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German Unification and the EMS: A Non-Parametric Approach to the Asymmetry Question

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Abstract

This paper attempts to analyze the impact of German unification on the position that Germany holds within the European Monetary System (EMS), i. e. on the asymmetry in the EMS. Applying kernel estimation and bootstrap methods we can corroborate the result of previous studies that Germany's position within the EMS was not weakened by German unification when speculative periods before and after German unification are compared. Moreover, our data exhibit significant non-linearities pointing to threshold cointegration or regime switching processes.

Keywords: European Monetary System (EMS), German unification, asymmetry, interest rates, speculative crises, kernel regression estimation, bootstrap.

JEL classification: F33, E43, C14.

Chapter 1

Introduction

The question whether the European Monetary System (EMS) is centered on the deutsche mark has been the object of various studies in the last years. Recently several new issues emerged in this connection. First, as plans for European Monetary Union (EMU) became material a debate about the monetary strategy for the European Central Bank and the transition to EMU started. One argument in this literature is that under the current arrangement only the stability of German money demand matters if Germany targets a monetary aggregate and the other EMS member countries target the exchange rate to the deutsche mark, i. e. if the EMS functions asymetrically. Hence, because of German money demand stability some authors, e. g. Fels and Mayer [1995], favour the current arrangement over targeting a European monetary aggregate.[1]

Second, German unification[2] is expected to have altered the functioning of

[1] Note that, once EMU is working, this line of argument is still valid for the relations between the EMS countries participating in the monetary union and the other member states which will not have met the Maastricht criteria and maintain a fixed parity to the newly created common currency.

[2] More particularly, we mean the German Economic, Monetary and Social Union (GEMSU) of 1 July, 1990. Another interesting issue in this context is the effect that the widening of the fluctuation bands from $\pm 2.25\%$ to $\pm 15\%$ in August 1993 had on the functioning of the EMS. Due to the few observations after that date we are unable to study this question with the non-parametric methods applied in this paper. See Henry and Weidmann [1995c] for a closer focus on the 1992-1993 crises and their impact on asymmetry.

the EMS. However, little research has been undertaken on that topic until now. Using simulation techniques VLAAR [1994] shows that under the current arrangement the raise in German interest rates due to unification induced nonoptimal interest rates in the other member states. Given the recession at that time the markets began to question the credibility of these countries' commitment to maintain a fixed parity to the Deutsche mark. These tensions then led to the EMS crises of 1992/93 and the de facto suspension of the exchange rate mechanism (ERM) in August 1993. To derive these results, however, the author must suppose that German leadership in the EMS exists and is not attenuated by German unification. Yet, several empirical studies suggest a regime shift to have occured at this date: GARDNER and PERRAUDIN [1993] and HENRY and WEIDMANN [1995a,b] analyze the impact of German unification on the asymmetry in the EMS by applying impulse-response and causality analysis to pre- and post-unification sub-samples. While the former find that in the long-run Germany's position in the EMS is not weakenend the latter state that asymmetry has even grown after 1990.

The purpose of this paper is to provide additional empirical evidence on the question whether Germany's leadership role has eroded or grown in the years since German unification. Furthermore, we ask whether the apparent break in the system's structure, i. e. the twist towards more asymmetry, is due directly to the shock associated with German unification or rather indirectly via the 1992/93 EMS crises. It could be argued that the change in German monetary and economic policy after GEMSU — characterized by an increased German insensitivity to the policy requirements of the other ERM members — can be held responsible for the increased asymmetry. However, given our empirical evidence and the asymmetry concept explained in the next chapter this seems to be a precipitate conclusions, since the finding of asymmetry is apparently linked to the occurence of speculative crises rather than GEMSU. Instead of using parametric causality tests, we apply non-parametric methods to high frequency data, thus explicitly allowing for nonlinearities attributable to band effects (cf. KRUGMAN [1991]) or threshold cointegration (cf. KUNST [1992]).

The rest of the paper is organized as follows: In the next chapter we briefly review the existing literature and the principal concepts. A short description of the econometric methods used to address the asymmetry question is given in chapter 3. The final chapter contains the empirical results and a short conclusion.

Chapter 2

Asymmetry

What is meant by asymmetry? Throughout this paper asymmetry designates a unilateral influence of the anchor country's, i. e. Germany's, monetary policy on the other EMS member countries' monetary policy. In other words: In an exchange rate union there remains one degree of freedom for the determination of monetary policy, which in an asymmetric system is managed solely by the centre country.[1]

The empirical evidence on asymmetry in the EMS is rather inconclusive and ranges from finding a more or less symmetric system, e.g. De Grauwe [1989] and Cohen and Wyplosz [1989], to a system dominated by German monetary policy, e.g. Karfakis and Moschos [1990] and Weber [1991]. Yet, this conflicting evidence seems not surprising since the bulk of these studies resorts to causality analysis using monthly data of interest rates — the use of a sampling frequency which is much lower than the expected transmission time between the variables of interest might make the system look more symmetric and favours the finding of bi-directional and instantaneous causality. Recent studies (cf. Gardner and Perraudin [1993] and Henry and Weidmann [1995a,b]) avoid this source of bias by using high frequency data and generally confirm Germany's dominant position in the EMS.

[1] Note that this definition of asymmetry is much stronger than saying that in one way or another German monetary policy has more influence on its French counterpart than vice versa, since it excludes any significant impact of French monetary policy on the German one.

There are essentially three ways to empirically address the asymmetry question, though the first two are closely linked. First, one could resort to causality analysis. The bulk of the relevant literature (e.g. De Grauwe [1989], Fratianni and von Hagen [1990a,b], Kirchgässner and Wolters [1993], and Weber [1991]) uses this approach applying various causality concepts ranging from Granger [1969] causality to neutrality as defined by Stock and Watson [1989]. Second, one could follow Gardner and Perraudin [1993] and analyze the impulse-response functions in a vector-autoregression (VAR). Evidently, this is similar to neutrality analyses without formal testing. Third, one could measure how national macro-variables, e.g. interest rates or money stocks, react to speculative disturbances (cf. De Grauwe [1989]). This boils down to the question which country manages to sterilize its foreign exchange interventions (see also Giavazzi and Giovannini [1989]). In this paper we take the third approach.

The underlying mechanism, how and by whom the speculative shocks are absorbed, can be roughly summarized as follows: When speculative crises occur the concerned central banks — say the French and the German one — have to intervene in the foreign exchange market in order to keep the exchange rate inside the exchange rate band. This leads to an increase of the money stock in the hard currency country (Germany) and to a decrease of the like in the weak currency country (France). In the asymmetric setting Germany immedeately sterilizes this effect. Thus the whole burden of adjustment is borne by France, leading to a raise in French interest rates and leaving German interest rates unchanged. On the other hand, in a symmetric system both interest rates adapt.

The empirical implementation now simply consists in measuring the impact of speculative disturbances on German and French interest rates. More formally, we can write the uncovered interest rate parity as

$$i_t^{FF} = i_t^D + \dot{e}_t + \pi \tag{2.1}$$

where i_t^{FF} and i_t^D denote the French and German interest rates, respectively. $\dot{e}_t$ stands for the expected change in the logarithm of the exchange rate and π denotes a risk premium which we assume to be constant over each of the subperiods in regard (cf. chap. 3). Speculative shocks are reflected in de-or revaluation expectations, i. e. in an $\dot{e}_t$ different from zero. Consequently, the interest rates have to move to restore the equilibrium determined by the uncovered interest rate parity (2.1). In the symmetric case we expect both rates to change by equal amounts in absolute terms, whereas in the asymmetric case there will be a one to one correspondence between $\dot{e}_t$ and changes in i_t^{FF}.

Chapter 3

Econometric Methods

Given a data set $\{y_t, \mathbf{x_t}\}_{t=1,\ldots,T}$ stemming from a strictly stationary stochastic process in $\mathbf{R} \times \mathbf{R}^k$ with $E|y_t| < \infty$, the kernel estimator of the regression function $m(\mathbf{x}) = E\{y_t|\mathbf{x}_t = \mathbf{x}\}$ is a random variable of the form

$$\hat{m}_T(\mathbf{x}) = \frac{\sum_{t=1}^T y_t K\left(\frac{\mathbf{x}-\mathbf{x}_t}{\gamma_T}\right)}{\sum_{t=1}^T K\left(\frac{\mathbf{x}-\mathbf{x}_t}{\gamma_T}\right)}, \tag{3.1}$$

where the kernel K is any k-dimensional density and $\{\gamma_T\}_{T\in\mathbf{N}}$ is a sequence of positive bandwidths (converging to zero with growing sample size T) that regulates the degree of smoothness of the regression output. The suitably standardized denominator of 3.1,

$$\hat{f}_T(\mathbf{x}) = \frac{1}{T}\gamma_T^{-k}\sum_{t=1}^T K\left(\frac{\mathbf{x}-\mathbf{x}_t}{\gamma_T}\right), \tag{3.2}$$

serves at the same time as an estimator for the density of $\mathbf{x_t}$. We note that the formulation of the dependence between y_t and $\mathbf{x}_t$ by a regression (or mean response) function without stipulating an explicit regression model is the most general one and avoids any kind of misspecification, apart from a possible violation of the strict stationarity and integrability of the underlying stochastic process which are required in order to ensure the existence of a time-invariant regression function $m(\mathbf{x})$.

The asymptotic behaviour of (3.1) and (3.2) in time-series applications has been studied extensively in literature (cf. Andrews [1995], Bierens

[1983,1990], COLLOMB and HÄRDLE [1986], SINGH and ULLAH [1985]). The main requirement for the data-generating process to meet in order to ensure the consistency of the estimators (3.1) and (3.2), respectively, is a mixing condition or a near-epoque-dependence (NED) condition in L_2 and L_1, respectively. A result based on the latter condition is given in CRON [1995].

The motivation for using an estimator of type (3.1) is to let the data themselves decide about their classification instead of stipulating any functional form of the dependence in advance, as is required in parametric estimation. At the same time, the requirement of flexibility causes two disadvantages of the nonparametric approach: At first, since no structure is given *a priori*, many observations are needed to identify the underlying regression function. The number of data required is growing dramatically with the dimension k of the regression. Secondly, the interpretation of the estimation output is not straightforward, since there are no coefficients and consequently no testable parametric hypotheses.

A main problem with (3.1) and (3.2) is the choice of the bandwidth γ_T. Various procedures for finding optimal bandwidths have been proposed in literature. Among them, the most established one is the cross-validation procedure (cf. EFRON [1982]). For our purposes this leave-out technique proves to be unsuitable, however, since the resulting bandwidths tend to oversmooth the data. Therefore we prefer a simple trial-and-error procedure. To make the respective regressions (based on equally sized data sets) comparable, the selections has to be the same.

The bootstrap (cf. EFRON [1982]) enables us to determine confidence bands for the estimated regression surfaces. Thus we can run informal tests on structural breaks identifying regions (i. e. subsets of the domain of the regression functions) where significant differences in the mean-response functions are apparent. The bootstrap procedure, however, is a computer intensive resampling technique that requires the specification of a regression model, thus reducing the generality of the nonparametric approach. In the sequel we stipulate additive regression models of the form

$$y_t = m(\mathbf{x}_t) + \epsilon_t, \tag{3.3}$$

with pure white noise ϵ_t. Assuming that $E\{\epsilon_t|\mathbf{x}_t\} = 0$ a. s. , we can estimate $m(\mathbf{x})$ using (3.1). We draw a random sample with replacement of length T from the estimated residuals $\hat{\epsilon}_t = y_t - \hat{m}_T(\mathbf{x}_t)$ to generate a pseudo-sequence $\{\tilde{y}_t\}_{t=1,\ldots,T}$ in accordance with (3.3) that is used to re-estimate $m(\mathbf{x})$. This procedure is repeated independently B times. From these B pseudo-estimates $m^{(b)}(\mathbf{x})$ we determine the confidence bands we are interested in. To obtain uniform and conservative confidence bands, we calculate the suprema of the deviations from the original estimator on the grid where $m(\mathbf{x})$ is estimated.

Provided that (3.3) is indeed the data-generating model, being well represented by the original sample $\{y_t, \mathbf{x_t}\}_{t=1,\ldots,T}$, and that B is chosen large enough such that the law of large numbers applies, the bootstrap procedure should supply reliable approximations of the true confidence bands. It seems to be worth mentioning that the properties of (3.1) as an estimator of the regression function $E\{y_t|\mathbf{x}_t = \mathbf{x}\}$ remain unaffected by a possible misspecification of the data-generating model (3.3).

We use daily data on one-month interest rates collected on the Euromarkets by the Banque de France at 11:30 a.m. for the French franc and the deutsche mark (respectively denoted i_t^{FF} and i_t^{D}), covering the period between October 1984 and December 1993. Daily data were chosen so as to have a sufficiently high observation frequency for the kernel estimation to work. Although intra-daily data are available, the daily frequency seems to suffice as the rates hardly vary within the space of one day.

Analyses are made for the periods well before (placebo-period: 12 October 1984 to 6 November 1987), before (pre-unification period: 7 November 1987 to 30 November 1990) and after (post-unification period: 31 November 1990 to 24 December 1993) German unification [1]. Each subsample consists of 800 data points, thus meeting the requirement mentioned above for the kernel estimation (in one dimension) to yield reliable results. The quite arbitrary division is justified by our formulation of the question: Since two subsamples (placebo- and post-unification period) cover periods of speculative crises, [2]

[1]The beginning of our post-unification sample was chosen somewhat after GEMSU was accomplished since an apparent structural break in the German series occured at this date (cf. Henry and Weidmann [1995a]).

[2]The placebo period includes the speculative crisis of December 1986/January 1987

whereas pre-unification period is characterized by little motion, we can test the hypothesis that German unification itself has caused some structural break against the alternative that it had no direct impact. Confusion could have been caused by the coincidence of German unification with another sort of structural break, namely the change of periods with and without large motion.

In a first step, we estimate the density of the excess interest rate differential (cf. Fig. 3.1)

$$\dot{e}_t := i_t^{FF} - i_t^D - \pi,$$

i. e. the variable that might serve as regressor, with the risk premium π estimated by $\frac{1}{T}\sum_{t=1}^{T}(i_t^{FF} - i_t^D)$, for the sub-periods in regard using (3.2) in order to make a preselection of the regression grid. In the second step we estimate, assuming a lag of one day, the mean responses

$$m_{FF,1}(x) = E\{\Delta i_t^{FF} | \dot{e}_{t-1} = x\}$$

and

$$m_{D,1}(x) = E\{\Delta i_t^{D} | \dot{e}_{t-1} = x\},$$

respectively, for the three subsamples using (3.1) on the preselected grid. Uniform confidence bands for all regressions are obtained by the bootstrap procedure with 500 replications stipulating additive models as in (3.3). Finally, the preceding steps are repeated with the first difference of the interest rate differential as independent and Δi_t^{FF} as dependent variable to estimate

$$m_{FF,0}(x) = E\{\Delta i_t^{FF} | \dot{e}_t = x\}$$

and to clarify the causes of changes in the interest rate differential.

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which was caused initially by speculation against the US-Dollar and finally led to the realignment of 12 January 1987. The post-unification period covers the EMS-turmoils of September 1992/August 1993 which culminated in the suspension of the narrow bands.

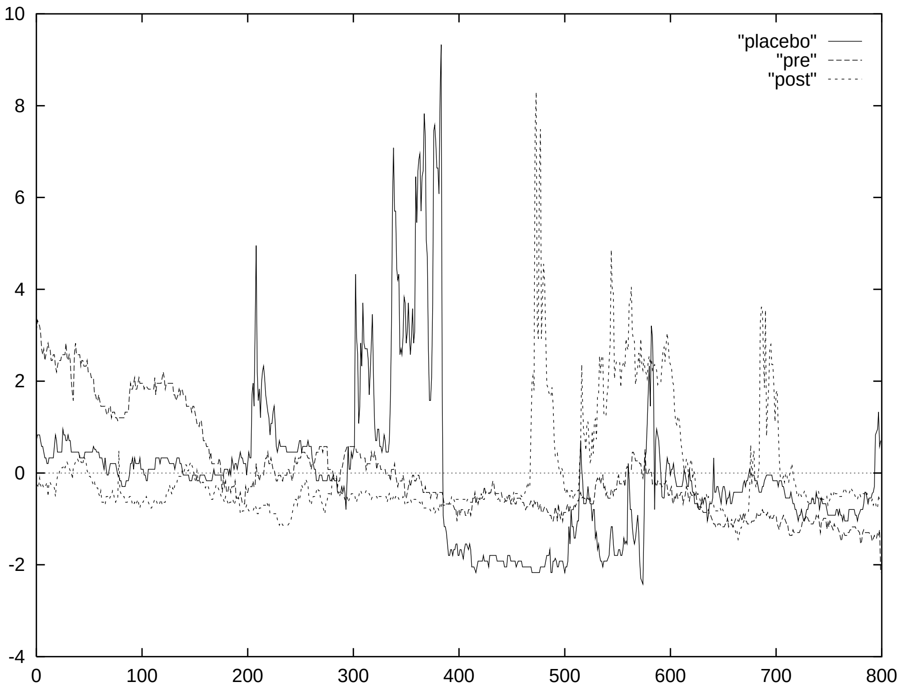

Figure 3.1: Time Series Plots of $\dot{e}_t$

Chapter 4

Empirical Results

In accordance with the previously described procedure we start presenting the estimated densities of the excess interest rate differential for the three subsamples in regard (cf. Fig. 4.1): Note that for the first and last subsample,

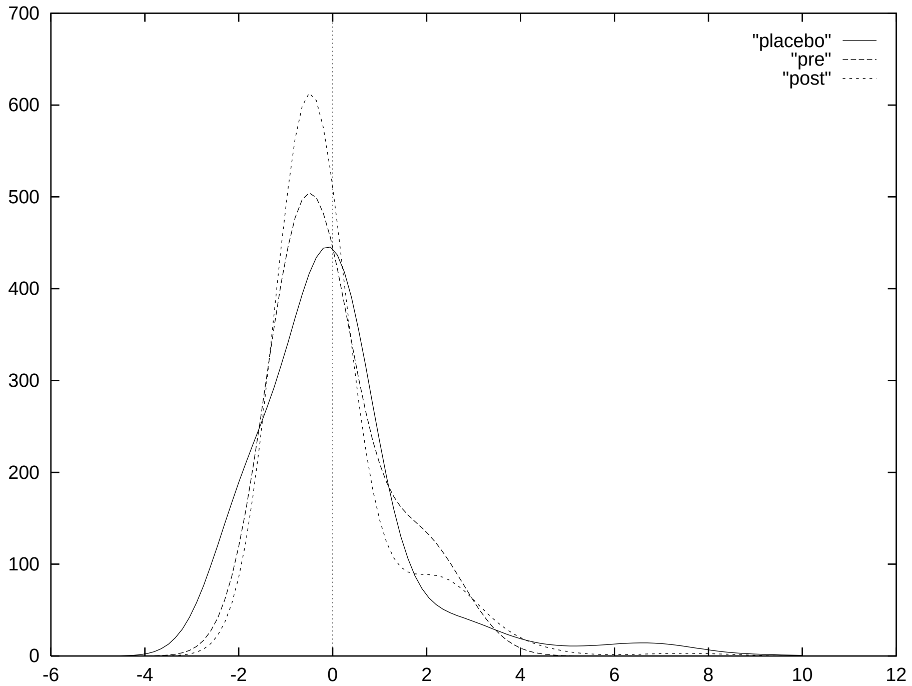

Figure 4.1: Estimated Densities of $\dot{e}$ $[\gamma = .5]$

the estimated densities are not unimodal, but exhibit slight elevations in the region about 5-10% . This seems to be the common size of excess interest rate differential movements during speculative crises (cf. Fig. 3.1). Therefore we estimate the the mean responses $m_{FF,1}(x)$ and $m_{D,1}(x)$ on a grid covering $[-2; 10]$. The estimated functions are shown in Figures 4.2 and 4.4, respectively. The most striking observation in Fig. 4.2 is the similarity between

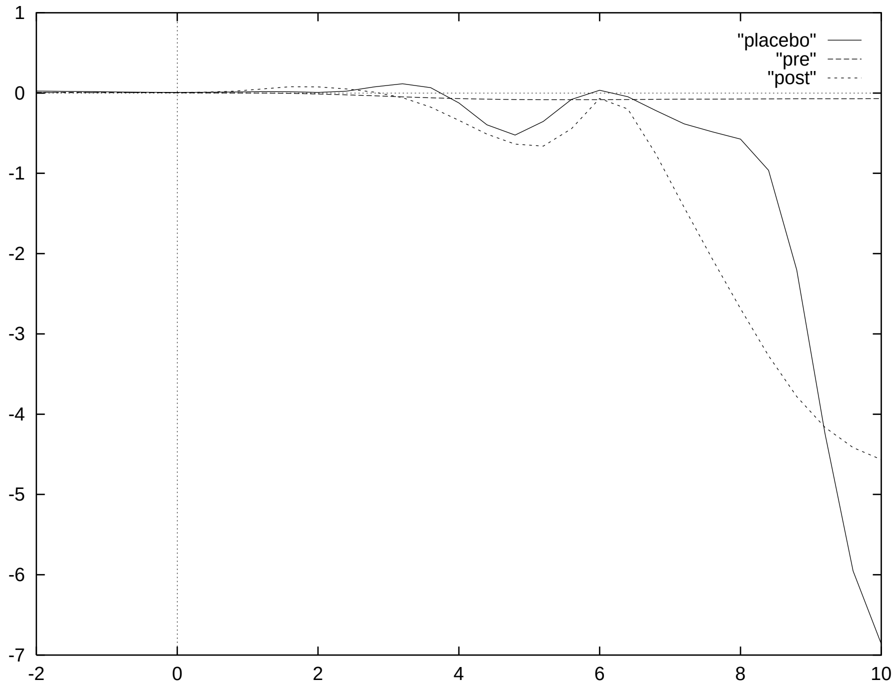

Figure 4.2: $m_{FF,1}(x) \quad [\gamma = .5]$

the post-unification subsample and the placebo subsample. Given the bootstrap confidence bands at 90% level, the figure shows a significant negative reaction for Δi^{FF} in the region where the excess interest rate differential exceeds 7%, whereas in the pre-unification period no reaction is observable. Therefore we interpret the 7%-mark as a threshold that must be exceeded before mean-reversion enters. For more clarity, we show the confidence bands for the post-unification period only (cf. Fig. 4.3).

This is in a line with the "threshold cointegration model" developed by Kunst [1992]. Threshold cointegration combines a cointegrated, nonsta-

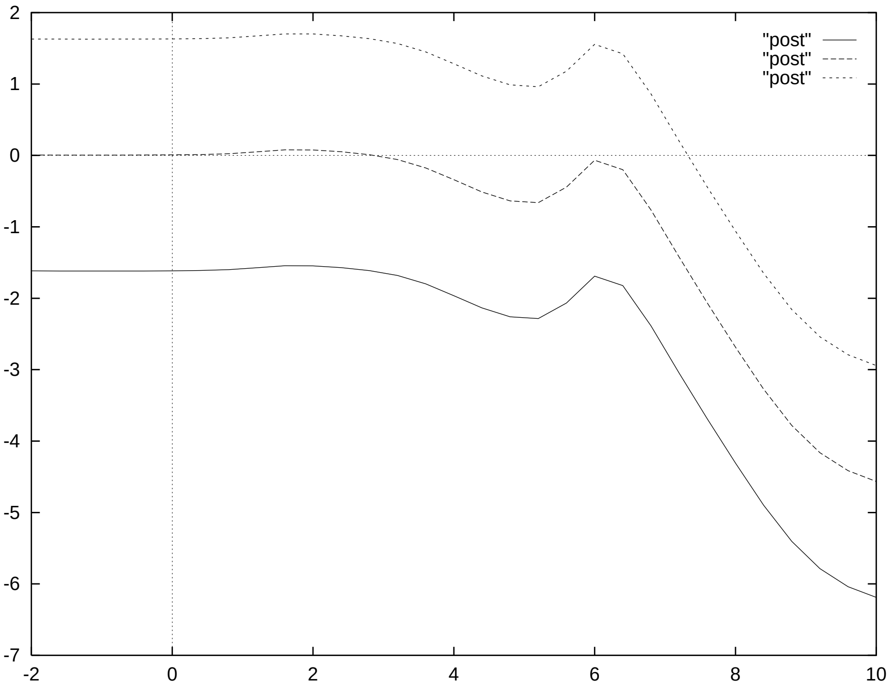

Figure 4.3: $m_{FF,1}(x)$

tionary and a stationary model: Outside some boundaries the series exhibit mean reversion, while inside they behave as integrated and potentially cointegrated variables. Thus, this model reconciles the random walk behaviour of interest rates commonly found in the empirical literature with the stylized fact — and economic intuition — that interest rates cannot grow indefinitly or become negative [1].

With respect to the German rate (cf. Fig. 4.4), we can confirm a significant reaction, being about 15 times weaker than the French one, only in the post-unification period.

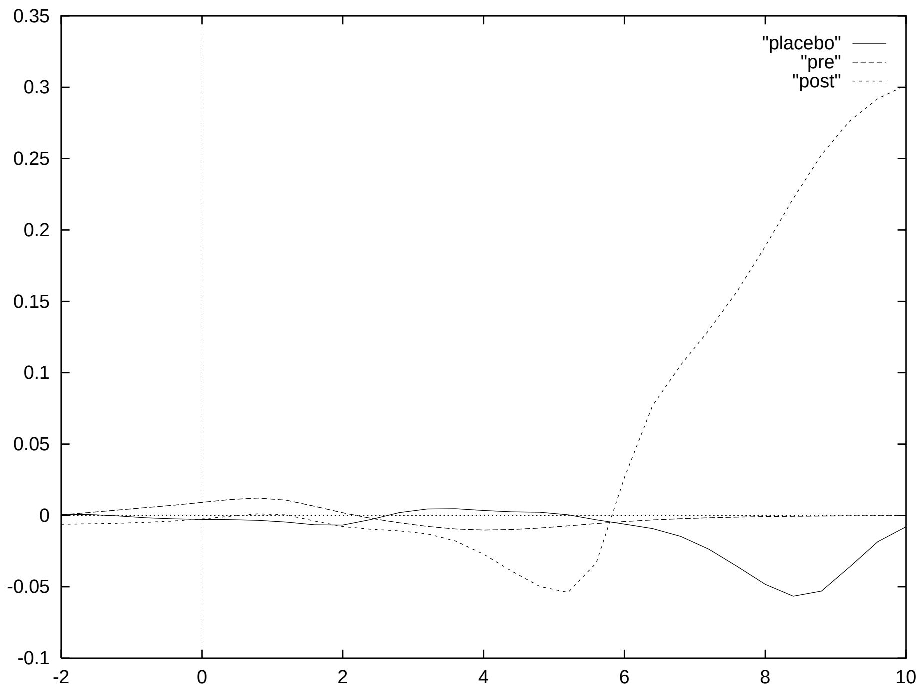

Figure 4.4: $m_{D,1}(x) \quad [\gamma = .5]$

On balance, we can confirm the asymmetry hypothesis for both speculative periods, but not for the pre-unification period. Hence, German unification did not affect Germany's position within the EMS. With the asymmetry

[1] Though negative interest rates occasionally occured, e. g. in Switzerland (cf. KUNST [1992]), they were not persistent.

concept used here, a comparison of the pre- and post-unification periods, however, would have led to a different conclusion, i. e. a grown asymmetry[2]. This signifies that the choice of the sample periods is of crucial importance with this asymmetry concept. Only agitated periods should be compared.

The estimated densities of $\Delta \dot{e}_t$ do not exhibit any irregularities. Therefore, their image is supressed. However, the estimation of the contemporary effects reveals some interesting features (cf. Fig. 4.5): During the agitated periods

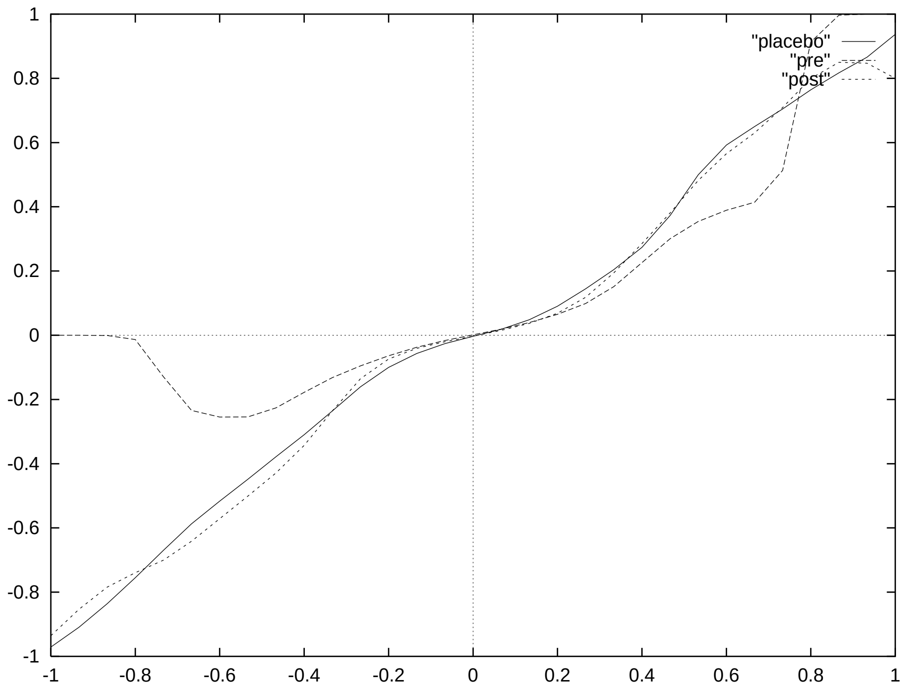

Figure 4.5: $m_{FF,0}(x), \quad [\gamma = .01]$

well before and after unification, only the Franc seems responsible for changes

[2]Note that this does not call in question the results of GARDNER and PERRAUDIN [1993] and HENRY and WEIDMANN [1995], since both papers resort to causality analysis. Furthermore, they include the American interest rate as supplementary variable. One of the main differences between their pre- and post-unification periods stems from the emergence of a monopolistic German influence, i. e. the absence of any causality from the American rate to the French rate since German unification.

in the interest rate differential. For the calm period before unification, however, the statement remains true only for positive changes. This seems to confirm the hypothesis that the reduction in the risk premia is not caused by declining French but by growing German rates.

Our findings suggest that

- interest rates should be modelled by regime-switching models, i. e. threshold cointegration models,

- non-parametric techniques are a powerful tool to reveal nonlinearities which seem to be important in interest rate models.

The above points could help to explain parts of the confusion in the discussion concerning the integration order of interest rates, since mean-reversion depends crucially on the regime and cannot be described by a linear response curve.

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