



Working Paper

## **A Challenge to the ECB's First Pillar? The French Experience the french experience**

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# Arbeitspapiere/ Working Papers

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The French Experience



Comments welcome.

# A Challenge to the ECB's First Pillar ? The French Experience\*

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## **Abstract**

This paper argues that the exchange rate regime between France and Germany prior to the EURO has led to a destabilized French demand for money function. To that aim the German dominance hypothesis is re-stated and empirically confirmed in a covered interest rate parity model with non-stationary variables and time varying parameters. Furthermore, in a multivariate time series model a stable long-run money supply but no money demand function is found. This casts doubt on the reliability of the so-called first pillar of the European Central Bank's monetary policy because a stable demand for money function is a prerequisite for the monetary targeting approach defined therein.

*JEL classification:* C32, E43, E51

*Keywords:* money supply, time varying parameter, covered interest rate parity

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# 1 Introduction

In consideration of its size and economic power, France has so far commanded surprisingly little attention in the money demand literature. Out of the studies available, a heterogeneous picture emerges. While e.g. Boughton and Tavlas (1990), Clausen (1998) and Lecarpentier-Moyal and Renou-Maissant (2000) have difficulties to establish reasonable estimates for the normalisation variable income, Fagan and Henry (1999) and Müller (1999) obtain coefficient estimates whose sign are at odds with the theoretical priors. Bordes, Chevrou-Séverac and Marimoutou (2001) even come to completely contradictory results depending on the estimation technique applied. In contrast to that e.g. Cassard, Lane and Masson (1995) and Wesche (1998) explicitly estimate a long-run money demand relation which widely conforms with what could be expected according to the money demand literature. Their estimation procedure or choice of variables respectively are not in line with the more recent concepts, though. Finally, Tullio, Souza and Giucca (1996) compare the money demand functions for the core European Monetary Union (EMU) countries (Italy, France and Germany), but fail to find a satisfactory relationship for France.

The diversity of results is the more striking if it is compared to the literature on German demand for money, for example. There, discussion mainly focusses on the magnitude of parameters but does not question the existence of a demand for money function as such. However, taken together, France and Germany produce far more than half of the EMU countries' output with France ranking second in terms of economic size. That's why it seems difficult to ignore the difficulties when it comes to formulating monetary policy for the whole area. This is especially relevant because the official strategy of the European Central Bank (ECB) is built on two pillars one of which being a monetary targeting approach.<sup>1</sup> For that a stable money demand relationship with known properties is assumed a pre-requisite.

This paper tries to shed some light on the reasons for the mixed results. It therefore investigates French monetary policy during the convergence period towards

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<sup>1</sup>The term for monetary targeting used by the European Central Bank (1999) is 'monetary referencing'.

EMU.

The structure of the paper is as follows. First, the historic events characterizing monetary policy conditions are briefly looked at. Their relevance for money demand investigations will be checked and finally, instead of a money demand function it will be argued that the monetary policy conditions in France can be best characterized by a stable money supply function. This result is the outcome of an investigation which aims at explaining the special economic links between Germany and France. The theoretical framework for that is the covered interest rate parity, which will be set in the context of the so-called German dominance hypothesis (see e.g. Artus and Salomon (1996), Hagen and Fratianni (1990), Kirchgässner and Wolters (1996), Henry and Weidmann (1996)).

The subsequent hypothesis states that a money demand investigation has to consider the history of the exchange rate regime, too. Moreover, having found support for the German dominance hypothesis, economic theory implies that monetary policy has very likely to be built on interest rate targeting. According to models of monetary policy in Europe (Bofinger, Reischle and Schächter (1998), Nautz (2000)) this tends to destabilize money demand functions and raises the probability of statistically identifying a stable long-run supply function. Taking a presumed break in 1987 into account eventually helps to explain the disappointing results reported e.g. in Tullio et al. (1996) and Müller (1999).

Throughout the paper, a system cointegration approach is applied to identify the relevant long-run relationships underlying the observations.

## **2 Preparing for the Euro: Implications for Money Demand Analysis**

### **2.1 Brief Summary of the Events**

For France as for most other western European countries, the break down of the Bretton-Woods system in 1973 meant the free floating of the domestic French cur-

rency against the US Dollar. At the same time the French market for capital remained moderately regulated, though. This first period culminated in 1982 in the so-called socialist experiment by the Mauroy government. In 1982 France even tightened capital controls. Finally, between 1986 and 1992 most restrictions were abolished. Thus, the time before 1986 can be characterized by regulated foreign exchange markets and the lack of a credit market.

In 1979 the European Monetary System (EMS) was founded in order to reduce the volatility of the exchange rates by, in principle, fixed rates. This meant controlled devaluations of currencies according to some macroeconomic criteria. This second period can thus be associated with working credit markets while the foreign exchange market remained subject to systematic Central Bank (CB) interventions. The latter aspect became even more important in the last sub-period when the French Franc (FF) for the last time was officially devaluated against the Deutschmark at the beginning of 1987 following a year of repeated adjustments of the central EMS parity.<sup>2</sup> Since then these two currencies have basically maintained their nominal values against each other. Figure 5 on page 37 in the Appendix illustrates this with a plot of the nominal prices of the Deutschmark quoted in FF and a purchasing power indicator.

## 2.2 Once Again - The German Dominance Hypotheses

The sequence of events could imply the following economic reasoning. For the first period up to the beginning of the 1980s no domestic credit market was working. Therefore, there is no reason to believe a demand for money function to rule either because credits are the very basis for the creation of and the demand for money. On the other hand, the exchange rate was subject to changes due to, albeit somewhat distant, market forces. All that became different in the third phase when the FF stayed fixed against the currency basket of the EMS members.

It is likewise noteworthy that nominal interest rates will become the more de-

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<sup>2</sup>To be more precise, this move was an appreciation of the Deutschmark against all currencies in the then ECU-basket including the French Franc. See also Loupiaz, Savignac and Sevestre (2001) on reforms within the banking sector at that time.



pendent on foreign rates the tighter the links are to the foreign markets. Naturally, the less flexible the exchange rate is, the less useful it is as a tool for cushioning asymmetric shocks.

The systematic links between interest rates and exchange rate regimes is the cornerstone of the so-called covered interest rate parity (CIP) condition. In a generalised version it can be expressed as equation (2.1)<sup>3</sup>

$$i_{F,t} = \beta_{i_G} i_{G,t} + \beta_E E_t + k_t \quad (2.1)$$

$$k_t = \kappa_0 + f(\vartheta, t - T_1)' z_t + a(L) \varepsilon_t \quad (2.2)$$

where  $i$  denote interest rates at time  $t$  with the subscripts indicating the country they refer to ( $F$  for France,  $G$  for Germany). The  $E_t$  is the expected depreciation of the French currency against the Deutschmark. It is defined as  $\frac{(fx_t^{t+1} - s_t)}{s_t}$  where  $s_t$  is the spot rate,  $fx_t^{t+1}$  measures the forward rate at time  $t$  for contracts due at  $t+1$ . All variables are prices of the Deutschmark (DM) quoted in French Franc. The  $\varepsilon_t$  are serially uncorrelated error terms,  $a(L)$  is a polynomial in the lag operator and the  $\kappa_i$ , ( $i = 0, 1$ ) are parameters. In (2.1) the term  $z_t = \gamma \otimes z_t^*$  denotes a subset vector of  $y_t$  variables,  $z_t^*$ , or constants whose long-run coefficient may change according to  $f(\vartheta, t - T_1)$ . The function  $f$  will be defined below. The  $\gamma$  is a  $2 \times 1$  vector with finite elements, and we have  $\lim_{(T-T_1) \rightarrow \infty} \gamma' f \rightarrow \kappa_1$ . If (2.1) holds and  $z_t^* = 1$ , then  $k_t$  yields an interpretation as a risk premium. As Marston (1995) (p.43) outlines, the risk premium may cover a number of different market inefficiencies like default risk, taxes, regulated markets for domestic securities, market segmentation due to capital controls, political or sovereign risks. Naturally, if markets work efficiently these risks should follow a stationary process.<sup>4</sup>

The CIP hypothesis in (2.1) can be analysed in the time series context by assuming all of these variables to be integrated of order one, but forming a cointegration

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<sup>3</sup>See e.g. Pentecost (1993) pp. 172ff who derives the CIP hypothesis from a portfolio approach. The restricted version is obtained if  $\beta_{i_G} = \beta_E = 1 \wedge k_t = 0, \forall t$ .

<sup>4</sup>Among others McCallum (1994), p. 109 has a strong economic prior that (2.1) is a definition equation with  $\beta_{i_G} = \beta_E = 1 \wedge k_t = 0, \forall t$ . However, he also states that the implicit expectation formation ( $\mathcal{E}_t(s_{t+l}) = fx_t^{t+l}$ ) only holds if agents are risk neutral. Therefore, we treat the economic prior as a hypothesis which also allows for (yet not imposes) risk aversion.

relationship. In other words, while each of the variables may be driven by stochastic trends, at least one of these trends they should have in common. This ensures that deviations from the theoretical CIP condition are short-lived. Since in 1987 the FF was officially devaluated for the last time, it is not only the cointegration relationship which is of interest, it is also the potential changes in this relationship due to the changing exchange rate regime. To see this, the following point might be considered. As long as there is no binding commitment to fixing the exchange rate (be it within certain bands of fluctuation), the interest rates might be set by the national monetary authorities at their will. International capital trade will then ensure that no risk free arbitrage is possible by adjusting the exchange rate in a way to offset potential differences between the interest rates. In the framework of non-stationary time series analysis, this scenario can be described as the presence of two independent stochastic trends within the trivariate system, where each of the trends originates from the national monetary policies. The forward rate (as well as the exchange rate) will then be driven by a combination of both these trends, while a (linear) combination of all of them has to yield a stationary relationship if arbitrage is ruled out in the long-run.

In case of interdependency between the two national interest rates, it can be assumed that either both of them adjust to each other to some extent, or that one of them will be set to account for changes of the other one. In the latter case, the interest rate in question will not only be affected by the domestic interest rate policy, but also by the foreign monetary policy decisions. In the time series context, this hypothesis can be cast as follows. Define

$$\begin{aligned} Y_t &= (i_{F,t}, i_{G,t}, E_t)', \\ Y_t^* &= (i_{F,t}, i_{G,t}, E_t, 1, f(\vartheta, t - T_1)'z_t)', \\ \varepsilon_t &\sim i.i.d(0, \Sigma) \end{aligned}$$

and find ( $\Delta x_t = x_t - x_{t-1}$ ),

$$\Delta Y_t = \alpha\beta'Y_{t-1}^* + \sum_{i=1}^{p-1} \Gamma_i \Delta Y_{t-1} + \varepsilon_t$$

**Table 1:** Hypotheses about  $\alpha\beta'$  and their Interpretation

Hypothesis	Interpretation (valid for the long-run)
$\mathcal{H}_1 : \text{rank}(\Pi) = 1, \wedge \beta_{i_G} = 1, \wedge \beta_E \geq 0$	CIP holds
$\mathcal{H}_2 : \mathcal{H}_1 \wedge \alpha_{i_G} = \alpha_E = 0$	French interest rate policy not independent of the German. German interest rate policy independent of the French.
$\mathcal{H}_3 : \mathcal{H}_2 \wedge \kappa_1 \neq 0 \mid T_1 \approx 1987$	Change in exchange rate regime affects CIP condition (and risk premium is not zero).

$$= \begin{pmatrix} \alpha_{i_F^s} \\ \alpha_{i_G^s} \\ \alpha_E \end{pmatrix} \begin{pmatrix} 1 & -\beta_{i_G} & -\beta_E & -\kappa_0 & -\gamma \end{pmatrix} Y_{t-1}^* + \sum_{i=1}^{p-1} \Gamma_i \Delta Y_{t-1} + \varepsilon_t \quad (2.3)$$

where the  $\alpha$ ,  $\beta$ ,  $\Gamma$  are coefficient matrices of appropriate dimension. In this setting,  $\beta$  defines the linear relationship representing the no-arbitrage, or CIP condition, and the vector  $\alpha$  provides information about the extent to which the endogenous variables adjust to short-run deviations from CIP. If indeed only one of the interest rates is subject to the stochastic trend of the other one, this should show up as a single significant coefficient in  $\alpha$  in the equation for the respective endogenous variable. If however, one of the interest rates is *not* subject to the stochastic trend of the other one, the corresponding coefficient must be zero. Note also that (2.1), (2.2) can easily be related to (2.3) by noting that  $E(\beta' Y_t^* + f(\vartheta, t - T_1)' z_t) = \kappa_0$  and hence, a decrease in the inevitable risk after fixing the exchange rate can be modelled as a shift in the mean of the cointegrating vector ( $\kappa_1 \neq 0$ ). This said, the appropriate choice of  $z_t^*$  would be 1 in this case.

From the foregoing discussion it follows that the subsequent analysis will focus on the derivation of the risk premium,  $k_t$  whose properties will provide the information being necessary for drawing conclusions. Table 1 surveys the various hypotheses on  $\Pi = \alpha\beta'$  and their economic interpretation.

The hypotheses  $\mathcal{H}_1$  to  $\mathcal{H}_3$  will be investigated according to the following strat-

egy. First the unit root properties of the variables in question will be investigated, second the cointegrating rank will be established, and finally a stability analysis performed. This stability check will very much focus on the stability of the cointegrating relationship.<sup>5</sup> The aim is to identify a valid representation of the data generating process (DGP) according to (2.3), which provides reliable information on the structural and economically meaningful long-run relationships between the variables.

## 2.3 A Small Econometric Study

In the econometric implementation of the three hypotheses, interest rates on credits referring to contracts over three months are considered along with the forward rate due in three months time. Thus, the maturities of the various assets are in line with each other. The data definitions and their sources are reported in the Appendix, while Figure 1 provides a graphical impression.

For investigating the time series properties, univariate as well as multivariate analyses will be employed. The hypothesis of no unit root is tested in the framework of the Augmented Dickey-Fuller test. All test results unanimously point to non-stationarity of the variables involved and will not be reported in order to save space. They are available on request, however.

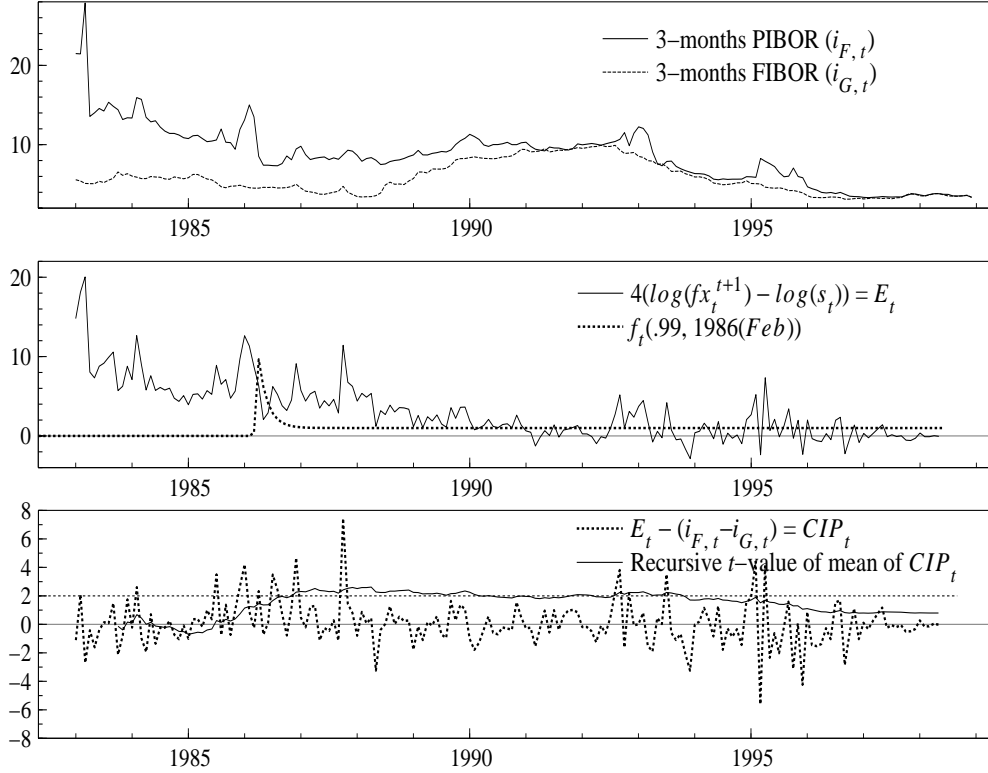
### 2.3.1 CIP System Analysis

Having found all three series non-stationary, the CIP hypothesis can be tested in the cointegration framework. The difficulty there is to cope with the possibility of a break in the cointegration relation under  $\mathcal{H}_3$  while not knowing exactly when such a break might actually occur. Unfortunately, having to tackle unknown timing of a break and determining the cointegrating rank has not yet been thoroughly dealt with in the literature. The only exception to that rule is, to my best knowledge, the work by Lütkepohl, Saikkonen and Trenkler (2001). In this paper, an itera-

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<sup>5</sup>Throughout the paper calculations are performed with the author's GAUSS routines and PcFiml 9.0 (see ?).

**Figure 1:** Data Input for the CIP Analysis



NOTE: McCallum (1994) argues that the mean of  $CIP_t$  is zero. The bottom panel shows the  $t$ -value of the estimate of the mean which turns out to be significantly different from zero between 1987 and 1994.

tive strategy will be applied. First, a cointegration test that does not take into account the possibility of a break will be applied, in the second step a search over many potential break points assuming the cointegrating rank found previously is performed. Finally, the cointegration rank test will be repeated, this time taking into account the possible shift. The cointegration test by Saikkonen and Lütkepohl (2000a) (henceforth S&L) will be employed. It has the advantage that it can be generalised to situations where the cointegration mean shifts without affecting the distribution of the test statistic.<sup>6</sup>

<sup>6</sup>See Saikkonen and Lütkepohl (2000b) and Lütkepohl, Saikkonen and Trenkler (2003) for a comparison of competing approaches.

The functional  $f(\vartheta, t - T_1)$  is defined as

$$d_t := \begin{cases} 0, & t < T_1 \\ 1, & t \geq T_1 \end{cases}$$

$$f = \left[ \frac{d_t}{1 - \vartheta L}, \frac{d_{t-1}}{1 - \vartheta L} \right]', \quad \vartheta \neq 0, |\vartheta| < 1$$

and  $\lim_{(T-T_1) \rightarrow \infty} \gamma' f \rightarrow \kappa_1$  with  $|\kappa_1| < \infty$ . Here,  $L$  stands for the lag operator that lags a variable by one period:  $Lx_t = x_{t-1}$ .

The Tables 2 and 3 collect the results of the steps one to three. In the first step, the test results suggest rank one. Under this assumption the break point selected by information criteria as well as by the maximum value of the likelihood is February in 1986 (see Table 3). The details for calculating the log-likelihood in this case can be found in Hansen (1992) and Seo (1998). Returning to Table 2 and observing the test statistic of step 3, the previous outcome can be re-stated. In an additional exercise, it was assumed that the  $\beta$  coefficients either of  $i_{G,t}$ , or  $i_{F,t}$ , or  $E_t$  may change instead of the cointegration mean but still, the estimated break point is always February in 1986. Therefore, this result appears robust.

The test find only one cointegration relationship. When the cointegrating rank is one, identification of the relationship is trivial. It just requires to normalise the cointegration vector. Table 4 reports the various hypotheses on the cointegrating relationship starting with the just identified relationship. These hypotheses are chosen such that they directly refer to  $\mathcal{H}_2$  and  $\mathcal{H}_3$ . While estimating the cointegration relationship with a smooth shift in the cointegrating relation is no problem at all, the literature provides little guidance with respect to the distribution of the likelihood-ratio statistic given in column 8 of Table 4. Therefore, the model was re-estimated assuming a once-off shift for which this problem has been solved by Hansen (1992). Although not being reported, the outcome of both, the estimates of the long-run coefficients as well as the values of the test statistics are qualitatively as well as quantitatively the same compared to those given in the table. Nevertheless, the smooth shift has been kept because doing so resulted in favourable residual properties.

**Table 2:** CIP Analysis, Steps 1 and 3: Cointegration TestsSample: 1983(Jan) - 1998(May),  $T = 179$ 

Step <sup>+</sup>	Hypotheses	Test statistic	5% critical value
1	$Y_t^* = (Y_t', 1)'$		
	$r_0 = 0$ vs. $r_0 \geq 1$	38.97	24.08
	$r_0 = 1$ vs. $r_0 \geq 2$	1.04	12.21
	$r_0 = 2$ vs. $r_0 = 3$	.78	4.14
3	$Y_t^* = (Y_t', 1, f(.99, t, 1986(Feb))'z_t)'$ , ( $z^* = 1$ ; mean shift)		
	$r_0 = 0$ vs. $r_0 \geq 1$	40.27	24.08
	$r_0 = 1$ vs. $r_0 \geq 2$	1.15	12.21
	$r_0 = 2$ vs. $r_0 = 3$	.92	4.14

<sup>+</sup> Note: the 2nd step is the search for the break point whose outcome is reported in Table 3.

A graphical evaluation of the long-run estimates confirms approximate constancy of the parameters over time. Figure 2 plots the recursively determined point estimates as well as the recursively calculated test statistic of the overidentifying restrictions. Under the hypothesis  $H_0^2$  of Table 4 the following long-run equilibrium relationship has been found (standard errors in parenthesis)<sup>7</sup>

$$i_{F,t} = i_{G,t} + .83 E_t + .91 - .69 f(.99, t, 1986(Feb)) + ec_t^1. \quad (2.4)$$

(.02)      (.11)      (.07)

All its parameters are highly significant, and they also have the signs suggested by economic theory. Here,  $ec_t^1$  stands for error correction term, which is a stationary process with mean zero. The error correction term only enters the equation for the French short-term interest rate which gives rise to the interpretation as the dependent variable in the long-run. In other words, German interest rate policy had a decisive long-run impact on the French monetary policy but not vice versa.

This dependency arose from the CIP condition given in (2.4). It is a relationship

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<sup>7</sup>To simplify the exposition, (2.4) reports the more informative estimated value of  $\kappa_1$  instead of  $\gamma$  (4th term on the r.h.s.).

**Table 3:** CIP Analysis, Step 2: Break Point Search

$$\Delta Y_t = \Pi Y_{t-1}^* + \sum_{j=1}^4 \Gamma_j \Delta Y_{t-j} + \varepsilon_t, \quad t = p+1, p+2, \dots, \quad rk(\Pi) = 1$$

$$Y_t = (i_G, i_F, E)_t', \quad Y_t^* = (Y_t', 1, f(\vartheta, t - T_1)' z_t)'$$
,  $12 \leq T_1 \leq T - 12$

Sample: 1983(Jan) - 1998(May)

$z_t^*$	jump ( $j$ ) <sup>+</sup>	$\hat{\vartheta}$	Loglikelihood	<i>AIC</i>	<i>HQ</i>	<i>SC</i>	$\hat{T}_1$
0	no		-479.934	-2.737	-2.398	-1.901	
1	yes	.99	-448.101	-3.132	-2.780	-2.262	1986(Feb)

<sup>+</sup> jump indicates whether or not a break in the cointegration relationship is assumed.

that underwent a transformation when the exchange rate regime was adjusted for the dawn of the Euro currency. From 1987 onwards the official parity of the currencies did not change but moved within narrow bands instead. In the long-run relationship this change is reflected in the variation of the cointegrating mean. As outlined before, the mean value of the long-run relation can be regarded a risk premium on cross border credits hedged by forward contracts. While this premium was approximately .91 before 1987 it decreased to a value of  $(.91 - .69) = .22$  afterwards.

To sum up the results of the CIP analysis Table 1 may be considered. In due course, all of the hypotheses  $\mathcal{H}_1$  to  $\mathcal{H}_3$  have found support providing evidence for the view that French monetary policy was structurally dependent on the German interest rate policy but not vice versa. The transmission channel was the CIP condition, a link which tightened after 1987 as it is reflected in the lowering risk premium.

### 2.3.2 Implications for Monetary Policy

The study of French interest rates reveals a change in the structure of the determination of short-term interest rates. It turns out that, in line with economic theory, the exchange rate regime is the crucial point. When in early 1987 the Deutschmark was appreciated against the EMS basket, this marked the last change in the official



**Table 4:** CIP Analysis: Hypotheses about the Cointegrating Relations

$$\Delta Y_t = \Pi Y_{t-1}^* + \sum_{j=1}^{p-1} \Gamma_j \Delta Y_{t-j} + \varepsilon_t, \quad t = p+1, p+2, \dots$$

$$\Pi = \alpha \beta', \quad Y_t = (i_{F,t}, i_{G,t}, E_t)',$$

$$Y_t^* = (Y_t', 1, f(\vartheta = .99, t, 1986(Feb)))' z_t' \quad (z_t^* = 1; \text{smooth shift of the cointegration mean})$$

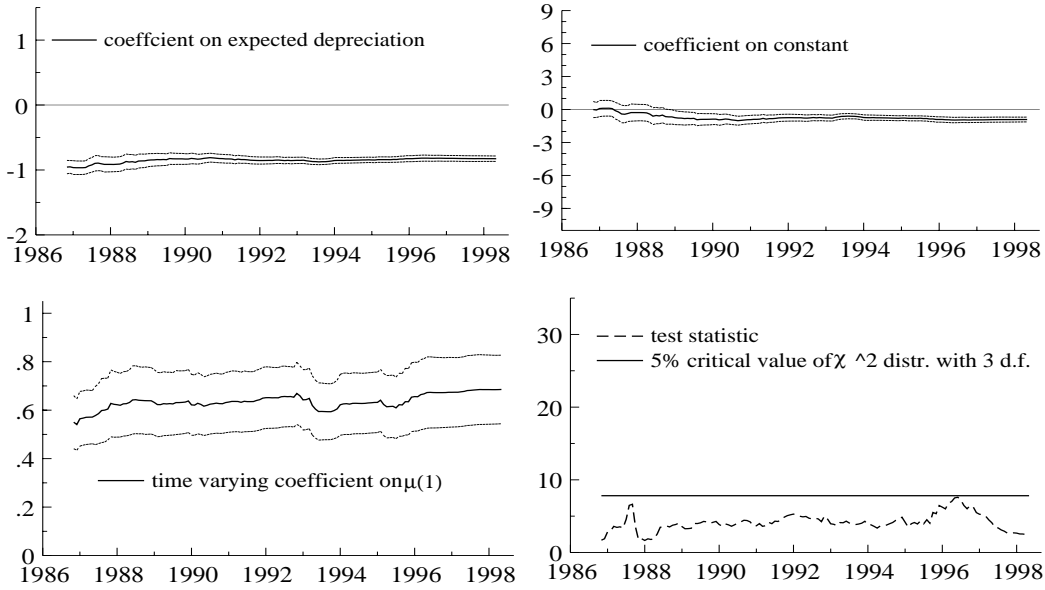
Hypotheses	vector	coefficient estimates on					LR statis- tic	d.f. <sup>#</sup>
		$i_{F,t}$	$i_{G,t}$	$E_t$	$\kappa_0$	$\kappa_1$		
just identified	$\hat{\beta}'$	1	-.96 (.03)	-.82 (.02)	-1.13 (.22)	.68 (.07)		
relationship	$\hat{\alpha}$	-.56 (.06)	.02 (.03)	-.05 (.21)				
$H_0^1 \quad \beta_{i_G^s} = 1$	$\hat{\beta}'$	1	-1	-.83 (.02)	-.90 (.11)	.68 (.07)	1.51 [.22]	1
	$\hat{\alpha}$	-.55 (.06)	.02 (.03)	-.04 (.21)				
$H_0^2 \quad H_0^1 \wedge \alpha_{i_G^s} =$ $\alpha_E = 0$	$\hat{\beta}'$	1	-1	-.83 (.02)	-.91 (.11)	.69 (.07)	2.34 [.50]	3
	$\hat{\alpha}$	-.55 (.05)	0	0				

+ likelihood ratio statistic which would have a  $\chi^2$  distribution in case of a once-off shift, #d.f. stands for degree of freedom of the  $\chi^2$  distribution. The p-value is given in brackets [].

exchange rate between FF and the Deutschmark. This date also coincides with the observations by Marston (1995) who quote the year 1986 as a starting point for financial market deregulations in France. Thus, the January 1987 can be seen as the beginning of a new era rather than as a single event and the timing of the shift identified appears reasonable. Since then the exchange rate basically stayed the same within narrow bands and even sustained the speculative attacks during the early 90s.

Notably, the introduction of the Euro, and hence the abolishment of the two currencies, does not alter the situation with respect to the relationship between the exchange rate and the short-term interest rate. This is because the situation of zero variations in the parity as of the time after 1998 is approximately equivalent to maintaining the central parity with very small, temporary deviations. This can be considered a first indication that a monetary policy analysis which certainly considers interest rates, and which is restricted to the period prior to the EMU should deliver results which might also be valid afterwards.

**Figure 2:** CIP Analysis: Recursive Long-run Estimates



Sample: 1983-1998, restrictions according to  $H_0^2$  imposed, short-run re-estimated. A smooth shift function is applied ( $f(\vartheta, T_1)$ ). Point estimates with two-standard error bands.

Against this rather optimistic interpretation the analysis also demonstrated that in the long-run French monetary policy, that is interest rate policy, was dependent on the German conditions. A transmission of policy shocks in the opposite direction could not be found as far as the long-run structure is concerned. Therefore, the present results support the hypothesis of the German dominance referred to by several authors like for example Artus and Salomon (1996), Kirchgässner and Wolters (1996). As a consequence, the conditions under which the Central Banks in these two countries conducted their policies were quite different. While the Bundesbank was seemingly able to act like being independent of foreign monetary policy (at least independent of its largest neighbour France), this does not hold for the French Central Bank. Consequently, while the Bundesbank was able to pursue a much more independent monetary policy according to the economic conditions of the domestic German economy, this seems to be very unlikely for the Banque de France. Thus, the claim made by Cassard et al. (1995), namely that France also applied a money targeting policy appears to be questionable.

If the latter statement was taken seriously, naturally the question arises, what the Central Bank did instead, and what the implications for the standard money demand analysis are likely to be. The most trivial answer to the first part is interest rate targeting, or to be more precise, exchange rate targeting by means of adjusting the Central Bank interest according to the deviations from target.

As to the second question, the answer is less straightforward. While on the one hand, the agents can be supposed to demand money as agents can generally be assumed to, it could well be that the econometric approaches aiming at revealing the properties of a money demand function are less likely to discover the related behavioural function. The central argument to be put forward in that respect has already been produced by Working (1927) and is the subject of any modern macroeconomic textbook; namely the necessity to identify both supply and demand if either of these curves is subject to statistical analysis.

Usually, identification of the supply curves remains a rudimentary procedure which can be justified by the assumption of exogeneity of money supply. The intuition is, that in case of monetary targeting as in the standard money multiplier model, say, the authorities focus on the supply of money. The exact amount offered depends on the forecast for the most important argument which is income and an implicit target inflation rate like the approximately two percent in the ECB reference growth rate of money. This commitment in turn implies a degree of exogeneity for the money supply process which most of the time allows to conveniently identify the demand curve. This is because in such a case the CB interest rate is used to achieve the monetary target and can be linked to the exogenous money supply process. If successful, this strategy will lead to a situation where the money supply function will shift along the demand curve in accordance with the targets set by the authorities. The observations made under these circumstances will recollect points on the demand curve, and therefore identify the demand function.

The foregoing is not a priori applicable though, if the money supply process itself is subject to exogenous influence. For France this might well be the case because the short-term interest rate could be shown to be ruled by the German rate. Therefore,

the interest rate cannot serve as a tool for achieving a money growth target when aiming at the exchange rate at the same time. Therefore, the identification of money demand and supply becomes an issue again and it is far from clear that the standard approach will necessarily disclose the features of the demand rather than the supply function. This argument is particularly relevant, because in the usual setup not only those variables are included which potentially enter the demand function but also factors being important in the money supply process. Therefore, if any of the candidate relationships (money supply or demand) dominates in the long-run, it will be identified and the actual coefficient estimates of the long-run relationship will inform about the prevailing relation.

Finally, having discovered a change in the channel transmitting foreign shocks to the French economy, it seems reasonable to consider that other relationships which are relevant for monetary policy transmission may change, too.

## **3 Money Supply in France Prior to EMU**

### **3.1 On French Money Demand**

In striking contrast to its size, political and economic weight among the European monetary system's member countries, France has attracted few interests in the empirical analysis of monetary policy, and money demand in particular. When considered, French money demand has typically been treated within a group of countries (see e.g. Tullio et al. (1996), Wesche (1998), Fagan and Henry (1999), Clausen (1998)) surveying the demand for money functions for various European countries.

A natural drawback of such approaches is the inevitable loss of information about valuable details like stability of parameter estimates, adjustment mechanisms other than adjustment of money to the money demand relation, complementary long-run relationships and the like.

The following analysis extends the previous approaches to French money demand in a number of ways. First, while the focus will also be on the money demand function, other long-run relationships which might be present will likewise be identified.

Acknowledging the relevance of the adjustment parameters for policy analysis, their properties will also be looked at. Finally, in the light of the CIP analysis, it will be assumed that a break in the cointegration relationships may have occurred around 1987. Modelling this event helps to identify a long-run money supply function for France and provides an explanation for the apparent failure of preceding studies on French money demand.

### **3.2 The Model and the Data**

The starting point for the following analysis is the notion of money demand being a function of those factors enhancing the desire for holding money and those working against that. The market to focus on is the market for bank deposits and securities in that commercial banks offer credits to the public and take deposits in return. This setting has been extensively studied e.g. by Bofinger et al. (1998) who emphasize the distinction between the market for monetary base and broad money. One of the remarkable outcomes is that the credit market, which is the market for broad money, can be dominated by either a money demand or a money supply function depending on what policy strategy is pursued by the central bank. In one instance, money supply will dominate if the central bank solely focusses on interest rate goals.

In particular, their model market for broad money has an interest rate for loans to the public as the price variable. Naturally, this price is negatively linked to the demand for broad money while it affects supply positively. The dynamics within the model crucially depend on the actual monetary policy strategy of the central bank. A monetary targeting strategy implies not only the money demand function for broad money to be stable in the credit interest rate - money dimension, but also a long-run one-to-one relationship between the monetary policy rate and credit market interest rate. When the central bank uses the commercial bank's interest rate as their intermediate target however, the money supply function appears to be stable and no long-run relationship between interest rates for credits and interest on central bank money should be observed.

These hypothesis are made operational in the following way. First, cointegra-

**Table 5:** Auxiliary Cointegration Analyses of Interest Rates

$$Y_t = \begin{pmatrix} \alpha_1 \\ \alpha_2 \end{pmatrix} (\beta_1 \ \beta_2 \ \mu_0) (Y'_{t-1} : 1)' + \sum_{i=1}^k \Gamma_i \Delta Y_{t-i} + \mathcal{E}_t$$

Sample: 1983 - 1998, quarterly data

endogenous variables ( $Y_t$ )	$k + 1$	Johansen Statistic	Test	critical values	Hypothesis <sup>+</sup> : $\beta_1 = -\beta_2 = 1$
call money, PI-BOR (3mth)	2	$H_0 : rank = 0$	35.53**	20.0	.33
		$H_0 : rank = 1$	3.31	9.2	[.56]
call money, repo rate	2	$H_0 : rank = 0$	16.76	20.0	.12
		$H_0 : rank = 1$	4.05	9.2	[.73]

\*\* indicates significance at the one percent level of significance. The Johansen test statistic refers to Johansen (1991). The corresponding critical values (5% level of significance) are available e.g. in Johansen (1995) Table 15.2.

<sup>+</sup> Likelihood-ratio statistic of the hypothesis having a  $\chi^2$  distribution with 1 degree of freedom.

tion tests are used to shed light on the long-run relationships between interest rates related to various assets. In particular, interest rates on central bank money, represented by the repurchasing rate, interest rate for very short-term credits on the interbank money market and medium term credits also on the interbank market. The latter two refer to contracts of one day (call money) and three months (PI-BOR) to maturity respectively. Table 5 reports the results of cointegration analyses in two-dimensional vector error correction models. The hypothesis of cointegration is tested and so is the hypothesis that the cointegration relationships are of the structure  $i^j = i^k$ , with superscripts  $j \neq k$  indicating the various definitions of the interest rates. The data is quarterly data from 1983 through 1998.

Table 5 tells that the call money rate is cointegrated with the interest rate for three months credits with the coefficients suggested e.g. by the expectation hypothesis about the term structure of interest rates. At the same time, the repurchasing rate (short repo) is not found cointegrating with the call money rate. The repo rate can be viewed as the actual central bank interest rate measuring the costs of commercial banks for borrowing central bank money. The other two interest rates relate to costs of credits between banks and are thus related to what banks are able

to charge from their customers. According to the model by Bofinger et al. (1998), the lack of a systematic long-run link between the central bank rate and the market rates for loans, indicates a situation in which monetary targeting is unlikely to be the dominating monetary policy strategy. Nevertheless, there are many other circumstances which could lead to not finding a cointegration relation between certain interest rates. Therefore, the following analysis will be designed such to encompass both hypotheses, existence of a long-run money supply and a long-run money demand relationship.

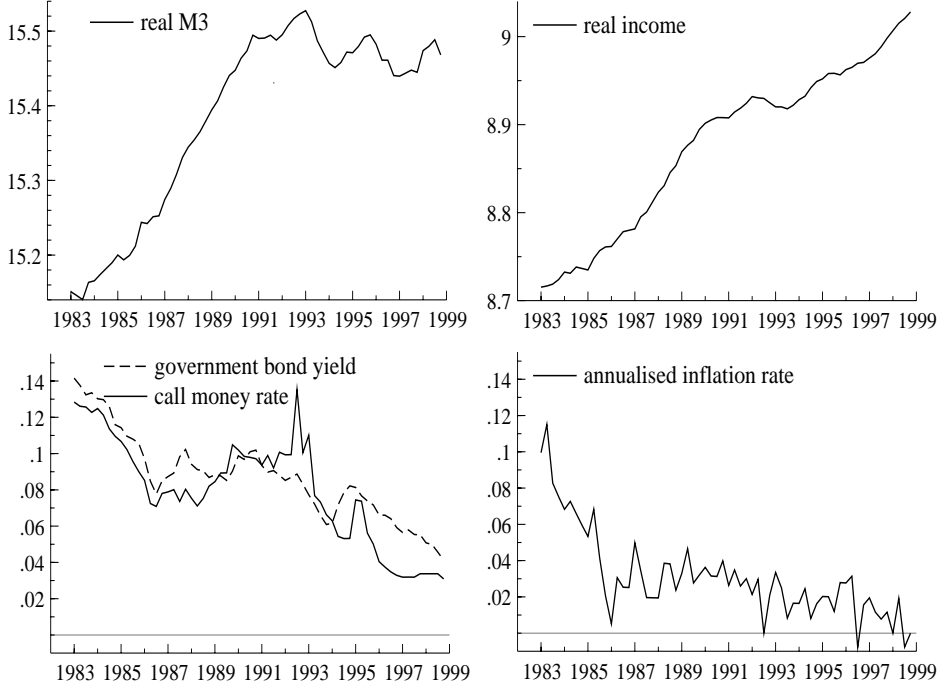
To that aim, those variables are modelled which are usually included in a money demand analysis such as long-term interest rates (bond rate, denoted  $i_t^l$ ) for measuring the opportunity costs of holding money instead of government bonds, a transaction variable which is real gross domestic product ( $y_t$ ). Another opportunity cost variable is inflation measured by the changes in the harmonised index of consumer prices ( $hicp_t$ ), which accounts for holding money instead of real goods. Technically, inflation is defined as  $\pi_t = 4\Delta hicp_t$  with  $\Delta$  being the difference operator. In addition, the call money market rate ( $i_t^s$ ) is included which will account for the own rate of money or, in other words for the price banks may charge when lending money. The dependent variable is real broad money ( $(m - hicp)_t$ ) in the definition of real M3. Figure 3 plots the time series which are analysed.

Notably, these variables may, in principle, yield both relationships long-run money supply and long-run money demand. In the framework of cointegration analysis, this should turn out as at least two cointegration relationships, one with real money depending on the short-term interest rate in a positive way, and in the other the real monetary aggregate should be negatively related to the long-term interest rate. For identification, the two interest rates must not be cointegrated themselves.

### 3.3 Empirical Analysis

All variables have been subjected to a unit root test and for all of them the unit root hypothesis could not be rejected. To save space, details are not reported but

**Figure 3: Data Graphics**



Money and income in logs, interest and inflation rates in annual decimals.

are available on request. The model underlying the estimation procedure is taken from Seo (1998) and Hansen (2003), it reads

$$\Delta Y_t = \nu_0 + \alpha \beta(t)' Y_{t-1} + \sum_{j=1}^{p-1} \Gamma_j \Delta Y_{t-j} + \varepsilon_t, \quad t = p+1, p+2, \dots \quad (3.5)$$

with  $Y_t = (m - hicp, y, i^l, i^s, hicp)'_t$  and where  $\varepsilon_t$  is a sequence of i.i.d. Gaussian variables with mean zero and covariance Matrix  $\Sigma$ . The term  $\beta(t)$  is given by

$$\beta(t) = \begin{cases} \beta_1 & t = 1, \dots, \tau_1 \\ \beta_2 & t = \tau_1 + 1, \dots, \tau_2 \\ \vdots & \\ \beta_q & t = \tau_{q-1} + 1, \dots, T. \end{cases} \quad (3.6)$$

which leaves the adjustment coefficients ( $\alpha$ ) and the remaining short-run coefficients ( $\Gamma_j$ ) unchanged. Thus, the number  $q - 1$  of possible breaks in the cointegration relation is finite. While Hansen (2003) introduces a test statistic which allows to



test for the significance of a potential break against a constant coefficient model if the break point is known, Seo (1998) and Hansen (2000) suggest a testing procedure for the case  $q - 1 = 1$  and unknown break point. Unfortunately, the latter statistic lacks a standard distribution and requires simulation of the critical value. To circumvent the latter problem, in what follows, the break point will be assumed known although it will be chosen according to some auxiliary analysis. In this case again, the selected break point is chosen according to model selection criteria. The maximum likelihood procedure for estimation is described in the Appendix.

As in the analysis of the CIP hypothesis, considering modelling a shift in the cointegration relation again invokes the difficulty of having to determine the cointegrating rank and the date of the break jointly. We deal with this problem as in the CIP analysis.

Various cointegration rank tests have been employed for the whole model as well as for many sub-models with fewer than the full set of the five variables.<sup>8</sup> The most plausible outcome appears to be rank two.

In the next step the break point will be determined under the assumption that the rank of  $\beta$  is two. In order not to deviate too far from the assumptions underlying the cointegration test applied before, only coefficients on the short-term interest rate and on the inflation rate are allowed to vary either separately or jointly. The motivation for a time varying interest rate follows directly from the change in the exchange rate regime which implies a changing behaviour of the central bank and hence a change in the behaviour of the agents in the money market.

In Table 6 the variables  $Z_{11t}$  and  $Z_{12t}$  indicate the variables whose long-run parameters are allowed to vary. It turns out that among the possibilities considered, a change of the parameters on the call money rate and inflation, results in the best model. This conclusion is also supported by the model selection criteria which do also account for the additional flexibility in the model. Thus, the long-run relationships feature time varying coefficients on the call money interest rate and on the inflation rate with the changes occurring in the fourth quarters in 1986 and 1985

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<sup>8</sup>The results are available on request.

**Table 6:** Break Point Search and Model SelectionModel (3.5),  $p = q = rank = 2$ ,  $1985(1) \leq \tau_i \leq 1989(4)$ ,  $i = 1, 2$ 

no.	$Z_{11t}$	$Z_{12t}$	Loglikelihood	$AIC$	$HQ$	$SC$	$\hat{\tau}_1$	$\hat{\tau}_2$
$\mathcal{H}_0$			284.796	-21.976	-21.360	-20.417		
$\mathcal{H}_1$	$i_t^s$		294.536	-22.353	-21.711	-20.728	1986(4)	
$\mathcal{H}_2$	$\pi_t$		296.732	-22.420	-21.777	-20.794	1985(4)	
$\mathcal{H}_3$	1		295.792	-22.322	-21.680	-20.687	1989(3)	
$\mathcal{H}_4$	$i_t^s$	$\pi_t$	304.333	<b>-22.724</b>	<b>-22.056</b>	<b>-21.032</b>	1986(4)	1985(4)

Bold face signifies the minimal value of the respective column.

respectively. In case of the interest rate, the optimal break point almost perfectly coincides with the previously mentioned change in exchange rate policy at 1987, first quarter. Thus, the prior knowledge about when the data generating process is likely to change, appears to be confirmed endogenously. With respect to inflation, the argument is not as straightforward. The most plausible explanation for the shift to occur in the fourth quarter of 1985 is that at about this time the inflation rate started to settle at a much lower level than before. This however, is more a valid description rather than a reasoning as to *why* it became so small.

The identification of the long-run relationships is reported in Table 7. It should be remembered that the matrices  $\alpha$  and  $\beta$  both consist of two columns each. That is why in Table 7 the estimates for their coefficients are reported line by line where each line corresponds to one column. In the first step, the estimates of the just identified cointegration relationships are displayed ignoring the estimates of  $\alpha$  for the moment. Then, restrictions on the cointegrating vectors (marked by bold face) are imposed and their validity is tested. The corresponding  $\chi^2$  statistics are reported in the last column. From the fifth step onwards the estimates for the  $\alpha$  coefficients are also provided and some of them are restricted, too.

**Table 7:** Identifying the Time Varying Cointegration Relationships

Sample: 1983-1998

$$\Delta Y_t = \nu_0 + \alpha \beta' Y_{t-1}^* + \Gamma_1 \Delta Y_{t-1} + \varepsilon_t$$

$$Y_t = ((m - hicp)_t, y_t, i_t^l, i_t^s, \pi_t)', \quad Y_t^* = (Y_t', I_{86(4)} i_t^s, I_{85(4)} \pi_t)'$$

<i>step</i>	<i>vector</i>	$(m - hicp)_t$	$y_t$	$i_t^l$	$i_t^s$	$\pi_t$	$I_{86(4)} i_t^s$	$I_{85(4)} \pi_t$	<i>d.f.</i>	$\chi^2$
1	$\beta_1$	1	-1.17 (.07)	0	-1.24 (.25)	.72 (.25)	-.35 (.19)	-1.99 (.35)		
	$\beta_2$	0	.002 (.03)	1	-.36 (.12)	-.72 (.12)	.19 (.09)	-.65 (.16)		
2	$\beta_1$	1	-1.17 (.05)	0	-1.25 (.22)	.72 (.25)	-.35 (.16)	-2.00 (.33)	1	.003 [.96]
	$\beta_2$	0	<b>0</b>	1	-.36 (.09)	-.72 (.12)	.20 (.07)	-.66 (.14)		
3	$\beta_1$	1	-1.17 (.05)	0	-.98 (.20)	.27 (.18)	-.38 (.17)	-.38 (.38)	2	2.39 [.30]
	$\beta_2$	0	0	1	-.20 (.06)	-1	.18 (.07)	-.88 (.19)		
4	$\beta_1$	1	-1.18 (.05)	0	-.88 (.18)	<b>0</b>	-.35 (.17)	-2.60 (.43)	3	3.22 [.36]
	$\beta_2$	0	0	1	-.20 (.06)	-1	.19 (.07)	-.88 (.19)		
5	$\beta_1$	1	-1.18 (.05)	0	-.87 (.17)	0	-.33 (.15)	-2.64 (.41)	4	3.24 [.52]
	$\beta_2$	0	0	1	<b>-.20</b>	-1	<b>.20</b> (.05)	-.91 (.18)		
	$\alpha'_1$	-.40 (.05)	.04 (.03)	.03 (.03)	-.01 (.06)	-.01 (.08)	-	-		
	$\alpha'_2$	-.43 (.10)	-.10 (.06)	-.16 (.05)	-.05 (.12)	.62 (.14)	-	-		
6	$\beta_1$	1	-1.14 (.05)	0	-.69 (.16)	0	-.47 (.14)	-2.35 (.36)	10	8.88 [.54]
	$\beta_2$	0	0	1	-.16	-1	.16 (.04)	-.79 (.16)		
	$\alpha'_1$	-.43 (.05)	<b>0</b>	<b>0</b>	<b>0</b>	<b>0</b>	-	-		
	$\alpha'_2$	.45 (.09)	<b>0</b>	-.12 (.04)	<b>0</b>	.65 (.11)	-	-		

Columns 3-9: Parameter estimates and their corresponding standard errors (in parentheses). Column 10: Degrees of freedom of the  $\chi^2$  statistic. Column 11:  $\chi^2$  statistic and the corresponding marginal level of significance [in brackets].

Where no number in brackets is given, the corresponding coefficient value had been imposed.

**Bold face** signifies the restriction additionally imposed in the current step.

Denoting the implied error correction terms  $ec_{MS,t}$  and  $ec_{f,t}$  respectively, and defining  $Y_t^* = (Y_t', I_{86(4)}i_t^s, I_{85(4)}\pi_t)'$ , the final specification of the long-run relationships can be written as:

$$\begin{aligned} \alpha\beta'Y_t^* &= (\alpha_1 \ \alpha_2) \begin{pmatrix} \beta_1 \\ \beta_2 \end{pmatrix} Y_t^* \\ &= (\alpha_1 \ \alpha_2) \begin{pmatrix} ec_{MS} \\ ec_f \end{pmatrix}_t \\ &= \begin{pmatrix} -.43 & .45 \\ (.05) & (.09) \\ 0 & 0 \\ 0 & -.12 \\ & (.04) \\ 0 & 0 \\ 0 & .65 \\ & (.11) \end{pmatrix} \begin{pmatrix} 1 & -1.14 & 0 & -.69 & 0 & -.47 & -2.35 \\ & (.05) & (.16) & (.14) & (.36) \\ 0 & 0 & 1 & -.16 & -1 & .16 & -.80 \\ & & (0) & (.04) & (.16) \end{pmatrix} \begin{pmatrix} (m - hicp) \\ y \\ i^l \\ i^s \\ \pi \\ I_{86(4)}i^s \\ I_{85(4)}\pi \end{pmatrix}_t \end{aligned}$$

The  $I_{85(4)}$  is an indicator function which assumes the value zero before 1985, fourth quarter and one afterwards. The estimation results have been further investigated by means of stability analysis. This included a recursive estimation of the  $\alpha$  and  $\beta$  coefficients (see Figure 4) as well as out-of-sample forecasting for the final periods of the sample. None of them pointed to problems with the robustness of the results over time.

Re-writing the cointegration relationships leads to the following interpretations. The first ( $\beta_1$ ) can be given an interpretation as a money supply function where the inflation enters with a positive sign

$$(m - hicp)_t = 1.14y_t + .69i_t^s + .47I_{86(4)}i_t^s + 2.35I_{85(4)}\pi_t + ec_{MS,t},$$

and for the end of the estimation period one finds

$$(m - hicp)_t = 1.14y_t + 1.16i_t^s + 2.35\pi_t + ec_{MS,t}. \quad (3.7)$$

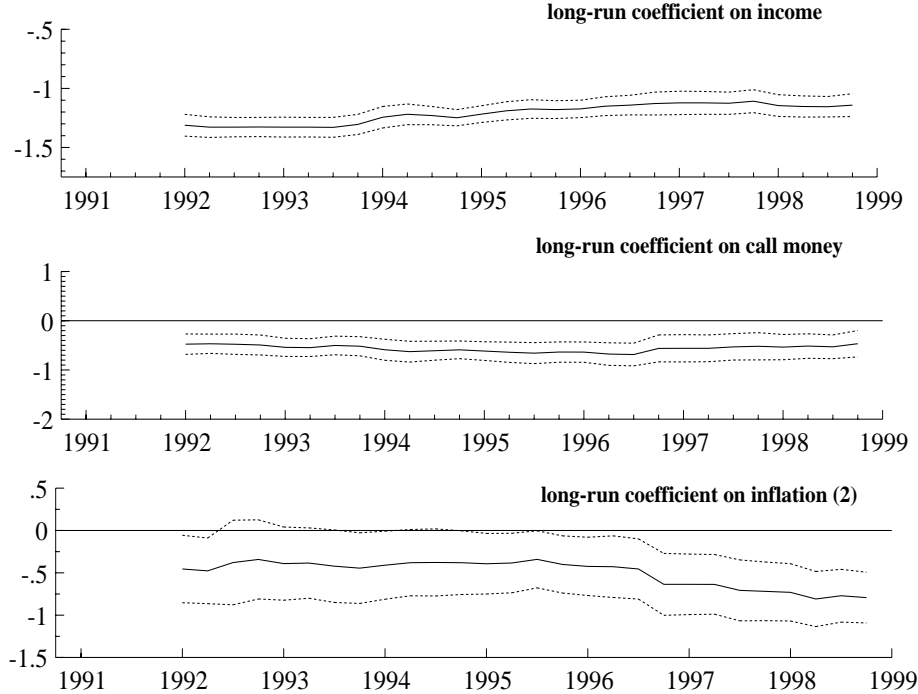
The reason for regarding this relationship as a money supply function mainly rests with the fact that the real money aggregate positively depends on changes of the interest rate. Because the second relationship given by

$$i_t^l = .16i_t^s + \pi_t - .16I_{86(4)}i_t^s + .79I_{85(4)}\pi_t + ec_{f,t}.$$

and after 1986 by

$$i_t^l = 1.79\pi_t + ec_{f,t}, \quad (3.8)$$

**Figure 4: Stability Analysis**



Top to bottom panels: Recursive estimates on the long-run income coefficient, the call money rate in the money supply relation, and long-run coefficient on the inflation rate in the 'Fisher' type relationship.

All estimates are obtained by estimating both  $\alpha$  and  $\beta$  vectors recursively.

relates the inflation rate positively to the long-term interest rate, no linear combination of the two relationships exists which would imply real money stock to shrink following a decrease in the long-term interest rate. Without that, the relation cannot be regarded a money demand function. Scrutinizing the second cointegration relationship a little further, it turns out that the short-term interest rate drops out after 1986 while the impact of the inflation rate increases. At the end of the estimation period a variant of the Fisher relation can be observed. The Fisher hypothesis states that the nominal interest rate is the sum of the real rate and expected inflation. If true, it follows that a theoretical cointegration relationship can be derived with the cointegrating vectors  $(1 \ -1)$  relating the long-term interest rate to inflation. In (3.8) the estimated coefficients do not correspond to this hypothesis. In the literature,

coefficients on inflation in excess of unity are often explained by the so-called *Peso problem* which might have been present in Europe during the run-up period towards EMU. It means that during this time, the remarkable decline in the inflation rate across Europe has led to a systematic failure of the assumption that past inflation is a good predictor for future inflation. If so, the hypothesised cointegration coefficients have to be different from  $(1 - 1)$ . In particular, the coefficient on inflation would be greater than one in absolute terms. This is, what can be observed in this case.

The estimated adjustment coefficients ( $\alpha$ ) imply that only real money strongly reacts to deviations from the long-run money supply equilibrium. Thus, if the actual supply is above its long-run equilibrium value given by the interest rate, inflation rate and national real income, money stock adjusts very quickly. The half-life of a deviation is approximately only one period. In the framework of the economic model mentioned before, this can be understood as an adjustment of credit supply by commercial banks in order to secure profits.

## 4 Comparison to Other Studies

In contrast to the large number of studies considering Germany, or more recently Europe for example, the evidence for money demand functions to exist in France appears rather thin.<sup>9</sup> Table 8 reports some results available for France, of course without claiming completeness of the selection.

A common feature of a French money demand investigation seems to be that it is conducted simultaneously within a group of countries. In the survey of Table 8 this is true for all but the Müller (1999), Lecarpentier-Moyal and Renou-Maissant (2000), Goux (2000) studies.

As an example, Fagan and Henry (1999) and Wesche (1998) solely focus on the existence of the money demand relation and derive feasibility of monetary targeting from its parameters and weak exogeneity of money. They both report the existence

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<sup>9</sup>Hubrich (2001), pp. 98ff lists about 20 studies on German money demand alone; a European aggregate has been analysed at least as often as that.

**Table 8:** Empirical Demand for Money Functions for France

sample	variables in CI relation <sup>‡</sup>	coefficient estimates*					reference <sup>‡</sup>
		$y_t$	$i_t^{out}$	$i_t^{in}$	$\pi_t$	$\alpha_{1,1}$	
61(4)-83(3)	nominal M2R, per capita income, income deflator, gov. bond yield	1.42	-.008	-	-	-.21	Dooley and Spinelli (1989)
79(1)-90(2)	M3 (deflated by hicp), GDP, 10 yrs gov. bond yield, 3mths interbank rate	1.31	-1.4	1.4	-	-0.27	Cassard et al. (1995)
79(1)-98(4)	M1, GDP, long and short term interest rate	.43	-1.40	.69	-	n.a.	Boughton and Tavlas (1990)
79(2)-90(2)	M3 (deflated by hicp), GDP, 10 yrs gov. bond yield, 3mths interbank rate	1.59	-3.0	3.0	-	-0.116	Cassard, Lane and Masson (1997)
73(1)-94(4)	M3H (deflated by hicp), GDP, gov. bond yield	1.27	-1.4	-	-	-0.07	Wesche (1998)
79(1)-96(4)	M1, GDP, gov. bond yield, short rate	-.05	-1.22	-	-	n.a.	Clausen (1998)
78(12)-94(6)	nominal M2, GDP, 3mths interbank rate, exch. rate	1.51	-	-.04	-	n.a.	Elyasian and Zadeh (1999)
81(1)-94(4)	M3H, GDP, long-term interest rate (not specified)	1.7	.19	-	-	n.a.	Fagan and Henry (1999)
83(1)-97(4)	M3, GDP, long-term bond yield call money rate, GDP inflation	1.23	-	2.88	-2.88	0	Müller (1999)
82(1)-97(4)	nominal M3, CPI, GDP, call money rate, trend	.40	-	-.0047	.60	-.61	Lecarpentier-Moyal and Renou-Maissant (2000)
72(1)-95(4)	M3, GDP, GDP defl., real call money rate, dummy 85-92	1.0	-	-1.99	1.99	-0.06	Goux (2000), T. 4
85(1)-95(4)	see above, no dummy	1.41	-	2.98	-2.98	0	Goux (2000), T. 6
85(1)-95(4)	see above	.87	-	-2.48	2.48	-0.1	Goux (2000), T. 7 <i>MD relation</i>
	see Goux (2000), T. 4	1.19	-	1.42	-1.42	-0.4	Goux (2000), T. 7 <i>IS relation</i>
79(1)-98(4)	M3, GDP, GDP deflator, gov. bond yield (>5 yrs.), call money rate	1.0	-8.42	8.42	-	-0.026	Bordes et al. (2001), single eq. result
	see above	1.42	11.35	-11.35	-	-0.018	Bordes et al. (2001), VECM (appendix 3)
83(1)-98(4)	M3 (deflated by <i>hicp</i> ), GDP, bond rate, call money rate, <i>hicp</i> inflation	1.14	-	1.16	2.35	-.41	this study

\*:  $i_t^{out}$  and  $i_t^{in}$  denote interest rates referring to assets *outside* and *inside* the respective monetary aggregate investigated.  $y_t$  stands for the transaction variable and  $\pi_t$  signifies inflation.

‡: Money (deflated by GDP deflator) and income in real terms if not indicated otherwise.

\*\* : The T. abbreviates Table.

The long-run coefficient estimates in the Clausen (1998) investigation are all insignificant, the estimates on the short-term interest rate are not reported.

of a second long-run relationship without telling their properties, however. Elyasian and Zadeh (1999)'s and Dooley and Spinelli (1989)'s estimates cannot be considered reasonable competitors to the current results because they use inappropriate econometric tools.<sup>10</sup> Most drastically, Tullio et al. (1996) simply state that they "were unable to estimate satisfactory dynamic versions of the demand for money functions" (see page 318).

The study of Goux (2000) can be compared best to the current exercise. There are two reasons for that. First, the estimation technique is more or less identical to the one used here. Second, the work also distinguishes between the two sub-periods before and after 1985. The difference is though that Goux (2000) does not explicitly relate this sample split to a change in the exchange rate regime but to the "mutation" of the financial system. Furthermore, Goux (2000) implicitly assumes a change in all coefficients of the model while it has been shown in the previous section that a break in some of the long-run parameters is sufficient.

Comparing the final outcomes, the differences appear less pronounced than the similarities, though. In particular the result of Müller (1999) and Goux (2000) Table 6 are more or less identical. The same holds true for the money supply relationship estimated in this paper and the *IS* relation of Goux (2000). Unfortunately, the long-run relationships reported by the latter are not identified which makes it impossible to tell anything about their significance. Therefore, the labels *money demand* and *IS curve* relationships are ultimately arbitrary. Therefore, it can be argued that Goux (2000) obtains more or less identical results but fails to interpret them appropriately.

## 5 Interpretations

Summing the new estimation outcomes, the central argument to explain the difficulties with a French demand for money function is that it cannot be identified econometrically because the market for broad money has been dominated by the long-run money supply function. The model by e.g. Bofinger et al. (1998) which has

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<sup>10</sup>See e.g. the comments by Thornton (1990).



also been used by Nautz (2000) for the setting of central banks in Europe can help to understand why the long-run money supply much better represents the money market outcome. They identify exactly one combination of monetary policy targets and shocks (demand or supply) to the market for broad money that results first, in a destabilizing effect on the demand for money function, and second, in a stabilizing impact on money supply. Such a situation occurs if the monetary policy objective is the money market interest rate (see for example Nautz (2000) pp. 95 ff.). Thus, combining this result with the insight gained in the first analysis yields exactly that outcome. The French Central Bank was bound to target interest rates in order to maintain the parity with the Deutschmark and hence shocks to the money market were accommodated in such a way to keep the desired interest rate level which eventually established a more stable long-run money supply function.

This said, even the money supply function cannot be viewed absolutely stable in the long-run due to the change in the exchange rate regime in 1987. This is reflected in the time varying coefficients of the long-run relationships.

## 6 Summary and Conclusion

During the period prior to EMU a stable long-run money supply function was found. Adding this finding to the results of a CIP analysis, evidence accumulates which implies that the demand for money function was destabilized at the expense of stabilizing the supply function. This is part of the consequences of the exchange rate regime introduced in 1987 when the French Franc was devaluated against the Deutschmark for the last time. Some newly developed tools which are designed to account for changes in the cointegration relationships had to be applied in order to cope with the impact of the 1987 events on the data generating process.

When comparing the recent results to those found by other authors, the largely contradictory outcomes can be explained by the differences in the definition of the sample period, estimation techniques, and by the infeasibility of a just comparison on grounds of missing information about the properties of the competing results and

by the widespread ignorance of the change in the FF/DM exchange rate regime.

The long-run money supply function derived from the attempt to model a money demand function is consistent with e.g. the theoretical model by Bofinger et al. (1998) and found coinciding with the implications of their framework.

Finally, with respect to the prospects of France for the European Monetary Union, it has to be pointed out that so far, French monetary policy was not built on the existence of a stable money demand relationship. This is a direct consequence of having to care for the exchange rate which led to the infeasibility of targeting money at the same time. Thus, underlining the positive aspects of that setting, it could be argued that due to the introduction of the Euro and the delegation of the monetary policy decisions to the ECB, France will regain some sovereignty about monetary policy because it can make its voice heard at the board meetings. But even if the demands by France would not play a role, the situation can hardly worsen because France has already sustained a long period during which the exchange rate had been fixed de facto with all its consequences for monetary policy. The introduction of the EURO currency merely prolongs that situation, indefinitely though.

Nevertheless, in case the ECB puts emphasis on its first pillar of monetary policy, which is monetary targeting, the lack of knowledge about the features of a French money demand will introduce some need for experimenting before the uncertainty originating from the missing facts will be eradicated.

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## A Appendix

### A.1 Estimation Procedure in the Presence of a Break in the Cointegration Coefficients

For estimating the model in (3.5), Hansen (2000) suggests the following variation to the standard reduced rank regression proposed by Johansen (1988). Defining

$\tilde{Z}_{0t} = \Delta Y_t$ ,  $Z_{1t} = Y_{t-1}$ ,  $\tilde{Z}_{2t} = (\Delta Y'_{t-1}, \dots, \Delta Y'_{t-k}, 1)'$ ,  $\Psi = (\Gamma_1, \dots, \Gamma_k, \nu_0)$  model (3.5) can be re-formulated as

$$\tilde{Z}_{0t} = \alpha\beta' \tilde{Z}_{1t} + \Psi \tilde{Z}_{2t} + \varepsilon_t \quad (\text{A.9})$$

and for  $q > 1$  we define  $Z_{11t} = Y_{t-1}I(t \leq \tau_1)$ ,  $Z_{12t} = Y_{t-1}I(\tau_1 + 1 \leq t \leq \tau_2)$ ,  $\dots$ ,  $Z_{1qt} = Y_{t-1}I(\tau_{q-1} + 1 \leq t \leq T)$  and  $\tilde{Z}_{1t} = (Z'_{11t}, \dots, Z'_{1qt})'$  which will result in a model equivalent to (A.9). In this new model,  $\beta$  is substituted with  $B = (\beta'_1, \dots, \beta'_q)'$  and  $\tilde{Z}_{1t}$  replaces  $Z_{1t}$ . Here,  $I(t \leq \tau_1)$  are matrices which indicate what elements of the  $Z_{1t}$  matrices are active. When  $\alpha\beta'$  is of reduced rank, the Johansen (1988) reduced rank regression can be straightforwardly generalised to the case of (A.9) by defining

$$\begin{aligned} \tilde{M}_{ij} &= \frac{1}{T} \sum_{t=1}^T \tilde{Z}_{it} \tilde{Z}_{jt}, \quad i, j = 0, 1, 2 \\ \tilde{R}_{it} &= \tilde{Z}_{it} - \tilde{M}_{i2} \tilde{M}_{22}^{-1} \tilde{Z}_{2t}, \quad i = 0, 1 \\ \tilde{S}_{ij} &= \frac{1}{T} \sum_{t=1}^T \tilde{R}_{it} \tilde{R}'_{jt}, \quad i, j = 0, 1 \end{aligned}$$

which leads to the estimates

$$\begin{aligned} \hat{B} &= (\tilde{v}_1, \dots, \tilde{v}_r) \\ \hat{\alpha} &= \tilde{S}_{01} \hat{B} \\ \hat{\Sigma} &= \tilde{S}_{00} - \hat{\alpha} \hat{\alpha}' \\ \hat{\Psi} &= \tilde{M}_{02} \tilde{M}_{22}^{-1} - \hat{\alpha} \hat{B}' \tilde{M}_{12} \tilde{M}_{22}^{-1} \end{aligned}$$

where  $(\tilde{v}_1, \dots, \tilde{v}_r)$  are the eigenvectors corresponding to the  $r$  largest eigenvalues  $\tilde{\lambda}_1, \dots, \tilde{\lambda}_r$  as solutions of the eigenvalue problem

$$|\lambda \tilde{S}_{11} - \tilde{S}_{10} \tilde{S}_{00}^{-1} \tilde{S}_{01}| = 0$$

(see Johansen (1988), Hansen (2000)). The maximum of the likelihood function is given by

$$L_{max}^{2/T}(\hat{\alpha}, \hat{\beta}, \hat{\Psi}, \hat{\Sigma}) = (2\pi e)^n |\tilde{S}_{00}| \prod_{i=1}^r (1 - \tilde{\lambda}_i)$$

of which a proof is available in Johansen (1988).<sup>11</sup>

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<sup>11</sup>In the first term on the r. h. s.  $e$  denotes Euler's constant.

Thus, the maximum of the likelihood has an explicit formula which simplifies computation of the estimates. These estimates are obtained for given values of  $q$  and  $\tau_1, \dots, \tau_q$ . When the latter of these is not known for sure, then the likelihood value can be calculated for all likely  $\tau_i$ . The estimate of the break point ( $\hat{\tau}_i$ ) will then be the one that corresponds to the maximum of the likelihood over all  $\tau_i$  scrutinised given the observations and the parameter estimates. The same strategy has also been used by Saikkonen and Lütkepohl (2001) in a uni-variate setting. The optimisation procedure is borrowed from Boswijk (1995), pp. 10f.



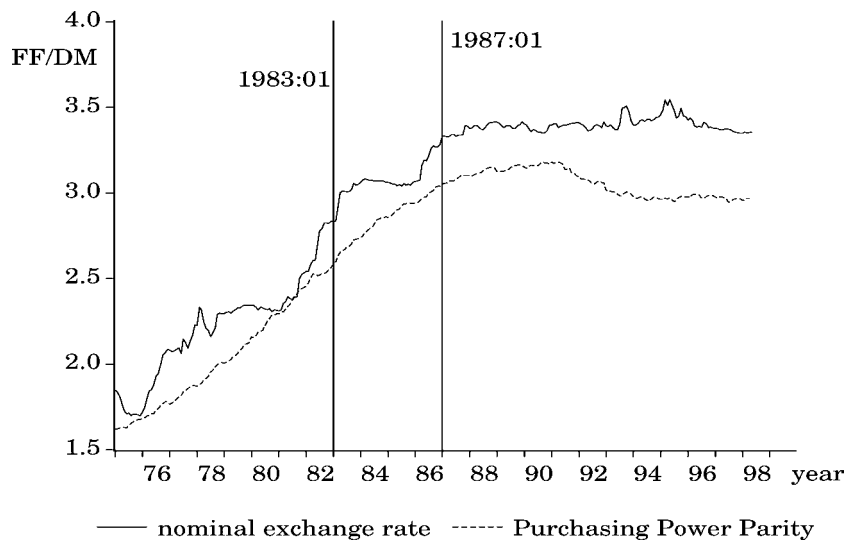
## A.2 Data and Figures

The data has been downloaded from Humboldt University's copy of the 1998 International Statistical Yearbook which includes the IMF data stream and OECD data bases.

**Table 9:** Data Sources

Data description and source code			
item	description	code	source
nominal money	M3 seasonally adjusted , billions FF	132 39MCC	IMF
real output	Gross domestic product, 1995 prices, seasonally adjusted, 1000 billions FF	141037NSA	OECD
French price level (1)	harmonized index of consumer prices; base year 1995	133 64...	IMF
French price level (2)	index of consumer prices; base year 1990	132 64	IMF
repo rate	repurchasing agreements	132 60A	IMF
call money rate	call money rate	146207D	OECD
3-mth interest rate	PIBOR 3-month or 90 day rates (Inter-bank)	146225D	OECD
bond rate	10 yrs bond guaranteed by the government	146281D	OECD
forward rate	Terminkurse am Frankfurt/M Markt; 3-Monatstermine (Mittelkurse), FF/DM	WU5046	Bundesbank
spot rate	official exchange rate, monthly averages (FF/DM)	132 RH	IMF

**Figure 5:** Nominal Exchange Rate and Purchasing Power Parity Line



The purchasing power parity is the ratio of the French CPI and the German CPI based on 1990 prices times the purchasing power indicator for the same year according to the OECD.