to German policy long before the creation of a formal monetary union. Further this interdependence seems to have involved more than just attempts to stabilize exchange rates, but instead involved more direct responses to German policy shifts. We also find evidence of a distinct European periphery, including Italy, Switzerland and the UK, which did not respond to German policy. Finally, some countries respond to U.S. policy shifts, notably Germany and the UK. But we did not find any evidence that Japan responds to the U.S. Our results also shed light on the debate whether countries should include the exchange rate in their policy rules – quite a number of countries in practice do assign a very prominent weight on this variable.

References

Batten, Dallas S. and Mack Ott (1985) "The Interrelationship of Monetary Policies Under Floating Exchange Rates," *Journal of Money Credit and Banking*, 17, 103-110.

Benigno, Gianluca and Pierpaolo Benigno (2001a) "Implementing Monetary Policy Cooperation through Inflation Targeting" mimeo, New York University.

Benigno, Gianluca and Pierpaolo Benigno (2001b) "Price Stability as a Nash Equilibrium in Monetary Open Economy Models" mimeo, New York University.

Bensidoun, Isabelle, Virginie Coudert, and Laurence Neyman (1997) "Interest Rates in East Asian Countries: Internal Financial Structures and International Linkages," CEPII Working Paper 97/02.

Biltoft, Karsten and Christian Boersch (1992) "Interest Rate Causality and Asymmetry in the EMS," *Open Economies Review* 3, 297-306.

Borio, Claudio E. V. (1997) "Monetary Policy Operating Procedures in Industrial Countries," *Bank for International Settlements* working paper.

Bryant, Ralph C., Dale W. Henderson, Gerald Holtham, Peter Hooper, and Steven A. Symansky, eds. (1988) *Empirical Macroeconomics for Interdependent Economies*. Washington, DC: Brookings Institution.

Burdekin, Richard C. K. and Paul Burkett (1992) "The Impact of U.S. Economic Variables on Bank of Canada Policy: Direct and Indirect Responses," *Journal of International Money and Finance*, 11, 162-187.

Canzoneri, Matthew B., Robert E. Cumby and Behzad T. Diba (2001) “The Need for International Policy Coordination: What’s Old, What’s New, What’s Yet to Come?” mimeo , Gergetown University.

Chung, Jae Wan (1993) “Monetary Interdependence Among G-3 Countries,” *Applied Economics*, 25, 681-688.

Clarida, Richard, Mark Gertler and Jordi Gali (2001) “A Simple Framework for International Monetary Policy Analysis,” mimeo, New York University.

DeGennaro, Ramon P., Robert A. Kunkel and Junsoo Lee (1994) “Modeling International Long-Term Interest Rates,” *Financial Review,* 29(4), 577-97.

Dominguez, Kathryn M. E. (1997) “Do the G-3 Countries Coordinate Monetary Policy?,” in Cohen, Benjamin J ed.: *International Trade and Finance: New Frontiers for Research: Essays in Honor of Peter B. Kenen*, New York: Cambridge University Press, 280-315.

Evans, Charles L. and Kuttner, Kenneth N. (1998) “Can VARs Describe Monetary Policy?,” Federal Reserve Bank of Chicago Working Paper 98-19.

Fischer, Stanley (1988) “Macroeconomic Policy” in: *International Economic Cooperation*, Martin Feldstein ed., Chicago: University of Chicago Press, 11-43.

Frankel, Jeffrey A., and Katherine Rockett (1988) “International Macroeconomic Policy Coordination When Policy-makers Do Not Agree on the True Model,” *American Economic Review* 78, 318-40.

Furman, Jason and John Leahy (1996) “International Transmission of Monetary Policy: The US and Canada,” Harvard University discussion paper, December.

Garcia-Herrero, Alicia and John Thorton (1996) “Additional Evidence on EMS Interest Rate Linkages,” IMF working paper, October.

Hamilton, James D. (1996) “The Daily Market for Federal Funds,” *Journal of Political Economy,* 104(1), 26-56.

Hamilton James D. and Òscar Jordà (2002) “A Model for the Federal Funds Rate Target,” *Journal of Political Economy,* forthcoming.

Karfakis, Costas J. and Demetrios M. Moschos (1990) “Interest Rate Linkages Within the European Monetary System: A Time Series Analysis,” *Journal of Money, Credit and Banking* 22, 388-394.

Katsimbris, George M. and Stephen M. Miller (1993) “Interest Rate Linkages Within the European Monetary System: Further Analysis,” *Journal of Money Credit and Banking* 25, 771-779.

Nogues, Julio and Martin Grandes (2001) “Country Risk: Economic Policy, Contagion Effect or Political Noise?” *Journal of Applied Econometrics,* 4(1), 125-162.

Obstfeld, Maurice and Kenneth Rogoff (2001) "Global Implications of Self-oriented National Rules," forthcoming in *Quaraterly Journal of Economics.*

Oudiz, Gilles and Jeffrey Sachs (1984) "Macroeconomic Policy Coordination Among the Industrial Economies," *Brookings Papers on Economic Activity* 1, 1-64.

Rudebusch, Glenn D. (1995) "Federal Reserve Interest Rate Targeting, Rational Expectations and the Term Structure," *Journal of Monetary Economics 35, 245-74.*

Rudebusch, Glenn D. (1998) "Do Measures of Monetary Policy in a VAR Make Sense?" *International Economic Review* 39, 907-31.

Rudebusch, Glenn D. (2002) "Term Structure Evidence on Interest Rate Smoothing and Monetary Policy Inertia," *Journal of Monetary Economics,* forthcoming.

Sims, Christopher A. (1998) "Comment on Glenn Rudebusch's 'Do Measures of Monetary Policy in a VAR Make Sense?' " *International Economic Review* 39, 933-41.

Taylor, John, ed. (1999) *Monetary Policy Rules*, Chicago: University of Chicago Press.

van Gaasbeck, Kristin (2001) "Is Timing Everything? Investigating U.S. Monetary Policy with Interest-Rate Rules," mimeo: U.C. Davis.

von Hagen, Juergen and Michele Fratianni (1990a) "German Dominance in the EMS: The Empirical Evidence," *Open Economies Review* 1, 67-87.

von Hagen, Juergen and Michele Fratianni (1990b) "German Dominance in the EMS: Evidence from Interest Rates," *Journal of International Money and Finance* 9, 358-375.

Table 1 – Monte-Carlo Experiments

Changing the Frequency of Target Changes – No correlation in the timing, β = 0

Theoretical Frequency	Empirical Frequency A/and B	Directional Correlation - ψ	B Granger-causes A p-value	A Granger-causes B p-value
0.4	0.403/0.400	0.99	0.0185	0.6570
		0.75	0.0393	0.6700
		0.50	0.1316	0.6927
		0.25	0.4045	0.7001
		0.00	0.6140	0.6894
0.3	0.304/0.301	0.99	0.0864	0.6264
		0.75	0.1280	0.6446
		0.50	0.2573	0.6679
		0.25	0.4982	0.6908
		0.00	0.6322	0.6768
0.2	0.202/0.199	0.99	0.2658	0.6306
		0.75	0.3302	0.6256
		0.50	0.4498	0.6363
		0.25	0.5989	0.6428
		0.00	0.6560	0.6348

Note: $\rho^A = \rho^B = 0.5$, $\gamma^A = \gamma^B = 0.5$, $\sigma_A = \sigma_B = 0.25$,

Changing the Correlation in the Timing – Directional Correlation ψ = 0.75

Timing Correlation - β	B Granger-causes A p-value	A Granger-causes B p-value	Empirical Frequency A
0.25	0.0392	0.6638	0.403
0.5	0.0410	0.6553	0.402
0.75	0.0408	0.6488	0.401
0.99	0.0427	0.6453	0.401

Note: $\rho^A = \rho^B = 0.5$, $\gamma^A = \gamma^B = 0.5$, $\sigma_A = \sigma_B = 0.25$, *Empirical Frequency for B = 0.400*

Changing the Variance of the Errors around the Target - β = 0.99, ψ = 0.99

$\sigma_A = \sigma_B =$	B Granger-causes A: p-value	A Granger-causes B: p-value
1	0.4387	0.5016
0.5	0.1583	0.5682
0.25	0.0207	0.6204
0.1	0.0151	0.6293

Note: $\rho^A = \rho^B = 0.5$, $\gamma^A = \gamma^B = 0.5$, $\sigma_A = \sigma_B = 0.25$, *Empirical Frequency for A = 0.400, B = 0.400*

Table 1 (Contd.)

Changing the Autocorrelation of the Timing Process – β = 0.75, ψ = 0.75

$\rho_A = \rho_B =$	**B Granger-causes A p-value**	**A Granger-causes B p-value**	**Empirical Frequency A**	**Empirical Frequency B**
0.9	0.1227	0.5967	0.332	0.302
0.75	0.1304	0.6117	0.306	0.301
0.5	0.1231	0.6270	0.304	0.301
0.25	0.1225	0.6282	0.303	0.301
0	0.1325	0.6215	0.303	0.301

Note: $\gamma^A = \gamma^B = 0.5$, $\sigma_A = \sigma_B = 0.25$.

Changing the Autocorrelation of the Directional Process - β = 0.75, ψ = 0.75

$\gamma^A = \gamma^B =$	**B Granger-causes A: p-value**	**A Granger-causes B: p-value**
0.9	0.1084	0.4319
0.75	0.1096	0.5320
0.5	0.1231	0.6270
0.25	0.1941	0.6500
0	0.2374	0.6609

Note: $\rho^A = \rho^B = 0.5$, $\sigma_A = \sigma_B = 0.25$, *Empirical Frequency for A = 0.303, B = 0.301*

Equations of the Monte Carlo Simulation

$$r_t^i = \bar{r}_t^i + \varepsilon_t^i \qquad \varepsilon_t^i \sim N(0,\sigma_i^2) \qquad i = A, B$$

$$\bar{r}_t^i = x_t^i \bar{r}_{t-1}^i + (1 - x_t^i)(\bar{r}_{t-1}^i + y_t^i)$$

$$P(y_t^i = k_j \mid x_t^i = 1, \Omega_{t-1}) = P(c_{j-1} < y_t^{i*} < c_j \mid x_t^i = 1, \Omega_{t-1}) \qquad j = 1,2,\ldots,m$$

with $m = 5$, $k_1 = -0.50$; $k_2 = -0.25$; $k_3 = 0.25$; $k_4 = 0.50$; and $c_1 = -\infty$; $c_2 = -.75$; $c_3 = 0$; $c_4 = 0.75$; $c_5 = \infty$

Timing Model

$$h_t^i = \frac{1}{1 + e^{\lambda_t^i}}$$

$$\begin{cases} \lambda_t^A = \omega_A + \rho_A \lambda_{t-1}^A + \beta \lambda_{t-1}^B + e_t^A \\ \lambda_t^B = \omega_B + \rho_B \lambda_{t-1}^B + e_t^B \end{cases} \qquad e_t^i \sim N(0, 0.1)$$

$x_t^i = 1$ if $h_t^i > u_t^i$, 0 otherwise, where u_t^i is a uniform random variable in [0,1]

Directional Model

$$\begin{cases} y_t^{A*} = \gamma_A y_{t-1}^{A*} + \varphi y_{t-1}^{B*} + v_t^A \\ y_t^{B*} = \gamma_B y_{t-1}^{B*} + v_t^B \end{cases} \qquad v_t^i \sim N(0,1)$$

Table 2 – Summary of the Data Definitions

COUNTRY	Operational Monetary Data	Codes	Sample Begins	Macro-International Data	Freq.
Australia	Unofficial cash rate	ONAU	1/1/86	GDP, CPI, NEER	Q/M
	Target Rate	***PRAU***	1/23/90		
Austria	Day-to-day money	ONAT	1/6/89	IP, CPI, NEER	M
	Short-term operations (GOMEX)	***PRAT***	6/5/85		
	Lombard Rate	LRAT	1/24/80		
Belgium	Overnight interbank deposits	ONBE	1/2/91	IP, CPI, NEER	M
	Central Rate	***PRBE***	1/29/91		
Canada	Overnight rate	ONCA	1/2/80	IP, CPI, NEER	M
	Operating target band	***PRCA***	4/15/94		
France	Day-to-day rate	ONFR	1/1/80	IP, CPI, NEER	M
	Tender rate	***PRFR***	1/4/82		
	5-10 day repurchase facility	OCFR	10/12/88		
Germany	Day-to-day rate	ONDE	1/2/80	IP, CPI, NEER	M
	Repo rate	***PRDE***	4/1/80		
	Lombard rate	LRDE	1/1/80		
Italy	Overnight interbank deposits	ONIT	10/1/87	IP, CPI, NEER	M
	3-month interbank	M3IT	2/27/90		
	Tender rate	***LRIT***	5/13/91		
Japan	Overnight call money rate	ONJP	7/3/85	IP (up to 5/98), CPI, NEER	M
	Discount rate	***DRJP***	1/1/85		
Netherlands	Call money	ONNL	1/2/80	IP, CPI, NEER	M
	Rate on special advances	***PRNL***	1/2/80		
	Rate on advances (quota scheme)	DRNL	1/2/85		
Spain	Overnight interbank deposits	ONES	1/2/80	IP, CPI, NEER	M
	10-day repo purchases (marginal rate)	***PRES***	5/14/90		
Sweden	Day-to-day money	ONSE	11/21/88	IP, CPI, NEER	M
	Repo Rate	***PRSE***	6/1/94		
Switzerland	Day-to-day money (tomorrow next)	ONCH	1/4/80	GDP. CPI, NEER	Q/M
	Flexible Lombard rate	***LRCH***	5/26/89		
U.K.	Overnight sterling interbank deposits	ONGB	1/2/80	IP (up to 7/98), CPI, NEER	M
	Band 1 bank bill purchases	***PRGB***	1/2/80		
U.S.	Federal funds rate	ONUS	3/4/84	IP, CPI, NEER	M
	Federal funds rate target	***PRUS***	3/4/84		

Note: Target rates in bold and italic. *NEER* is the Nominal Effective Exchange Rate, *IP* is the Industrial Production Index, *CPI* is the Consumer Price Index

Table 3 – Hazard Rates for Policy Rate Adjustments as a Function of Adjustments in Germany, Japan and the U.S.

	Austria	**Australia**	**Belgium**	**Canada**	**Switz.**	**Germany**	**Spain**
Baseline	0.12**	0.18**	0.17**	0.44	0.83**	0.48	0.38*
Germany	0.53**	0.14	0.70**	0.53	0.90	0.48	0.49
Japan	0.04*	0.32	0.13	0.99	0.87	0.69*	0.23
U.S.	0.06*	0.30*	0.29	0.65	0.88	0.58	0.44
Obs.	111	98	89	48	111	150	94

	France	**Italy**	**Japan**	**Nether.**	**Sweden**	**U.K.**	**U.S.**
Baseline	0.25**	0.27**	0.06**	0.35**	0.41	0.35**	0.38**
Germany	0.42**	0.33	0.22	0.73**	0.55	0.26	0.37*
Japan	0.16	0.28	0.06**	0.47	0.36	0.27	0.35
U.S.	0.25	0.23	0.02**	0.25*	0.48	0.46	0.38**
Obs.	118	91	144	150	48	147	150

Comments:

- *Baseline*: measures the probability that in any given month, the corresponding country's policy rate will be adjusted.
- *Germany*: measures the probability of a domestic adjustment when Germany adjusted its policy rate that month.
- *Japan*: measures the probability of domestic adjustment when Japan adjusted its policy rate that month.
- *U.S.*: measures the probability of adjustment when U.S. adjusted its policy rate that month.
- */** indicates significant at the 90%/95% confidence level.

Table 4 – Classification of Target Changes

Country	***k*₁ = strong decrease**	***k*₂ = decrease**	***k*₃ = increase**	***k*₄ = strong increase**
Australia	$\Delta\bar{r}_t < -0.75$	$-0.75 \leq \Delta\bar{r}_t \leq 0$	$0 < \Delta\bar{r}_t \leq 0.75$	$\Delta\bar{r}_t > 0.75$
Austria	$\Delta\bar{r}_t < -0.5$	$-0.5 \leq \Delta\bar{r}_t \leq 0$	$0 < \Delta\bar{r}_t \leq 0.5$	$\Delta\bar{r}_t > 0.5$
Belgium	$\Delta\bar{r}_t < -0.5$	$-0.5 \leq \Delta\bar{r}_t \leq 0$	$0 < \Delta\bar{r}_t \leq 0.5$	$\Delta\bar{r}_t > 0.5$
Canada	$\Delta\bar{r}_t \leq -0.5$	$-0.5 < \Delta\bar{r}_t \leq 0$	$0 < \Delta\bar{r}_t < 0.5$	$\Delta\bar{r}_t \geq 0.5$
France	$\Delta\bar{r}_t \leq -0.4$	$-0.4 < \Delta\bar{r}_t \leq 0$	$0 < \Delta\bar{r}_t \leq 0.4$	$\Delta\bar{r}_t > 0.4$
Germany	$\Delta\bar{r}_t \leq -0.4$	$-0.4 < \Delta\bar{r}_t \leq 0$	$0 < \Delta\bar{r}_t \leq 0.4$	$\Delta\bar{r}_t > 0.4$
Italy	$\Delta\bar{r}_t < -0.5$	$-0.5 \leq \Delta\bar{r}_t \leq 0$	$0 < \Delta\bar{r}_t \leq 0.5$	$\Delta\bar{r}_t > 0.5$
Japan	$\Delta\bar{r}_t < -0.5$	$-0.5 \leq \Delta\bar{r}_t \leq 0$	$0 < \Delta\bar{r}_t \leq 0.5$	$\Delta\bar{r}_t > 0.5$
Netherlands	$\Delta\bar{r}_t \leq -0.4$	$-0.4 < \Delta\bar{r}_t \leq 0$	$0 < \Delta\bar{r}_t \leq 0.4$	$\Delta\bar{r}_t > 0.4$
Spain	$\Delta\bar{r}_t < -0.5$	$-0.5 \leq \Delta\bar{r}_t \leq 0$	$0 < \Delta\bar{r}_t \leq 0.5$	$\Delta\bar{r}_t > 0.5$
Sweden	$\Delta\bar{r}_t \leq -0.3$	$-0.3 < \Delta\bar{r}_t \leq 0$	$0 < \Delta\bar{r}_t < 0.3$	$\Delta\bar{r}_t \geq 0.3$
U.K.	$\Delta\bar{r}_t < -0.5$	$-0.5 \leq \Delta\bar{r}_t \leq 0$	$0 < \Delta\bar{r}_t \leq 0.5$	$\Delta\bar{r}_t > 0.5$
U.S.	$\Delta\bar{r}_t \leq -0.5$	$-0.5 < \Delta\bar{r}_t \leq 0$	$0 < \Delta\bar{r}_t < 0.5$	$\Delta\bar{r}_t \geq 0.5$

Table 5 – ACH Estimates: The Timing Model

	Austria	**Australia**	**Belgium**	**Canada**	**Switz.**	**Germany**	**Spain**
Constant	0.001**	0.86*	0.11**	0.45	0.79	0.05**	0.52
Inflation	0.001	0.85	0.08	0.58	0.86	0.06	0.47
Growth	0.002	0.65**	0.12	0.38	0.81	0.07**	0.51
Ex. Rate	0.99**	0.74	0.78**	0.39	0.31	0.69	0.68
Germany	0.10*	0.55**	0.58**	0.77*	0.82	0.05**	0.65
Japan	0.001	0.97**	0.09	0.45	0.99	0.65*	0.42
U.S.	0.000	0.89	0.18	0.45	0.70	0.53**	0.60
α	0.08**	0.15**	0.09*		0.00	0.09**	0.25
	(0.04)	(0.04)	(0.06)		(0.1)	(0.04)	(0.33)
β	0.73**				0.63**	0.75**	
	(0.40)				(0.25)	(0.31)	
OBS.	99	86	77	36	99	138	83
Log-Lik.	-44.51	-34.44	-35.89	-23.03	-37.74	-75.17	-54.52

	France	**Italy**	**Japan**	**Nether.**	**Sweden**	**U.K.**	**U.S.**
Constant	0.54	0.14	0.10**	0.05**	0.55	0.09*	0.38
Inflation	0.24**	0.15	0.06	0.05	0.61	0.09	0.54**
Growth	0.50	0.16**	0.10	0.05	0.49*	0.11*	0.36
Ex. Rate	0.72	0.18	0.16*	0.99**	0.49	0.36*	0.31
Germany	0.76*	0.14	0.33**	0.52**	0.60	0.02**	0.27
Japan	0.26	0.14	0.10**	0.07	0.55	0.01	0.24
U.S.	0.62	0.10	0.05	0.01**	0.55	0.34*	0.38
α	0.07**		0.03*	0.00		0.00	0.12**
	(0.03)		(0.02)	(0.14)		(0.13)	(0.03)
β	0.26**			0.45**		0.67*	
	(0.12)			(0.17)		(0.48)	
OBS.	106	79	132	138	36	135	138
Log-Lik.	-61.31	-45.21	-39.40	-73.08	-23.15	-79.07	-79.16

Notes: *Constant* refers to the baseline hazard; *Inflation, Growth, Ex. Rate* refer to the hazard when there is a 1% deviation in the corresponding variable from its norm; *Germany, Japan, U.S.* refer to the hazard when the corresponding G-3 bank adjusts its policy rate that month; α,β are the actual coefficient estimates from the ACH (standard errors in parenthesis); */** indicates significant at the 90/95% confidence level; *OBS.* Is the number of observations, which differ from those reported in Table 3 because we loose the first 12 observations to compute annual growth rates of the explanatory variables; *Log-Lik* is the log-likelihood value.

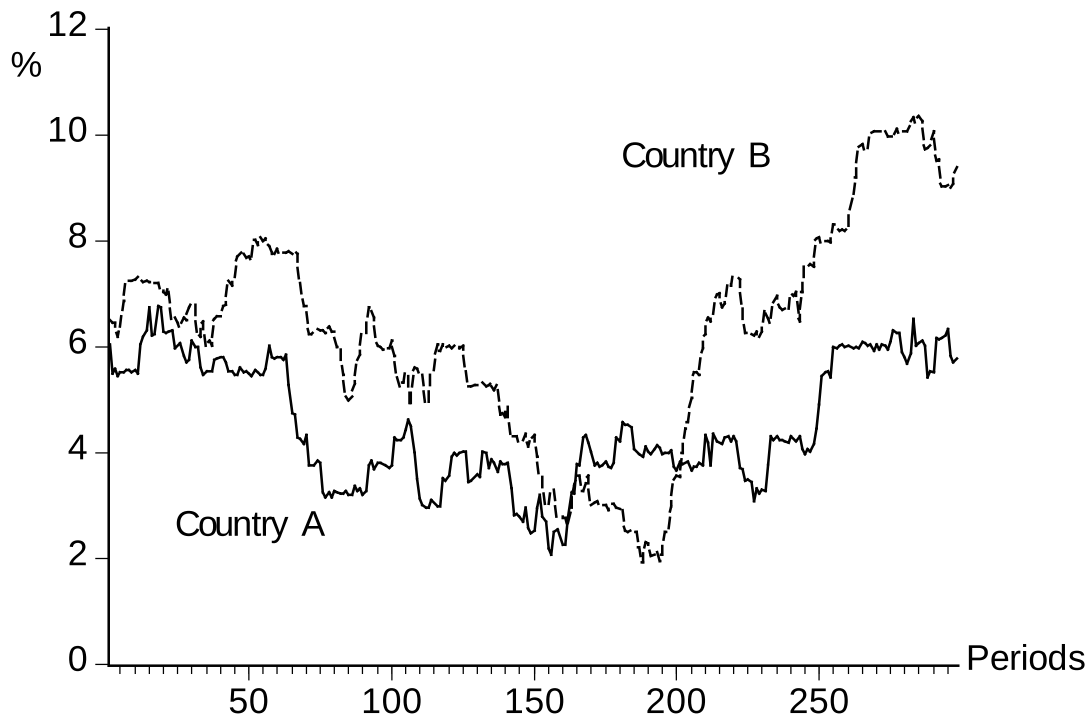
Table 6 – Ordered Probit Estimates

	Austria	**Australia**	**Belgium**	**Canada**	**Switz.†**	**Germany**	**Spain**
Inflation	6.56**	-1.38	12.14**	-0.54	0.32**	0.68**	10.10*
	(2.66)	(0.97)	(4.72)	(0.54)	(0.01)	(0.24)	(5.92)
Growth	0.53	2.84	0.34	5.20**	0.02**	-0.19*	-0.27
	(0.39)	(1.90)	(0.24)	(1.50)	(0.01)	(0.11)	(0.37)
Ex. Rate	2.23	-0.21	1.68	11.71**	-0.01	0.94**	-1.93
	(2.37)	(0.82)	(1.19)	(4.20)	(0.02)	(0.46)	(3.23)
Germany	5.75**		2.72*		-0.04		9.94*
	(2.91)		(1.64)		(0.06)		(5.70)
Japan	0.92	-3.80**	-0.44		-0.06**	1.16**	
	(4.05)	(1.96)	(0.79)		(0.03)	(0.32)	
U.S.	1.16		-1.97		0.14**	0.22	0.73
	(1.81)		(1.63)		(0.04)	(0.41)	(1.93)
Log. Lik.	-9.65	-4.72	-12.22	-22.06	67.71	-54.90	-5.22
Pseudo R^2	0.48	0.57	0.59	0.29	0.89	0.25	0.80
Obs.	28	16	35	23	72	75	35

	France	**Italy**	**Japan**	**Nether.**	**Sweden**	**U.K.**	**U.S.**
Inflation	39.39		3.80*	-1.66**	3.28	1.85**	0.37
	(31.20)		(2.03)	(0.36)	(3.65)	(0.41)	(0.32)
Growth	-0.78		0.59	-0.02	0.27	-0.56**	0.64**
	(0.76)		(0.39)	(0.08)	(0.35)	(0.20)	(0.14)
Ex. Rate	4.05		-0.30	-0.51	-0.04	0.15	0.32*
	(2.89)		(0.40)	(0.55)	(0.76)	(0.24)	(0.18)
Germany	12.10*			0.68	1.06	0.87	
	(7.41)			(0.55)	(2.71)	(0.72)	
Japan	5.85			0.35		-0.32	-0.70
	(5.74)			(0.32)		(0.48)	(1.12)
U.S.	4.00			0.86*		0.53	
	(3.25)			(0.46)		(0.61)	
Log. Lik.	-6.82		-6.33	-53.49	-10.24	-28.14	-69.74
Pseudo R^2	0.77		0.51	0.29	0.12	0.47	0.17
Obs.	36		13	72	14	43	64

Note: */** indicates significant at a 90/95% confidence level. Standard Errors in parenthesis. † Switzerland's model estimated by OLS due to the large variation in policy adjustments. There were not enough observations on Italy to estimate a viable model.

Figure 1 – Example of Monte Carlo Generated Interest Rate Series



Notes:

- p-value B Granger-causes A = 0.1231 for this example
- Timing Model Parameters:

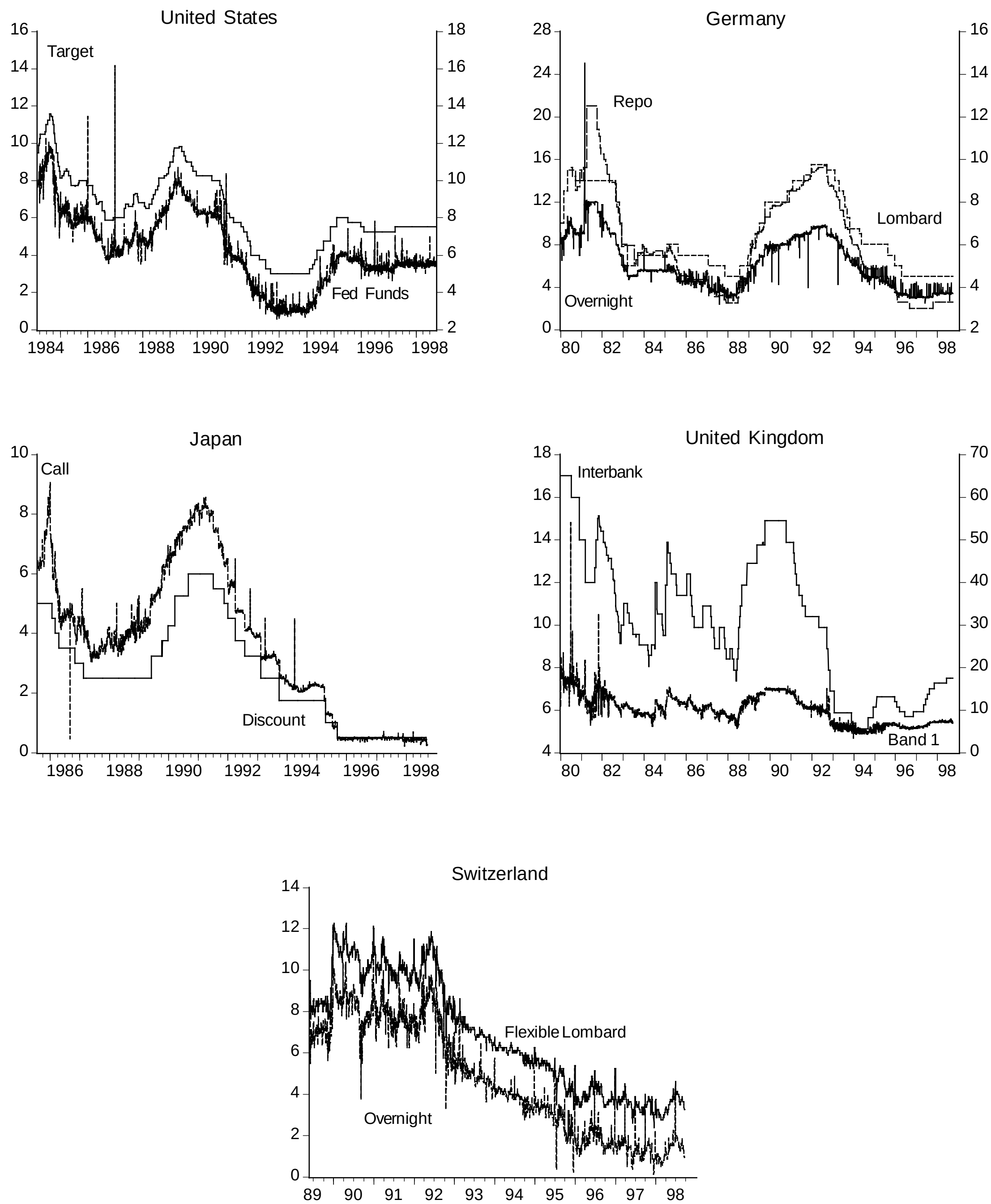
$$\rho_A = \rho_B = 0.5; \quad \beta = 0.75$$

- Directional Model Parameters:

$$\gamma_A = \gamma_B = 0.5; \quad \varphi = 0.75$$

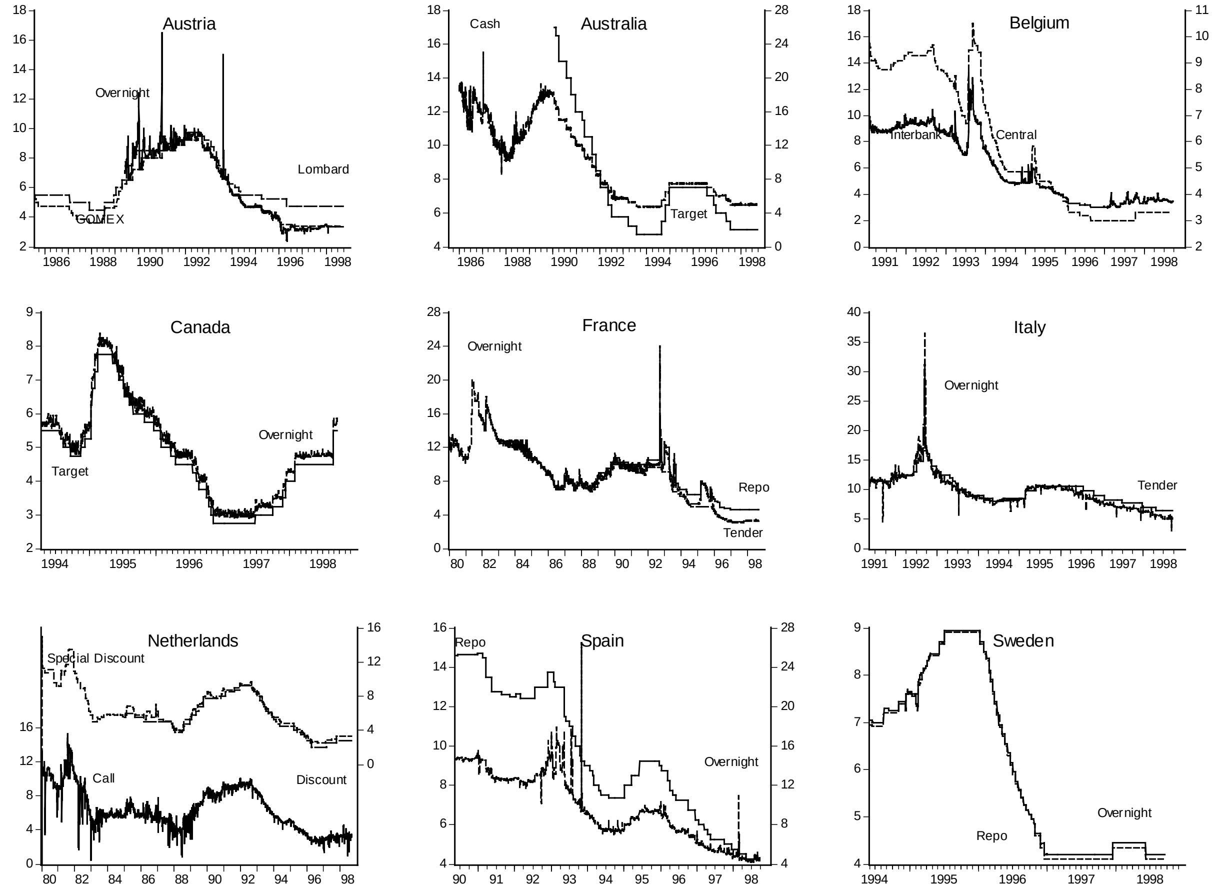
- Timing Frequency = 0.3

Figure 2 – Targets and Overnight Rates



Note: Whenever target rates and overnight rates are indistinguishable, we use dual scales for clarity.

Figure 2 (contd.) – Targets and Overnight Rates



Note: Whenever target rates and overnight rates are indistinguishable, we use dual scales for clarity.