<table>
<thead>
<tr>
<th><strong>Title</strong></th>
<th>The influence of Sterling on Irish interest rates</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Authors(s)</strong></td>
<td>Rodney, Thom</td>
</tr>
<tr>
<td><strong>Publication date</strong></td>
<td>1994-11</td>
</tr>
<tr>
<td><strong>Series</strong></td>
<td>UCD Centre for Economic Research Working Paper Series; WP94/24</td>
</tr>
<tr>
<td><strong>Publisher</strong></td>
<td>University College Dublin. School of Economics</td>
</tr>
<tr>
<td><strong>Item record/more information</strong></td>
<td><a href="http://hdl.handle.net/10197/1766">http://hdl.handle.net/10197/1766</a></td>
</tr>
<tr>
<td><strong>Notes</strong></td>
<td>A hard copy is available in UCD Library at GEN 330.08 IR/UNI</td>
</tr>
</tbody>
</table>
"The Influence of Sterling on Irish Interest Rates".

by

Rodney Thom
Department of Economics
University College, Dublin

November 1993

Working Paper WP94/24

DEPARTMENT OF ECONOMICS
UNIVERSITY COLLEGE DUBLIN
BELFIELD DUBLIN 4
The Influence of Sterling on Irish Interest Rates:

Rodney Thom.

Department of Economics
University College Dublin

Key Words: Interest rates. Exchange rates. Time Varying Parameters. Credibility

JEL Classification: C3, C32
The Influence of Sterling on Irish Interest Rates:

1. Introduction.

Recent empirical research has focused on the credibility of Irish exchange rate policy in the period between the August 1986 devaluation and the exchange rate crisis of late 1992 with emphasis given to the correlation between Irish interest rates and Sterling exchange rates. In the wake of the 1986 devaluation Irish policy makers appeared to move towards a 'hard ERM' strategy targeted at severing links between movements in Sterling and expectations of an Irish Pound realignment within the Exchange Rate Mechanism (ERM).¹

In general, when deteriorating competitiveness against the U.K. fuels expectations of an Irish Pound devaluation, the authorities may be forced to accept higher domestic interest rates in order to stem speculative outflows. However the extent to which such expectations materialise may depend on the perceived credibility of exchange rate policy. If, as intended by the 'hard ERM' strategy, market participants accept the credibility of the DM peg then a loss of competitiveness against Sterling does not necessarily imply an increased probability of a realignment and, as a consequence, a rise in interest rates. Hence a gain in credibility should be associated with a lower correlation between movements in domestic interest rates and the Sterling exchange rate.

Walsh (1993) considers the relationship between Irish, German and British short-term interests rates and the real Sterling exchange rate. He concludes that although the relative importance of the U.K. variables diminished over 1990 to 1992 they continued to exert a significant influence on Irish interest rates, suggesting that the policy of no realignments.

¹ For example the, then, Governor of the Central Bank describes exchange rate policy since 1987 an being "expressed clearly in terms of the ERM commitment and the Government remains committed to a policy similar to the 'franc fort' policy pursued in France." Doyle (1992, page 45)
while partially successful in moderating the influence of Sterling, was not fully credible.

A stronger conclusion is reached by Honohan and Conroy (1994) who analyse the influence of Irish pound exchange rates against Sterling and the DM on the Irish-German interest rate differential. They report that most of the variation in the latter can be explained by the Sterling exchange rate and that its influence is relatively constant over 1982 to 1992. In particular, their results cannot identify a significant role for the DM exchange rate or support the hypothesis that the 'hard ERM' policy pursued between the August 1986 devaluation and the currency crisis of September 1992 increased the credibility of the DM-peg as a means of narrowing the Irish-German interest rate differential.

Frain (1993) criticises both Walsh and Honohan and Conroy on the grounds that their results must be considered spurious in the sense of Philips (1986). His major concern is that most of the data series used in their empirical work appear to contain unit roots but are not cointegrated. Hence it is spurious to consider equilibrium relationships between them and to use the results to evaluate the long-term effects of interest and exchange rate policy.

Given Frain's rigorous testing procedures it would be difficult to contest his conclusions on econometric grounds. However the object of the present paper is not to refute Frain's findings, but to take an alternative approach which may go some way towards addressing the issues he raises. Rather than assuming that shifts occur at discrete and identifiable points in time this paper uses a specification based on time-varying parameters which permit gradual, or evolving, changes in the model's structure. The rationale underlying this approach can be explained by the following hypothetical example. Suppose the monetary authorities announce a new policy such as a firmer commitment to the DM-peg and that this is, correctly, interpreted as meaning that they will not respond to exchange rate changes against non-ERM currencies. Hence a researcher using time series data and least
squares techniques to investigate the influence of the Sterling exchange rate on Irish interest rates might reasonably expect to find a structural break coinciding with the period in which the policy change takes place and a decline in the exchange rate coefficient. However, if the policy shift is not fully understood or not fully explained by the authorities, then standard testing procedures may be unable to correctly identify structural changes in the underlying interest rate process because market participants, even if they make decisions on all available information, will not necessarily adjust their behaviour to accommodate the new, but uncertain, policy. Rather, the full impact of the policy shift may not be evident in the researcher’s estimates until the market undergoes a learning process with the appropriate information set changing as the policy is understood.

The following section discusses the Walsh and Honohan and Conroy models in more detail and illustrates how the latter can be modified to incorporate parameter variation which can be estimated with the Kalman Filter. Non-stationarity in the data is recognised by using a random walk specification for the model’s parameters which goes some way towards meeting Frain’s reservations. Results for 1986 to 1992, presented in Section 4, provide support for Walsh’s general conclusions and, to a lesser extent, for those of Honohan and Conroy. Specifically, the estimates suggest a decrease in the importance of the U.K. interest rate and a reduced responsiveness in the Irish-German interest rate differential to changes in the real Sterling exchange rate. However, as the latter retains statistical significance throughout the sample period we cannot conclude that exchange rate policy was the dominant factor driving the convergence of Irish interest rates with those in Germany.


The models used by Walsh and Honohan and Conroy both start with an uncovered
interest parity condition which equates the domestic interest rate to the difference between the 'foreign interest rate' and the expected appreciation of the Irish pound. Walsh models the former as a weighted average of the corresponding U.K. and German interest rates while Honohan and Conroy use a parity condition against the DM only. However both assume that the nominal rate of appreciation is a function of the deviation of the real exchange rate from its long-run, or target, level. Walsh’s approach yields a long-run model which expresses the Irish interest rate as a linear function of German and British rates and the real exchange rate against sterling.\textsuperscript{2} Walsh estimates a dynamic error correction version of this model using monthly data for two periods, 86:8 to 89:7 and 89:8 to 92:8. His estimates suggest an increase in the relative importance of the German interest rate and a decline in the responsiveness of Irish interest rates to changes in the real Sterling exchange rate. The latter is, however, statistically significant in both sub-periods. Hence Walsh concludes that although Ireland’s ERM commitment gained credibility over 1986 to 1992, "the continued role of both the sterling exchange rate and the UK interest rate indicates that the EMS peg was not wholly successful in breaking the influence of developments in the British economy on Irish interest rates" (page 448).

Honohan and Conroy use an alternative approach in that they decompose the DM nominal exchange rate into its central rate and spot position within the ERM band. However both the probability of a realignment and the expectation on a change in the spot position are modelled as simple linear functions of the current competitiveness position which, as in Walsh, is defined, as the deviation of the real exchange rate from its equilibrium level. Although their model suggests a time-varying parameter approach they resort to least squares

\textsuperscript{2} Walsh uses the real exchange rate against Sterling whereas Honohan and Conroy use a weighted average of Sterling and DM rates. However in each model the long-run real exchange rate is assumed constant.
by assuming a constant risk premium and equal devaluations at realignments. Using monthly
data over 1982:2 to 1992:7 they conclude that "overall, the results support a remarkably
simple specification, with the nominal sterling exchange rate as the key variable." (page 218)
Honohan and Conroy also use slope dummies to assess the influence of the sterling and DM
exchange rates over 1986:8 to 1992:7 but find that these dummies are insignificant suggesting
no significant gain in credibility during this period which is characterised by an absence of
realignments and a firmer commitment to ERM parities.³

It is of interest to note that they interpret this latter result as indicating that "there is
no evidence of an increase in the responsiveness of the interest rate to the Sterling exchange
rate after the 1986 unilateral devaluation, contrary to the fears of some market commentators
at the time of the devaluation and thereafter." (page 215) However in most of their
regressions the Sterling dummy has a negative sign and in one case, using the real exchange
rate, approaches significance with a t-statistic of 1.60. This suggests an alternative
interpretation: as the slope dummies span the period of the no devaluation policy, it is
plausible to expect a negative sign for these coefficients on the hypothesis that the objective
was to increase credibility and, if successful, reduce the responsiveness of domestic interest
rates to the Sterling exchange rate. Hence their results may indicate that there was no

These papers share a common methodology in that they each allow for structural
change at a specific point in the data sample. For example, although Walsh’s results suggest
some credibility gain over 1990 to 1992 his methodology implies that the model’s parameters
are constant within each period and that changes between the end of the first period and the

³ Walsh’s results differ in that his tests reject the hypothesis of constant parameters.
This may be due to different sample periods and/or different data. Honohan and Conroy use
on-shore interest rates whereas Walsh uses off-shore Euro-sterling and Euro-DM rates.
start of the second are correctly anticipated by market participants. Hence the implications of a policy change announced in period 1, such as a firmer commitment to ERM parities, are assumed to be fully understood in terms of their effects on the structure of the interest rate process. On the other hand, if a policy change is not fully understood, or is implemented but not announced, then market participants may have to "guess" or "estimate" its impact and the model's parameters which may vary over the learning period. Also, it is possible that unanticipated events, such as a rapid depreciation against Sterling, may lead to uncertainty about the feasibility of existing policy commitments and, as a consequence, to variations in the structure of the interest rate process. Again this suggests that interesting results might be obtained from a specification which permits the parameters to change over time.

To implement a procedure which allows for parameter variation I follow Honohan and Conroy by starting with an uncovered interest parity condition against the DM with the log spot exchange rate decomposed into the central rate \( c_i \) and the spot position within the ERM band \( x_i \). That is:

\[
\begin{align*}
\int r_t \, rdm_t \, s_t \, E(s_{t+1}) \\
\end{align*}
\]

(1)

where \( r_t \) denotes the domestic interest rate, \( rdm_t \) the corresponding German rate, \( s_t \) the log of the spot rate, DM per £IR, and \( s_t = c_i + x_i \). If \( \theta \) denotes the probability of a change in the central rate and \( \delta_{t+1} \) the size of the devaluation then the expected change in the spot rate is:

\[
\begin{align*}
s_t \, E(s_{t+1}) \, \theta \delta_{t+1} \, x_t \, E(x_{t+1}) \\
\end{align*}
\]

(2)

\(^4\) Walsh reports that this results are not sensitive to the period in which the structural break is assumed.
Honohan and Conroy model $\theta$ and the expected change in $x_i$ as linear functions of the deviation of the real exchange rate ($q_t$) from its target level. That is:

$$\theta = \theta_1 + \theta_2(q_t - \bar{q}_t)$$  \hspace{1cm} (3)$$

$$x_i - E(x_{i+1}) = \phi_1 + \phi_2(q_t - \bar{q}_t)$$ \hspace{1cm} (4)$$

where $\bar{q}_t$ is the target, or equilibrium, level. Assuming $\delta_{i+1}$ constant and combining (1), (2), (3) and (4) gives:

$$r_i - rdm_i = \beta_1 + \beta_2(q_t - \bar{q}_t)$$ \hspace{1cm} (5)$$

Where the $\beta$'s are functions of the parameters of (3) and (4). Following both Walsh and Honohan and Conroy the present paper also assumes $\bar{q}_t$ constant but amends (5) as follows. First, the Sterling interest rate is included on the grounds that competitiveness against the U.K. depends on financial market conditions as well as movements in the real exchange rate. For example, other things equal, a decline in U.K. interest rates gives a relative advantage to sectors of the British economy specialising in the production of import substitutes.\(^6\)

Second, (5) implies that a given change in $q_t$ generates the same probability of a realignment irrespective of the exchange rate policy in force and, by implication the same change in $r_i$ relative to $rdm_i$. However, as $\theta$ reflects the market's judgement of official reaction to changes

\(^5\) Honohan and Conroy define $q_t$ as a weighted average of the real exchange rates against Sterling and the DM. However as the 1986 and 1993 devaluations were primarily responses to competitive losses on U.K. markets the Sterling rate is used in this paper. Hence $q_t = \log(S_{Eg}P_t/P_{Eg})$ where $S_{Eg}$ is the nominal rate (£ per US$) and $P_t$, $P_{Eg}$ are the respective price levels.

\(^6\) Including the U.K. interest rate in a simple dynamic specification goes some way towards reconciling (6) with Walsh's model.
in competitiveness it cannot be assumed invariant with respect to exchange rate policy. For example, if the policy regime shifts from one which accommodates competitive losses against Sterling to a strategy based on a firmer commitment to the DM peg then we might expect to observe a decline in \( \theta \) for a given deviation of \( q_i \) from its target level. This type of change in the interest rate process can be accommodated by allowing the parameters of (5) to vary over time. Finally, adding a lagged dependent variable to allow for 'policy smoothing' gives the following model specified in matrix form:\(^7\)

\[
y_t = Z_t \beta_t + u_t, \quad u_t \sim N(0, \sigma^2)
\]  

(6)

where \( y_t = r_t - r_{dm_t}, \ Z_t = (1, \ q_t, \ rsg_t, \ y_{t-1}) \), \ rsg \ is the Sterling interest rate and \( \beta_t = (\beta_{1t}, \ \beta_{2t}, \ \beta_{3t}, \ \beta_{4t})' \).

If a policy shifts are well understood and considered credible then the parameters of (6) may, as in Walsh and Honohan and Conroy, be estimated by standard least squares methods with slope and intercept dummies used to control for the policy shift. Note that in this case, for given \( Z_{t+1} \), the conditional expectation of \( y_{t+1} \) is \( E_t y_{t+1} = Z_{t+1} \beta_{t+1} \). On the other hand, if the policy is not well understood with market participants having to estimate its implications for changes in the interest rate differential then the conditional expectation is \( E_t y_{t+1} = Z_{t+1} E_t \beta_{t+1} \), where \( E_t \beta_{t+1} \) is the estimate of \( \beta_{t+1} \). As the latter requires a stochastic specification for \( \beta_t \) this paper models the parameter vector as a random walk. That is:

\[
\beta_{t+1} = \beta_t + \epsilon_t, \quad \epsilon_t \sim N(0, \Omega_t)
\]  

(7)

---

\(^7\) Ó Cofaigh (1983) suggests that "although the Central Bank does not attempt to affect the broad trend of interest rates, it does pursue ... the important function of smoothing ... fluctuations in interest rates."
While the choice of an appropriate specification for $\beta_t$ is often arbitrary in time-varying parameter models, the use of a random walk model in the present context may be justified as follows.\footnote{See Evans (1991) who uses a similar argument to justify a random walk specification for the parameters of the U.S. inflation process.} If changes in exchange rate and/or interest rate policy are the major source of parameter variations and if policy changes at time $t+1$ cannot be predicted at time $t$ then $E_t \beta_{t+1} = \beta_t$ as implied by a random walk. Further, if, as is demonstrated in Frain (1993), the variables under consideration contain unit roots then the random walk specification will allow for this characteristic as shocks to $\beta_t$ will have a permanent effect on $r_t$.\footnote{Note that although a random walk model for $\beta_t$ is consistent with non-stationarity in the data it does not meet Frain's criticism that long-run inferences cannot be made when the series under consideration are not cointegrated. Little, however, is known about cointegrating relationships in time-varying structures.}

4. Estimation and Results.

Equations (6), measurement, and (7), transition, constitute a state-space model with time-varying parameters which can be estimated with the Kalman Filter.\footnote{See Chow (1983) for details.} If $b_{t-1}$ denotes a prior estimate for $\beta_{t-1}$ based on observations through $t-1$, the optimal linear predictor for $b_t$ is given by $h_{t,t} = h_{t,1} \quad h_{t,1}$ with covariance matrix $\Sigma_{t,t-1} = \Sigma_{t-1} + \Omega_t$. When a new observation becomes available the Kalman filter gives the updating equations:

$$h_t = h_{t,1} + K_{t-1} [y_t - Z_t h_{t,1}]$$

$$\Sigma_t = \Sigma_{t,t-1} - K_t Z_t \Sigma_{t,t-1}$$
where $K_t = \Sigma_{t+1}Z_t'[Z_t\Sigma_{t+1}Z_t' + \sigma^2]^{-1}$ is the Kalman gain. Hence the updated estimate $b_t$ is computed as the prior estimate plus the prediction error $[y_t - Z_t b_{t-1}]$ weighted by the gain of the Kalman Filter. Setting $\Omega_t = 0$ implies that $\beta_t = \beta_{t-1}$ and gives the classical fixed coefficient model. While this approach permits investigation of structural change it requires that the researcher has prior knowledge of when the change occurs.\textsuperscript{11} In the present context where market participants may assign credibility to a policy shift only when they learn how the authorities react to variations in competitiveness it is more appropriate to permit structural change to evolve over time and set $\Omega_t \neq 0$ which allows $\beta_t$ to change randomly from one period to the next.

The parameter vector $\beta_t$ was estimated using monthly data over 1986:10 to 1992:8. The Irish interest rate is the end of period three-month interbank rate while those for Germany and the UK are three-month Euro-rates. The Sterling real exchange rate was computed using the end of month nominal rate and consumer price indices for Ireland and the U.K.\textsuperscript{12} Following Chavas (1983) the process noise $\Omega_t$ was assumed proportional to $\Sigma_{t-1}$, the variance of the parameter estimates at time $t-1$ with the factor of proportionality estimated by minimising the one-step-ahead prediction error. To initialise the estimation starting values for $b$ and $\Sigma$ were taken from a least squares regression using data for 1985:1 to 1986:8 with estimates for 1986:10 to 1992:8 computed via the Kalman Filter.

\textsuperscript{11} When $\Omega_t = 0$ the population parameters are assumed constant and Kalman estimates are asymptotically equivalent to OLS. Although the Kalman Filter may give varying parameters due to sampling error, estimation by least squares over $t = 1 \ldots T$ will give the results identical to Kalman estimates for period $T$.

\textsuperscript{12} With the exception of the price indices all series were taken from the Central Bank of Ireland \textit{Quarterly Bulletin}. The former are from OECD \textit{Main Economic Indicators}. The data set is similar to that used by Walsh except that the monthly Irish CPI series was interpolated from the original quarterly series. The RATS software package was used for all computations.
Table 1. Kalman Estimates.

<table>
<thead>
<tr>
<th>Period</th>
<th>Constant</th>
<th>$q_t$</th>
<th>rsg$_t$</th>
<th>$y_{t+1}$</th>
<th>SSE</th>
</tr>
</thead>
<tbody>
<tr>
<td>1986:10</td>
<td>-5.077</td>
<td>0.168</td>
<td>1.021</td>
<td>0.402</td>
<td>0.684</td>
</tr>
<tr>
<td></td>
<td>(2.295)</td>
<td>(0.040)</td>
<td>(0.253)</td>
<td>(0.137)</td>
<td></td>
</tr>
<tr>
<td>1987:06</td>
<td>-1.926</td>
<td>0.151</td>
<td>0.655</td>
<td>0.538</td>
<td>0.720</td>
</tr>
<tr>
<td></td>
<td>(1.712)</td>
<td>(0.035)</td>
<td>(0.167)</td>
<td>(0.109)</td>
<td></td>
</tr>
<tr>
<td>1988:06</td>
<td>-0.424</td>
<td>0.115</td>
<td>0.448</td>
<td>0.584</td>
<td>0.657</td>
</tr>
<tr>
<td></td>
<td>(1.146)</td>
<td>(0.031)</td>
<td>(0.127)</td>
<td>(0.105)</td>
<td></td>
</tr>
<tr>
<td>1989:06</td>
<td>0.480</td>
<td>0.113</td>
<td>0.226</td>
<td>0.776</td>
<td>0.633</td>
</tr>
<tr>
<td></td>
<td>(1.247)</td>
<td>(0.033)</td>
<td>(0.107)</td>
<td>(0.075)</td>
<td></td>
</tr>
<tr>
<td>1990:06</td>
<td>1.928</td>
<td>0.075</td>
<td>-0.009</td>
<td>0.861</td>
<td>0.664</td>
</tr>
<tr>
<td></td>
<td>(1.284)</td>
<td>(0.029)</td>
<td>(0.082)</td>
<td>(0.077)</td>
<td></td>
</tr>
<tr>
<td>1991:06</td>
<td>1.836</td>
<td>0.075</td>
<td>0.012</td>
<td>0.839</td>
<td>0.606</td>
</tr>
<tr>
<td></td>
<td>(1.367)</td>
<td>(0.031)</td>
<td>(0.104)</td>
<td>(0.076)</td>
<td></td>
</tr>
<tr>
<td>1992:06</td>
<td>1.710</td>
<td>0.077</td>
<td>0.022</td>
<td>0.842</td>
<td>0.664</td>
</tr>
<tr>
<td></td>
<td>(1.316)</td>
<td>(0.033)</td>
<td>(0.121)</td>
<td>(0.076)</td>
<td></td>
</tr>
<tr>
<td>1992:08</td>
<td>1.753</td>
<td>0.078</td>
<td>0.022</td>
<td>0.838</td>
<td>0.546</td>
</tr>
<tr>
<td></td>
<td>(1.323)</td>
<td>(0.034)</td>
<td>(0.125)</td>
<td>(0.074)</td>
<td></td>
</tr>
</tbody>
</table>

Note: Standard errors are in parenthesis.
Figure 1. Normalised Estimate For $\beta_{2t}$: Real Exchange Rate
Figure 2. Normalised Estimate For $\beta_{3k}$: U.K. Interest Rate
Kalman estimates for June of each year together with the starting and final values are given in Table 1. The exchange and interest rate coefficients both decline over the estimation period with the former retaining statistical significance throughout and the latter becoming insignificantly different from zero by mid-1990. A more comprehensive picture is given by Figures 1 and 2 which display normalised Kalman coefficients for the real exchange rate and the Sterling interest rate over the estimation period. These normalised estimates are computed as:

\[ i_{it} - \frac{\hat{\beta}_0 - \hat{\beta}_{i0}}{SE(\hat{\beta}_{i0})} \quad i = 2,3 \]  

(10)

where \( \hat{\beta}_{i} \) is the Kalman estimate, \( \hat{\beta}_{i0} \) is the starting value and SE is the latter’s standard error. Mc Nelis and Neftçi (1982) define these normalisations as a "time-varying version of the standard t-test, where the hypothesised value and the standard error are held constant at the initial starting points." (page 301) They should not, however, be interpreted as indicating statistical significance as in constant parameter models. Rather, they indicate the "significance" of the deviation in each parameter from its initial value relative to what may be considered normal variation in the latter. Note also that as the initial value and its standard error are constants, the patterns of the normalised coefficients in Figures 1 and 2 exactly replicate those of the Kalman estimates. With the exception of the vertical axis scale, graphs for the actual estimates are identical to those for the normalised estimates.

Figure 1 indicates that the exchange rate coefficient increases over the second half of 1986, probably as a response to the August devaluation, and then decreases to a relatively

---

13 Kalman estimates start in 86:10 because \( y_{t,1} \) is included in (6). Ljung-Box test statistics for serial correlation in the Kalman residuals are: \( \chi^2(1) = 0.186, \chi^2(2) = 0.250, \chi^2(3) = 0.938, \chi^2(4) = 0.994 \) which are insignificant at the 95% level.
constant value between mid-1988 and the second half of 1989. The estimate then declines from just over 0.1 in June 1989 to a stable level of approximately 0.07 which persists from mid-1990 to August 1992. The normalised t-ratio of approximately -2.0 suggests a significant change relative to the initial value with the "jump" in late 1989 - early 1990 indicating a structural shift. Further, as this shift coincides with weakness of Sterling in the latter part of 1989 which put upward pressure on domestic interest rates but did not induce an accommodating adjustment in exchange rate policy it is possible to interpret the decline in the exchange rate coefficient as evidence of a gain in credibility. The fact that the authorities did not opt for a devaluation despite Sterling’s appreciation can, using Honohan and Conroy’s model, be interpreted as a decline in \( \theta \), the probability of a realignment, resulting in a reduced responsiveness of the interest rate differential to changes in the real exchange rate.
As with the real exchange rate, the U.K. interest rate coefficient, Figure 2, declines steadily from early 1987 and reaches a stable value in mid-1990. Given a normalised t-ratio of -4.0 this movement appears to be highly significant relative to the starting value. However, unlike the exchange rate coefficient which also stabilises in mid-1990, the U.K. interest rates ceases to have a significant influence on the Irish-German interest differential from late 1989 onwards.


This paper reconsidered the relationship between the Irish-German short term interest differential and the Sterling real exchange over a period between the unilateral devaluation in August 1986 and the onset of the ERM currency crisis in September 1992. As with previous papers the approach used allows for structural change by permitting the model’s parameters to vary over time. However, rather than assuming prior knowledge on when
structural shifts may occur, the estimation procedure assumes a stochastic random walk process which permits the parameters to evolve over time.

The general picture which emerges from the results is that the influence of U.K. variables on the Irish-German interest rate differential appears to have diminished over the sample period. In particular, the decline in the estimate for $\beta_2$ from 0.168 in October 1986 to 0.078 in August 1992 suggests a gain in credibility in the sense that the differential becomes much less responsive to movements in the real exchange rate. To this extent the 'hard ERM' policy introduced in the wake of the August 1986 devaluation appears to have been partially successful. However at least two factors must be borne in mind when interpreting the results. First mid-1990 to August 1992 was a period characterised by relatively stable foreign exchange markets with little movement in the real exchange rate against Sterling and an absence of pressure on Irish interest rates. In fact the interest rate differential narrowed significantly over this period and was close to zero in August 1992. Second the exchange rate coefficient, while decreasing over the estimation period, retains statistical significance throughout. In terms of the model outlined in Section 2 this implies that the market did not completely discount the probability of a realignment even though exchange rate movements in the latter part of the sample period posed little threat to the Irish pound's position within the ERM band. This suggests that although there is evidence of modest gains in credibility, the downward trend in the interest rate differential was primarily a consequence of stable currency markets, with the real exchange rate close to its target level, rather than total credibility of exchange rate policy.\(^{14}\)

Finally, although the results in this paper are broadly consistent with those suggested

\(^{14}\) An improving Irish fiscal situation together with an upward trend in the German interest rate may also have been responsible for narrowing the differential.
by Walsh and, to a lesser extent, by Honohan and Conroy who report no change in Sterling’s influence over the ‘hard ERM’ period, they appear to be contrary to recently expressed views of the Central Bank. For example:

the strong defence of the currency ... underlined Ireland’s commitment to the ERM and market perception of this has undoubtedly contributed to the rapid reduction in interest rates since end-January (Central Bank of Ireland, Annual Report 1992, p.9)

It is significant that this view was expressed in mid-1993 and that the "rapid reduction in interest rates" refers to the period following the February 1993 devaluation. Hence the Bank appears to be claiming that the no devaluation strategy followed from September 1992 to January 1993 was a significant factor underlying the post-devaluation fall in interest rates. While the reasoning behind this position is far from clear, it surely implies that the authorities strategy could not have diminished credibility during the months of the currency crisis. However when the Kalman estimates are extended to include data to December 1992 the exchange rate coefficient increases to 0.167, with a standard error of 0.032, which is exactly equal to the initial value in September 1986 and suggests a loss of credibility over September to December when policy was characterised by a "strong defence of the currency".
REFERENCES.


