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Abstract

The US Federal Reserve's new relaxed monetary policy (the so-called *quantitative easing*) has triggered controversy among economists and policy makers about its effectiveness. This paper investigates the role of monetary policy in managing the euro – dollar exchange rate *via* alternative cointegration tests and impulse response functions. It is found that monetary fundamentals have neither long- nor short-run impact on the exchange rate. This implies that the Fed's quantitative easing schemes are unlikely to have any significant impact on the euro – dollar rate.

Keywords: Exchange rates; Monetary model; Cointegration; Impulse response functions
JEL classification: F31; E52

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1. Introduction

The euro – dollar exchange rate is of great importance given the leading international status of both currencies. Euro depreciated since its introduction (on January 4, 1999) steadily and without any major interruption until February 2002. Then it began to rise against dollar almost smoothly and reached a height of €0.6254 to 1 US$ on July 15, 2008. With the outburst of the global financial crisis, euro initially depreciated against the dollar, but it almost regained its value by November 2009 (€0.665/US$). Since then, the value of the euro first depreciated and then picked-up again (see Figure 1).

This paper attempts to examine the behaviour of the euro – dollar exchange rate by utilising the monetary approach to exchange rate (MAER) determination. The MAER emerged as the dominant exchange rate model at the outset of the recent float in the early 1970s and remains an important exchange rate paradigm. Despite its questionable forecasting ability (as first suggested by Meese and Rogoff 1983), this approach retains its appeal as it provides the theoretical framework for investigating the role of monetary policy in managing exchange rates.[1] This is a topical issue since the talk of a *currency war* is still making headlines as various economies attempt to find ways to weaken their currencies, and thus make their exports more competitive in the struggle for economic recovery. There is a widespread view, for example, that the US Federal Reserve's policy of *quantitative easing* (i.e., printing money to buy bonds) in the post-2009 period has reduced the value of the dollar relative to other currencies (such as the euro), whose volume remains constant or rises more slowly.[2] Although this may be true in the very short-term, visual inspection of Figure 1 suggests that the influence of the Fed's quantitative easing schemes of the last two years on the euro – dollar exchange rate is less than clear cut for longer horizons.

[Figure 1]

Our major concern in this paper is to empirically estimate the all-out dynamic of the euro – dollar exchange rate in relation to its fundamentals by discerning the long-run effects of monetary policy on the exchange rate from its short-run impacts. In this regard, we employ both cointegration techniques and impulse response functions (IRFs) to analyse the long- and short-term dynamic model for the euro – dollar rate. This paper possesses two main novelties. First, it investigates not only the long-run relationship between the

[1] For a review of the recent literature on the empirical validity of the MAER see Beckmann *et al.* (2010).

[2] The Fed launched its quantitative easing programme in January 2009. The European Central Bank officially denies embarking on quantitative easing.

euro – dollar rate and its fundamentals, but also the short-run dynamics as traced by the time-profile and response trajectories of the exchange rate to the shocks to the innovations of variables under consideration. This is particularly important since the Fed has only recently adopted quantitative easing schemes whose impact on the euro – dollar exchange rate has not been yet formally tested. Second, the paper uses exclusively post-1999 data. A number of articles that utilise the MAER approach use either euro synthetic data (van Aarle *et al.* 2000; Frenkel and Koske 2004; Altavilla 2008), or Deutschmark data (Beckmann *et al.* 2010) for the period before the introduction of the common currency. According to our opinion, this strategy masks the fact that the euro area countries did not necessarily pursue a common monetary policy prior to 1999. The exclusive use of actual euro market data overcomes the problem of diversified monetary policies, and thus allows us to accurately examine the (possible) linkages between the euro – dollar exchange rate and its fundamentals.

The rest of the paper is structured as follows: The MAER and the methodologies employed are sketched in Section 2. The data set is described and the empirical results are discussed in Section 3. Section 4 concludes.

2. Methodological considerations

The MAER was developed after the collapse of the fixed exchange rate system in the 1970s. Several versions have been put forward that can be broadly classified into two main types of models: (a) the flexible-price monetary model (Frenkel, 1976; Bilson, 1978), and (b) the sticky-price monetary model (Dornbusch, 1976) and the real interest rate differential model (Frankel, 1979). Whichever model that one adheres to, the clear implication is that monetary policy is the most effective means of managing the exchange rate. The MAER can be expressed through the following reduced form:

$$s_t = c + \gamma_1 \left(m_t - m_t^*\right) + \gamma_2 \left(y_t - y_t^*\right) + \gamma_3 \left(i_t - i_t^*\right) + \gamma_4 \left(\pi_t^e - \pi_t^{e*}\right) + \varepsilon_t \qquad (1)$$

where s_t is the spot exchange rate (price quotation), m_t $\left(m_t^*\right)$ is the domestic (foreign) money supply, y_t $\left(y_t^*\right)$ is the domestic (foreign) real income, i_t $\left(i_t^*\right)$ is the domestic (foreign) interest rate, π_t^e $\left(\pi_t^{e*}\right)$ is the domestic (foreign) expected inflation rate, c is a constant and ε_t is a white noise error. Table 1 summarizes the direction in which the explanatory variables are expected to influence the exchange rate with respect to the aforementioned versions of the MAER.

[Table 1]

Equation 1 outlines the long-run relationship among the variables under consideration. To examine the long-run validity of the MAER, we implement three alternative cointegration techniques. The first technique, developed by Johansen and Juselius (1990) and Johansen (1991), applies maximum likelihood to a vector autoregressive (VAR) model assuming that the errors are Gaussian. Defining $X = [s, (m - m^*), (y - y^*), (i - i^*), (\pi^e - \pi^{e*})]'$ as the vector of endogenous variables, the error-correction form of the VAR can be written as:

$$\Delta X_t = \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Pi X_{t-1} + \varepsilon_t \qquad (2)$$

The rank of matrix *Π* (which shows the number of cointegrating vectors) is determined by means of two likelihood ratio tests: the *trace* statistic and the *maximum eigenvalue* statistic.

The second technique, known as the Autoregressive Distributed Lag (ARDL) bounds approach to cointegration, was introduced by Pesaran *et al*. (2001). The advantage of this approach is that it does not require any unit root pretesting of the variables and it can be applied whether the variables under consideration are integrated of order one or zero or even fractionally cointegrated. To implement the bounds test let a vector $\xi_t = (s_t, Z_t')'$ where s_t is the exchange rate and Z_t is the vector of regressors. The error correction representation of the ARDL specification model is given by:

$$\Delta s_t = a + \lambda_1 s_{t-1} + \sum_{i=2}^{5} \lambda_i' Z_{t-1} + \sum_{i=1}^{p} \theta_i \Delta s_{t-i} + \sum_{j=0}^{q} \varphi_j' \Delta Z_{t-j} + \varepsilon_t \qquad (3)$$

where λ_1 and λ_i are long-run multipliers, a is the constant, and ε_t are white noise errors. The test for the absence of a long-run relationship between s_t and Z_t entails the following null hypothesis:

$H_0 : \lambda_1 = 0,\ \lambda_i = 0$ for i=2, … , 5

Pesaran *et al*. (2001) provide the critical values for this *F*-test. If the computed *F*-statistic is above (below) the upper (lower) bound critical value the null hypothesis of no cointegration is rejected (accepted). If the *F*-statistic falls within the lower and upper bounds, the results are inconclusive.

A major shortcoming of both techniques is that they do not account for structural changes in the cointegrating vector. Thus, we also use the Gregory and Hansen (1996) approach that tests for cointegration with a one shift in the cointegrating vector at some unknown date. Here we consider a level shift model (model 2 in Gregory and Hansen) which (following the former notation) takes the form:

$$s_t = \mu_1 + \mu_2\varphi_t + \sum_{i=1}^{4} a_i' Z_t + u_t \qquad (4)$$

where $\varphi_t = 0$ if $t \leq [n\tau]$ and $\varphi_t = 1$ otherwise, τ is an unknown parameter denoting the timing of the change point and **[ ]** denotes the integer part. In equation (4), μ_1 is the intercept before the shift and μ_2 is the change in the intercept due to the shift. Gregory and Hansen provide three tests which are modified versions of the Z_a and Z_t (Phillips, 1987) and the *ADF* statistics. It should be noted here that the underlying motivation of Gregory and Hansen's methodology is not the estimation of the break date *per se*; instead the focus is on improving the power of conventional cointegration tests by allowing for a structural change.

Information regarding the existence of cointegration is crucial if one wants to accurately estimate the short-run dynamic relationships between the euro – dollar rate and its fundamentals. The most convenient way to characterise these relationships is to simulate their responses to unanticipated shocks in each of the variables (impulse response functions - IRFs). IRFs capture dynamic behaviour as they trace the effect of an exogenous shock to a variable on current and future values of another variable. IRFs are calculated from the moving average representation of the VAR model:

$$X_t = \sum_{i=0}^{\infty} A_i \varepsilon_{t-i} \qquad (5)$$

where X_t is the vector of the jointly determined dependent variables and the coefficient matrices A_i are recursively calculated using the following expression:

$$A_i = \varphi_1 A_{i-1} + \varphi_2 A_{i-2} + \ldots + \varphi_\rho \mathrm{A}_{i-\rho}, \quad i = 1, 2, \ldots,$$

with $A_0 = I_X$ and $A_i = 0$ for $i<0$.

Following Pesaran and Shin (1998), the scaled generalized IRF (GIRF) of variable X_i with respect to a standard error shock in the j^{th} equation can be defined as:

$$GIRF\left(X_{it}, X_j, h\right) = \frac{A_h \sum e_j}{\sqrt{\sigma_{jj}}}, h = 0, 1, 2, \ldots$$

where σ_{jj} is jj^{th} element in the variance – covariance matrix and e_j is $m \times 1$ vector with unity at its j^{th} row and zeros elsewhere. GIRFs are unique and do not require the prior orthogonalisation of the shocks. In contrast, the widely used orthogonalised IRFs are dependent on the choice of the *a priori* ordering for the variables in the Choleski decomposition.

GIRFs can be derived from two types of VAR models. One is a standard VAR in levels (if all variables are stationary), or in first differences (if all variables are non-stationary but not cointegrated). The other is a vector error-correction model (VECM) that explicitly models non-stationary variables and cointegrating relationships that are present in the data.

3. Empirical findings

The data originate from the IMF's International Financial Statistics (IFS) database. The data are of monthly frequency spanning from 1999M01 to 2010M11. Exchange rates are monthly averages in terms of euro/US$. The chosen monetary aggregates are narrow money stock (M1). Industrial production indices are used as proxies for real income. Interest rates are monthly averages of short-term market rates. Preceding twelve months growth in consumer price indices is used as a measure of the unobservable expected inflation rate. All variables, except for interest rates and expected inflation rates, are expressed in natural logarithms.

Table 2 reports the outcome of the Augmented Dickey – Fuller (ADF) tests (Dickey and Fuller, 1979). The tests are carried out in models with and without a (linear) trend, in addition to the intercept term. The trend appears to be both numerically and statistically insignificant in almost all instances. Therefore, we confine our attention to the intercept only case. The test results show that all, but one, variables are stationary in first differences (i.e. *I(1)*). The expected inflation differential is the only variable that exhibits strong stationarity in levels (*I(0)*).[3]

[Table 2]

Next, we test for cointegration using Johansen and Juselius (1990) procedure (Table 3 – Panel A). Both the trace and the maximum eigenvalue statistics indicate the existence of at most one cointegrating equilibrium relationship between the euro – dollar exchange rate and its fundamentals at the 10% significance level. Nevertheless, we suspect that this single cointegrating vector is spanned by the stationarity of the expected inflation differential. Formal testing of this hypothesis verifies our *a priori* conjecture (χ^2 = 3.044, *p*-value = 0.550). In essence, this finding indicates that there is no long-run relationship among the variables under consideration. The ARDL approach to cointegration further supports this outcome (Table 3 – Panel B). The computed *F*-statistic (=2.582) is less than the corresponding lower bound

[3] We also performed Elliott *et al*. (1996) DF-GLS tests with corrected mean. The test results verify the outcome of the reported ADF tests. For brevity the results are not reported but they are available upon request.

critical value (=2.86). Therefore, we accept the null of no long-run relationship. However, this conclusion might be misleading if the cointegrating relationship has shifted over time due to a structural change. To this end, we employ the procedure proposed by Gregory and Hansen. The alternative test results are reported in Table 3 – Panel C. Again, all tests fail to reject the null of no cointegration.[4,5] Conclusively, the alternative cointegration test results uniformly suggest that there is no long-run equilibrium relationship between the euro – dollar exchange rate and its fundamentals in the post-1999 period.

[Table 3]

In the final stage of our analysis, we trace the short-run impact of a one time shock to the macroeconomic fundamentals on the euro – dollar exchange rate. For it, we proceed by estimating the corresponding GIRFs. Given the absence of cointegration, GIRFs are derived from the residuals of a stable VAR(2) model where all variables (except expected inflation differential) appear in first-difference form.[6] Figure 2 shows the response of the euro – dollar exchange rate (along with their respective bootstrapped 95% confidence intervals) to innovations of one standard deviation in the fundamentals. A 20-month horizon is considered. The upper left graph in Figure 2 illustrates the GIRF of the exchange rate subject to a shock in the relative money supply. It is evident that the effect is initially negative (i.e. euro appreciates against the US$) but becomes positive from the third month onwards. Nevertheless, this exchange rate response is statistically insignificant since it lies within the confidence intervals which uniformly evolve around the zero axis. Similarly, the responses of the exchange rate to a shock to relative income, interest rate differential and expected inflation differential are quite small and non-significant in almost all cases. These findings suggest the absence of any significant short-run effects of shocks in macroeconomic fundamentals to the euro – dollar exchange rate.

[Figure 2]

4. Concluding remarks

This paper uses the MAER to investigate the role of monetary policy in managing the euro – dollar exchange rate during the recent past. Contrary to

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[4] We also tested for cointegration using Gregory and Hansen's Models 3 (level shift with trend) and 4 (regime shift). Again, all tests indicate no cointegration.

[5] Beckmann *et al.* (2010) use the Bai and Perron (1998, 2003) sequential procedure and find one structural break in the 2000s (2004:11). The ADF* test of the Gregory and Hansen (1996) procedure (Table 3 – Panel C) identifies one structural break in 2005:04. Nevertheless, accounting for this break does not seem to alter the evidence of no-cointegration.

[6] The estimated VAR(2) does not suffer from serial correlation and satisfies the stability condition since all inverse roots of the characteristic AR polynomial have modulus less than one.

previous studies, we also consider short-run dynamics together with the (more conventional) long-run relationship between the exchange rate and its underlying monetary/macroeconomic fundamentals. The cointegration test results do not provide any support for the long-run properties of the monetary approach. Similarly, the generalized impulse response functions show that the exchange rate response to monetary shocks is small and insignificant. Thus, the Fed's new relaxed monetary policy is unlikely to have any significant impact on the euro – dollar rate.

Overall, the outcome of our analysis corroborates the so-called exchange rate disconnect puzzle. Therefore, further research is needed in order to explore the way that expectations of exchange rates are formed. Alternatively, modelling the euro – dollar exchange rates in a linear fashion may be inadequate. In this sense the results from this study set the stage for future research.

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Table 1. Alternative hypotheses on the coefficients of the MAER

Coefficients	*(m-m*)* γ_1	*(y-y*)* γ_2	*(i-i*)* γ_3	$(\pi^e-\pi^{e*})$ γ_4
Frenkel	+1	-	0	+
Bilson	+1	-	+	0
Dornbusch	+1	-	-	0
Frankel	+1	-	-	+

Table 2. ADF test results

I. *Levels*	Intercept	Intercept and trend	
	t-ADF	t-ADF	Trend
s	-1.107	-2.697	$4\text{x}10^{-5}$ [0.01]
$(m - m^*)$	-2.036	-0.104	$5\text{x}10^{-5}$ [0.61]
$(y - y^*)$	-2.401	-2.429	$-7\text{x}10^{-6}$ [0.43]
$(i - i^*)$	-2.553	-2.601	$3\text{x}10^{-6}$ [0.56]
$(\pi - \pi^*)$	-3.749*	-3.879*	$3\text{x}10^{-6}$ [0.31]

II. *First differences*	Intercept
	t-ADF
s	-8.785*
$(m - m^*)$	-10.900*
$(y - y^*)$	-15.871*
$(i - i^*)$	-7.911*

Notes: The ADF tests are based on parsimonious ADF models that were derived by minimising the SIC, starting from a generous lag length of 13. Panel I: Columns 2 and 3 report the t-ADF values with intercept and intercept & trend, respectively. Column 4 presents the estimated coefficients of the (linear) trend and the associated p-values (in brackets).
* Rejection of the null at the 5% level.

Table 3. Cointegration test results

Panel A: Johansen and Juselius procedure

	λ_{trace}	$\lambda_{\max}$	LM(1)	H_0: [0, 0, 0, 0, 1]
r=0	68.876 (0.059)	32.700 (0.068)	27.991 (0.308)	3.044 (0.550)
r<=1	36.176 (0.387)	16.472 (0.625)		

Panel B: ARDL approach

$F[s \mid (m-m^*),(y-y^*),(i-i^*),(\pi^e-\pi^{e*})]$	LM(1)
2.582	0.790 (0.373)

Panel C: Gregory and Hansen procedure (Model 2)

*ADF**	Z_a^*	Z_t^*
-3.893 [2005M04]	-26.413 [2002M08]	-3.924 [2002M07]

Notes: The figures in parentheses denote *p*-values. All computations are carried out in EViews 5.0.

Panel A: The *trace* and *max. eigenvalue* statistics are based on an unrestricted VAR(2). We allow for the presence of an intercept, but no time trend, in the cointegrating equations. LM(1) tests for serial correlation of 1st order in the unrestricted VAR. H_0 tests for the stationarity of the expected inflation differential and is distributed as $\chi^2(4)$.

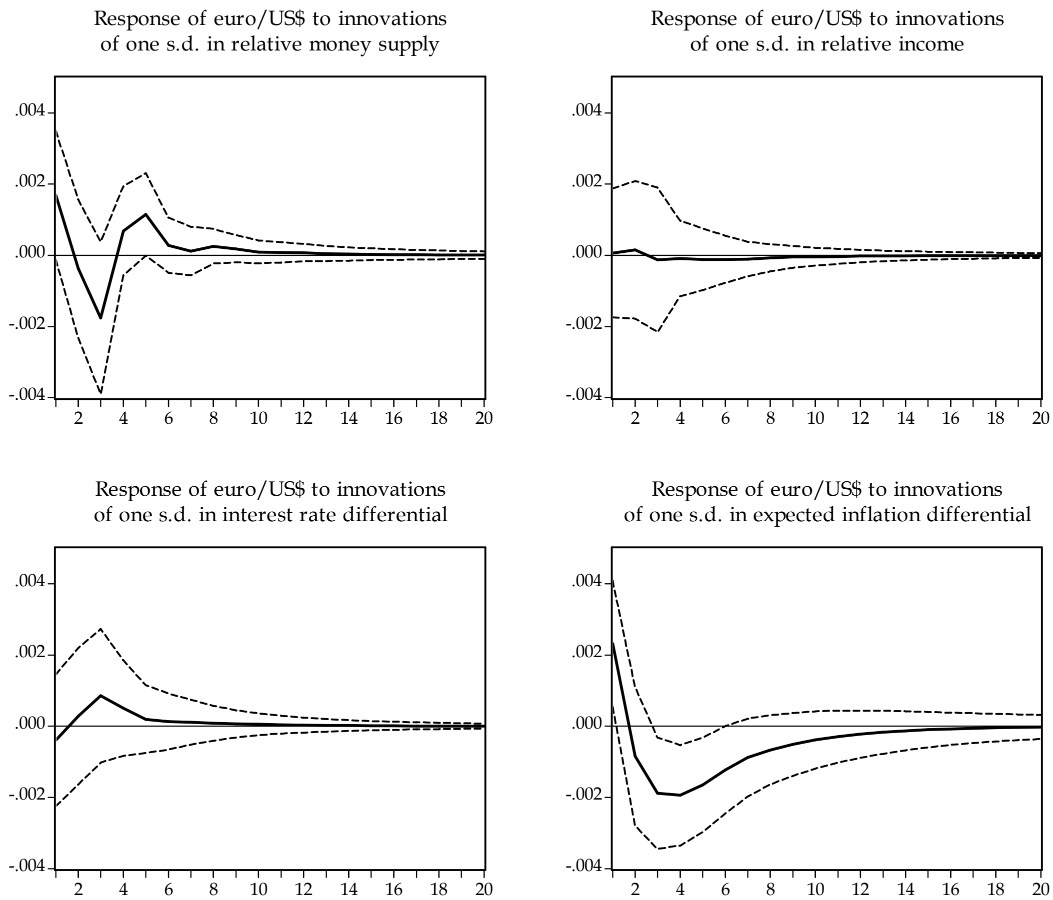
Panel B: The number of lags in eq. (3) is set equal to 1. LM(1) tests for serial correlation of 1st order. The 5% critical values (Case III: unrestricted intercept and no trend) for k=4 are 2.86 (lower bound) and 4.01 (upper bound).

Panel C: The 5% critical values for m=4 are -5.56, -59.40 and -5.56 for ADF*, Z_a* and Z_t*, respectively. The numbers in brackets are the estimated structural break dates [year/mm].

Figure 1. The euro – dollar exchange rate: 1999M01 – 2010M11

Notes: The shaded area represents the period of the Fed's embarking on quantitative easing.

Figure 2. Generalised Impulse Response Functions



Notes: The dotted lines represent the 95% confidence intervals. Monte Carlo simulations using 1,000 draws are performed to compute the error confidence bands.