CARDIFF BUSINESS SCHOOL
WORKING PAPER SERIES

Cardiff Economics Working Papers

Laurence Copeland and Saeed Heravi

Structural Breaks in the Real Exchange Rate Adjustment Mechanism

E2006/21

Cardiff Business School
Cardiff University
Colum Drive
Cardiff CF10 3EU
United Kingdom
t: +44 (0)29 2087 4000
f: +44 (0)29 2087 4419
www.cardiff.ac.uk/carbs

ISSN 1749-6101
July 2006

Structural Breaks in the Real Exchange Rate Adjustment Mechanism

Laurence Copeland and Saeed Heravi[1]

Cardiff Business School

July 11, 2006

[1] Subject to the usual disclaimer, the authors wish to acknowledge the comments of participants in seminars at De Nederlandsche Bank, Amsterdam and the University of Hannover for their helpful comments.

Abstract

Structural Breaks in the Real Exchange Rate Adjustment Mechanism

Laurence Copeland and Saeed Heravi

We show that the behaviour of the real exchange rates of the UK, Germany, France and Japan has been characterised by structural breaks which changed the adjustment mechanism. In the context of a Time-Varying Smooth Transition AutoRegressive of the kind introduced by Lundbergh et al (2003), we show that the real exchange rate process shifted in the aftermath of Black Wednesday in the case of the Pound, in 1984-5 in the case of the Franc and, more tentatively, during the Asian crisis of 1997-8 in the case of the Yen

1 Introduction[1]

The last few years have seen a remarkable revival of Purchasing Power Parity (PPP) as a theory of long run exchange rate equilibrium. The view that the real exchange rate is a random walk seemed almost unshakeable twenty years ago (see, for example, **?**) and had been given a theoretical basis by a number of authors, notably Roll (1979). However, more recently a large number of papers have appeared showing that PPP may be a credible description of the steady state. In particular, it has been shown that nonlinear models may help to explain the riddle posed by Rogoff (1996), that while nominal exchange rates exhibit extremely high volatility, adjustment to PPP shocks derived from the typical linear equation estimates stretch to a half-life of 3 years or more. By contrast, the evidence in the nonlinear models of Michael, Nobay and Peel (1997), Taylor, Peel and Sarno (2001), Baum, Barkoulas and Caglayan (2001), Paya, Venetis and Peel (2003), among others, supports the view that real exchange rates are driven by an arbitrage process, such that the speed of reversion to PPP is an increasing function of the scale of the shock and hence of the divergence from equilibrium. In this context, while small shocks are only very slowly reversed, since they create only small profit margins for arbitrageurs, large shocks trigger a scramble to exploit the free lunch, quickly driving prices and/or nominal exchange rates, if they are freely floating, back towards equilibrium.

The explanation for this type of mean reversion can take a number of different forms (e.g Dumas (1992), Sercu, Uppal and van Hull (1995)). One could imagine potential arbitrageurs having different trading costs, depending on their location, the scale of their

[1] Subject to the usual disclaimer, the authors wish to acknowledge the comments of participants in seminars at De Nederlandsche Bank, Amsterdam and the University of Hannover for their helpful comments.

operations, their competitive position in the market etc. Then, if they are thought of as ranged along a continuum from lowest to highest (transaction) cost, the further the real exchange rate from PPP, the greater the proportion of traders in a position to exploit arbitrage opportunities. Equally, the mechanism described here could be interpreted in terms of different goods, with deviations from the Law of One Price being rectified in ascending order of transportation costs etc. In any case, one implication is that the disequilibrium behaviour of the real exchange rate is best described by a nonlinear process which divides the space into three regimes: a central zone in the neighbourhood of the PPP exchange rate, surrounded by upper and lower zones defining the outer regimes, with smooth transition between the interior and exterior (Michael, Nobay and Peel (1997), Taylor, Peel and Sarno (2001)). In the current paper, we generalise this nonlinear approach to allow for time variation in the real exchange rate transition process itself, an extension which seems highly desirable for a number of reasons.

In the first place, while we accept the view expressed by Lothian and Taylor (1996) that using a dataset which covers a number of different regimes has the virtue of providing a more stringent test of mean reversion (in addition, of course, to more degrees of freedom), we nonetheless feel it may be unsafe to ignore major structural breaks altogether. The issue here is not simply a matter of the number or scale of the regime changes during the data period, but also their protracted nature, which suggests that it may take months or even years for agents to adjust fully to the change. For example, neither the breakdown of Bretton Woods nor the start and subsequent collapse of the European Exchange Rate Mechanism can be represented as clean once-and-for-all structural breaks. The Bretton Woods system collapsed on August 15th, 1971, and was succeeded from 1972 onwards by a series of short–lived fixed rate regimes (the Smithsonian Agreement, the so-called Snake, etc) before more or less free floating became the norm at some point in 1973. Likewise, the ERM collapse occurred in stages from mid-1992 onwards as fluctuation bands were first widened, then abandoned altogether. Secondly, even when regime change is instantaneous, we would expect the arbitrage mechanism underpinning the adjustment model to involve a learning process over a period of months or even years as traders in both the currency and goods markets adjust to the new policy environment.[2] This is even more true of structural changes occurring without any official or explicit regime change, as may have been the case with the development of Britain's North Sea oil fields in the early 1980's,[3] for example, or when the regime change is less obvious, like the Louvre Agreement of February 1987, or involves a completely unannounced change in intervention policy. To ignore the possibility that the adjustment process itself may have evolved over any given dataperiod is to risk ending up with a fitted model which is some kind of average of the different regimes, probably with highly unstable parameters.

In this paper, we confront this problem by fitting a model which makes explicit allowance for the evolution of the adjustment process over time, the Time-Varying Smooth Transition Autoregressive (TV-STAR) Model introduced by Lundbergh et al (2003), to a dataset consisting of real exchange rates of the Pound, Deutsche Mark, Yen and French Franc over a period starting in 1957. The next section sets out the model in its most general form, alongside a number of special cases yielding simpler formulations. Subsequent sections describe

[2] For the role of expectations in smooth transition models, see Peel and Venetis (2005)

[3] O'Connell (1998) cites this as an example of the kind of structural change which may explain his counter-intuitive results.

the dataset, specify the estimation methodology to be used, and discuss the results.

2 The Time-Varying Smooth Transition Autoregressive (TV-STAR) Model

The TV-STAR model of Lundbergh et al (2003) can be written in the general form:

$$\begin{aligned}\Delta q_t &= \{\boldsymbol{\alpha}'\mathbf{x}_t[1 - F_2(q_{t-d})] + \boldsymbol{\theta}'\mathbf{x}_t F_2(q_{t-d})\}\,[1 - F_1(\bar{t})] \\ &\quad + \{\boldsymbol{\beta}'\mathbf{x}_t[1 - F_2(q_{t-d})] + \boldsymbol{\phi}'\mathbf{x}_t F_2(q_{t-d})\}\,F_1(\bar{t}) + \varepsilon_t \qquad (1)\end{aligned}$$

In this equation, $\Delta q_t = \Delta s_t + \Delta p_t^* - \Delta p_t$ is the change in the log of the real exchange rate, where s_t is the log of the US dollar price of foreign currency (Pounds, Deutsche Mark, Yen and French Francs), and p_t^*, p_t are logs of the non-US and US consumer price index respectively, $F_1(\bar{t})$ is a logistic STAR (LSTAR) function of time, normalised so that $\bar{t} = t'/T$, where t' is the number of periods elapsed at any point in the time series and T is the total number of observations available for fitting. Hence $0 \leqq \bar{t} \leqq 1$ at all points. $F_2(q_{t-d})$ is an Exponential STAR (ESTAR) or LSTAR function of the transition variable, which in this case is the level of the real exchange rate itself, lagged d periods, and $\boldsymbol{\alpha}, \boldsymbol{\beta}, \boldsymbol{\theta}, \boldsymbol{\phi}$ are lag polynomials of order l in the vector of pre-determined variables, $\mathbf{x}_t$, which in this case consists only of past values of q_t and/or Δq_t.[4]

This formulation represents the adjustment process itself as evolving smoothly from the STAR in the first curly bracket parameterised by $\boldsymbol{\alpha}$ and $\boldsymbol{\theta}$ to the one in the second bracket

[4] See Akram, Eitrheim and Sarno (2005) for an example of a TV-STAR process fitted to the time series of nearly two centuries of annual Norwegian real exchange rates.

parameterised by $\boldsymbol{\beta}$ and $\boldsymbol{\phi}$, as the passage of time takes the value of the logistic $F_1(\bar{t})$ from its initial value of zero to its terminal value of 1.0 Hence, writing the LSTAR function of time as :

$$F_1(\bar{t}; \gamma_1, c_1) = \left[1 + e^{-\gamma_1(\bar{t} - c_1)}\right]^{-1} \qquad 0 \leqslant \bar{t} \leqslant 1 \quad \gamma_1 \geqslant 0 \tag{2}$$

makes the value of $F_1(\bar{t})$ exactly $\frac{1}{2}$ at the point $\bar{t} = c_1$, where the regime changes. On either side of this break, the process changes smoothly from its asymptotic starting point at:

$$\Delta q_t = \boldsymbol{\alpha}'\mathbf{x}_t[1 - F_2(q_{t-d})] + \boldsymbol{\theta}'\mathbf{x}_t F_2(q_{t-d}) \tag{3}$$

to its ultimate steady state value of:

$$\Delta q_t = \boldsymbol{\beta}'\mathbf{x}_t[1 - F_2(q_{t-d})] + \boldsymbol{\phi}'\mathbf{x}_t F_2(q_{t-d}) \tag{4}$$

when the effect of the regime change dies away. The size of γ_1 determines the speed of the transition from pre- to post-structural break regimes. Notice that where this parameter is extremely large, the associated shift can be regarded as equivalent to a one-off structural break of the kind which could be adequately modelled by conventional econometric methods (dummy variables etc). The time-varying model by contrast allows for evolutionary change around a structural break.

As far as the within-regime adjustment process is concerned, we allow for two alternatives. One possibility is the ESTAR mechanism:

$$F_2(q_{t-d}; \gamma_2, c_2) = e^{-\gamma_2(q_{t-d} - c_2)^2} \qquad d \geqslant 1 \quad \gamma_2 \geqslant 0 \tag{5}$$

which has generally been favoured by researchers in this area.[5] Since F_2 takes a value of 1.0 when $q_{t-d} = c_2$, and asymptotes to zero as $(q_{t-d} - c_2) \to \pm\infty$, it implies a symmetrical

[5] Though Sarantis (1999) which finds LSTAR rather than ESTAR best explains the behaviour of the real effective exchange rates of the Franc and Deutsche Mark.

reversion pattern. In other words, it imposes the condition that the extent of mean reversion in any period is the same whether the dollar is over- or undervalued relative to PPP, since the adjustment process at time t is dependent only on the absolute size of disequilibrium d periods ago (i.e. at $t-d$), irrespective of whether the disequilibrium was negative or positive.

Unlike many researchers in this area, we also allow for the possibility of asymmetric adjustment via an LSTAR process of the same general form as (2):

$$F_2(q_{t-d};\gamma_2,c_2)=\left[1+e^{-\gamma_2(q_{t-d}\,-c_2)}\right]^{-1}\qquad d\geqslant 1\quad \gamma_2\geqslant 0 \tag{6}$$

which allows the pattern of mean reversion to differ, depending on whether the currency is above or below its PPP equilibrium level.

The choice between the two specifications must ultimately be decided empirically, but it is worth considering the different implications of (6) and (5). To see why asymmetry cannot be ruled out a priori, consider the example of a shock to the nominal exchange rate of the Pound against the Dollar. If the shock is positive for the Pound (i.e makes it overvalued), then in the absence of any concomitant change in the equilibrium real exchange rate, UK exports become uncompetitive if the pass-through of costs is immediate and complete. This scenario is unlikely, however. Given imperfect competition, it is far more probable that, initially at least, UK exporters will absorb some of the exchange rate appreciation and accept lower margins in order to hold on to market share in the USA, especially if they expect the shock to be reversed in the near future. On the other hand, in the aftermath of a shock of the opposite sign, the incentive for US exporters to reduce their margins would in most cases be far smaller, given the relatively small UK market, so the pass-through from dollar to sterling prices would probably be far quicker.[6]

[6] For a discussion of some of these issues, see Campa and Goldberg (2002)

Against this, it could be argued that the opposite applies to British and American importers, if they tend to smooth prices by absorbing part of the short run impact of exchange rate fluctuations. However, it is not clear that the two effects need cancel out. In the first place, in both fixed and floating rate eras, bilateral current account balances were nonzero for long periods, especially in the cases of Japan and Germany. Secondly, since exports and imports are not necessarily the same good produced by the same industries, they may well have different associated trading costs, be sold in a different competitive environment and therefore adjust to disequilibrium at different speeds.

In summary, we see no reason to rule out LSTAR adjustment a priori and in fact prefer to let the data speak for itself.[7]

Interpretation of the results of TVSTAR estimation is often easier when (1) is rewritten as follows:

$$\begin{aligned}
\Delta q_t &= \boldsymbol{\alpha}'\mathbf{x}_t[1 - F_2(q_{t-d})][1 - F_1(\bar{t})] + \boldsymbol{\theta}'\mathbf{x}_t F_2(q_{t-d})[1 - F_1(\bar{t})] \\
&\quad + \boldsymbol{\beta}'\mathbf{x}_t F_1(\bar{t})[1 - F_2(q_{t-d})] + \boldsymbol{\phi}'\mathbf{x}_t F_2(q_{t-d}) F_1(\bar{t}) \\
&= \boldsymbol{\alpha}'\mathbf{x}_t - (\boldsymbol{\alpha}' - \boldsymbol{\beta}')\mathbf{x}_t F_1(\bar{t}) - (\boldsymbol{\alpha}' - \boldsymbol{\theta}')\mathbf{x}_t F_2(q_{t-d}) \\
&\quad + (\boldsymbol{\alpha}' - \boldsymbol{\theta}' - \boldsymbol{\beta}' + \boldsymbol{\phi}')\mathbf{x}_t F_1(\bar{t}) F_2(q_{t-d}) \qquad (7)
\end{aligned}$$

[7] In particular, we do not accept the argument of Taylor, Peel and Sarno (2001) that:

"Since....the LSTAR model implies asymmetric behaviour ...... we regard that model a priori as inappropriate.....it is hard to think of economic reasons why the speed of adjustment should vary according to whether the dollar is overvalued or undervalued....against the currencies of other ....industrialized countries" (p.1021)

They do nonetheless test for LSTAR processes in their dataset, presumably finding ESTAR ultimately preferable.

A restricted version of the model would eliminate the final interaction term by assuming:

$$\boldsymbol{\alpha}' - \boldsymbol{\theta}' - \boldsymbol{\beta}' + \boldsymbol{\phi}' = 0 \tag{8}$$

(e.g. Sensier, Osborn and Ocal (2002)). Although this formulation simplifies matters considerably, and makes estimation (and interpretation) a little more straightforward, it has no very obvious justification in the present context, and in fact experiments suggested it provided an inferior fit to the fully specified model.

3 Data

The data used here are monthly, starting in January 1957 and ending in December 2001 in the case of the Pound and Yen (about 540 observations), and with the advent of European Monetary Union in December 1998 in the case of the Deutsche Mark and Franc (about 500 observations).[8] Exchange rates are defined as the dollar price of foreign currency, and consumer price indices are seasonally adjusted. The series analysed is the log of the real exchange rate in each case. The data are originally taken from the Datastream database.

4 Estimation Method and Results

As is made clear by Lundbergh et al (2003), estimation of the TV-STAR model is not straightforward, with arguments in favour of a specific-to-general approach balanced by considerations favouring the opposite. (See also Kesriyeli, Osborn and Sensier (2004),

[8] Hopefully, enough observations to be immune to the upward bias in estimates of the adjustment speed in ESTAR documented in Paya and Peel (2006), though whether their conclusions apply equally to TV-STAR models and to LSTAR is unclear.

Sensier, Osborn and Ocal (2002)). The methodology adopted here involved a number of stages. First, a linear AR model was fitted to the data

$$\Delta q_t = a_0 + a_1 q_{t-1} + \sum_{i=1}^{m} b_i \Delta q_{t-i} + \varepsilon_t \tag{9}$$

with the order of the lag polynomial determined by the standard AIC criterion. This procedure served to identify as the best possible linear models $m = 1$ for Germany and Japan, and $m = 3$ for France and UK (Table 1). As expected, the estimated models are consistent with stationarity of Δq_t. Perhaps surprisingly, the diagnostics are not unfavourable, though the distribution of the residuals is, as usual with exchange rate data, a long way from normality.

Based on the optimal lag order established at this stage, we implemented the explicit tests for nonlinearity summarised in van Dijk, Terasvirta and Franses (2002),[9]where they are also advocated as a guide to identifying the appropriate model (ESTAR or LSTAR) and the best-fitting value of the delay in the transition variable. (see also Escribano and Jorda (1999)). The results, as summarised in Table 2, are suggestive rather than conclusive. The evidence of nonlinearity in the reported tests was unambiguous for Germany and Japan, far less so for France and UK. As far as the transition variable was concerned, repeating all tests for different values of the delay d gave no clear reason to prefer a lag greater than 1.

What is much clearer from the sequence of tests is that LSTAR is preferred to ESTAR, a conclusion which flies in the face of most of the published literature.[10] While the LM test results reported in the first panel of the table can only be said to favour LSTAR on balance, the hypothesis tests which follow them are quite unambiguous when we apply the decision

[9] We are grateful to the authors of this survey for the use of their software.

[10] One of the few exceptions is Sarantis (1999)

rule advocated by van Dijk, Terasvirta and Franses (2002): choose ESTAR if and only if the p-value for H2 is smallest, otherwise prefer LSTAR. As can be seen in the table, on this basis the condition for preferring ESTAR is not satisfied for any of the four countries, whether we consider the standard or the robust version of the statistics.

The bottom half of the Table compares the results of fitting a series of single-transition models by a process of two-dimensional grid-search over values of the transition parameters, γ_1 or γ_2 and the constant in the nonlinear function, c_1 or c_2. Judged by the AIC criterion, the model with a time transition alone (i.e. TV-AR) is preferrable to both ESTAR and LSTAR, though it is noticeable that while all the estimates of γ are large, indicating rapid transition between regimes, none is significantly different from zero. Compared on the same basis, LSTAR dominates ESTAR, albeit by a tiny margin, though ESTAR yields far more significant estimates of γ, especially for France and Germany. Moreover, the coefficient estimates in the TV-AR model look unstable, whereas the ESTAR and LSTAR equations both appear far more stable.

Estimating the TV-STAR model in its general form (7) is challenging, requiring a four-dimensional grid search over the γ_i and c_i with the full complement of m lags in each of the four polynomials. At the final stage, an attempt was made to find a more parsimonious representation by suppressing terms in the lag polynomials wherever possible. In the event, the interaction term seemed to be required in each case, in the sense that imposing the restriction given in (8) resulted in an inferior fit. The results reported in Table 3 are for the best-fitting models from the TV-STAR class, as selected by the AIC criterion. Notice that in the majority of cases, there is no evidence of any remaining autocorrelation, nor of any instability in the coefficient estimates. The c_i are all highly significant, and the estimates of the transition speed parameter, γ_i are always positive and mostly significantly greater than zero. Moreover, the AIC's for the best TV-STAR equations dominate not only those for linear but also for both classes of single-transition models.

There are a number of noteworthy features of the results. First, the coefficients of q_{t-1} are nearly all negative, with the few positives insignificantly different from zero, which means that the estimated equations are broadly stable, in the sense that, while the process may well be a random walk in the neighbourhood of equilibrium, it is clearly mean reverting in the outer regimes. (see e.g. Sarno and Taylor (2002))

The most surprising aspect of the results is in the estimates of the breakpoints (measured as a proportion of the dataperiod), $c_1 = \widehat{t}/T$. At the outset, our presupposition was that the regime change from fixed to floating rates at the end of the Bretton Woods era must have marked a decisive break in each time series. In the event, the outcome can be seen by reading Table 3 alongside Table 4, which, in order to help in interpreting the size of the regime shifts, shows computations of the implied steady values of q for the three limiting cases $F_2(q) = 0, \frac{1}{2}, 1$.

It can be seen that the TV-STAR process only picks out the start of the floating era in the case of Germany, for which the move to a floating rate regime seems to have been associated with an appreciation in the long run real exchange rate of the DM of about 1.5% (See Table 4).

As far as the other countries are concerned, although for the sake of completeness we have presented computations for all three regimes in Table 4, in reality the results reflect the limitation that, apart from Germany, each of the other three countries seems to have resided for most of the period in a single regime, leaving too few observations in the others to make for stable estimates. Specifically, as can be seen from the Figures, our results suggest France and Japan were almost always in the upper regime i.e. with $q >> c_2$ and hence $F_2(q) = 1$, which probably explains the implausibly large estimates of the the shift in the equilibrium in the two other states.

With that caveat in mind, we conclude that there is strong evidence of a shift in the Pound's steady state value around or soon after Black Wednesday (16th September 1992), probably associated with an appreciation of about 3% against the Dollar. For Japan, the 1972-3 break appears to be susbsumed in the second LSTAR function, $F_2(q)$. Examination of the graph (Figure 3) suggests that, since the Yen underwent a more or less unbroken appreciation over the first half of our data period, the 1972-3 episode appears as a steady state level of the real exchange rate. The time function, $F_1(t)$, then picks out a second break in the early stages of the Asian crisis, when Japan began to experience deflation for the first time. Some indication of the scale of the shift can be gained from Table 4, which shows that the implied appreciation in the long run steady-state level of the Yen against the dollar may have amounted to about 30%.

To ensure this conclusion was not simply caused by the fact that we have relatively few data points after the 1997 break,[11] we fitted the same equation to the floating rate subperiod only. As can be seen, the results are broadly similar to those for the full period, with the break identified a few months later, in early 1998, suggesting the shift is quite robust.

[11] In fact, as was clear from the graph of $F_1(t)$, (not shown here, but available from the authors on request) the dataset ends well before the time transition function gets close to its upper asymptote of 1.0. The scale of the apparent shift for Japan may also be due to other factors. For example, markets may have extrapolated the falling Japanese price level into the indefinite future. Also, the model implemented here ignores any possible Balassa-Samuelson productivity-growth effect.

For the Franc, the regime change is also at an unexpected point. Presumably, the break in 1984-5 reflects the previous year's turning point in French monetary policy with the abandonment of reflation and the adoption of the *franc fort,* generating a rise in the steady-state real exchange rate of nearly 6%.[12] It is also notable that only for France does it look as though the regime change may have been sharp enough to be proxied by a step function or dummy variable. For Germany, the UK and, especially Japan, the shift in the mean reversion process was far more gradual.

5 Generalised Impulse Response Functions

The task of assessing the response of nonlinear models to shocks is nontrivial, as the fitted equations cannot be straightforwardly inverted to yield impulse response functions of the kind typically used to analyse ARMA and VARMA processes. Instead, we are forced to use so-called Generalised Impulse Response Functions (GIRF's), (Koop, Pesaran and Potter (1996)) defined as:

$$G_{h,\delta}(h, \delta, \omega_{t-1}) = E(\Delta q_{t+h} \mid \varepsilon_t = \delta, \omega_{t-1}) - E(\Delta q_{t+h} \mid \varepsilon_t = 0, \omega_{t-1}) \tag{10}$$

where $h = 1, 2.....60$ is the forecast horizon in months, δ is the size (in units of standard deviations) of the shock to the residual in the basic equation 1 for which the simulated path is being generated and ω_{t-1} is the history of the process up to and including the time $t-1$.

[12] taking account of the fact that France was in the upper zone of the exchange rate transition function for almost the whole period (see Figure 1). The implausibly large size of the implied devaluation in the other two states is almost certainly a result of the fact that we have virtually no observations in the lower regime and not many in the transition phase between the two.

It is well known that nonlinear systems are path-dependent, in the sense that the effect of a random shock at time t depends on the whole past history of the process. In other words, the function in 10 is a random variable whose realisation depends not only on the shock, δ simulated at t, but also on the particular path taken from time 0 until $t-1$ summarised by ω_{t-1}. It follows that our conclusions have to be based on the means of artificially generated histories of length $T-1$, where T is the length of our dataset, supplemented by simulated post-sample paths for the shocks from $T+1$ to $T+h$. For each currency, we computed the mean[13] of 1000 simulations bootstrapped from the estimated error process, with $\delta = i\sigma_\varepsilon$ for $i = \pm 1, \pm 4, \pm 8$. (It should be noted, of course, that for highly nonlinear systems, the effect of a shock of $k\sigma$ is not necessarily k times the effect of a 1σ shock, nor can we be sure the outcome will be the negative of a shock of $-k\sigma$).

The results of the GIRF exercise are best viewed as regime-dependent. Since the in-sample observations predominantly relate to one or two regimes (e.g. for Germany $F_1 < 0.5, F_2 < 0.5$ rather than $F_1 < 0.5, F_2 > 0.5$, which is never actually observed), we only present a subset of the possibilities for each currency. The GIRFs are summarised in Table 5 which presents the half-lives of shocks i.e. the number of months required for one-half of the initial disturbance, δ, to be eliminated. The results are only given for $i = \pm 1, \pm 4$ to save space[14], and for the regimes which predominated during our sample period. Graphs of the GIRFs for $i = +1$ are given in Figures 2A to 2D.

[13] In principle, we need not confine attention to the mean. In fact, van Dijk, Terasvirta and Franses (2002) consider three different percentiles of the realized distribution of outcomes. Here, we examined the medians whenever we suspected the means may have been distorted by a small number of extreme realizations. The results, however, were not qualitatively different, so are not reported here.

[14] For other values, the results (which were qualtitatively similar) are available from the authors.

There are a number of notable points. First, the results for Germany displayed remarkable stability at every level of shock, from -8 to +8. Not only did the GIRFs invariably tail off, but the rate of convergence seemed to vary little across fixed and floating regimes. In fact, a half-life of around 2 years for a $\delta = 1\sigma$ shock is not far off the consensus in the published literature. One puzzling aspect of the results for Germany is that half lives of large shocks are barely any shorter than for small shocks, contrary to the logic of the model. The same is true for Japan and UK, though post-1983 France does appear to fit the pattern of more rapid convergence in the aftermath of larger shocks. Prior to the 1992 break, the UK results suggest instability (often in the form of unstable cycles), whereas post-1992 the process looks slightly more consistent with very slow convergence. Both for France and Japan, the real mean reversion process appears to have speeded up dramatically after the respective breaks. In the French case, this could be a consequence of tying the nominal exchange rate more closely to the DM, though this might have been expected to cause both countries to mean revert at the same speed, rather than cause France to actually be faster than Germany. As far as Japan is concerned, the dramatic fall in convergence speed after 1997 is harder to interpret.

6 Conclusions

The evidence presented in this paper suggests the real exchange rate adjustment process has itself been subject to regime shifts even within the floating rate era. Taking them into account explicitly through a TV-STAR model allowed us to identify these shifts, while at the same time estimating the mean reversion process on either side of the breaks. The analysis does however leave a number of obvious loose ends. First, it remains to be seen whether the models fitted here have any forecasting power. Also, this paper has not addressed the question of the existence or otherwise of a Balassa-Samuelson effect, which has already been mentioned as possibly distorting our results, especially for Japan. It is not obvious how to introduce a productivity-growth proxy into a model which already includes a nonlinear function of time, (Paya and Peel (2003) suggest this is problematic even in single-transition models) and the few initial experiments we conducted (not reported here) were not very promising. However, this is a topic which will be investigated in future work.

References

Adler M and Lehman B (1983) Deviations from Purchasing Power Parity in the Long Run, Journal of Finance, 38, 1471-87

Akram Q F, Eitrheim O and Sarno L (2005) Nonlinear Dynamics in Output, Real Exchange rates, and Real Money Balances: Norway 1830-2003, Norges Bank W.P. ANO 2005/2

Baum, C F., Barkoulas, J T. and Caglayan, M (2001) Nonlinear adjustment to purchasing power parity in the post-Bretton Woods era, Journal of International Money and Finance, Elsevier, vol. 20(3), pages 379-399, June

Campa, J M and Goldberg L (2002) Exchange Rate Pass-Through into Import Prices: a Macro- or Micro Phenomenon? NBER Working Paper 8934

Dumas B (1992) Dynamic Equilibrium and the Real Exchange Rate in a Spatially Separated World, Review of Financial Studies, 5, 153-80

Escribano, A and Jorda, O (1999) Improved Testing and Specification of Smooth Transition Models, in Nonlinear Time Series Analysis of Economic and Financial Data, Kluwer Academic Publishers, USA

Koop, G, Pesaran M H and Potter S M (1996) Impulse Response Analysis in Nonlinear Multivariate Models, Journal of Econometrics, Vol 74, pp 119-47

Kesriyeli M, Osborn D R and Sensier M (2004) Nonlinearity and structural change in interest rate reaction functions for the US, UK and Germany, CGBCR Discussion Paper Number 44, University of Manchester

Lothian, J R and Taylor, M P (1996) Real Exchange Rate Behavior: The Recent Float from the Perspective of the Past Two Centuries, Journal of Political Economy, 104 (3), 488-509

Lundbergh, S, Terasvirta, T and van Dijk, D (2003) Time-Varying Smooth Transition Autoregressive Models, Journal of Business and Economic Statistics, 21, 1, 104-21

Michael, P, Nobay, A R and Peel, D A (1997) Transactions Costs and Nonlinear Adjustment in Real Exchange Rates: an Empirical Investigation, Journal of Political Economy, 105, 862-79

O'Connell P G J (1998) Market Frictions and Real Exchange rates, Journal of International Money and Finance, 17, 71-95

Paya, I and Peel, D A (2003) Purchasing Power Parity Adjustment Speeds in High Frequency Data when the Equilibrium Real Exchange Rate is Proxied by a Deterministic Trend, Manchester School (Supplement), 39-53

Paya, I and Peel, D A (2006) On the Speed of Adjustment in ESTAR Models when Allowance is Made for Bias in Estimation, Economics Letters, Vol 90, 272-7

Paya, I, Venetis, I A and Peel D A (2003) Further Evidence on PPP Adjustment Speeds: the Case of Effective Real Exchange Rates and the EMS, Oxford Bulletin of Economics and Statistics, 65, 4, 421-37

Peel D A and Venetis I A (2005) Smooth Transition Models and Arbitrage Consistency, Economica, 72, 413-430

Rogoff, K (1996) The Purchasing Power Parity Puzzle, Journal of Economic Literature, 34, 647-68

Roll R (1979) Violations of Purchasing Power Parity and their Implications for Efficient International Commodity Markets, in Sarnat, Szego (eds) International Finance and Trade, Vol 1, Ballinger, Cambridge, Mass, pp 133-76

Sarantis, N (1999) Modelling Non-Linearities in Real Effective Exchange Rates, Journal of International Money and Finance, Elsevier, vol. 18, pages 27-45

Sarno, L and Taylor M P (2002) Purchasing Power Parity and the Real Exchange Rate, IMF Staff Papers, vol 49, Number 1, 65 -105

Sensier M, Osborn D R and Ocal N (2002) Asymmetric interest rate effects for the UK real economy, Oxford Bulletin of Economics and Statistics, 2002, vol. 64, pp.315-339

Sercu P, Uppal R and van Hulle C (1995) The Exchange Rate in the Presence of Transactions Costs: Implications for Tests of Purchasing Power Parity, Journal of Finance, 50, 1309-19

Taylor, M P, Peel, D A and Sarno, L (2001) Nonlinear Mean-Reversion in Real Exchange Rates: Towards a Solution to the Purchasing Power Parity Puzzles, International Economic Review, 42(4), 1015-42

van Dijk, D, Terasvirta, T and Franses, P H (2002) Smooth Transition Autoregressive Models - a Survey of Recent Developments, Econometric Reviews, 21(1), 1-47

Figure 1A: France

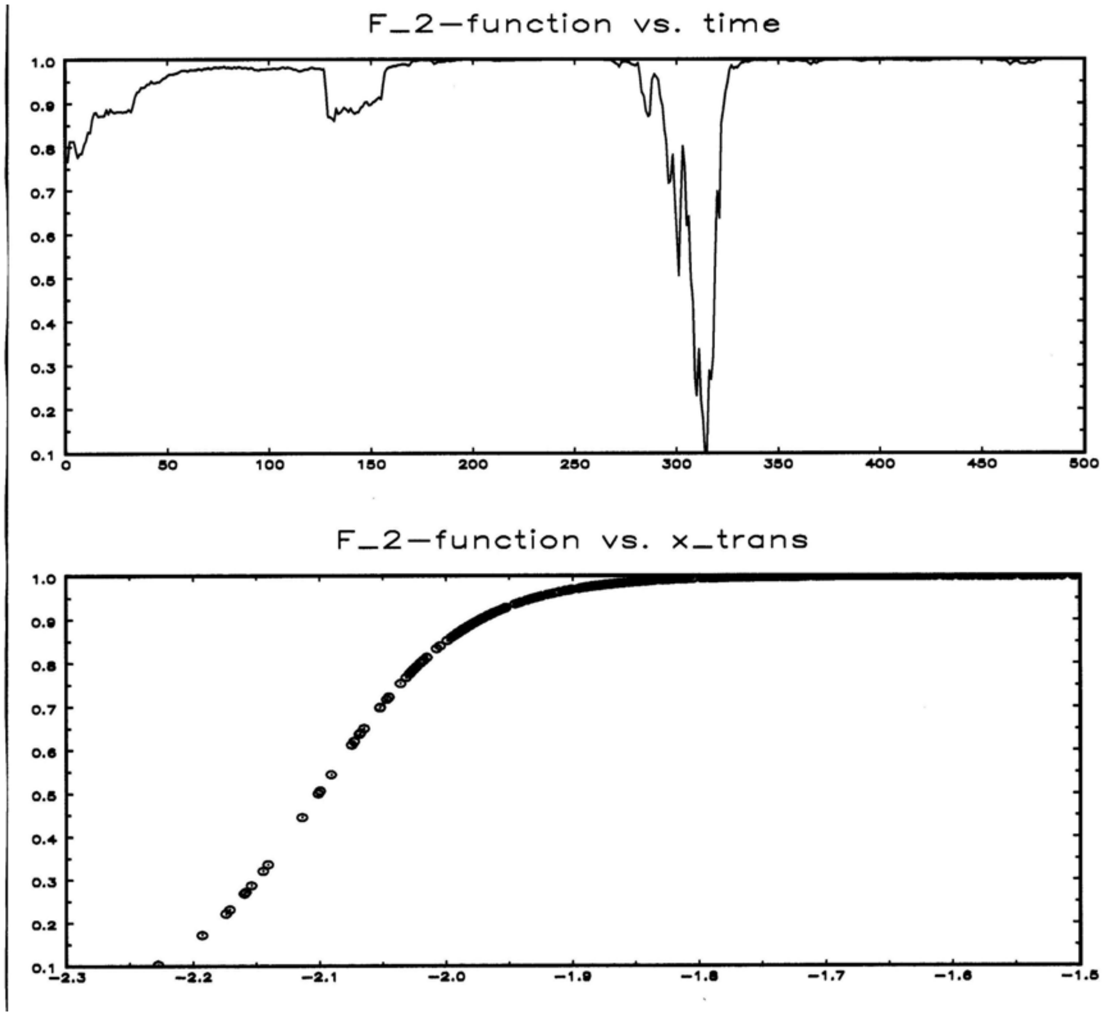


Figure 1B: Germany

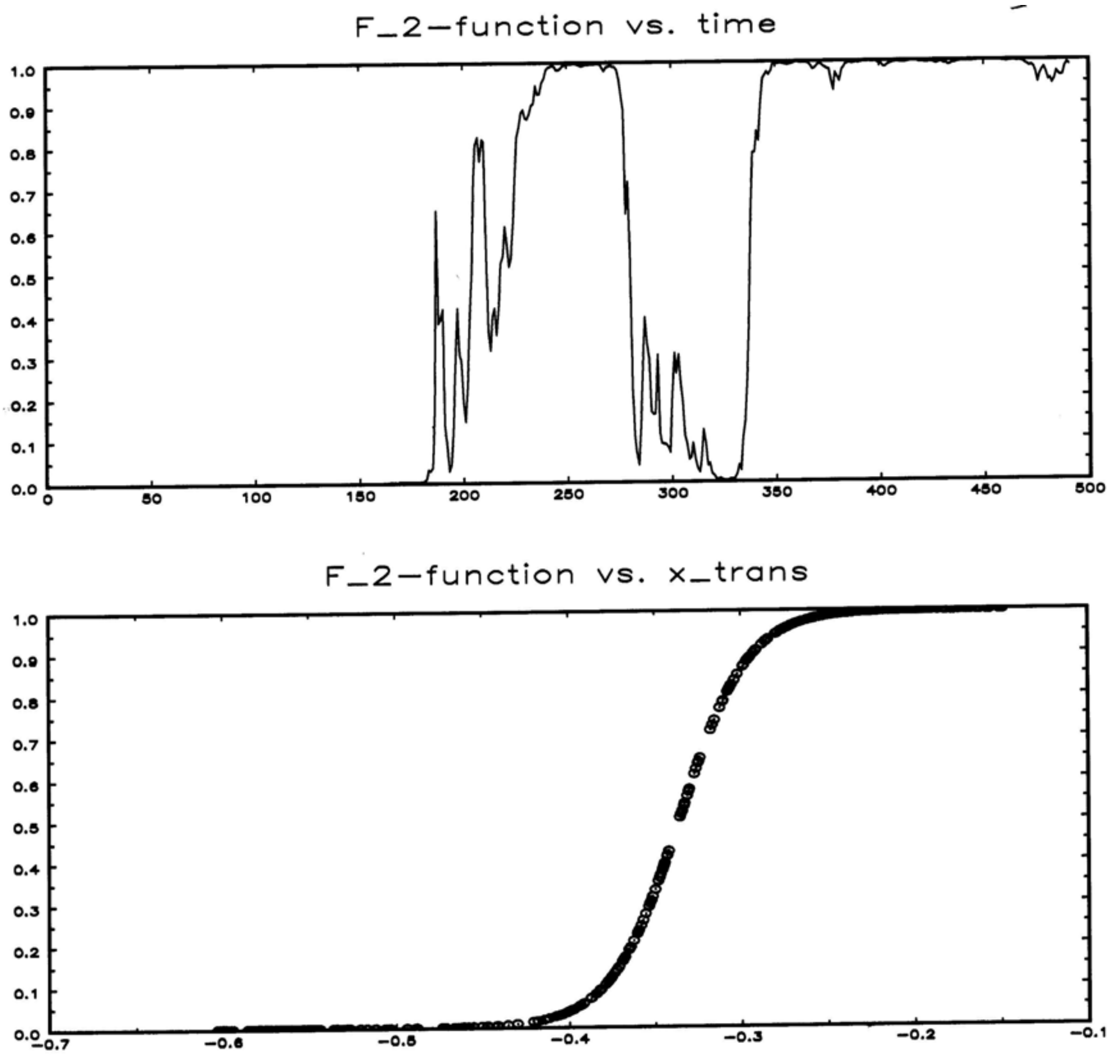


Figure 1C: Japan

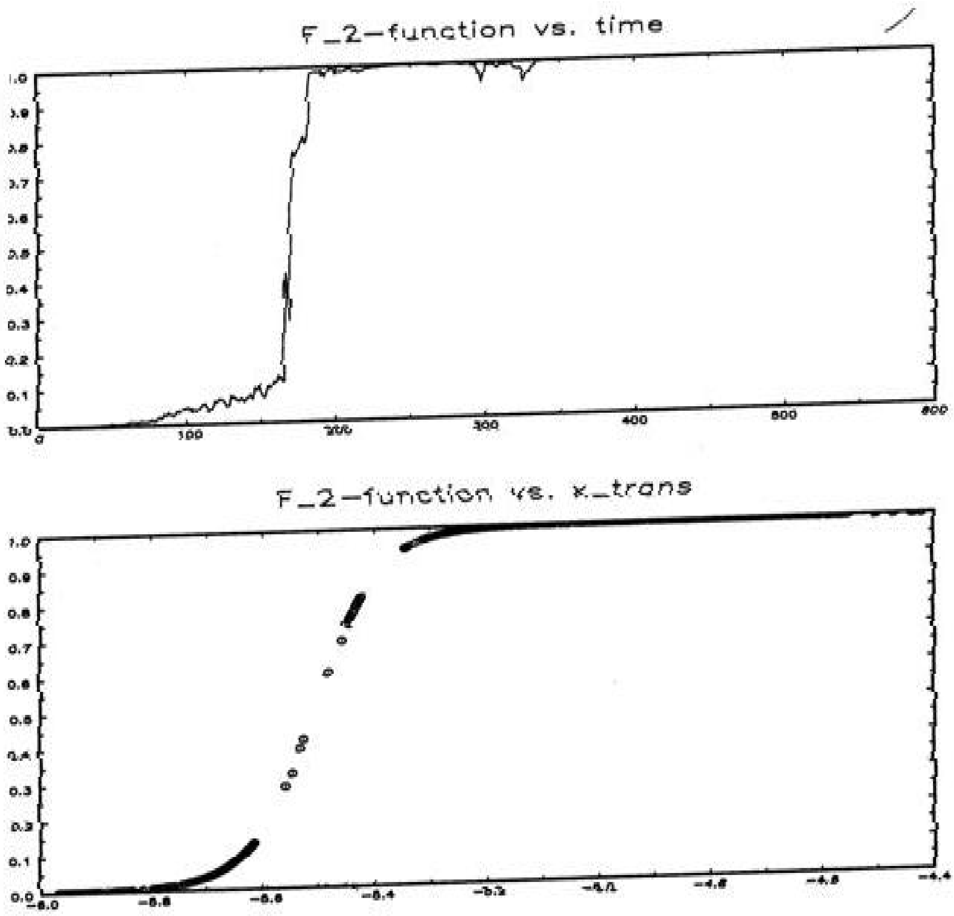
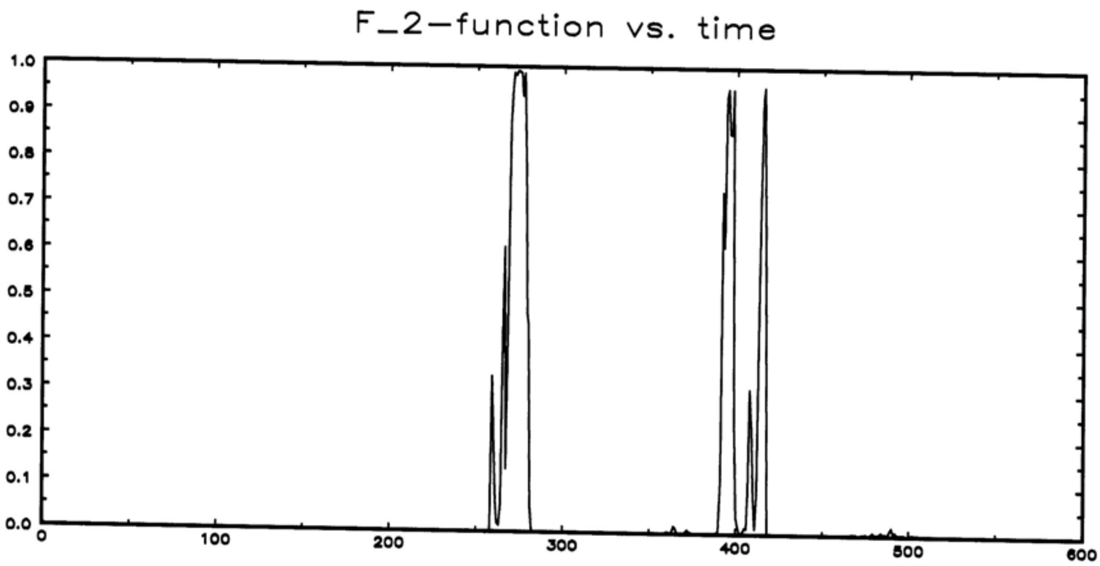


Figure 1D: UK


F_2-function vs. time
1.0
0.9
0.8
0.7
0.6
0.5
0.4
0.3
0.2
0.1
0.0
0
100
200
300
400
500
600

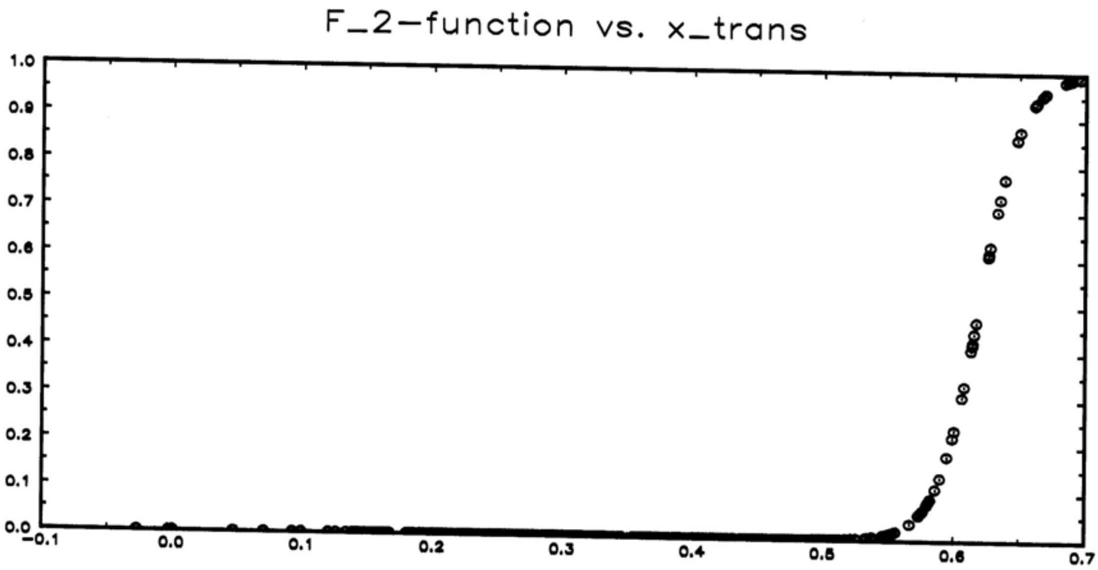
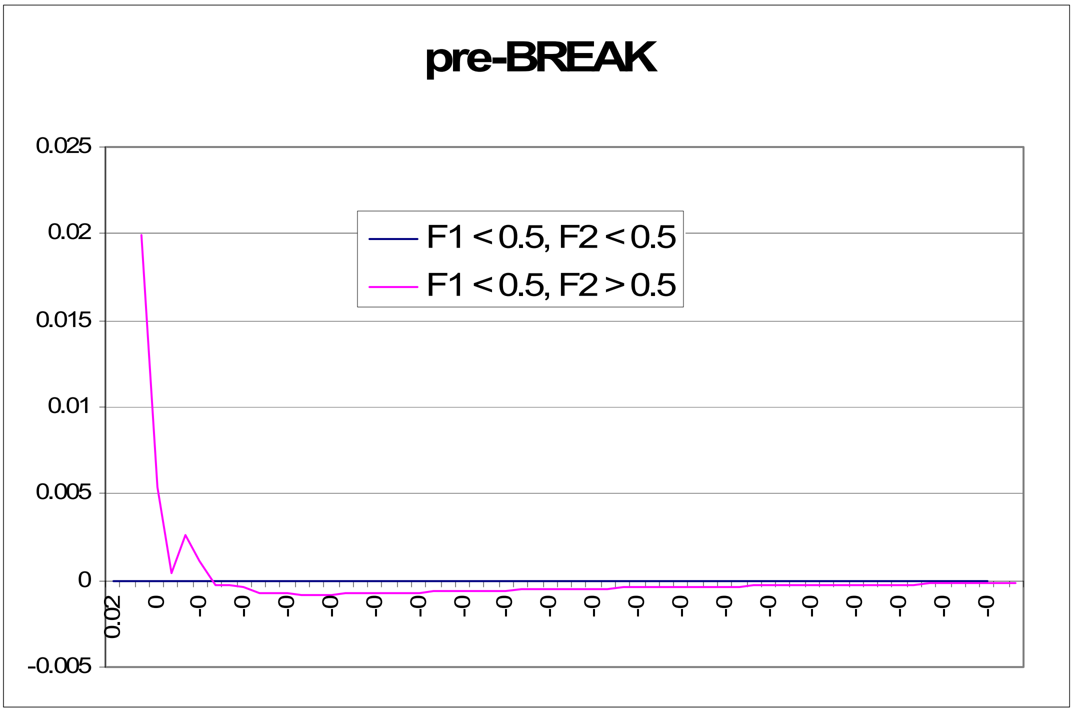

F_2-function vs. x_trans
1.0
0.9
0.8
0.7
0.6
0.5
0.4
0.3
0.2
0.1
0.0
-0.1
0.0
0.1
0.2
0.3
0.4
0.5
0.6
0.7

Figure 2A: France
Generalised Impulse Response Function h = + 1σ



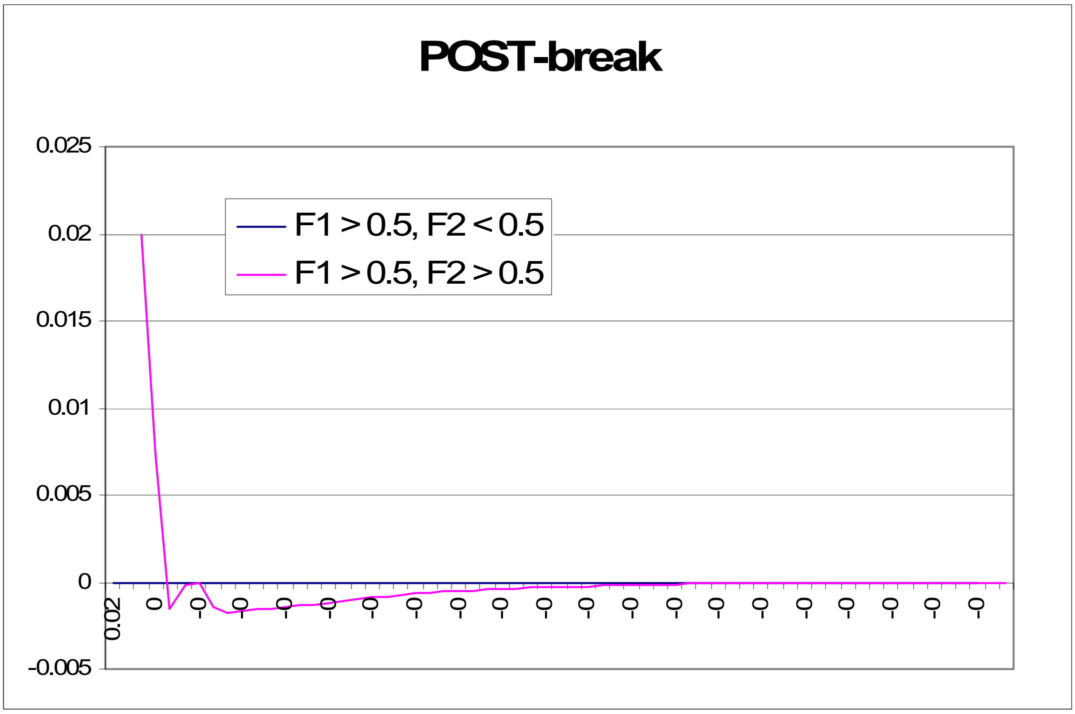
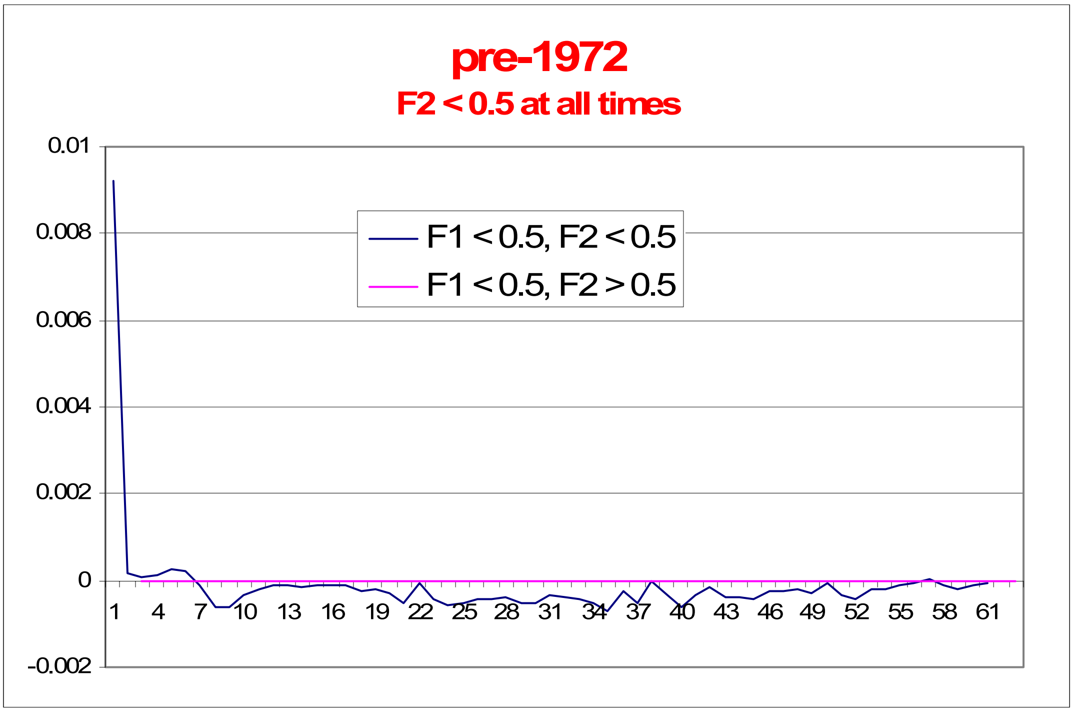


Figure 2B: Germany

Generalised Impulse Response Function h = + 1σ



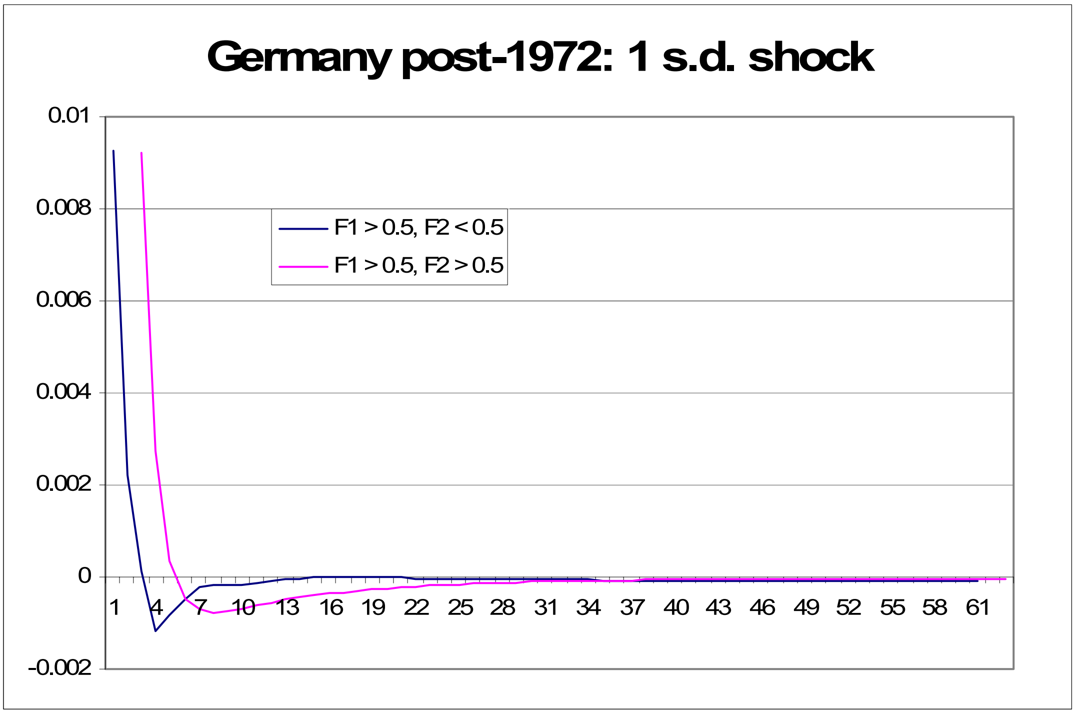
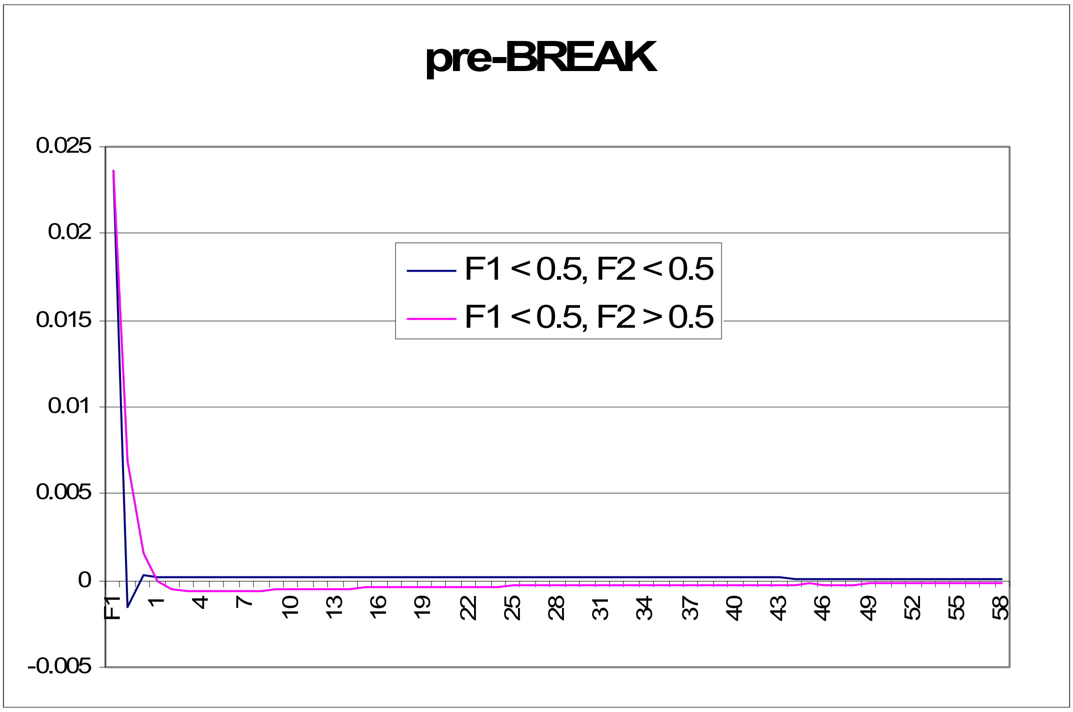


Figure 2C: Japan
Generalised Impulse Response Function h = + 1σ



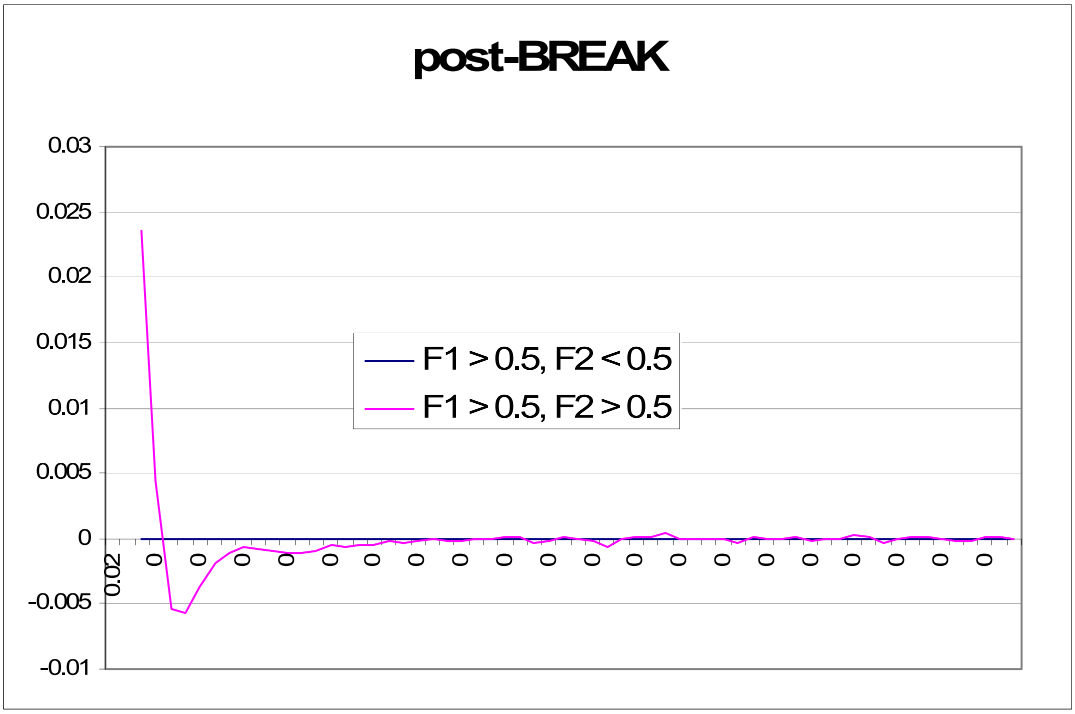
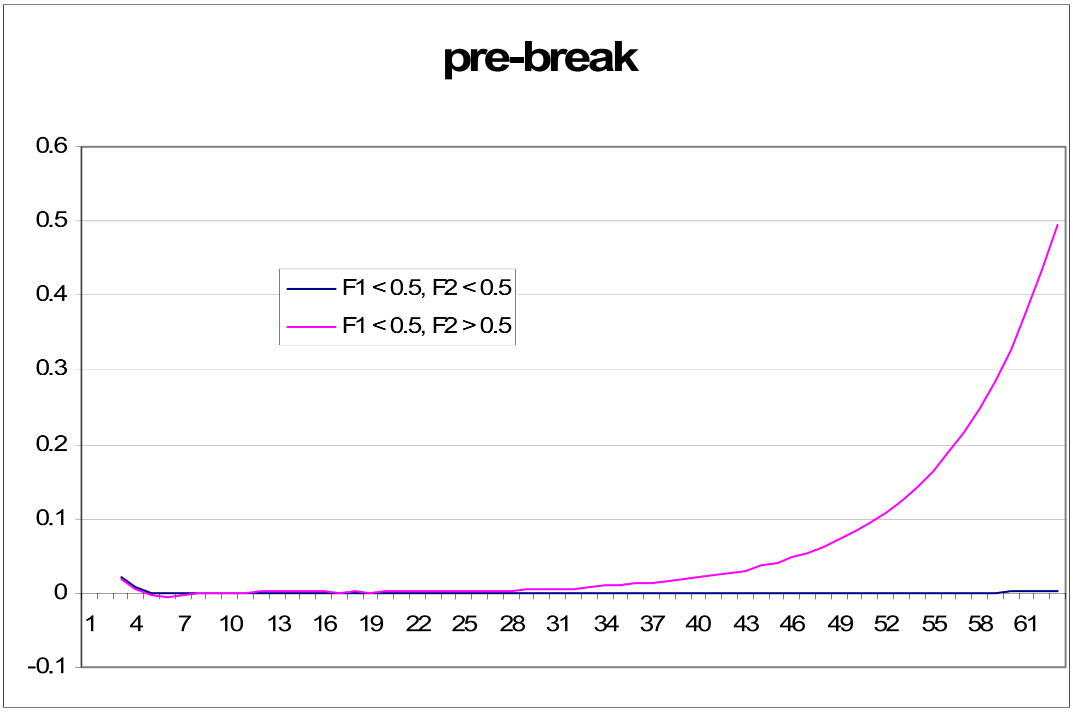


Figure 2D: UK
Generalised Impulse Response Function h = + 1σ



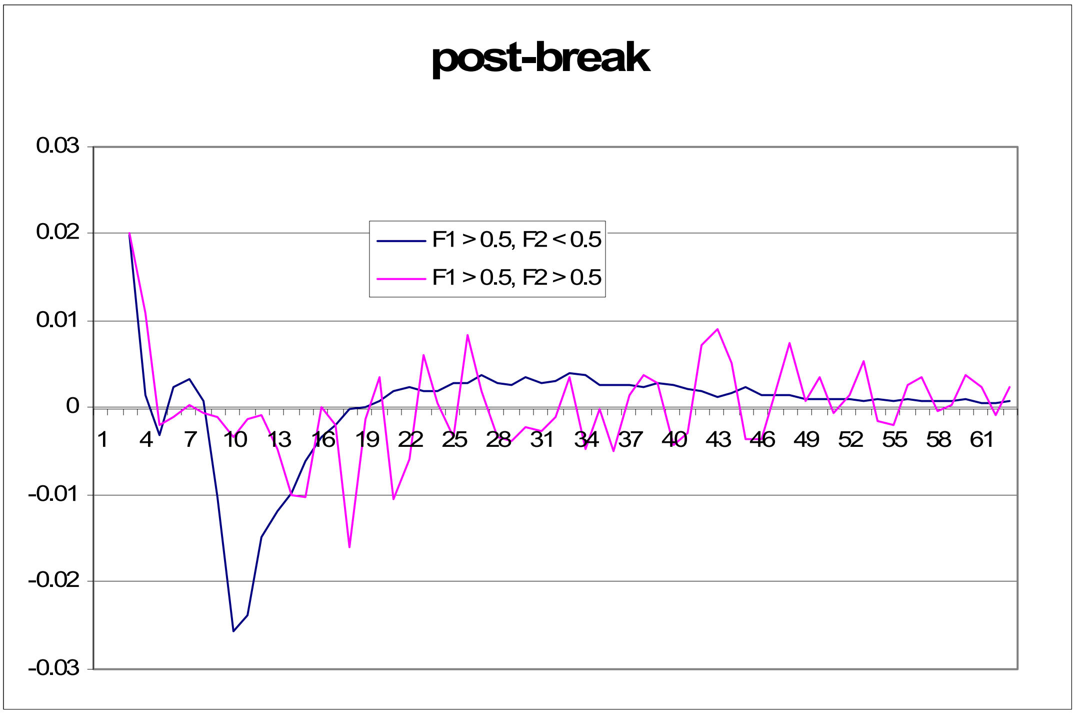


TABLE 1: LINEAR MODELS

Component	France		Germany		Japan		UK	
	Estimate	Robust t-statistic	Estimate	Robust t-statistic	Estimate	Robust t-statistic	Estimate	Robust t-statistic
Constant	-0.029	-2.178	-0.001	-0.787	-0.020	-1.681	0.006	2.176
$\Delta q(t-1)$	0.308	6.134	0.290	4.902	0.285	5.719	10.370	6.021
$\Delta q(t-2)$	-0.090	-1.958					-0.150	-2.778
$\Delta q(t-3)$	0.132	2.809					0.100	1.744
$q(t-1)$	-0.016	-2.236	-0.004	-1.579	-0.004	-1.886	-0.016	-2.032
AIC	-7.692		-9.262		-7.423		-7.729	

p -values

Component	France	Germany	Japan	UK
Q(4)	0.67	0.73	0.14	0.93
Q(8)	0.78	0.88	0.36	0.93
Q(12)	0.92	0.71	0.03	0.51
Stability 1	0.77	0.49	0.84	0.72
Stability 2	0.60	0.66	0.43	0.39
Skewness	0.001	0.028	0.000	0.080
Jarque-Bera	0.0000	0.0000	0.0000	0.0000

TABLE 2: TEST RESULTS

	FRANCE		GERMANY		JAPAN		UK	
	p-values							
d =1	**Standard**	**Robust**	**Standard**	**Robust**	**Standard**	**Robust**	**Standard**	**Robust**
LM 1: γ = 0 (LSTAR) 1st order	0.529	0.739	0.0173	0.0485	0.0323	0.0494	0.894	0.976
LM 2: γ = 0 (ESTAR) 1st order	0.335	0.335	0.0446	0.1430	0.102	0.0615	0.181	0.741
LM 3: γ = 0 (LSTAR) 3rd order	0.362	0.488	0.0763	0.1510	0.201	0.033	0.00148	0.769
LM 3e: γ = 0 (LSTAR) 3rd order*	0.148	0.159	0.0327	0.0913	0.0684	0.103	0.105	0.473
LM 4: γ = 0 (ESTAR) 3rd order	0.0897	0.248	0.0440	0.2200	0.294	0.0483	0.00138	0.846
H4	0.0287	0.16	0.1080	0.1540	0.579	0.651	0.164	0.521
H3	0.404	0.669	0.4350	0.4230	0.664	0.723	0.000431	0.354
H2	0.207	0.503	0.4340	0.4840	0.642	0.733	0.0366	0.469
H1	0.528	0.739	0.0173	0.0485	0.0323	0.0494	0.894	0.976
HL	0.0161	0.351	0.0082	0.0082	0.057	0.00616	0.0448	0.976
HE	0.126	0.389	0.2020	0.1590	0.615	0.533	0.0498	0.731
d = 2: LM 3: γ = 0 (LSTAR) 3rd order	0.234	0.457	0.0882	0.1200	0.21	0.0367	0.00742	0.662
d = 3: LM 3: γ = 0 (LSTAR) 3rd order	0.171	0.443	0.0520	0.0702	0.164	0.0431	0.018	0.71
d = 4: LM 3: γ = 0 (LSTAR) 3rd order	0.304	0.553	0.1460	0.3330	0.281	0.0265	0.0207	0.566
d = 5: LM 3: γ = 0 (LSTAR) 3rd order	0.152	0.376	0.1630	0.3420	0.213	0.0377	0.0184	0.474
d = 6: LM 3: γ = 0 (LSTAR) 3rd order	0.0133	0.273	0.1950	0.3640	0.348	0.101	0.00007	0.215
Jarque-Bera	0.0000		0.0000		0.0000		0.0000	
AIC	-7.692		-9.2620		-7.423		-7.729	
	Coefficient Estimate	**Robust t-ratio**	**Coefficient Estimate**	**Robust t-ratio**	**Coefficient Estimate**	**Robust t-ratio**	**Coefficient Estimate**	**Robust t-ratio**
	p-values		**p-values**		**p-values**		**p-values**	
TV-AR								
γ	78.220	0.658	68.411	0.592	500.000	0.625	500.000	0.617
c	0.580	73.340	0.672	94.322	0.654	590.280	0.490	459.700
AIC	-7.704		-9.2890		-7.434		-7.755	
Stability 1	0.737		0.2000		0.0614		0.945	
Stability 2	0.203		0.0010		0.0033		0.0785	
Skewness	0.002		0.037		0.000		0.062	
LSTAR								
γ	428.990	0.027	168.160	0.186	81.110	0.063	1.057	0.559
c	-1.732	-634.200	-0.293	-70.344	-5.488	-75.820	0.143	0.288
AIC	-7.649		-9.269		-7.420		-7.729	
Stability 1	0.947		0.3930		0.648		0.776	
Stability 2	0.956		0.5090		0.023		0.288	
Skewness	0.00149		0.0170		0.000		0.379	
ESTAR								
γ	27.220	9.011	25.012	10.250	2.507	0.663	2.673	1.706
c	-1.698	-158.550	-0.544	-79.766	-5.906	-10.020	0.191	1.915
AIC	-7.644		-9.2640		-7.418		-7.728	
Stability 1	0.808		0.7570		0.678		0.666	
Stability 2	0.975		0.7250		0.837		0.138	
Skewness	0.00149		0.4900		2.194E-11		0.183	

TABLE 3: TV-STAR ESTIMATES

	France						Japan				UK			
	Full Period		Post-1973		Germany		Full Period		Post-1973		Full Period		Post-1973	
Component	**Estimate**	**Robust t-statistic**	**Estimate**	**Robust t-statistic**	**Estimate**	**Robust t-statistic**	**Estimate**	**Robust t-statistic**	**Estimate**	**Robust t-statistic**	**Estimate**	**Robust t-statistic**	**Estimate**	**Robust t-statistic**
LINEAR														
Constant	0.211	0.438	-0.165	-0.955	-0.002	-0.110	0.041	0.702	-2.339	-0.445	0.004	1.116	0.009	1.154
Δq(t-1)	-0.423	-0.853	0.121	0.560	0.005	0.030	-0.089	-0.906	0.356	0.941	0.432	5.958	0.447	5.436
Δq(t-2)	-0.548	-0.956	-0.265	-1.915							-0.157	-2.898	-0.155	-2.604
Δq(t-3)	-0.602	-1.200	-0.170	-1.003							0.110	1.910	0.111	1.773
q(t-1)	0.119	0.541	-0.074	-0.883	-0.004	-0.150	0.007	0.659	-0.444	-0.448	-0.010	-0.819	-0.019	-0.922
F(t)														
Constant	2.134	0.630	0.512	1.876	-0.044	-1.230	-98392.597	-3.050	-1074.600	-0.302	0.018	1.150	0.011	0.737
Δq(t-1)	-7.141	-0.978	-2.046	-5.240	0.303	1.270	11190.850	0.029	81.534	0.300	-0.326	-2.489	-0.324	-2.656
Δq(t-2)	-5.371	-1.019	-1.424	-6.642										
Δq(t-3)	-3.606	-0.922	-0.906	-2.965										
q(t-1)	0.765	0.546	0.164	1.258	-0.112	-1.330	-19647.867	-0.304	-213.416	-0.302	-0.040	-1.099	-0.027	-0.766
F(q)														
Constant	-0.244	-0.523	0.137	0.787	-5.660	-0.470	-0.059	0.833	1.703	0.340	0.124	0.754	0.070	0.539
Δq(t-1)	0.711	1.376	0.091	0.363	4.068	0.410	0.409	3.591	-0.025	0.060	-0.684	-1.881	-0.564	-2.030
Δq(t-2)	0.515	0.859	0.229	1.311										
Δq(t-3)	0.760	1.523	0.306	1.457										
q(t-1)	-0.137	-0.653	0.058	0.676	-16.331	-0.460	-0.011	-1.055	0.430	0.420	-0.188	-0.777	-0.108	-0.556
F(t)F(q)														
Constant	-2.188	-0.654	-0.610	-2.247	5.690	0.470	98390.707	3.047	1070.470	-0.941	1.975	0.375	4.196	0.249
Δq(t-1)	7.266	0.990	2.241	5.226	-4.022	-0.410	-11191.086	-0.028	-81.840	0.287	16.555	0.470	14.967	0.329
Δq(t-2)	5.245	0.991	1.264	5.101										
Δq(t-3)	3.590	0.906	0.923	2.680										
q(t-1)	-0.796	-0.580	-0.222	-1.695	16.378	0.470	19647.471	0.304	212.570	-0.934	-3.430	-0.389	-6.745	-0.253
γ_1	242.506	1.476	500.000	0.044	23.238	3.450	9.858	3.730	6.957	0.796	23.740	1.742	21.205	0.729
c_1	0.660	534.001	0.458	82.653	0.369	13.700	0.909	19.190	0.877	7.064	0.807	24.260	0.691	18.309
DATE	Jan 1985		Nov 1984		Jan 1973		Sep 1997		Feb 1998		Mar 1993		Nov 1992	
γ_2	2.600	2.983	9.023	0.744	7.545	2.610	7.364	2.810	2.799	1.548	8.892	1.742	16.976	0.567
c_2	-2.101	-23.338	-1.945	-57.636	-0.336	-12.170	-5.508	-47.430	-5.301	-18.023	0.619	34.370	0.608	53.533
AIC	-7.743		-7.393		-9.319		-7.443		-7.111		-7.770		-7.424	
Diagnostic Tests (Robust p-values)														
Q(4)	0.44		0.33		0.35		0.06		0.05		0.59		0.660	
Q(8)	0.79		0.45		0.72		0.20		0.13		0.55		0.640	
Q(12)	0.79		0.52		0.63		0.04		0.07		0.28		0.280	
Stability 1	0.86		0.51		0.31		0.56		0.72		0.71		0.960	
Stability 2	0.52		0.58		0.54		0.06		0.50		0.37		0.240	

TABLE 4: REGIMES

Break at c_1	FRANCE Jan-85	GERMANY Sep-72	JAPAN Oct-97	UK Mar-93
		Pre-break:	**t << c_1**	
Low q << c_2				
Constant	0.211	-0.002	0.041	0.004
Δq(t-1)	-0.423	0.005	-0.089	0.432
Δq(t-2)	-0.548			-0.157
Δq(t-3)	-0.602			0.110
q(t-1)	0.119	-0.004	0.007	-0.010
Long run level of $q = q^*$	-1.773	-0.432	-5.843	0.410
High q >> c_2				
Constant	-0.033	-5.661	-0.018	0.128
Δq(t-1)	0.288	4.073	0.320	-0.252
Δq(t-2)	-0.033			-0.157
Δq(t-3)	0.158			0.110
q(t-1)	-0.018	-16.335	-0.004	-0.198
Long run level of $q = q^*$	-1.833	-0.347	-5.076	0.647
Steady state q* = c_2				
Constant	0.089	-2.831	0.011	0.066
Δq(t-1)	-0.068	2.039	0.116	0.090
Δq(t-2)	-0.291			-0.157
Δq(t-3)	-0.222			0.110
q(t-1)	0.051	-8.169	0.002	-0.104
Long run level of $q = q^*$	-1.762	-0.347	-6.632	0.636
		Post-break:	**t >> c_1**	
Low q << c_2				
Constant	2.345	-0.046	-98392.556	0.022
Δq(t-1)	-7.564	0.308	11190.761	0.106
Δq(t-2)	-5.919			-0.157

Δq(t-3)	-4.208			0.110
q(t-1)	0.884	-0.116	-19647.860	-0.050
Long run level of *q* = *q**	-2.653	-0.393	-5.008	0.444
Change in *q** after shift	-87.961%	3.899%	83.506%	3.378%
High q >> c_2				
Constant	-0.087	-0.016	-1.908	2.121
Δq(t-1)	0.413	0.354	0.085	15.977
Δq(t-2)	-0.159			-0.157
Δq(t-3)	0.142			0.110
q(t-1)	-0.049	-0.069	-0.400	-3.668
Long run level of *q* = *q**	-1.776	-0.225	-4.775	0.578
Change in *q** after shift	5.782%	12.193%	30.099%	-6.867%
Steady state q* = c_2				
Constant	1.129	-0.031	-49197.232	1.072
Δq(t-1)	-3.576	0.331	5595.423	8.042
Δq(t-2)	-3.039			-0.157
Δq(t-3)	-2.033			0.110
q(t-1)	0.418	-0.092	-9824.130	-1.859
Long run level of *q* = *q**	-2.704	-0.330	-5.008	0.577
Change in *q** after shift	-94.18%	1.61%	162.41%	-5.91%

TABLE 5: HALF-LIVES

GIRF's with bootstrapped errors for *t+1….t+60*

Shock size	France	Germany	Japan	UK
+1σ/-1σ				
Whole period	29/n.a.	25/25	>60/>60	n.a./n.a.
Pre-break	38/n.a.	24/25	>60/>60	n.a./n.a.
Post-break	15/10	26/26	4/4	n.a./n.a.
+4σ/-4σ				
Whole period	>60/n.a	18/32	>60/>60	n.a./n.a.
Pre-break	>60/n.a	16/31	>60/>60	n.a./n.a.
Post-break	12/6	23/34	6/7	n.a./48