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ABSTRACT

Is Official Exchange Rate Intervention Effective?*

I examine the effectiveness of exchange rate intervention within the context of a Markov-switching model for the real exchange rate. The probability of switching between stable and unstable regimes depends non-linearly upon the amount of intervention, the degree of misalignment and the duration of the regime. Applying this to dollar-mark data for the period 1985-98, I find that intervention increases the probability of stability when the rate is misaligned, and that its influence grows with the degree of misalignment. Intervention within a small neighbourhood of equilibrium will result in a greater probability of instability.

JEL Classification: C10, F31 and F41
Keywords: mean reversion, non-linearity, official intervention and real exchange rate

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1 Introduction

Official exchange rate intervention in the foreign exchange market occurs when the authorities buy or sell foreign exchange, normally against their own currency and in order to affect the exchange rate. Given the policy importance of the issue of whether or not intervention is effective, it is not surprising that a large empirical literature has grown up examining this issue. In a recent survey of this literature, Sarno and Taylor (2001) conclude that the evidence on the effectiveness of official intervention is still somewhat mixed. Nevertheless, as these authors argue, it is tempting to view studies of intervention operations published during the 1990s and later, which are largely supportive of the effectiveness of intervention, as more convincing than previous studies, which largely find little support for the effectiveness of intervention, since the later studies were able to employ better data denied earlier researchers - in particular data on intervention itself.[1] Alongside the later literature on intervention, a literature has also emerged which appears to demonstrate that there are important nonlinearities in exchange rate movements.[2] In this paper, I report research which utilises actual intervention data by the US and German authorities and uses an explicitly nonlinear framework to analyse the effectiveness of intervention operations.

In particular, I examine the effectiveness of intervention within the context of a regime-switching model for the real exchange rate. The real exchange rate is postulated to be governed by two regimes, one of which displays unstable, random-walk behavior and the other of which displays stable, mean-reverting behavior. The probability of switching between regimes is assumed to be endogenous and dependent on a number of factors. One of these factors is the deviation of the real exchange rate from its long-run mean - as suggested by recent work on nonlinearities in real and nominal exchange rate movements, the real exchange rate is likely to behave more like a random walk the closer it is to its long-run equilibrium, and is more likely to mean revert as it moves further away from that equilibrium. A second factor I allow for is the *duration* of the regime, on the supposition that it is well known that trend-following techniques are widely employed by technical analysts in the major foreign exchange markets.[3] Thirdly, however, I also analyze the influence of the degree of intervention on the probability of switching between regimes, and in particular on the probability of switching from the non-mean reverting regime to the mean-reverting regime. Further, I test whether this influence of intervention on the probability of mean reverting becomes more marked as the deviation of the real exchange rate from equilibrium becomes more marked.

The key innovations of the present analysis are thus as follows. First, I judge the effectiveness of intervention by analyzing whether it increases the *probabil-*

[1] See, for example, the influential studies of Dominguez and Frankel (1993a, 1993b, 1993c).

[2] See, for example, Peel and Speight (1997); Coakley and Fuertes (2000, 2001); Taylor, Peel and Sarno (2001); Kilian and Taylor (2002); and Sarno and Taylor (2002). For more general issues on nonlinear estimation in macroeconometrics, see, e.g. Peel and Speight (1998) and Coakley, Fuertes and Perez (2001). For a survey treatment, see Sarno and Taylor (2003).

[3] See, in particular, Frankel and Froot (1986, 1990a, 1990b), Allen and Taylor (1990) and Taylor and Allen (1992).

ity of exchange rate stability. Given the volatility of international financial markets, this seems to be a reasonable criterion. Second, I analyze stability with respect to simple, purchasing power parity (PPP) fundamentals by seeing whether intervention tends to induce stability in the *real* exchange rate. Third, I examine nonlinearities in the relationship: does intervention become more effective the further the real exchange rate is from its equilibrium level?

The remainder of the paper is set out as follows. In the following section I describe the econometric model while in Section 3 I briefly describe the data. The estimation results are reported and discussed in Section 4 and I make some concluding comments in a final section.

2 A Markov-Switching Model of the Effectiveness of Intervention

The real exchange rate, q_t, may be expressed in logarithmic form as

$$q_t \equiv e_t - p_t + p_t^* \tag{1}$$

where e_t is the logarithm of the nominal exchange rate (domestic price of foreign currency), and p_t and p_t^* denote the logarithms of the domestic and foreign price levels respectively. A Markov-switching first-order autoregressive, MS-AR(1), model for the real exchange rate, where the autoregressive parameter ϕ_i $(i = 1, 2)$ takes on one of two values depending on the realization of a discrete-valued unobserved state variable S_t, may be written:

$$q_t = \phi_{S_{t-1}} q_{t-1} + \varepsilon_t \tag{2}$$

with $\varepsilon_t \sim N(0, \sigma^2)$, $S_t \in \{1, 2\}$, $\phi_1 = 1$, $|\phi_2| < 1$. I assume further that the state variable S_t evolves according to a two-state Markov process. Thus state 1 corresponds to random-walk behavior and state 2 to mean-reverting behavior (assuming that the mean of the logarithm of the real exchange rate has been normalized to zero).[4]

A key component of the approach taken here is that the probabilities of switching from one regime to another are modelled endogenously: by modelling the transition probabilities as a function of, *inter alia*, regime duration,

[4] A variation of this model would be to allow for shifts in volatility across regimes. This would be particularly attractive in the case of nominal exchange rates at, say, the weekly frequency, since evidence of fat tails in the distribution of nominal exchange rate innovations at relatively high frequencies (which is often detected as autoregressive conditional heteroscedasticity or ARCH effects) is quite common, and this might be modelled in a Markov-switching framework (see e.g. Clarida, Sarno, Taylor and Valente, 2002, for a recent example of exactly such a modelling framework). Researchers have generally found much less evidence of ARCH effects or fat tails in the distribution of innovations to *real* exchange rates at the monthly frequency, however, and formal tests for ARCH effects with our real exchange rate data could not reject the hypothesis of homoscedasticity at standard levels of significance, either in terms of the log-changes of the real exchange rate or in terms of the innovations from the estimated Markov-switching model (see Table 1).

I allow the state transition probabilities to be functions of both the inferred current state and the associated number of periods the system has been in the current state. In particular, following Durland and McCurdy (1994), I define the duration of the current regime, $D(S_t)$, to be the length of a run of consecutive states, so that

$$D(S_t) = \begin{cases} D(S_{t-1}) + 1 & \text{if } S_t = S_{t-1} \\ 1 & \text{otherwise} \end{cases}. \tag{3}$$

The effect of duration on the transition probabilities allows the model to capture positive feedback or momentum effects in the foreign exchange market.[5]

In the light of recent evidence of nonlinearity in real exchange rate behavior (Peel and Speight, 1997; Taylor, Peel and Sarno, 2001; Kilian and Taylor, 2002), whereby the degree of mean reversion of the real exchange rate appears to be positively related to the degree of misalignment, I also allow the transition probabilities to be a function of the deviation from equilibrium - or more precisely the sample mean of the real exchange rate with the implicit assumption that the mean value is a reasonable approximation to the long-run equilibrium level. I allow intervention to enter the transition probability functions in two ways. I include the lagged value of an index of intervention activity I_{t-1} and an interaction term formed by taking the product of lagged intervention with the lagged deviation of the real exchange rate from its mean.

Having decided on the variables for which I will test for an effect on the transition probabilities, then following Filardo (1994), Durland and McCurdy (1994), and Diebold, Lee and Weinbach (1994), I employ a logistic function to ensure that the probabilities lie in the unit interval. Thus, if we denote the transition probability of switching from regime j to regime i at time t as p_t^{ij} for $i, j \in \{1, 2\}$, then we can write the postulated functions as:

$$\begin{aligned} p_t^{i1} &\equiv \Pr[S_t = 1 | S_{t-1} = i,\, D(S_{t-1}),\, q_{t-1},\, I_{t-1}] \\ &\equiv \frac{\exp[\beta_{i1} + \beta_{i2}|q_{t-1}| + \beta_{i3} I_{t-1} + \beta_{i4} I_{t-1}|q_{t-1}| + \beta_{i5} D(S_{t-1})]}{1 + \exp[\beta_{i1} + \beta_{i2}|q_{t-1}| + \beta_{i3} I_{t-1} + \beta_{i4} I_{t-1}|q_{t-1}| + \beta_{i5} D(S_{t-1})]}, \end{aligned} \tag{4}$$

for $D(S_{t-1}) \leq \tau$, and

$$\begin{aligned} p_t^{i1} &\equiv \Pr[S_t = 1 | S_{t-1} = i,\, D(S_{t-1}),\, q_{t-1},\, I_{t-1}] \\ &\equiv \frac{\exp(\beta_{i1} + \beta_{i2}|q_{t-1}| + \beta_{i3} I_{t-1} + \beta_{i4} I_{t-1}|q_{t-1}| + \beta_{i5} \tau)}{1 + \exp(\beta_{i1} + \beta_{i2}|q_{t-1}| + \beta_{i3} I_{t-1} + \beta_{i4} I_{t-1}|q_{t-1}| + \beta_{i5} \tau)}, \end{aligned} \tag{5}$$

for $D(S_{t-1}) > \tau$, where the β_{ij}s denote unknown parameters. While this defines p_t^{11} and p_t^{21}, clearly we also have the implicit definitions $p_t^{12} \equiv 1 - p_t^{11}$ and $p_t^{22} \equiv 1 - p_t^{21}$.

[5] Duration dependence has previously been applied to stock market returns data by McQueen and Thorley (1994) and Cochran and Defina (1995).

3 Data

The data set comprises monthly observations on consumer price indices for the US and Germany, end-of-period spot exchange rates for the German mark against the US dollar and daily intervention data in the dollar-mark foreign exchange market by the US and German authorities. All data cover the sample period from August 1985 through to December 1998. The real exchange rate series was constructed with these data in logarithmic form as in equation (1), with e_t taken as the logarithm of the dollar price of marks, p_t as the logarithm of the US consumer price level and p_t^* as the logarithm of the German consumer price level. The (log) real exchange rate series was then normalized by subtracting out the mean of the sample. Daily data on intervention in the dollar-mark exchange rate market - the sale or purchase of dollars against marks, expressed in millions of dollars (negative amounts indicating sales) - by the US and German authorities over the same sample period was very kindly supplied by the Federal Reserve System and by the Bundesbank.[6] A monthly index of real intervention activity operations was constructed from this data as follows. First, the total amount of intervention for each month undertaken by both authorities was added together to obtain a series for net intervention by the US and German authorities, and this was then deflated by the US consumer price index to express it in real terms. I then took the logarithm of the absolute value of the total real dollar interventions in each period and scaled the resulting series to map it on to the unit interval.[7] The series for the real exchange rate and the intervention index are graphed in Figures 1 and 3 respectively. An interesting feature of Figure 3 is that it shows a relative decline in the frequency of intervention operations during the 1990s and, indeed, that intervention operations appear to have ceased altogether from late 1995 until the end of the sample period.

4 Estimation Results

Estimation of the model was carried out using a modified version of the EM algorithm due to Diebold, Lee and Weinbach (1994). Parameter estimates, estimated standard errors and diagnostics for the model are given in Table 1. Since the state variable S_t is unobservable, it follows that the residuals from the fitted model are also unobservable. However, following Maheu and McCurdy (2000), I constructed the standardized expected residuals as a weighted average of the residuals obtaining under each regime, with the weights equal to the smoothed probabilities of being in each regime (i.e. the probabilities constructed using the whole sample data). These expected residuals may then be

[6] I am grateful to Chris Neely for help in obtaining the US data.

[7] I also experimented with alternative measure of intervention activity. In the first of these I used an index of *nominal* rather than real intervention. In the second, I deflated the nominal series by the German price level with an exchange rate adjustment, i.e. (in logarithms) $\widetilde{p_t} = e_t + p_t^*$. In each case, qualitatively similar results to those reported below were obtained.

used to construct residual-based diagnostic tests such as the Ljung-Box portmanteau test which applied to the residuals for j autocorrelations may be used to test for residual autocorrelation up to j-th order $[Q(j)]$ and applied to the squared residuals for j autocorrelations may be used to test for autoregressive heteroscedasticity of up to j-th order $[Q^2(j)]$. Applying these tests to the standardized expected residuals from the estimated model in neither case suggests misspecification (Table 1).[8]

The smoothed probabilities of the mean reverting regime - shown in Figure 2 - illustrate the ability of the model to identify long-swings in the dollar-mark real exchange rate. They show the period since the Plaza Accord as being characterized by a series of periods of non-stationarity, followed by sharp mean reversions. In particular it is possible to identify three periods of prolonged non-stationary behavior: one from January to December 1987, another from October 1989 until September 1990, and a third from July 1993 until April 1995.[9]

The endogenous mean-reverting properties of the real exchange rate are captured by the estimates of the parameters β_{12} and β_{22}, which measure the effects of deviations from equilibrium on the transition probabilities. The negative and statistically significant[10] value of $\widehat{\beta}_{12}$ indicates that the probability of remaining in the unstable regime falls as the real exchange rate moves further from equilibrium.[11] The large negative and strongly statistically significant value of $\widehat{\beta}_{22}$ has the effect that far from equilibrium the probability of switching from the stable regime to the unstable regime is small and increases as the exchange rate approaches its equilibrium level. Thus, the model is able to capture the notion that once the real exchange rate embarks on a mean-reverting path toward equilibrium, it is likely to continue mean reverting until it draws near to equilibrium.

The effects of duration on the transition probabilities are summarized by the coefficients β_{15} and β_{25}. The positive value of $\widehat{\beta}_{15}$ indicates negative duration dependence, i.e. the longer the real exchange rate persists in the *unstable* regime, the higher the probability of remaining in that state, suggesting a modicum of hysteresis or path dependence in the real exchange rate, although the fact that the equilibrium in the stable regime remains constant means that this effect will

[8] Note that these tests should be treated with caution however, since the asymptotic distribution of the Ljung-Box statistic is unknown for standardized expected residuals. Nevertheless, given the high p-values associated with the statistics we may perhaps safely conclude that no strong misspecification is suggested.

[9] As a supplementary check on these results, I applied the methods of Bai and Perron (1998, 2000, 2001a, 2001b) to test for multiple structural breaks in the $AR(1)$ process for q_t, with allowance for up to eight breaks. The results were, however, only partially in accordance with the identification of periods of prolonged non-stationarity suggested by the smoothed estimates of the transition probabilities. In particular, results obtained using the sequential $supF_T$ method of Bai and Perron (1998) suggested the presence of two significant breaks, in May 1993 and in June 1995. The autoregressive parameter was estimated at 0.903 prior to May 1993, then jumps to 1.002 and moves down to 0.910 in June 1995. (I am grateful to Jushan Bai and Pierre Perron for use of their *Gauss* program to obtain these results.)

[10] In the following discussion I implicitly use a five percent nominal test size.

[11] Circumflexes are used to indicate estimates of the parameters below the cirumflex.

disappear once the real exchange rate enters the mean-reverting regime; hence, it is perhaps best thought of as a kind of 'quasi-hysteresis'. Moreover, the estimated value of β_{25} is significant and quite large, so that the probability of switching into the random-walk state from the mean-reverting state increases as the duration of the mean-reverting state rises.

The effects of intervention on the transition probabilities are captured by the parameters β_{13}, β_{14}, β_{23} and β_{24}. The point estimate for β_{14} of -9.487 implies that intervention tends to increase the probability of switching into the stable regime when intervention is conducted at levels of the real exchange rate which are far from equilibrium. Note, however, that the small but significant value of $\widehat{\beta}_{13}$ is positive, indicating that, under certain conditions, intervention may actually *increase* the probability of remaining in the unstable regime next period. Whether or not this effect is realized depends on the size of the deviation from equilibrium. If the real exchange rate is *at,* or very close to equilibrium, the interaction term $\beta_{14}I_{t-1}|q_{t-1}|$ will be very small so that the net effect of the intervention will be to *increase* the probability of remaining in the unstable regime. The level of the real exchange rate at which this effect starts to occur is $|q_t| = -\beta_{13}/\beta_{14}$, and the point estimate of this threshold is 0.0355. Thus, conditional on the process being in state 1, interventions conducted when $|q_t| < 0.0355$ - i.e. when the real exchange rate is within about 3.5 percent of its equilibrium value - actually *increase* the probability of remaining in state 1. One way to interpret this is as follows. When the exchange rate is heavily misaligned, there will tend to be a greater degree of consensus among both policy makers and market participants concerning the *direction* in which it would be appropriate to engineer an exchange rate movement. In contrast, as the equilibrium is approached, there will be less consensus both in the market and among policy makers and advisers concerning the appropriate direction of exchange rate movements, and this will be reflected both in the nature of intervention operations and in the way intervention affects market behavior. Moreover, because the degree of consensus among traders and advisers who rely mainly on economic fundamentals is likely to be smaller for smaller deviations from equilibrium, the greater will be the influence of non-fundamentalist traders and advisers such as technical analysts, who may impart unstable behaviour into the market (see Kilian and Taylor, 2002, for further arguments along these lines).

A further policy implication of these results, therefore, is that attempting to 'fine tune' the exchange rate through intervention operations may be counterproductive.

With respect to β_{23} and β_{24}, estimates of both of these coefficients were found to be insignificantly different from zero at the five percent level in initial estimations and so were set to zero. Since the strongest effects of intervention would be expected to be with respect to the probability of switching from random-walk behaviour to mean-reverting behaviour rather than *vice versa*, this is perhaps not surprising.

5 Conclusion

In this paper I have provided some further evidence on the effectiveness of official exchange rate intervention operations by estimating a Markov-switching model of the real exchange rate in which the transition probabilities of switching between stable and unstable regimes are allowed to depend upon official intervention activity, as well as the extent of exchange rate misalignment and on the duration of the regime.

In common with an emerging parallel literature on nonlinearities in real exchange rate movements, the probability of switching from the unstable regime into the stable regime increases as real exchange rate deviates further from its equilibrium value and the size of misalignment grows, and the probability of switching into the unstable state from the stable state is higher the closer the real exchange rate is to its equilibrium level. The literature has tended to rationalize nonlinearities of this kind on the basis of arguments concerning transactions costs in international goods arbitrage (see Taylor, Peel and Sarno, 2001, for a discussion of this literature). More recently, however, Kilian and Taylor (2002) have suggested that nonlinearities in both the nominal and the real exchange rate may be due to diversity of opinion in the foreign exchange market, since as the exchange rate deviates further from the equilibrium level a greater degree of consensus will develop as to the appropriate direction of exchange rate movements, even if agents disagree about the equilibrium level.

I also discovered an important duration effect in the real exchange rate series: the longer the real exchange rate has been in the unstable regime, the more likely it is to remain in that regime. This may reflect 'band wagon' or herding effects which may generate irrational bubbles in the market, and which in turn may reflect prevalence in the use of trend-following techniques by technical or chartist analysts (Frankel and Froot, 1986, 1990a, 1990b; Allen and Taylor, 1990; Taylor and Allen, 1992)[12].

Perhaps most interestingly, however, I found that the probability of switching from the unstable state to the stable state is a function of the level of intervention activity. Once the real exchange rate is greater than about 3.5 percent from its equilibrium level, not only does intervention tend to increase the probability of exchange rate stability, its influence in this respect also tends to grow as the deviation from equilibrium grows. On the other hand, intervention operations when the real exchange rate is within a small neighborhood - ± 3.5 percent - of its equilibrium, intervention will result in a greater probability of persistence of the unstable regime.

The policy implications of these findings are interesting. First, interventions designed to 'fine tune' the level of the real exchange rate may turn out to be counter productive, perhaps because both policy makers and market participants will find it hard to agree precisely on the appropriate level of the exchange rate. Second, however, I have uncovered strong evidence that intervention will tend to be more effective the greater is the degree of misalignment, perhaps

[12] See also Menkhoff (1997), Vigfusson (1997), Levin (1997), Curcio *et al.* (1997) and Taylor (1997).

because there will in this case be a high degree of consensus concerning the appropriate direction of exchange rate movement required to move back towards equilibrium.

These findings may also be taken as indirect evidence supportive of the presence of a 'coordination channel' for official intervention as suggested by Sarno and Taylor (2001), whereby intervention operations may be seen as fulfilling a coordinating role in that they may organize the 'smart money' to enter the market at the same time. The further the exchange rate is from its equilibrium level, the more likely it is that many 'smart speculators' will have experienced substantial losses by relying on the fundamentals, thereby increasing their reluctance to enter the market in an uncoordinated fashion (see Shleifer and Vishny, 1997). Moreover, the further the rate is from the equilibrium, the more speculators there will be who agree on the appropriate *direction* of the exchange rate change needed in order to restore stability. Hence, the greater the degree of misalignment, the greater the opportunity for official intervention to play a coordinating role - and this appears to be what is being picked up in the present analysis.

Further work might usefully be addressed to separating out the effects of publicly announced and concerted intervention. If indeed we are picking up evidence of a coordination channel, we should find that publicly announced intervention in particular is effective when the exchange rate is misaligned.

Finally, note that the effectiveness of intervention which I have apparently highlighted in this research must be qualified. As noted by Sarno and Taylor (2001), for example, it is obvious that intervention operations must ultimately be doomed to failure if they are inconsistent with the underlying stance of monetary and fiscal policy and other economic fundamentals. Moreover, a feature of the present analysis is that I only claim that intervention when the exchange rate is clearly misaligned will tend to increase the *probability* of exchange rate stability. Nevertheless, the results do appear to support the view of Dominguez and Frankel (1993a, p.160) concerning the effectiveness of intervention operations: *'It may be that .... intervention can only have effects in the short term. But if "short-term effects" include the bursting of a nine-month bubble earlier than it would otherwise have burst, then such an effect may be all that is needed.'*

Table 1: Estimated Parameters for the Markov-Switching Model

Parameter	Estimate	Standard Error
$\boldsymbol{\phi}_2$	0.906	0.0020
$\boldsymbol{\sigma}$	0.034	0.0013
$\boldsymbol{\beta}_{11}$	−0.586	0.1235
$\boldsymbol{\beta}_{12}$	−1.057	0.2201
$\boldsymbol{\beta}_{13}$	0.337	0.0666
$\boldsymbol{\beta}_{14}$	−9.487	1.0302
$\boldsymbol{\beta}_{15}$	0.316	0.0236
$\boldsymbol{\beta}_{21}$	0.0	
$\boldsymbol{\beta}_{22}$	−83515.490	1024.6270
$\boldsymbol{\beta}_{23}$	0.0	
$\boldsymbol{\beta}_{24}$	0.0	
$\boldsymbol{\beta}_{25}$	4429.783	3.071
$\boldsymbol{\tau}$	10	
$\mathbf{Q(6)}$	7.030	(0.318)
$\mathbf{Q^2(6)}$	0.964	(0.986)

Notes: $Q(6)$ and $Q^2(6)$ denote the Ljung-Box statistic applied to the standardized expected residuals and the squared standardized expected residuals respectively, with p-values given in parentheses. Where a parameter estimate is entered as zero, this indicates that the estimated parameter was originally estimated as insignificantly different from zero at the five percent level and was therefore set to zero in subsequent estimation.

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Figure 1: Real Dollar Mark Exchange Rate

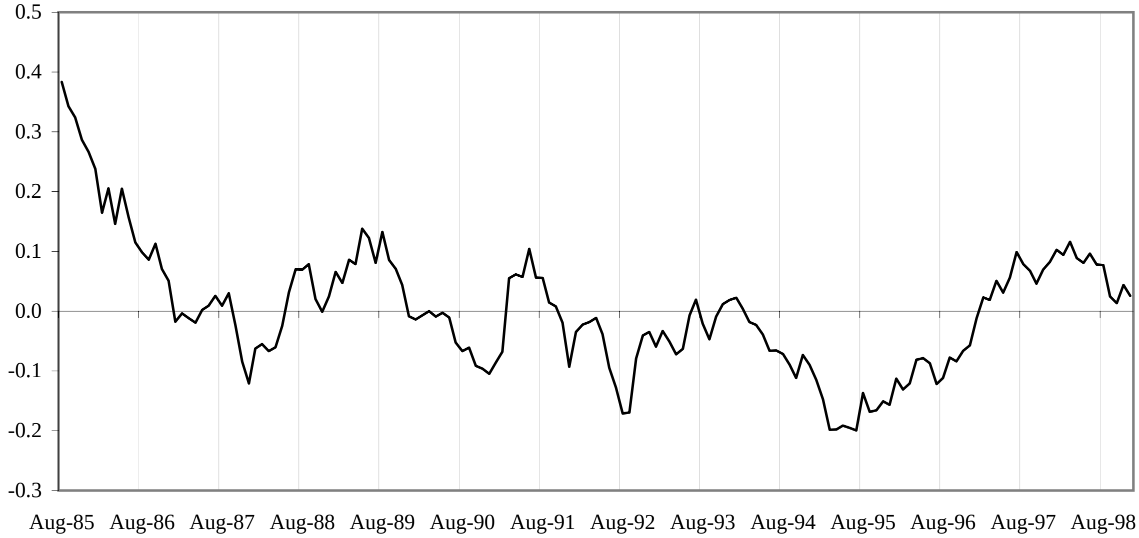

Figure 2: Smoothed Probability of Mean Reversion

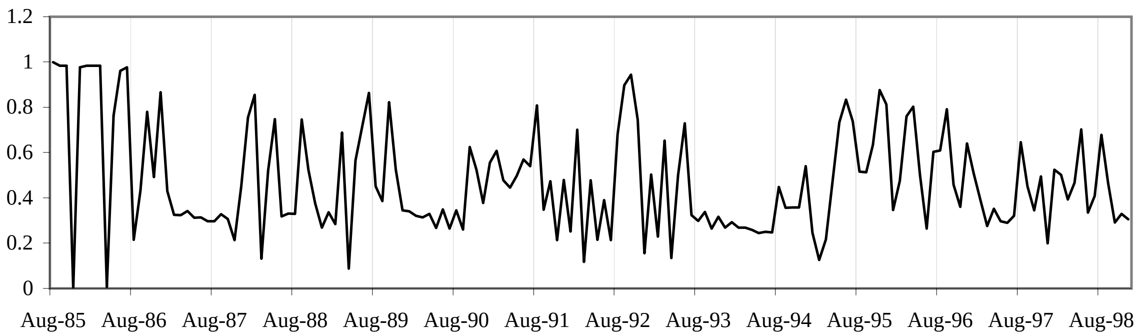

Figure 3: Intervention Index