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## The empirical determinants of the Euro: Short and long run perspectives

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# The Empirical Determinants of the Euro: Short and Long Run Perspectives

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## **Abstract**

The behavior of the dollar/euro exchange rate is modeled using a monetary model of the exchange rate. The econometric analysis is complicated by the short sample span of actual euro data available for analysis. Hence, data on a “synthetic” euro are used. The assumptions underlying the monetary approach are discussed. A cointegrating relationship involving the exchange rate, money stocks, industrial production, interest and inflation rates, augmented by a relative price of nontradables, is identified for the 1991M08-1999M12 period using the Johansen procedure. The model implies that the euro was undervalued by about 13-15% in January 2000. This finding is robust to the use of alternative sample periods, and alternative estimation methodologies such as single-equation error correction and first differences specifications.

A longer term perspective is provided by a productivity-based model of the real value of the euro. Some panel regression estimates of the relationship between intercountry relative productivity differentials and real exchange rates is presented. Using these estimates to conduct some calculations, one comes to the conclusion that unless drastic changes to productivity trends occur, there is little reason to believe that the real value of the euro will deviate from its current zero-drift path.

**JEL:** F31, F41, F47

**Keywords:** euro, exchange rate, monetary model, productivity

# 1. INTRODUCTION

The introduction of the euro is arguably one of the most important events in recent macroeconomic history. Not only did the process leading up to monetary union occupy much of the attention international economists during the 1990s, so too did the subsequent gyrations of the euro.

In this paper the behavior of the euro/dollar exchange rate is examined from several perspectives. The first is in the context of a monetary model of exchange rates. Both the long run and short run relationship between exchange rates and money supplies, interest and inflation rates, and relative prices are identified and estimated. The results conform, broadly, to monetary models of the exchange rate, although the results are far from satisfying, as several questions are left unanswered.

One is tempted to question the feasibility of estimating such structural models, given the well-known difficulties in predicting this asset price. The conventional wisdom concerning the futility of structural exchange rate modeling was established by Meese and Rogoff (1983a,b). However, this view has been cast in doubt by several studies of G-7 currencies that provide robust evidence that macroeconomic fundamentals do affect nominal exchange rates.

The second major portion of the paper investigates the low frequency behavior of the real value of the euro. This analysis takes as its starting point the findings of a link between the relative price of nontradables, and the real exchange rate. Given that relationship, a linkage between the real exchange rate and relative productivity differentials is implied by a supply side model of relative prices. Extrapolating from US-European trends in productivity

over the last decade, some scenarios for the euro are outlined.<sup>1</sup>

This paper proceeds in the following manner. In Section 2, preliminaries are discussed: in particular, the purchasing power parity building block is assessed. In section 3, the monetary model of the exchange rate is presented and the econometric methodology is discussed. The empirical estimates and out-of-sample performance of the models are examined in Section 4. In section 5, a productivity-based model is used to discuss long term prospects. Some concluding remarks are included in Section 6.

## 2. SOME PRELIMINARIES: PPP AND THE EURO

A basic building block of the monetary approach to exchange rates is some sort of purchasing power parity (PPP) concept, relating exchange rates and price levels. If relative PPP holds, then the real exchange rate (appropriately defined) should be constant.

$$\begin{aligned} s_t &= p_t^{US} - p_t^{euro} + \kappa \Rightarrow \\ q_t &\equiv s_t - p_t^{US} + p_t^{euro} = \kappa \end{aligned} \tag{1}$$

Where  $p^{euro}$  is a measure of the price level (such as a GDP deflator or the CPI) in the EU-11 countries,  $p^{US}$  is the US price level,  $s$  is measured in US\$/€, and  $\kappa$  is a constant relating to the base years of the price indices.

However, as pointed out in Rosenberg (2000) the trade weighted real euro has experienced a long term decline over the past 20 years. Corsetti and Pesenti (1999), argue that a similar argument holds for the bilateral euro/US\$ from 1990 onward. Over this period, the

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<sup>1</sup> I eschew discussion of other determinants of the euro's value, such as the use of the euro as a reserve and vehicle currency, since that has been discussed elsewhere (Demertzis and Hughes Hallett, 1999; Portes and Rey, 1998).

CPI deflated euro has been depreciating by about 1.4% per year, according to a simple linear regression on time. Figure 1 depicts the real euro over the past 20 years, and an estimated 1990M01-1999M12 trend. As observed by Breuer (1994), a significant trend in the real rate is not consistent with a literal interpretation of purchasing power parity.

It is possible to rationalize deterministic linear trends by appealing to either measurement error, or the presence of nontradables, in which case what becomes important is whether the real exchange rate  $q$  is  $I(0)$  (as argued by Cheung and Lai, 1993b). Interestingly, it is possible to obtain an essentially zero appreciation or depreciation in the real rate by the judicious choice of sample period. For instance, taking the Louvre Accord in April 1987 as a break point, a linear trend estimated for the 1987M06-1999M12 period yields a nearly flat trend term. Thus the estimated trend is sensitive to the sample period, as is to be expected when a difference stationary series is mistakenly treated as a trend stationary series.

In terms of *understanding* the behavior of the euro, however, it is rather troubling to have long term movements being adduced to essentially unknown factors -- especially if one wishes to make predictions about the future. Rather than merely accepting a linear trend in the real rate, an alternative measure is examined: the PPI deflated real rate. Figure 2 depicts this real rate, and a trend term estimated over the same post-Louvre sample. In this case, no slope is discernable.

These graphical results are interesting, but hardly conclusive. Since purchasing power parity is essential to almost any monetary model of exchange rates, it is crucial to verify that it holds for some price indices. If none obtains, then it might be desirable to resort to some type of "real" model, a la Stockman (1980). Table 1 displays results from estimating the long

run relationship in the top line of equation (1) over the 1987M06-99M12 period<sup>2</sup> using dynamic OLS (DOLS, see Stock and Watson, 1993). Provided there is a single cointegrating relationship, DOLS represents a simple and robust means by which to estimate the long run parameters.<sup>3</sup>

Overall, the results demonstrate that PPP holds for the narrow, but not broad, price indices. The CPI coefficient estimates are incorrectly signed, and are sensitive to the inclusion of a time trend; column [1] reports the baseline estimates, in which a one percent increase in the US price level induces a 1.7% *appreciation* of the dollar against the euro, while the same increase in the Euroland price level induces a 2.2% appreciation of the euro. Inclusion of a linear time trend (column [2]) moves the US price coefficient in the right direction, but not the Euroland coefficient. Even when the coefficients on US and Euroland price levels are constrained to be equal and opposite (columns [3] and [4]), the coefficient on relative prices is sensitive to the inclusion of a time trend (although one cannot reject the null of PPP in these cases, regardless of whether a time trend is included or not).

In contrast, the PPI-based results are more promising. If the coefficients are unconstrained, then the price levels enter with the expected signs (column [5]). While it is true that inclusion of a time trend yields poor results, it is also true that when the regressor is

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<sup>2</sup> Papell (1998) argues that the mid-1980s appreciation and depreciation of the dollar is not well explained by a single deterministic trend model or a simple unit root model. Instead, he argues that a segmented deterministic trends model works best for most bilateral exchange rates, with a flat trend at the ending portion, corresponding approximately to the sample used here. See in particular the figure for the DM/\$ rate.

<sup>3</sup> DOLS can be thought of as an Engle-Granger regression augmented with leads and lags of the first difference of the right hand side variables, to account for endogeneity. As is often the case, the point estimates obtained using the Johansen procedure are quite different than those obtained by DOLS. Stock and Watson (1993) present simulation evidence indicating that the Johansen estimates are typically more dispersed than their DOLS counterparts.

the *relative* price level (so that the US and Euroland price levels enter with equal and opposite signs) then the coefficient is correctly signed and statistically different from zero, irrespective of whether a time trend is included or not (columns [7] and [8]). Furthermore, a Horvath-Watson (1995) test indicates that relative PPIs are cointegrated with the nominal exchange rate.

Once one drops the PPP assumption for a broad index, it is reasonable to inquire what implications there are for trends in the real CPI deflated euro/dollar rate. A reasonable approach is to assume that the presence of nontradable goods inhibits equalization of some prices in common currency terms. To see this, consider the aggregate price level for country  $j$  as being defined by:

$$p_t^j = (1-\alpha)p_t^{T,j} + \alpha p_t^{N,j} \quad (2)$$

where  $\alpha$  is the share of nontradables in the aggregate index. Further assume PPP holds for traded goods,

$$s_t = p_t^{T,US} - p_t^{T,euro} \quad (3)$$

Then one obtains

$$q_t = -\alpha[(p_t^{N,US} - p_t^{T,US}) - (p_t^{N,euro} - p_t^{T,euro})] \quad (4)$$

where  $\kappa$  has been suppressed. One does not observe pure measures of nontradables and tradables, but proxies exist. If one assumes that the CPI basket incorporates half nontradables and half tradables and the PPI entirely tradables, then

$$\alpha[(p^{N,US} - p^{T,US}) - (p^{N,euro} - p^{T,euro})] \approx \log \left[ \frac{(CPI^{US}/PPI^{US})}{(CPI^{euro}/PPI^{euro})} \right] \quad (5)$$

and the CPI deflated real exchange rate should move approximately one for one (and inversely) with the intercountry relative CPI/PPI ratio in [brackets]. A DOLS regression yields a slope coefficient equal to -1.84 (standard error of 1.03), instead of the expected -1.0.<sup>4</sup> The results are not very sensitive to the inclusion of a time trend. In that case, the point estimate is 2.46 (standard error of 1.10), although the time trend is not statistically significant.

Taken together, these results indicate that the PPP does not hold for broad price indices, but that a substantial component of the *long term* movement in the real exchange rate comes from the relative price of nontradables. It will prove necessary to incorporate this finding in the subsequent sections.

### 3. ESTIMATING THE MONETARY MODEL

#### 3.1 The Model

The asset-based approach to monetary model of the exchange rate can be represented as

$$s_t = \beta_0 + \beta_2(m_t^{US} - m_t^{euro}) + \beta_3(y_t^{US} - y_t^{euro}) + \beta_4(i_t^{US} - i_t^{euro}) + \beta_5(\pi_t^{US} - \pi_t^{euro}) + \beta_6\omega \quad (6)$$

where  $s$  represents the value of the *euro* in US dollars,  $m_t$  is the (log) nominal money stock,  $p_t$  is the (log) price level,  $y_t$  is (log) income,  $i_t$  and  $\pi_t$  are the interest and expected inflation rates, respectively.  $\omega$  is the intercountry relative price of nontradables, as described above.

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<sup>4</sup> This point estimate is, however, close to the 1.92 obtained by Clostermann and Schnatz (2000) in their analysis of the synthetic euro, using a single equation error correction model on quarterly data over the 1975-99 period.

In the standard monetary model the coefficients have structural interpretations which may vary with the assumptions in effect. In monetary models,  $\beta_2$  equals unity, while  $\beta_3 < 0$ , where  $|\beta_3|$  represents the income elasticity of money demand. If prices are assumed to be flexible, then the interest rate and inflation differential are the same, and the conditions  $\beta_4 > 0$  and  $\beta_5 = 0$  holds.  $\beta_4$  is equal to the absolute value of the interest semi-elasticity of money demand. On the other hand, if prices are sticky and there is secular inflation, then  $\beta_4 < 0$  and the magnitude of this parameter is positively related to price stickiness; the more rapid price level adjustment is, the smaller this coefficient is, in absolute value terms; then the coefficient on inflation is positive ( $\beta_5 > 0$ ) and increases with the interest semi-elasticity of money demand and decreases with the degree of price stickiness.<sup>5</sup> Finally,  $\beta_6 < 0$ ; in words, this means that as the relative price of nontradables goes up in the United States, the dollar appreciates vis a vis the euro, and vice versa.

The negative findings of Meese and Rogoff (1983a,b), recently recounted in Rogoff (1999), provide a pessimistic backdrop for any econometric project involving structural models of exchange rates. More recently, however, work of MacDonald and Taylor (1994), Mark (1995) and Chinn and Meese (1995) suggest that there is some empirical content to the long run predictions of the monetary model. These authors conclude that structural models *can* outperform a random walk if long run relationships are incorporated into the econometric specifications.

### **3.2 Complications**

There are several key issues that must be addressed when applying the monetary

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<sup>5</sup> See Frankel (1979) for a discussion of how these monetary models are related.

model the euro: (i) how to address the short sample period; and (ii) how to account for the EMS crises of 1992 and 1993. These issues are addressed in descending order of difficulty

***The short sample span.*** Clearly, as of February 2000, there exist only at most 13 monthly observations on the euro; observations on other variables of interest – say prices and money stocks – are likely to have even fewer observations at hand. Because it is impractical to estimate structural monetary models, along with the associated dynamics, with such few observations, the researcher is forced to use of data from the pre-EMU period. Hence, in estimating the models, “synthetic euro” exchange rate data will be used. Counterfactual data for the EU-11 will also be used for money stocks, industrial production and prices. (More details on the construction of the variables are reported in the Data Appendix).

Although in a mechanical sense it is easy to generate these historical EU-11 series, the Lucas Critique is an obvious factor in interpreting the validity of these calculations. The relationships between money stocks and income, and money stocks and interest rates, are unlikely to be completely invariant to the exchange rate regime. Indeed, monetary union would hardly be worthwhile if all these relations did remain unaltered.

The manner in which this issue will be addressed is by examining the pre-1999 data separately from the post-euro data. Specifically, regressions will be estimated on both the full sample and pre-euro samples. In order to assess the robustness of the models, the estimates obtained from the pre-euro sample will be used to conduct out-of-sample exercises.

***The EMS Crises.*** In any set of estimates spanning the 1990s, one must account for possible breaks in the long run relationship due to the 1992-1993 currency crises. A commonplace observation is that movements in the dollar exchange rate have different impacts upon the constituent currencies of the EMS. One might think that such effects would be particularly pronounced during the crises episodes. Hence, the regressions will be

augmented by a dummy variable taking a value of unity from the beginning of the August 1992 crises to the September 1993 French devaluation. The dummy is treated as exogenous, which is plausible insofar as one considers the *timing* of the crises as independent of the included macro variables.<sup>6</sup>

### 3.3 Econometric Methodology

The series being analyzed are integrated of order one [ $I(1)$ ]. Hence the appropriate estimation methodology involves testing for cointegration. As discussed in Phillips and Loretan (1991), one can proceed along a number of avenues. The main choice is between single equation methods and multiple equation methods. In this study the both approaches are utilized. The multi-system approach is appropriate because one is interested not only in how exchange rates adjust, but also how the *other* macroeconomic variables in the system adjust. For instance, one might think it plausible that central banks adjust interest rates in order to affect exchange rates; or that inflation rates respond to changes in exchange rates; or both could be true.<sup>7</sup>

The systems approach is, however, susceptible to biases in all equations if but one equation is misspecified. Hence, the system approach puts a premium on adequately modeling each and every endogenous variable in the system. More confidence in the estimates can be derived from single equation procedures, in those cases where single equation procedures are

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<sup>6</sup> However, the conditions that make possible the crises are likely to involve the variables in the cointegrating relation. For instance, movements in relative nontradables prices may have been important factors; see Froot and Rogoff (1991). I thank Anders Vredin for making this observation.

<sup>7</sup> Expanding upon McCallum (1994), Meredith and Chinn (1998) suggest that central banks target interest rates in response to inflation and output gaps that are in turn affected by exchange rate movements.

valid. The key requirement is that the right hand side variables be weakly exogenous with respect to the parameters of the cointegrating vector. The systems-based estimates will be used to determine how valid this assumption is.

The standard in testing for cointegration in time series is the full-system maximum likelihood estimation technique of Johansen (1988) and Johansen and Juselius (1990). In the current context, this procedure involves estimating the system:

$$\begin{aligned}
\Delta s_t &= \gamma_{10} + \Phi_1 ECT_{t-1} + \sum_{i=1}^k \gamma_{1i} \Delta s_{t-i} + \sum_{i=1}^k \zeta_{1i} \Delta \hat{m}_{t-i} + \sum_{i=1}^k v_{1i} \Delta \hat{y}_{t-i} \\
&\quad + \sum_{i=1}^k \xi_{1i} \Delta \hat{i}_{t-i} + \sum_{i=1}^k \mu_{1i} \Delta \hat{\pi}_{t-i} + \sum_{i=1}^k \chi_{1i} \Delta \omega_{t-i} + \varepsilon_{1t} \\
\Delta \hat{m}_t &= \gamma_{20} + \Phi_2 ECT_{t-1} + \sum_{i=1}^k \gamma_{2i} \Delta s_{t-i} + \sum_{i=1}^k \zeta_{2i} \Delta \hat{m}_{t-i} + \sum_{i=1}^k v_{2i} \Delta \hat{y}_{t-i} \\
&\quad + \sum_{i=1}^k \xi_{2i} \Delta \hat{i}_{t-i} + \sum_{i=1}^k \mu_{2i} \Delta \hat{\pi}_{t-i} + \sum_{i=1}^k \chi_{2i} \Delta \omega_{t-i} + \varepsilon_{2t} \\
\Delta \hat{y}_t &= \gamma_{30} + \Phi_3 ECT_{t-1} + \sum_{i=1}^k \gamma_{3i} \Delta s_{t-i} + \sum_{i=1}^k \zeta_{3i} \Delta \hat{m}_{t-i} + \sum_{i=1}^k v_{3i} \Delta \hat{y}_{t-i} \\
&\quad + \sum_{i=1}^k \xi_{3i} \Delta \hat{i}_{t-i} + \sum_{i=1}^k \mu_{3i} \Delta \hat{\pi}_{t-i} + \sum_{i=1}^k \chi_{3i} \Delta \omega_{t-i} + \varepsilon_{3t} \\
\Delta \hat{i}_t &= \gamma_{40} + \Phi_4 ECT_{t-1} + \sum_{i=1}^k \gamma_{4i} \Delta s_{t-i} + \sum_{i=1}^k \zeta_{4i} \Delta \hat{m}_{t-i} + \sum_{i=1}^k v_{4i} \Delta \hat{y}_{t-i} \\
&\quad + \sum_{i=1}^k \xi_{4i} \Delta \hat{i}_{t-i} + \sum_{i=1}^k \mu_{4i} \Delta \hat{\pi}_{t-i} + \sum_{i=1}^k \chi_{4i} \Delta \omega_{t-i} + \varepsilon_{4t} \\
\Delta \hat{\pi}_t &= \gamma_{50} + \Phi_5 ECT_{t-1} + \sum_{i=1}^k \gamma_{5i} \Delta s_{t-i} + \sum_{i=1}^k \zeta_{5i} \Delta \hat{m}_{t-i} + \sum_{i=1}^k v_{5i} \Delta \hat{y}_{t-i} \\
&\quad + \sum_{i=1}^k \xi_{5i} \Delta \hat{i}_{t-i} + \sum_{i=1}^k \mu_{5i} \Delta \hat{\pi}_{t-i} + \sum_{i=1}^k \chi_{5i} \Delta \omega_{t-i} + \varepsilon_{5t} \\
\Delta \omega_t &= \gamma_{60} + \Phi_6 ECT_{t-1} + \sum_{i=1}^k \gamma_{6i} \Delta s_{t-i} + \sum_{i=1}^k \zeta_{6i} \Delta \hat{m}_{t-i} + \sum_{i=1}^k v_{6i} \Delta \hat{y}_{t-i} \\
&\quad + \sum_{i=1}^k \xi_{6i} \Delta \hat{i}_{t-i} + \sum_{i=1}^k \mu_{6i} \Delta \hat{\pi}_{t-i} + \sum_{i=1}^k \chi_{6i} \Delta \omega_{t-i} + \varepsilon_{6t} \\
ECT &\equiv [\beta_1 s + \beta_2 \hat{m} + \beta_3 \hat{y} + \beta_4 \hat{i} + \beta_5 \hat{\pi} + \beta_6 \omega]
\end{aligned} \tag{7}$$

where the carats (^) denote relative differences. For simplicity, (7) is written assuming only one cointegrating vector, although in principle nothing prevents the existence of multiple cointegrating vectors.

The procedure yields a trace statistic on which a likelihood ratio (LR) test can be

conducted for the null of  $r$  cointegrating vectors against the alternative of  $m$  cointegrating vectors. The asymptotic critical values for this test are reported in Osterwald-Lenum (1992). Cheung and Lai (1993a), among others, have shown that finite sample critical values may be more appropriate given the relatively small samples which are generally under study. In this study, inferences will be made using both sets of critical values.

As mentioned earlier, a key advantage of the multivariate approach is that it enables one to examine what variables do the adjustment to restore equilibrium. The typical focus on a single equation, normalized on the exchange rate, may miss such adjustment mechanisms. For instance, the exchange rate might not exhibit any response to the disequilibrium represented by the error correction term, and so the cointegrating relationship may not be detected.

An information based method is used to select the lag length for the Johansen procedure. VARs in levels are estimated, with lags extending up to 6. The optimal lag length  $k$  is selected on the basis of the Akaike Information Criterion; in general either 4 or 5 lags are indicated. Hence, the corresponding VECMs used in the Johansen procedure incorporate 4 lags of first difference terms.

## **4. EMPIRICAL ANALYSIS**

### **4.1 Data**

The analysis is conducted on monthly data over the period 1991M01 to 2000M01, with the US data drawn from the IMF's *International Financial Statistics* November 1999 CD-ROM. Exchange rates are period-average, in US\$/euro. Money is narrow money (*IFS* line 34) or broad money (M2 or M3). Income is proxied by industrial production or interpolated

real GDP. Three month offshore euro rates are used for EU-11 interest rates, and the Fed Funds rate for the United States. Inflation rates are calculated as the annualized first difference of the log-CPI. Since one does not observe a pure measure of either  $p^N$  or  $p^T$ , these measures are proxied by the CPI and PPI respectively (as discussed in Section 2). Further details are contained in the Data Appendix.

## 4.2 Empirical Results

The cointegration results are reported in Panel 1 of Table 2. In all cases, one can reject the no-cointegration null, and indeed can find multiple cointegrating vectors. However, using the finite sample adjustment suggested by Cheung and Lai (1993a) one typically finds only a single cointegrating vector.<sup>8</sup>

The results of estimating the system over the 1991M08-1999M12 period is reported in column [1]; a constant in the cointegrating vector, but no time trends in the data, is assumed. GDP (interpolated) is used as the income variable. The estimates are broadly in line with the monetary model. Increases in the money supply depreciate the currency, while higher income appreciates the currency.<sup>9</sup> Aside from money and income, the most statistically significant coefficient is on inflation. A one percentage point increase in inflation weakens the currency by 11%, slightly higher than Hayo's (1998) estimate of the Euroland semi-elasticity of money

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<sup>8</sup> Omitting one of the error correction terms will reduce the efficiency of the estimates, but will not bias the results.

<sup>9</sup> The coefficient estimate corresponding to the income elasticity of money demand is about twice as large as anticipated. This finding could be due to imperfect substitution between US and Euroland goods (Stockman, 1980), although the finding of PPP is not consistent with this interpretation. An alternative interpretation is that GDP (if viewed as being measured in units of "tradable" goods) incorporates the relative price variable. I thank Karlhans Sauernheimer for this observation.

demand of approximately 10.<sup>10</sup> Notably, interest rates do not enter into the equation in a significant manner. Furthermore, the nontradables price variable does not enter with the correct sign, although it is barely significant.

These results are virtually unchanged if deterministic time trends are allowed in the data (column [2]). However, the results are very sensitive to whether the relative price variable is included or not. Omission of this variable results in an essentially zero coefficient on money, and positive signs on both interest rates and inflation rates.

Qualitatively similar estimates are obtained if industrial production is used instead of GDP, although the money stock coefficient is now closer to the posited value of unity – 0.69. The income coefficient is also closer to the posited value of unity that is associated with income elasticity of money demand (see Hayo, 1998).<sup>11</sup> Finally, the relative price coefficient is no longer statistically significant. It does not appear that the allowance of time trends in the data changes the results in any appreciable way.

While the identified cointegrating vector does confirm the monetary model as a long run relationship, it is not informative about the manner in which the equilibrium is restored. Panel 2 of Table 2 presents the reversion coefficient estimates from the vector error correction models (i.e., the  $\Phi$ 's from equation 7). Interestingly, a consistent pattern appears: the bulk of the adjustment is undertaken by exchange rates and inflation. Further, if industrial production is used as a scale variable, then industrial production also responds. Thus a

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<sup>10</sup> Note that under the sticky price formulation, the inflation coefficient is the interest semi-elasticity *plus*  $1/\theta$ , where  $\theta$  is the rate of exchange rate reversion to its long run value.

<sup>11</sup> Coenen and Vega (1999) argue that a stable relationship between M3 and income and short and long term interest rates exist for the euro area. However, in all cases, use of M3 (as well as M2) yields a statistically significant *negative* coefficient on money. Hence, estimation of the exchange rate equations using this monetary aggregate is not pursued.

depreciated exchange rate (relative to the fundamentals) induces an expansion in industrial production. The contrasting effect on GDP (which is unresponsive to exchange rate disequilibria) can be rationalized by considering the fact that the exchange rate sensitive component of GDP is much smaller than that of industrial production.

A fascinating aspect of the results, in both the long term and short term, is that the interest differential fails to exhibit the expected effect upon the exchange rate. That is, one expects, in the context of the sticky-price monetary model, that an increase in the interest differential should appreciate the currency. No statistically significant evidence is found for this result, in contrast to earlier studies.

Since the landscape of the empirical literature is littered with failed models that fit well in-sample, it behooves the researcher to test the model out of sample. This wariness is heightened by the obvious regime shift which occurred with the actual introduction of the euro in January 1999. In this next subsection, the model is assessed in this light.

### **4.3 Is the Euro's Recent Behavior Inexplicable?**

The first step in answering the question posed is to assess whether omitting post-euro data changes the estimates substantially. Table 3 reports results for regressions restricted to the 1991M08-1998M12. The parameter estimates are similar to those obtained from the full sample; however, a more relevant test is whether the estimates obtained from the pre-euro period predict out of sample. In figure 3, the euro and predicted values from dynamic simulations (using GDP as the scale variable) are presented. One set of values (LXEU1F0) is from the model without a time trend assumed in the data, while the other (LXEU1F2) does allow for time trends. Assumption of a time trend clearly yields an improved fit, relative to the no-trend fit. Interestingly, in both cases, the value of the euro at the end of 1998 is not

unnaturally high. Furthermore, regardless of the specification selected, the euro's value is overpredicted by approximately 15% in January 2000.

If the dynamic simulation is started in January 1999, then the results depicted in Figure 4 are obtained. Using the no-trend assumption, dynamic forecast LXEU1F1 is the result, implying a 13% undervaluation. Matters are not helped if one allows for deterministic trends in the data (LXEU1F3): then the implied undervaluation is some 15%.

To investigate whether the implied undervaluation is sensitive to the manner in which the other endogenous variables are treated, the exchange rate is also modeled using a single equation error correction model (ECM). This approach is appropriate if the exchange rate responds to the disequilibrium embodied in the error correction term, as it does in this sample. It also requires that the other variables do not; in this case, this condition does not hold exactly. In particular, the inflation rate does appear to respond to the disequilibrium, as indicated in both Tables 2 and 3, so it is not strictly valid to estimate only the marginal ECM. Regardless, the results of estimating these ECMs are reported in the first two columns of Table 4.

A parsimonious specification estimated over the 1991M08-1998M12 period<sup>12</sup> yields an out-of-sample misprediction of 13% (LXEU1F5). Hence, the results do not appear to be dependent upon the particular estimation methodology. Interestingly, the results do not seem to vary depending upon whether one takes the run-up in value of the euro prior to its inception as exceptional or not.<sup>13</sup> If the ECM (and the corresponding error correction term) is

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<sup>12</sup> This ECM includes 2 lags of the exchange rate difference and the relative price, 3 lags of the other variables, and no lags of the inflation rate. The error correction term is that indicated in Table 3, column [1].

<sup>13</sup> Demertzis and Hughes Hallett (1999) argue that the 1:1 conversion rate of euro's for ecu's induced an artificial overvaluation in the run-up to EMU, given the value of the *FT*'s synthetic

instead estimated over the 1991M08-97M12 period, and the out-of-sample dynamic simulation initiated at 1998M01, then the model predicts an even greater overvaluation of 18% (LXEU1F7). It must be noted, however, that the long run restriction does not seem to exert a statistically significant reversion effect on the euro over the sample period, so that one should not place too much weight on this prediction. After all, visual inspection of the Figures 3-5 does seem to suggest that some of the euro's movements in 1998 are not being captured by the included macroeconomic variables.

Finally, one can assess whether the estimation error involved in identifying the long run relationships is at fault. In columns [3] and [4], the first-differences specification results are reported. The coefficients are correctly signed. Estimation using more appropriate instrumental variables techniques yields, as is commonly the case, insignificant estimates, so these OLS estimates are relied upon for the out of sample forecasts. The implied undervaluation, using the estimated equation in column [3], is 14% in December 1999, while that in column [4] is 10%.

To sum up, a carefully specified regression analysis tends to confirm the first impression that the euro is currently undervalued, at least when assessed using the conventional macroeconomic measures. Even treating the 1998 appreciation as an anomalous event does not overturn this result. A "best-case" scenario indicates that the October 1998 euro was only 6% overvalued, suggesting that the subsequent decline in the euro's value has placed it substantially below its long run value.<sup>14</sup>

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euro.

<sup>14</sup> As of 5 April 2000, the US-German differential on 2 year bonds was 195 basis points (*Economist*, April 8<sup>th</sup> - 14<sup>th</sup>, 2000). If uncovered interest parity holds, then this means the market is anticipating a 2% per annum appreciation of the euro over the next two years. This amount of appreciation would go less than halfway in eliminating the estimated undervaluation on the order of

#### 4.4 A Digression: Equity Markets and the euro

A recent vogue in financial research has attributed the weakness of the euro to the strength in the United States equity markets (e.g., Bogler, 2000; *Economist*, March 18<sup>th</sup>, 2000, page 81). Superficially, there does appear to be a correlation between the US equity price index and the euro/dollar exchange rate over the 1999M01-M12 period. The elasticity of the exchange rate with respect to the US equity index is -0.81, with a standard error of 0.15, and adjusted  $R^2$  of 0.77. However, even a cursory investigation indicates that the correlation is much less pronounced in the pre-1999 period. It is important to recall that the bulk of accumulated evidence indicates that these two variables are  $I(1)$ , so one must be on guard for spurious correlation. A Johansen test fails to reject the null hypothesis of no cointegration between these two variables at anywhere near the conventional levels.

At the very least, it is the *relative* performance of the US and Euroland equity markets should enter into the equation. In Figure 6, the equity indices (relative to the US) for Germany and for France are plotted against the euro/dollar rate. The negative relationship holds true for the US-German equity indices; that is, more rapid equity price increases in the US, relative to Germany, leads to a stronger dollar, vis a vis the euro. However, the story relies upon a specific choice of the equity market. Using the French equity market index yields a substantially different correlation – positive instead of negative.

Extensive searching did not reveal robust evidence of cointegration between the monetary fundamentals, augmented with relative returns on equities. To the extent that relative returns enter into the cointegrating relationship, they enter in *positively* for 1999. That is, higher US returns induce a *weaker* dollar. If this factor explains the behavior of the euro

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10%.

over its first year of existence, then alternative data sets and methods will probably be required to uncover the evidence.

## **5. A LONG-TERM PERSPECTIVE**

### **5.1 Overview**

The presence of long term swings in exchange rates far away from their (econometrically implied) fundamental values is a widely held belief, even if the econometric evidence is less than fully persuasive. If one believes that in the long term, the fundamentals will assert themselves, then one must take a stand on *which* fundamentals are relevant.

The long term analysis used here takes as its starting point the detection of a role of the relative price of nontradables in the real exchange rate. That is, the real exchange rate is modeled as, primarily, a function of the relative tradables/nontradables price, which in turn depends upon the relative tradables/nontradables productivity differential. This framework is popularly known as the Balassa-Samuelson approach.

While this approach may seem detached from the concerns of real-world analysts, this appearance is deceiving. For instance, Rosenberg (2000) ascribes the weakness of the euro to the deficiencies in Euroland manufacturing “competitiveness” and rigidities in labor markets. This focus is consistent with the Balassa-Samuelson approach, to the extent that higher productivity levels in manufacturing vis a vis nontraded services will tend to lead to a stronger currency.

Before proceeding, it may be useful to comment on alternative long term approaches. In the behavioral equilibrium exchange rate approach (e.g., Clark and MacDonald, 1999), relationships between net foreign asset positions and exchange rates are identified. In the

fundamental equilibrium medium-term approach of the Faruquee, Isard and Masson (1999), equilibrium exchange rates are backed out from medium term current account balances. Finally, in Aglietta, Baulant and Coudert (1997), an eclectic approach incorporating several of the above elements is used to model the behavior of the euro.

These approaches are not necessarily inconsistent with the Balassa-Samuelson framework. For instance, net foreign asset positions may move in anticipation of expected future changes in relative productivity levels. However, the Balassa-Samuelson approach would take productivity as exogenous, and the net foreign asset positions as endogenous. My view is that it is more aesthetically pleasing to relate the exchange rate to a more causal variable, than to another endogenous variable. Of course, there is a tradeoff between aesthetics and empirical realities, and others may justifiably opt for the use of net foreign asset positions.<sup>15</sup>

## 5.2 Framework and Estimates

The approach adopted here requires a statement about the determinants of the relative price of nontradables. Assume in the long run, supply side factors dominate so that the relative price of nontradables move inversely with the relative productivity levels:

$$p_t^{N,j} - p_t^{T,j} = a_t^{T,j} - a_t^{N,j} \quad (8)$$

Then substituting this expression in for the relative prices in (2) yields:

$$q_t = -\alpha[(a_t^{T,US} - a_t^{N,US}) - (a_t^{T,euro} - a_t^{N,euro})] \quad (9)$$

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<sup>15</sup> Wingle (2000) finds that in a panel regression for OECD countries, inclusion of both productivity levels and net foreign asset positions yields essentially a zero coefficient on the latter.

In estimating this relationship, one is once again constrained by the data limitations. Sectoral data on productivity, on a cross-country basis, is not widely available. Using information drawn from the OECD's 1997 edition of the *International Sectoral Data Base* (ISDB), the data on total factor productivity (TFP) levels are aggregated into two categories: tradables -- roughly manufacturing and other industries -- and nontradables. In the latter case, nontradables are taken to be services. (See the Data Appendix for details). The analysis is restricted to the post-Bretton Wood period (1975-1992).

In principle, one could use the Johansen procedure. However, estimation of the full system would rapidly exhaust the degrees of freedom, for each currency. Hence, resort is made to panel DOLS, as described by Kao and Chiang (1998). This means equation (9) is estimated, allowing for fixed effects, and incorporating currency-specific short run dynamics,

$$q_{it}^T = \alpha_i + \beta_1 \hat{a}_{it}^T + \beta_2 \hat{a}_{it}^N + \sum_{j=-2}^{+1} \xi_{ij} \Delta \hat{a}_{it+j}^T + \sum_{j=-2}^{+1} \zeta_{ij} \Delta \hat{a}_{it+j}^N + u_{it} \quad (10)$$

The equation is also augmented by one additional regressor – the real price of oil. The oil price variable is allowed country-specific slope coefficients.

The results of estimating this equation for France, Germany, Italy and the Netherlands (FR, GY, IT and NE) are reported in Table 5. In column [1], the basic specification is presented. A one percentage point increase in the US-Euroland tradable productivity differential results in a 1.7 percentage point appreciation in the US dollar. This slope coefficient is statistically significant,<sup>16</sup> although it is hard to justify in the context of the

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<sup>16</sup> The inferences are conducted using heteroskedasticity-consistent standard errors, although in principle one might want to use long-run standard errors (as in Mark and Sul, 1999). Since the amount of serial correlation remaining in the residuals is fairly small – usually not statistically significant – it is not clear that the conclusions would change after making the suggested

Balassa-Samuelson model. The nontradable productivity coefficient should have an opposite sign. In this case, the estimate is not statistically significant. The failure of nontradable productivity to evidence the posited sign is probably due to the large measurement errors entailed in measuring real nontradables output (specifically services).

Constraining the tradable and nontradable coefficients to have equal and opposite signs yields the estimates in column [2]; the implied nontradables share is 0.85. Apparently, most of the explanatory power of productivity comes from movements in tradable productivity. Omitting nontradable productivity (in column [3]) yields the statistically significant estimate of -1.528, which is very similar to that obtained in column [1].

These very high estimates for the impact of tradable productivity have been remarked upon in a more general context of OECD real exchange rates (and in particular, for G-7 rates) in Chinn and Johnston (1999). They speculate that productivity has impacts on the relative price of traded goods, above and beyond the effect on the relative price of nontradables.<sup>17</sup> To investigate whether this is the case, the tradables-price deflated real exchange rate is regressed upon tradable productivity. The results are presented in column [4]; the absolute value of the point estimate is substantially lower than the corresponding estimate for the real exchange rate: 0.916 vs. 1.703. Hence, tradable productivity appears to have independent effects upon both the relative price of traded goods in common currency terms, and the relative price of nontradables.<sup>18</sup>

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adjustments.

<sup>17</sup> For the G-7 countries, Chinn and Johnston (1999) report a point estimate for the tradables productivity coefficient of -0.73 (0.33 standard error) and, for nontradables, of 0.44 (0.32 standard error). Hence, the estimates for the productivity effect are larger for this set of countries than for the G-7.

<sup>18</sup> This conclusion may appear at odds with the finding of relative PPP for PPIs in Section 2. However, recall that the sample period used in this analysis extend back to before the 1987-99

For the purposes of this study, what is of most interest is the impact of productivity on the real exchange rate, regardless of the exact channel of effects. Roughly speaking, the estimates bracket [0.85, 1.70].

### 5.3 Conditional Forecasts

Using these parameter estimates to make forecasts is an exercise fraught with hazards. One key difficulty is that the publication of TFP figures -- disaggregated to sectors -- usually lags years behind events. Hence, the 1997 *ISDB* contained data on TFP only up to 1992 for most countries. For recent trends in productivity, the best one can do is to use labor output per man hour in manufacturing, and assume that the intercountry nontradables productivity differential ( $a^{N,US} - a^{N,euro}$ ) stays constant.

Over the 1990-98 period, US manufacturing labor productivity has risen by about 3.27% per year (in log terms), while the corresponding values for Germany, France, Italy and the Netherlands are 3.12%, 3.88%, 2.99% and 3.72% respectively (US Department of Labor, 1999). The economies of these four countries represent roughly three-quarters of Euroland GDP, so omission of the other Euroland economies should not be problematic. Aggregating by their shares in the euro, one finds that there is almost no difference in movements in US and Euroland manufacturing productivity (Figure 7). Based on historical patterns, the linearly extrapolated trends do not indicate a closing of the gap. Thus, in the absence of a dramatic change in productivity trends, the long term prospects of the real value of the euro are neither rosy, nor gloomy. That is, the real value of the euro will continue to “trend sideways”. On the

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monthly sample. Over the later period, manufacturing productivity trends were roughly equal (see Figure 7), so that even if such tradable productivity did have an effect on relative prices of traded goods, such effects would not necessarily be evident in the data.

other hand, if with EMU labor productivity in manufacturing picks up substantially relative to that in the US, such that the differential widens to 1%, then the euro should be expected to begin appreciating by between 1 and 1.5 percent per year.<sup>19</sup>

## 6. CONCLUSIONS

Recent commentary regarding the euro's behavior can be placed into two categories: those that ascribe its recent weakness to strong macroeconomic conditions in the United States, combined with a certain wariness regarding how expansionary future monetary policy will be in the -- present and future -- Euroland; and those that view the 1998 appreciation of the euro as an aberration and hence the subsequent depreciation as a resumption of the trend nominal depreciation in the euro.

This study indicates that there is an element of truth to both views. However, the most notable aspect of the simulations is that in almost all cases the euro appears approximately 10-15% below its equilibrium value (as of December 1999), when assessed using historical relationships. It must be stressed, however, that in the short history of the euro, the weakness of the currency may be due to expectations regarding events not easily captured by the historical behavior of macroeconomic variables such as money stocks and interest rates. Such factors include the new operating procedures of the ECB, expectations regarding expansion of EMU to Greece, and other EU candidates, and the differential impacts of the "New Economy" in the US and the "Neuer Markt" in Euroland.

Finally, there does not appear to be substantial evidence for a drastic change in the

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<sup>19</sup> Coppel, Durand and Visco (2000) interpret an acceleration in productivity growth due to more rapid restructuring as decreasing Euroland NAIRU by 2 percentage points (increasing potential GDP growth in Euroland by approximately 0.2 percentage points). However, they do not provide an estimate of the resulting appreciation of the euro.

trend in the euro's value. Even if the productivity growth differential were to widen to one percentage point in Euroland's favor – plausible but difficult to envision given recent US output performance – the best one could expect is between a 0.85 to 1.7 trend appreciation in the real value of the euro.

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## Data Appendix

The data for Sections 2 through 4 are from IMF, *International Financial Statistics*, November 1999 CD-ROM., and internal databases supplied by V. Coudert (Bank of France) and B. Schnatz (Bundesbank).

Bilateral exchange rate in US\$/euro, period average, from the Pacific website (<http://pacific.commerce.ubc.ca/xr/>), for 1993M01-2000M01. Pre-1999M01 data are “synthetic euro” exchange rates. For 1990M01-1994M12, “synthetic euro” exchange rates are created using the conversion weights, and period average exchange rates from *IFS*, line *ah*.

Narrow money is national concept M1 (not seasonally adjusted) for the EU-11, from the BIS. For the US, national definition M1 (seasonally adjusted) from *IFS*. EU-11 broad money is M2 (not seasonally adjusted), from the Annex to the February 1999 issue of the *ECB Monthly Bulletin* (Table 4) updated with the March 2000 *ECB Monthly Bulletin*. The alternative broad money is the harmonized concept M3 (not seasonally adjusted), from the BIS. For the US, national definition of M2 (seasonally adjusted), from *IFS*, and the alternative broad measure is the national definition of M3 (seasonally adjusted) from the St. Louis Federal Reserve Bank’s *FRED* system.. The EU-11 M1, M2 and M3 monetary aggregates are seasonally adjusted using multiplicative X-11.2.

EU-11 GDP is interpolated from quarterly data from the BIS, in billions of 1995 euros. Data is interpolated by use of quadratic matching to average.

EU-11 industrial production is from the BIS, *IFS*, line 66..c for 1998.01 onward. From 1980.01-1997.12, EU-11 industrial production is national industrial production (not seasonally adjusted) from *IFS* line 66, seasonally adjusted using multiplicative X-11.2, aggregated using conversion rates. US industrial production (seasonally adjusted), *IFS* line 66*eyc*.

The EU-11 interest rate are offshore 3 month deposit rates, from BIS up 1998.12 and Banque de France thereafter, and *ECB Monthly Report* from 1999M06 onward. The US interest rate is the Fed Funds rate, *IFS* line 60*b*.

Equity prices are indices from *IFS*, line 62.

The EU-11 CPI is the harmonized index of consumer prices, from the BIS, for 1995.01 onward. Prior to that, national CPIs (*IFS* line 64) are aggregated using the conversion weights, and spliced to the BIS series. The US CPI is *IFS* line 64.

The EU-11 PPI is the national producer price index, from BIS. For the *IFS* line 63, 1990 = 100.

Inflation is the first difference of  $\log(\text{CPI})$ , annualized.

The relative price variable is defined as:

$$\tilde{\omega} \equiv \log(CPI^{US}/PPI^{US}) - \log(CPI^{euro}/PPI^{euro})$$

The data for Section 5 were drawn from *IFS* and from the OECD's *International Sectoral Data Base (ISDB)*.

Exchange rates are bilateral, period average, *IFS* line *rf*. The real exchange rate is defined using aggregate GDP deflators, or (for the tradable real exchange rate  $q^T$ ) the tradable sector price deflator.

Price deflators and total factor productivity (TFP) series are from *ISDB*. Traded sectors: mining, manufacturing, agriculture and transportation. Nontraded: all other services, construction.

Oil Prices are crude petroleum export prices, in 1990 US\$, calculated using *IFS* line *76aad* and US CPI, *IFS* line 64.

**Table 1**  
**Tests for Purchasing Power Parity**  
**Dynamic OLS Estimates: 1987M06-1999M12**

$$s_t = \gamma_0 + \gamma_1 p_t^{US} + \gamma_2 p_t^{euro} + \sum_{k=-2}^2 \theta_k \Delta p_{t+k}^{US} + \sum_{k=-2}^2 \Theta_k \Delta p_{t+k}^{euro} + u_t$$

Var.	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]
Deflator	CPI	CPI	CPI	CPI	PPI	PPI	PPI	PPI
$p$ (1)	<b>-1.696</b> <b>(1.497)</b>	1.911 (1.299)	0.315 (0.872)	-0.456 (1.703)	1.251 (1.265)	0.957 (0.966)	1.608** (0.799)	1.958** (0.998)
$p^{euro}$ (-1)	<b>2.164</b> <b>(1.691)</b>	<b>1.474</b> <b>(1.236)</b>	-0.315 (0.872)	0.456 (1.703)	-1.105 (1.525)	<b>1.714</b> <b>(1.197)</b>	-1.608** (0.799)	-1.958** (0.998)
time (?)		-0.008*** (0.001)		0.000 (0.000)		-0.003*** (0.001)		-0.000 (0.000)
N	149	149	149	149	149	149	149	149

Notes: Point estimates from Dynamic OLS regressions(robust standard errors in parentheses, calculated using average of 4 autocorrelations and Bartlett window). \*(\*\*)[\*\*\*] indicates significance at the 10%(5%)[1%] marginal significance level for null of zero coefficient. Figures in **bold face italics** denote estimates significantly different at the 10% level from anticipated values.

**Table 2**  
**The Monetary Model of Exchange Rates**  
**Johansen Cointegration Results: 1991M08-1999M12**

Panel 2.1: Long Run Coefficients				
	[1]	[2]	[3]	[4]
LR	158.0	136.8	162.4	133.4
c.v.	111.0[144.3]	103.2[134.2]	111.0[144.3]	103.2[134.1]
CR's	2[1]	2[1]	3[1]	2[0]
$m^{US}-m^{euro}$	0.396*** (0.086)	0.396*** (0.086)	0.687*** (0.088)	0.658*** (0.082)
$y^{US}-y^{euro}$	-2.219*** (0.478)	-2.217*** (0.480)	-0.754** (0.291)	-0.703*** (0.278)
$i^{US}-i^{euro}$	0.968 (1.195)	0.947 (0.556)	0.129 (0.808)	-0.084 (0.791)
$\pi^{US}-\pi^{euro}$	10.797*** (2.302)	10.881*** (2.319)	13.626*** (3.181)	13.542*** (3.044)
$\tilde{\omega}$	2.057** (0.898)	2.070** (0.902)	1.323 (1.132)	1.268 (1.082)
Panel 2.2: Reversion coefficients				
$\Delta s$	-0.186*** (0.064)	-0.186*** (0.064)	-0.140*** (0.055)	-0.127*** (0.061)
$\Delta(m^{US}-m^{euro})$	-0.021 (0.033)	-0.021 (0.033)	-0.005 (0.028)	-0.033 (0.029)
$\Delta(y^{US}-y^{euro})$	0.003 (0.006)	0.003 (0.006)	-0.132*** (0.032)	-0.099*** (0.033)
$\Delta(i^{US}-i^{euro})$	0.011 (0.010)	0.011 (0.010)	-0.003 (0.008)	-0.002 (0.009)
$\Delta(\pi^{US}-\pi^{euro})$	0.033*** (0.006)	0.033*** (0.006)	0.033*** (0.006)	0.025*** (0.006)
$\Delta\tilde{\omega}$	0.001 (0.013)	0.001 (0.013)	0.001 (0.013)	-0.005 (0.011)
adj-R <sup>2</sup>	.27	.27	.19	.18
SER	0.019	0.019	0.020	0.020
trend	no	in data	no	in data
k	5	5	5	5

Notes: LR is the likelihood ratio test statistic for the null of zero cointegrating vector against the alternative of one. c.v. is the 1% asymptotic critical value for the test of zero cointegrating vectors against the alternative of one [finite sample critical values in brackets]. CR's is the number of cointegrating relations implied by the asymptotic critical values [finite sample critical values]. Coefficients are long run parameter estimates from the Johansen procedure described in the text.  $k$  is the number of lags in the VAR specification of the system. Panel 2 Reversion coefficients are the  $\Phi$  coefficients from equation (6). SER is the standard error of the regression. VECM Adj-R<sup>2</sup> and SER refer to the exchange rate error correction model.  
a/ One cointegrating vector at 5% MSL, using finite sample critical values.

**Table 3**  
**The Monetary Model of Exchange Rates**  
**Johansen Cointegration Results: 1991M08-1998M12**

Panel 3.1: Long Run Coefficients				
	[1]	[2]	[3]	[4]
LR	157.2	129.6	159.7	124.5
c.v.	111.0[167.5]	103.2[155.6]	111.0[167.5]	103.2[155.6]
CR's	3[0]a	1[0]	3[0]a	1[0]
$m^{US}-m^{euro}$	0.514*** (0.077)	0.507*** (0.077)	0.504*** (0.136)	0.489*** (0.117)
$y^{US}-y^{euro}$	-1.762*** (0.398)	-1.756*** (0.398)	-1.706** (0.561)	-1.225*** (0.423)
$i^{US}-i^{euro}$	1.310 (0.630)	1.284 (0.631)	0.348 (1.241)	-0.459 (1.133)
$\pi^{US}-\pi^{euro}$	11.113*** (2.384)	11.130*** (2.384)	16.248*** (4.212)	16.909*** (4.585)
$\tilde{\omega}$	2.135*** (0.777)	2.120*** (0.775)	1.658 (1.723)	1.622 (1.476)
Panel 3.2: Reversion coefficients				
$\Delta s$	-0.116 (0.079)	-0.142** (0.075)	-0.134*** (0.041)	-0.130*** (0.050)
$\Delta(m^{US}-m^{euro})$	-0.001 (0.030)	0.008 (0.029)	0.012 (0.016)	-0.006 (0.018)
$\Delta(y^{US}-y^{euro})$	0.017 (0.007)	0.016 (0.007)	-0.110*** (0.025)	-0.084*** (0.029)
$\Delta(i^{US}-i^{euro})$	0.007 (0.012)	0.007 (0.012)	0.004 (0.006)	0.005 (0.008)
$\Delta(\pi^{US}-\pi^{euro})$	0.035*** (0.007)	0.035*** (0.007)	0.012*** (0.004)	0.020*** (0.005)
$\Delta\tilde{\omega}$	0.002 (0.014)	0.003 (0.014)	0.006 (0.007)	0.004 (0.009)
adj-R <sup>2</sup>	.21	.29	.27	.25
SER	0.020	0.019	0.019	0.019
trend	no	in data	no	in data
k	5	5	5	5

Notes: LR is the likelihood ratio test statistic for the null of zero cointegrating vector against the alternative of one. c.v. is the 1% asymptotic critical value for the test of zero cointegrating vectors against the alternative of one [finite sample critical values in brackets]. CR's is the number of cointegrating relations implied by the asymptotic critical values [finite sample critical values]. Coefficients are long run parameter estimates from the Johansen procedure described in the text.  $k$  is the number of lags in the VAR specification of the system. Panel 2 Reversion coefficients are the  $\Phi$  coefficients from equation (6). SER is the standard error of the regression. VECM Adj-R<sup>2</sup> and SER refer to the exchange rate error correction model. a/ One cointegrating vector at 5% MSL, using finite sample critical values.

**Table 4**  
**The Monetary Model of Exchange Rates**  
**Alternative Specifications and Sample Periods**

Panel 4.1: Long Run Coefficients				
	91M08-98M12	91M08-97M12	91M08-98M12	91M08-97M12
	[1]	[2]	[3]	[4]
LR	157.2	166.3		
c.v.	111.0[167.5]	102.1[155.6]		
CR's	3[0]a	3[0]a		
$m^{US}-m^{euro}$	0.514*** (0.077)	0.188 (0.206)		
$y^{US}-y^{euro}$	-1.762*** (0.398)	-1.033** (0.502)		
$i^{US}-i^{euro}$	1.310 (0.630)	-2.580 (1.900)		
$\pi^{US}-\pi^{euro}$	11.113*** (2.384)	25.142*** (8.139)		
$\bar{\omega}$	2.135*** (0.777)	7.155** (3.586)		
Panel 4.2: Reversion coefficient and short run dynamics				
$\Phi$	-0.114*** (0.054)	-0.053 (0.038)		
$\Delta s$	0.232	0.430		
$\Delta(m^{US}-m^{euro})$	0.608	0.545	0.159 (0.306)	0.204 (0.324)
$\Delta(y^{US}-y^{euro})$	-0.125	-0.677	-1.924** (0.786)	-2.318*** (0.754)
$\Delta(i^{US}-i^{euro})$	-1.715	0.214	-2.417*** (0.719)	-2.274*** (0.751)
$\Delta(\pi^{US}-\pi^{euro})$	-	-	1.208 (1.304)	1.097 (1.337)
$\Delta\bar{\omega}$	-1.259	1.305	-1.197** (0.555)	-1.017 (0.658)
CRISIS	-0.016 (0.010)	-0.019** (0.009)	-0.013 (0.009)	-0.012 (0.009)
adj-R <sup>2</sup>	.35	.36	.21	.22
SER	0.018	0.018	0.020	0.020
trend	no	no		
k	5	5		

Notes: Columns [1] and [2]: LR is the likelihood ratio test statistic for the null of zero cointegrating vector against the alternative of one. c.v. is the 1% asymptotic critical value for the test of zero cointegrating vectors against the alternative of one [finite sample critical values in brackets]. CR's is the number of cointegrating relations implied by the asymptotic critical values [finite sample critical values]. Coefficients in Panel 1 are long run parameter estimates from the Johansen procedure described in the text.  $k$  is the number of lags in the VAR specification of the Johansen estimation (see text for lags in single-equation ECM). Panel 2 Reversion coefficients are the  $\Phi$  coefficients from the exchange rate equation of (6), Panel 2 short run dynamics are the sum of the statistically significant coefficients of lagged first difference terms. Columns [3] and [4]: short run dynamics are the coefficients for contemporaneous coefficients in the first difference specification. CRISIS is the coefficient on the indicator variable for EMS crises.

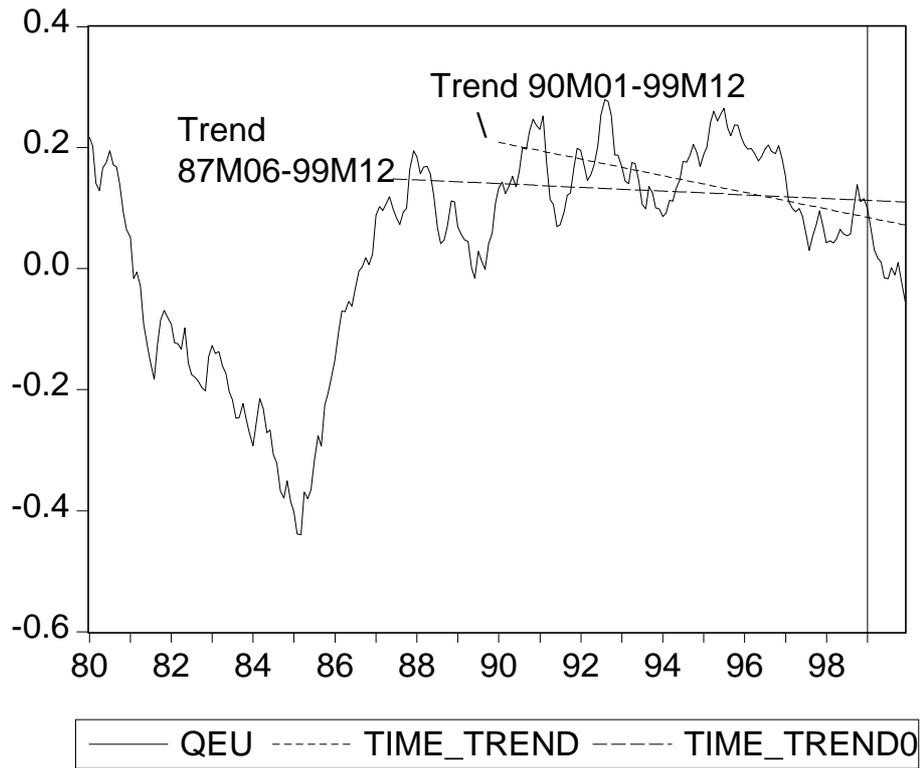
a/ One cointegrating vector at 5% MSL, using finite sample critical values.

**Table 5**  
**Panel Regressions:**  
**Real Exchange Rates and Relative Productivity Differentials**

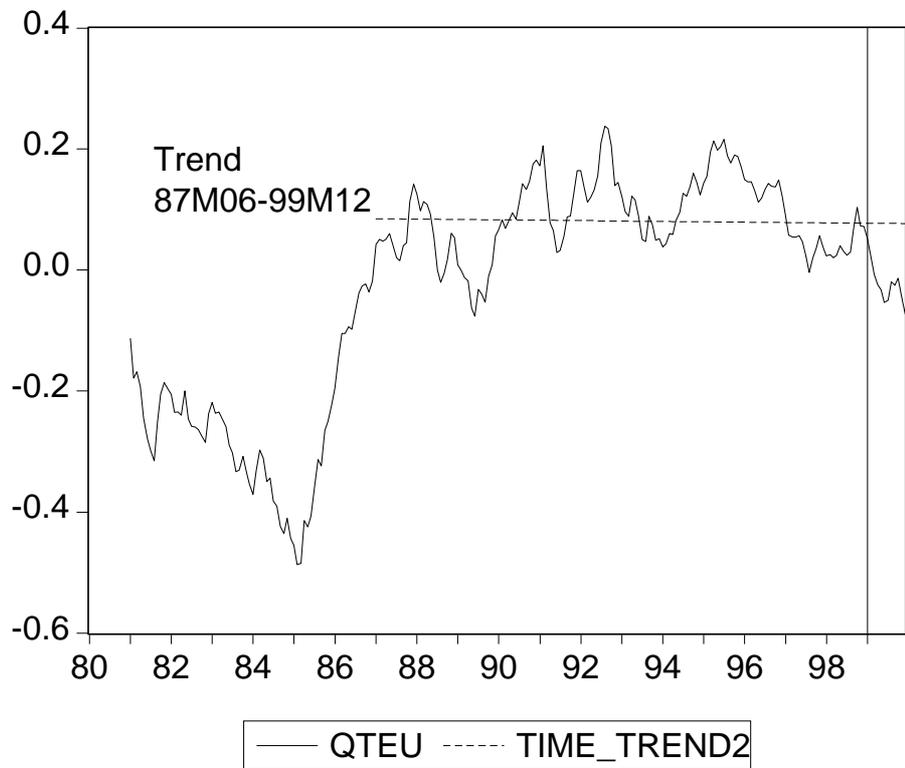
$$q_{it} = \alpha_i + \beta_1 \hat{a}_{it}^T + \beta_2 \hat{a}_{it}^N + \sum_{j=-2}^{+1} \xi_j \Delta \hat{a}_{it+j}^T + \sum_{j=-2}^{+1} \zeta_j \Delta \hat{a}_{it+j}^N + u_{it}$$

Variable	[1]	[2]	[3]	[4]
$a^T$	-1.703** (0.136)	-0.847*** (0.117)	-1.528*** (0.190)	-0.916*** (0.169)
$a^N$	-0.112 (0.888)	+0.847*** (0.117)	--	--
$p^{oil}$ (FR)	-0.075 (0.285)	-0.373*** (0.082)	-0.164* (0.083)	-0.319*** (0.087)
$p^{oil}$ (GY)	-0.064 (0.256)	-0.317** (0.113)	-0.187** (0.084)	-0.391*** (0.087)
$p^{oil}$ (IT)	-0.170 (0.248)	-0.430*** (0.116)	-0.147 (0.092)	-0.348*** (0.096)
$p^{oil}$ (NE)	-0.038 (0.240)	-0.274*** (0.105)	-0.134 (0.084)	-0.308*** (0.087)
N	72	72	72	72

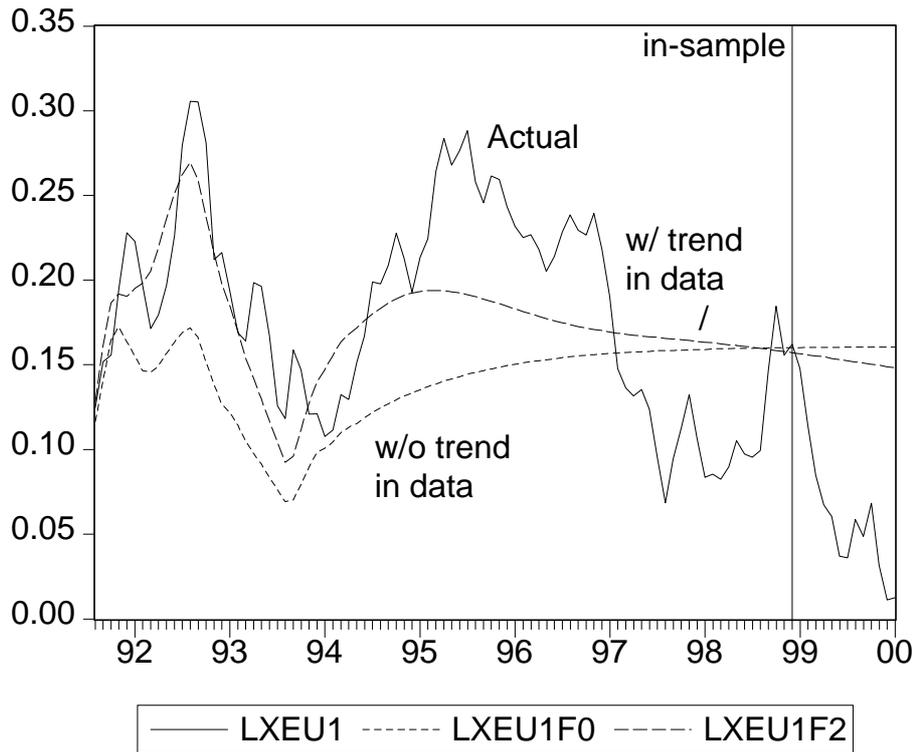
Notes: Dependent variable: GDP deflator real exchange rate (columns [1]-[3]); tradables price index real exchange rate (column [4]). Panel DOLS(1,2) parameter estimates. No leads and lags of real oil price included. Heteroskedasticity consistent standard errors in parentheses. \*(\*\*)[\*\*\*] denotes significance at 10%(5%)[1%] significance level.



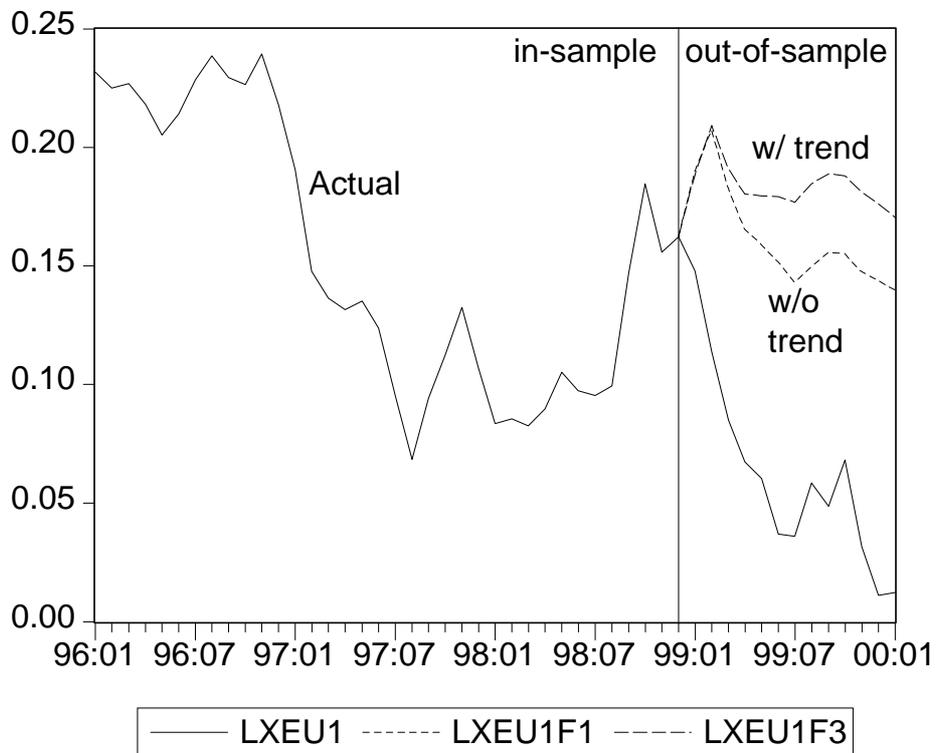
**Figure 1:** CPI deflated real euro and trends



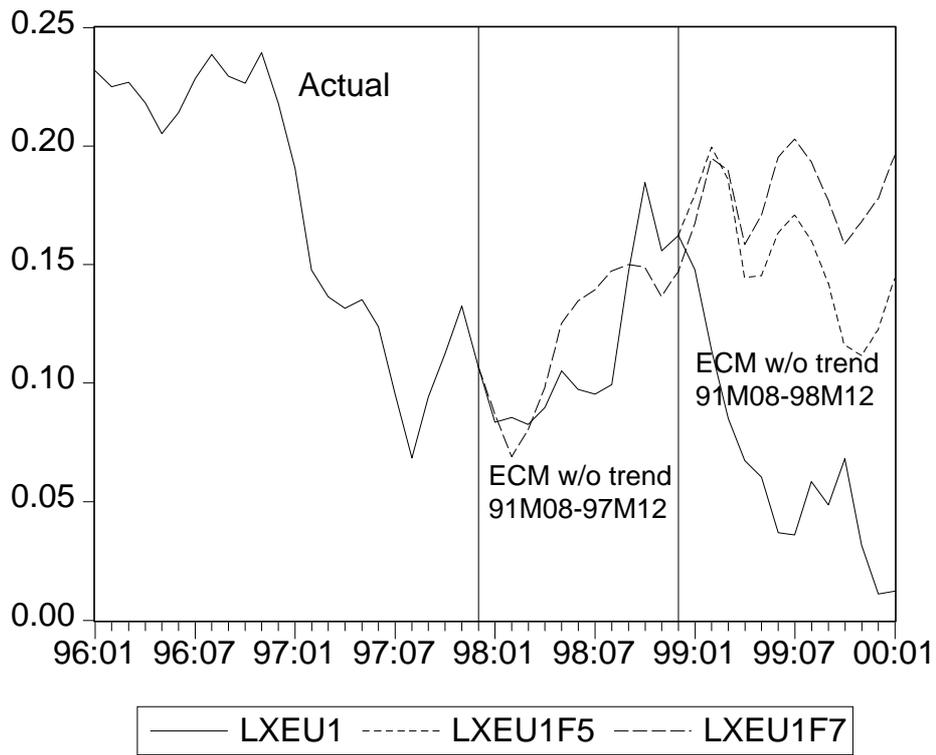
**Figure 2:** PPI deflated real euro and trend



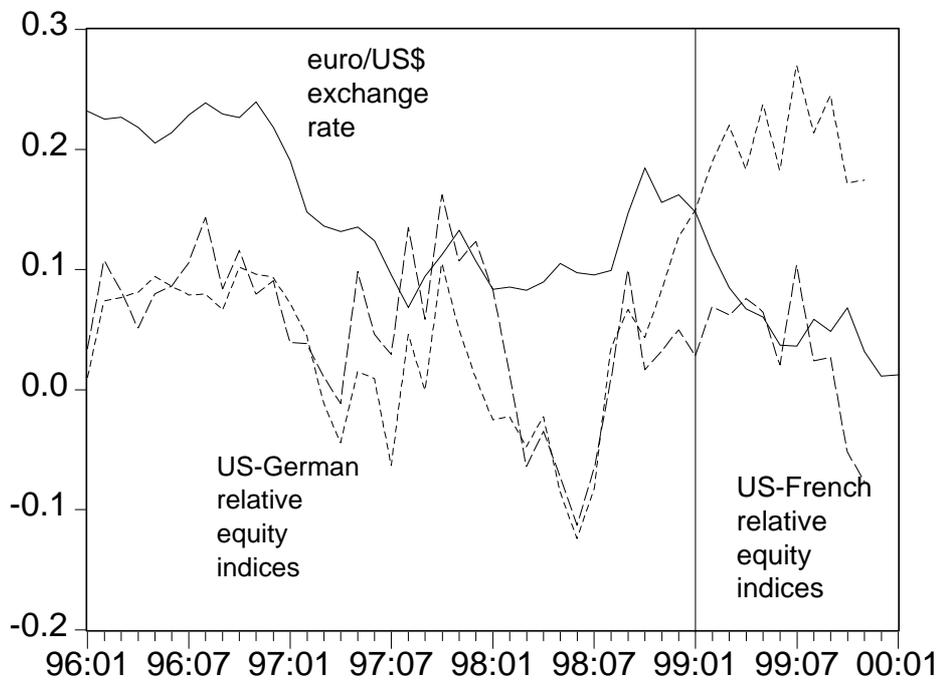
**Figure 3:** Dynamic simulations of monetary model (using GDP).



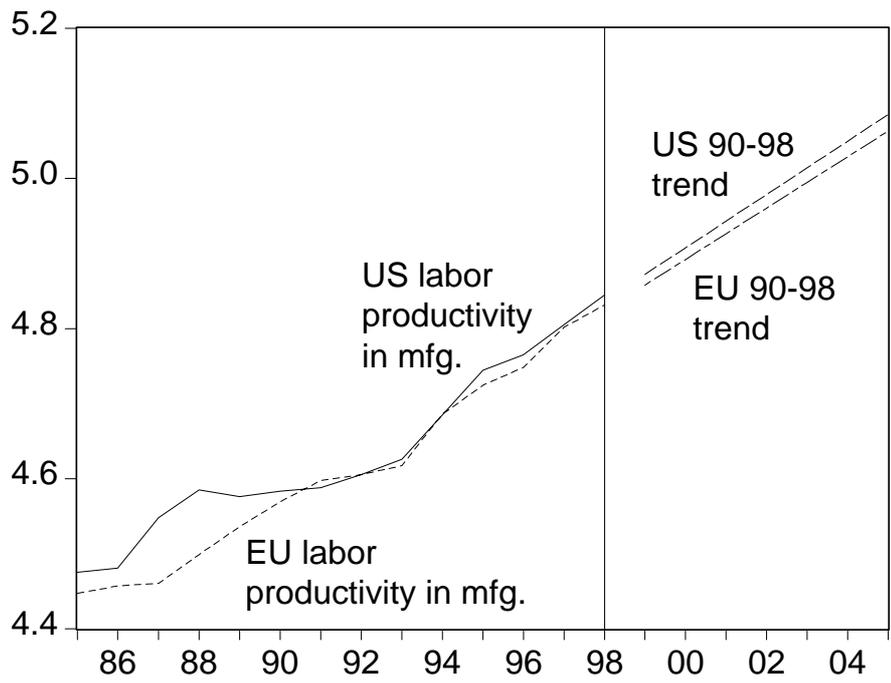
**Figure 4:** Dynamic simulations of monetary model (using GDP).



**Figure 5:** Single equation dynamic simulations



**Figure 6:** Euro/dollar exchange rate versus US-German and US-French relative equity indices



**Figure 7:** Trend labor productivity in manufacturing, US and Euroland.  
 Source: US Dept. of Labor and author's calculations.