Long-run money demand in the new EU member states with exchange rate effects

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Abstract: Besides the analysis of a wide range of other economic and financial indicators, money is highly relevant in the monetary strategy of the European Central Bank (ECB). However, the prominent role for money requires a stable money demand function. This analysis is of importance for the new member countries of the European Union (EU) as well, as they will are expected to join the euro area in some years. In this study a money demand analysis is conducted using panel cointegration methods. A well-behaved long run money demand relationship can be identified only if the exchange rate as part of the opportunity cost is included. In the long run vector, the income elasticity exceeds unity. Moreover, the US-Dollar exchange rate seems to be more appropriate than the euro.

1   Introduction

In May 2004 the European Union (EU) enlarged by eight Central and Eastern European countries and two mediterranean countries. These countries agreed in a treaty that they will join the exchange rate mechanism (ERM) II and introduce the euro, if a country fulfils the convergence criteria of the Maastricht treaty. These conditions include ceilings for inflation and interest rates, budget deficits and government debt, and exchange rate stability. After the euro introduction the Governing Council of the ECB will also get the responsibility of monetary policy for these countries. The ECBs monetary policy is embedded in a monetary strategy.

In 1999, the ECB (1999) published a monetary strategy, which contains two pillars to conduct its monetary policy. The ECB gave the monetary aggregate M3 a prominent role in the first pillar. In Spring 2003 the ECB redefined its strategy. The discussion in the ECB ended up with a confirmation of a two pillar strategy (ECB 2003). The first pillar contains the economic analysis of price risks in the short term. The second pillar includes the monetary analysis of price risks in the medium term and long-run, where the monetary aggregate M3 has again a prominent role. Hence, this monetary aggregate remain important for forecasting price risks. In line with the second pillar, money demand analysis of the new EU members will become more relevant. For the monetary strategy a stable long run money demand function is required, where money demand is linked to other macroeconomic variables like income and interest rates.

In the period of transition, foreign determinants can also play a crucial role in explaining money demand. During periods of high inflation, the Central and Eastern European countries experienced a partial replacement of domestic by foreign currencies, either as a store of value or a medium of exchange. Therefore the exchange rate is presumably important to money demand behaviour in these states. As the euro was introduced to the public not before 1999, euro and US-dollar rates are considered alternatively in the analysis.

In order to estimate long-run money demand functions for the new member countries, cointegration techniques are employed (see for example Engle, Granger 1987). As most countries are transition economies, they have to manage enormous structural changes. Hence it is difficult to obtain data for a long period. The estimated parameters, which

are based on a short period, are not very reliable. Evidently, estimates for long-run parameters require data for a long period. Alternatively, the sample can be extended, if the information of all countries is pooled. This is done by panel integration and cointegration techniques (see Banerjee 1999). Specifically, the procedures of Pedroni (2000), Mark and Sul (2002) and Breitung (2002) are used to get efficient estimates of the cointegration parameters.

The results indicate that a well-behaved money demand function can be justified only if the exchange rate is allowed to enter the specification. In reduced systems containing money, income and interest rates, a long run relationship cannot be detected at all. This principal finding is confirmed when euro exchange rates are part of the variable set, possibly due to the late introduction of this currency. Only if US-Dollar rates are considered, cointegration appears. In contrast to money demand in the euro area, the income elasticity seems to be significantly larger than 1.

The rest of the paper is organised as follows. Section 2 gives the specification of the long-run money demand function. Section 3 shows the econometric methods used. Section 4 describes the data and the empirical results. Finally, concluding remarks are presented in Section 5.

2   Money demand of transition countries

Only a few studies analyse money demand functions of transition countries (see Buch 2001). Earlier investigations cover only a short period of reform years (Dzwonik-Wrobel and Zieba 1994, International Monetary Fund 1998). Based on a correlation analysis, Antczak (2003) and Jorocinski (2003) have stressed the importance of money growth for stabilization the inflation rates. More recently Buch (2001) has specified money demand functions for Hungary and Poland, which account for the transition situation of these countries. Her money demand function includes an income variable, domestic and foreign interest rates and, changes of exchange rate expectations as well as inflations rates. Hence, she ends up with more than one variable measuring opportunity costs of money holding. The importance of exchange rates is also stressed by Orlowski (2004) for the Hungary, Poland and the Czech Republic as well as by Komarék, Melecký (2001) for the Czech Republic.

The analyses of money demand functions for the euro area contain not more than two opportunity cost variables (see for example Görgens et al. 2004, Bruggemann et al. 2003). Those studies suggest the following money demand function

(1) $M/P = f(Y, oc)$

where M is a broad monetary aggregate, P the consumer price index (CPI), Y income proxied by real GDP, and oc an opportunity cost indicator. According to textbook presentation the income variable should have a positive effect on money holdings. If the opportunity cost measures the earnings of alternative assets its coefficient should be negative. Regarding the development of the financial markets of the new member countries the opportunity costs are approximated by a short-term interest rate. The interest rate variable includes via the Fisher effect the inflation rate of these countries (see Orlowski 2004). At least for most industrial countries Crowder (2003) finds evidence in favour of the Fisher effect using panel cointegration methods.

All new member countries are small open economies. Due to foreign trade liberalisation in the transition process their agents could easily switch between foreign and domestic currencies. This may affect money holding in these economies. In this sense the exchange rate may reflect the return of foreign money. According to the analyses of Buch (2001) and Orlowski (2004) the money demand function in the new member countries additionally includes the exchange rate against the euro. A depreciation reduces the confidence in the domestic currency, thereby lowering money demand. Hence, the exchange rate coefficient should be negative, if the exchange rate is denoted as units of domestic currency per unit of foreign currency. Moreover, some of the new member countries gave the exchange rate policy a prominent role in achieving their monetary policy aims, its importance should be analyzed in the study (see Backé et al. 2004). For example Estonia introduced a currency board to the euro in 1992. Malta has followed a currency basket peg since 1971, where the weights in the basket are trade weighted. The current value of the euro is 70%. Cyprus and Hungary has a peg to the euro. However, the euro was introduced to the public not before 1999. To this reason we also consider US-dollar rates.

3   Panel unit root and cointegration tests

The integration and cointegration properties of the variables involved determine the specification of money demand. If the series are cointegrated, equation (1) should be viewed as a long run relationship. However, it has been widely acknowledged that standard unit root and cointegration tests can have low power against stationary alternatives for the important cases, see for example Campbell and Perron (1991). As an alternative, recently developed panel unit root and cointegration tests are applied. Since the time series dimension is enhanced by the cross section, the results rely on a broader information set. Therefore, gains in power are expected, and more reliable evidence can be obtained.

In the paper, the LLC (Levin, Lin and Chu, 2002), the IPS (Im, Pesaran and Shin, 2003) and the HD (Hadri, 2000) tests are applied. These procedures allow that deterministic and dynamic effects might differ across the panel members. The first two tests are generalizations of the ADF principle. The null of a unit root is investigated against the alternative of a stationary process for all (LLC) or at least for one cross section (IPS). The hypotheses are interchanged by the HD procedure, which adapts the KPSS test to panels. For the LLC and IPS test, the optimal lag length is selected using the general-to-simple procedure proposed by Campbell and Perron (1991). The consistent estimator of the long run residual variance relevant for the LLC and HD statistics is obtained using the Bartlett kernel and the automatic bandwidth parameter suggested by Newey and West (1994). Provided that the degree of cross section correlation is not substantial, the statistics

$$(2) \qquad Z_t = \frac{Z_t^* - \mu}{\sigma}$$

are asymptotically distributed as standard normal with a left (LLC, IPS) or right (HD) hand side rejection area. Standardization factors are obtained by simulation and depend on the deterministic components included in the testing procedure.

For panel cointegration, the tests suggested by Pedroni (1999) are employed. They extend the Engle and Granger (1987) two step strategy to panels and rely on ADF and PP principles. First, the cointegration equation is estimated separately for each panel member. Second, the residuals are examined with respect to the unit root feature. If the null

of is rejected, the long run equilibrium exists, but the cointegration vector may be different for each cross section. In addition, deterministic components are allowed to be individual specific. The residuals are pooled either along the within or the between dimension of the panel, giving rise to the panel and group mean statistics (see Pedroni, 1999). In the case of the panel statistics the first order autoregressive parameter is restricted to be the same for all cross sections. If the null is rejected, the parameter is smaller than 1 in absolute value, and the variables in question are cointegrated for all panel members. In the group statistics, the autoregressive parameter is allowed to vary over the cross section, as the statistics amount to the average of individual statistics. If the null is rejected, cointegration holds at least for one individual. Hence, group tests offer an additional source of heterogeneity among the panel members. Overall, 7 tests are proposed. In the limit, the statistics are distributed as standard normal with a left hand side rejection area, except of the variance ratio test, which is right sided. Standardization factors arise from the moments of Brownian motion functionals. The factors depend on the number of regressors and whether or not constants or trends are included in the cointegration relationships.

In addition, the Kao and McCoskey (1998) LM test for the null of cointegration is applied. The long run is estimated by efficient methods carried out separately for the panel members. Then, the cointegration residuals are pooled, and the test statistic is asymptotically Gaussian with a right hand side rejection area.

Generally panel cointegration tests do not provide an estimate of the long run relationship. However, the cointegration vector should be roughly common for the panel members, as fundamental economic principles are involved. Hypothesis testing of the cointegration parameters is also a critical issue. In fact, the asymptotic distribution of the OLS estimator depends on nuisance parameters. In a panel environment, this problem seems to be more serious, as the bias can accumulate with the size of the cross section. To overcome these deficits, efficient methods like fully modified (FM) and dynamic OLS (DOLS) are required. As these techniques control for potential endogeneity of the regressors and serial correlation, asymptotically unbiased estimates of the long run can be obtained. As the methods are asymptotically equivalent, their relative merits boil down to a comparison in finite samples (Banerjee, 1999). In the FM case, nonparametric

techniques are used to transform the residuals from the cointegration regression and get rid off nuisance parameters (Phillips, 1995, Pedroni, 2001). In the time series model

(3) $y_{it} = \alpha_i + \beta_i x_{it} + u_{it}$

$x_{it} = x_{it-1} + \varepsilon_{it}$ $, \varpi_{it} = (u_{it}, \varepsilon_{it})'$

the asymptotic distribution of the OLS estimator is conditioned to the long run covariance matrix of the joint residual process. The FM estimator for the i-th panel member is given by

(4) $\hat{\beta}^* = (X_i'X_i)^{-1}(X_i'y_i^* - T\hat{\delta})$

where y^* is the transformed endogenous variable and δ a parameter for autocorrelation adjustment. Appropriate correction factors are based on certain submatrices of the joint long run covariance matrix.

In the DOLS framework, the long run regression is augmented by lead and lagged differences of the explanatory variables to control for endogenous feedback (Saikkonen, 1991). Lead and lagged differences of the dependent variable can be included to account for serial correlation (see Stock and Watson, 1993). In particular, the equation

(5) $y_{it} = \alpha_i + \beta_i x_{it} + \sum_{j=-p_1}^{p_2} \delta_j \Delta y_{it-j} + \sum_{j=-q_1}^{q_2} \lambda_j \Delta x_{it-j} + u_{it}$

is run for the i-th panel member, where the appropriate choice of leads and lags is based on data dependent criteria (Westerlund, 2003). Standard errors are computed using the long run variance of the cointegration residuals.

In a panel setting, the cointegration relationship is homogeneous. Heterogeneity is limited to fixed effects, time trends and short run dynamics. The panel FM estimator is the average of the individual parameters (see Pedroni, 2001). According to Mark and Sul (2002) a panel DOLS estimator is obtained using a two step procedure. First, individual dynamic and deterministic components are regressed out separately for the panel members. Then, the residuals are stacked, and a pooled regression is run. As an alternative to

these methods, Breitung (2002) has suggested a two step procedure based on a cointegrated VAR model. In the VECM

$$(6) \qquad \Delta z_{it} = \alpha_i \beta' z_{it-1} + \varepsilon_{it}$$

the feedback coefficient α_i and the covariance matrix Σ_i of the residuals are allowed to vary across the individuals. As the information matrix of the Gaussian likelihood is asymptotically block diagonal with respect to the short run and cointegration parameters, the long run relation can be uncovered conditional on consistent estimates of the former. Hence, the short run parameters are revealed by individual VECM's, and the restriction that the individuals have a common cointegration vector is temporarily ignored. Then, the variables are transformed according to

$$(7) \qquad z_{it}^* = \beta' z_{it-1} + \varepsilon_{it}^*$$

$$z_{it}^* = (\hat{\alpha}_i' \hat{\Sigma}_i^{-1} \hat{\alpha}_i)^{-1} \hat{\alpha}_i' \hat{\Sigma}_i^{-1} \Delta z_{it} \ , \quad \varepsilon_{it}^* = (\hat{\alpha}_i' \hat{\Sigma}_i^{-1} \hat{\alpha}_i)^{-1} \hat{\alpha}_i' \hat{\Sigma}_i^{-1} \varepsilon_{it}$$

and a pooled regression is run. The long run parameters are asymptotically distributed as standard normal. According to simulation evidence provided by Breitung (2002), his estimator is preferable over FMOLS and DOLS alternatives, as it comes with a smaller finite sample bias.

As a major shortcoming, the panel tests for integration and cointegration presume that the cross sections are independent. However, this requirement is not met in the analysis presented here. For example, medical advancements are correlated across countries. In particular, the presence of cross section cointegration can distort the panel results, see Banerjee, Marcellino and Osbat (2001) and Urbain (2004). In these cases, either the endogenous variable or specific regressors cointegrate across the panel members. To control for this problem, cointegration tests based on nonstationary common factors are proposed, see Bai (2004), where factors are obtained as principal components. Compared to the individual country analysis, the procedure is likely to be more robust, because idiosyncratic (country specific) parts cancel out.

4   Data and empirical results

The analysis is done using quarterly seasonal adjusted data for the following countries: Cyprus, Czech Republic, Estonia, Hungary, Latvia, Lithunia, Malta, Poland, Slovak Republic and Slovenia. The sample ranges over the 1995.1-2004.2 period. Data include broad money, the consumer price index (CPI), the money market (3month) interest rate, real GDP, and nominal exchange rates, denoting the units of national currencies against the euro or the US-dollar, respectively. Money, prices, interest and exchange rates are usually taken from the International Financial Statistics of the IMF. As an exception Hungarian money has been obtained from the OECD Main Economic Indicators. Nominal money stocks are deflated by the CPIs. Real GDP series are from Eurostat. In case of missing values, national central bank information is used to generate these data points. All variables enter the analysis in their logs. Real GDP and real money are seasonally adjusted. The evolution of the individual time series is depicted in Figure 1.

-Figure 1 about here-

All countries experienced an increase of the real money stock over the period under consideration. The rise is most pronounced for Estonia and Slovenia. At the end of the sample, the series exceed their initial values by a factor of 2.5 or 3.5, respectively. Real GDP has grown in all countries, where the strongest acceleration is observed in the Baltic states (Estonia, Latvia, Lithunia). Interest rates have declined in all economies. But the decrease was not steadily, as phases of higher interest rates are also apparent for some countries (Estonia, Czech Republic, Slovak Republic). Compared to the starting level, interest rates are lower at the end of the sample for all countries, where the sharpest decline can be detected for Lithunia. Hungary, Poland and Slovenia experienced a remarkable depreciation of their currencies relative to the euro and the US-Dollar. In contrast, Lithunia experienced a strong appreciation against these currencies.

-Table 1 about here-

The stationarity hypothesis of the considered variables is checked by the IPS-test, the LLC-test the Breitung-test and the Hadri test. Table 1 includes the results for the levels and for the first differences of the variables considered. Only the LLC-test indicates stationarity for the exchange rate levels against the US-dollar. The nonstationarity of the first differences is always rejected at the five percent level for the LLC, IPS and Breitung test. The Hadri test rejects the stationarity hypothesis for broad money and exchange rate. In sum, it seems sensible to conclude that all variables are stationary in first differences. These results allow to test for cointegration among the variables and to estimate money demand functions. Table 2 presents the panel cointegration tests. The first system include real money, GDP and the interest rate. The second system additionally contains the exchange rate against the US-dollar.

-Table 2 about here-

The upper part of table 2 shows that the tests indicate that no cointegration exists among real broad money, GDP, and the interest rate for this period. All tests do not reject the null hypothesis of no cointegration in this panel. For the system including the euro-exchange rates only one test rejects the null of no cointegration. The results are more favourable, if the system contains the US-dollar exchange rate. Six of seven Pedroni tests reject the null hypothesis of no cointegration. This is a strong evidence for the existence of a cointegrating relationship among the variables.

-Table 3 about here-

Table 3 shows the results of the FM, DOLS and 2-step-approach. The values of the DOLS-method are determined under the assumption of one lead and two lags of the changes of the regressors. The income elasticity of the money demand function is significantly above unity in all cases. The panel income coefficient is much higher than values for euro area functions (see Bruggemann et al. 2003, Görgens et al. 2004, p. 179). It is worth noting that all methods obtain a higher income elasticity, if the system

includes the exchange rates via the US-dollar than via the euro. The interest rate elasticity is significantly negative and the exchange rate elasticity is negative as expected. The interest rate elasticity is relatively small and may reflect the fact that it is difficult to control the money holding. The exchange rate elasticity is not significantly different from zero, if euro rates are considered. All methods find significant impacts of exchange rates via the US-dollar. These results confirm to some extent the results of Buch (2001), Komarék, Melecký (2001) and Orlowski (2004) in the sense that an exchange rate variable is important. In contrast, to Orlowski (2004) the exchange rates via the US-dollar is necessary to obtain a cointegration relationship. Hence, the specification of money demand function for the new member countries differs from the specification for the euro area.

Finally, a cointegration analysis is performed using common factors, see table 4. Principal components are estimated separately for real money, income, interest and exchange rates. For each variable, the first principal component is considered. Then, the cointegration test is performed using standard methods. In carrying out this exercise, the ADF type cointegration test (MacKinnon, 1991) is considered. The cointegrating regression is estimated by the DOLS, and the residuals are checked for stationarity.

-Table 4 about here-

The cointegration result can be confirmed at the 5 percent level. In addition, the variables enter the long run money demand relation with the correct sign. However, the elasticities seem to be different from the panel evidence. Most strikingly, the income elasticity is not significantly larger than 1.

5 Conclusions

In this paper the long-run coefficients of a money demand function for the new member states of the European Union are estimated. The estimation is conducted with means of panel cointegration methods for the period 1995Q1 to 2004Q2 for 10 countries. It is necessary to account for the exchange rate variables via the US-dollar in obtaining a

long run money demand function. The panel income elasticity is around 1.70 and the interest rate elasticity is negative. Hence, the specification of the money demand function and especially, the income coefficient are different from euro area results.

A sudden introduction of the euro in all new member countries may introduce problems for the stability of the euro area money demand function. However, the introduction of the euro requires that the Maastricht convergence criteria are fulfilled. The probability is small that all countries would achieve the critera at the same time. Moreover, the number of inhabitants in these countries is small and GDP is markedly less than the average of actual EU-citizens. Hence, their weights are relatively small. In this sense the changes should be manageable for the ECB. Nevertheless, future research should analyse the stability of the euro area money demand function when accounting for new EU members.

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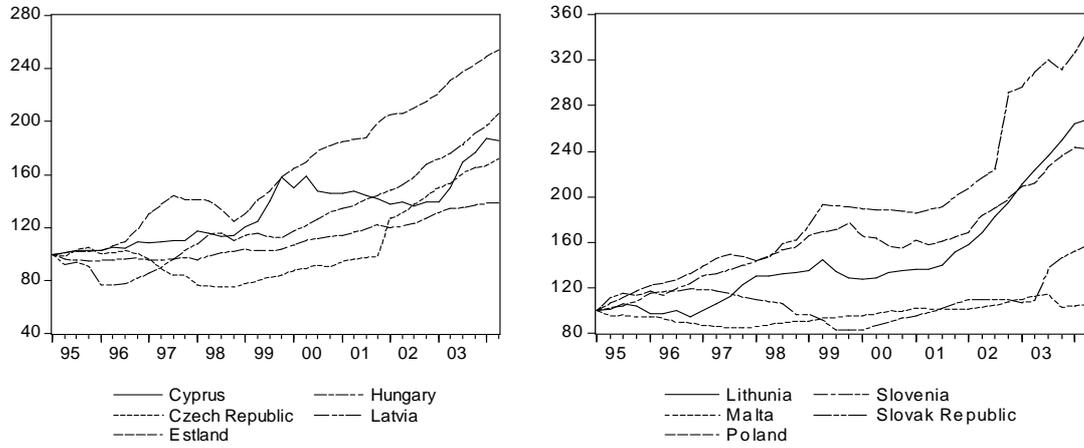
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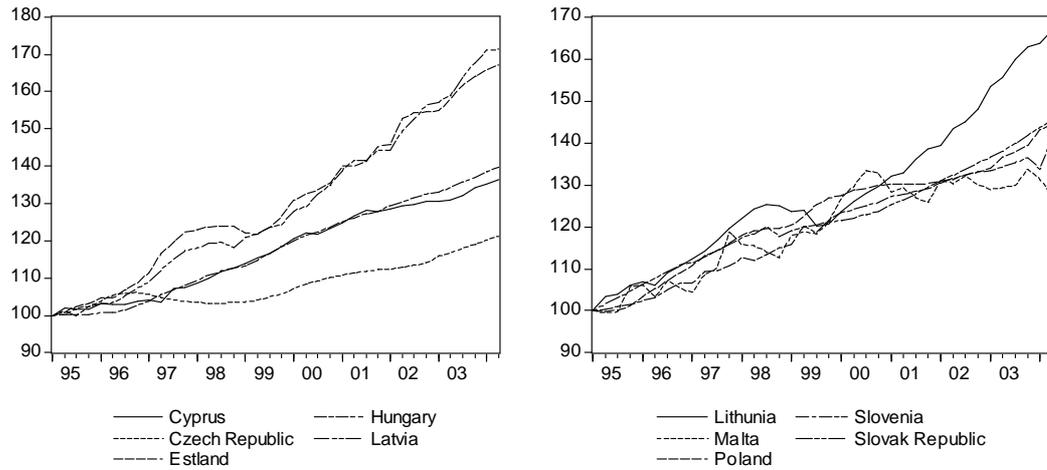
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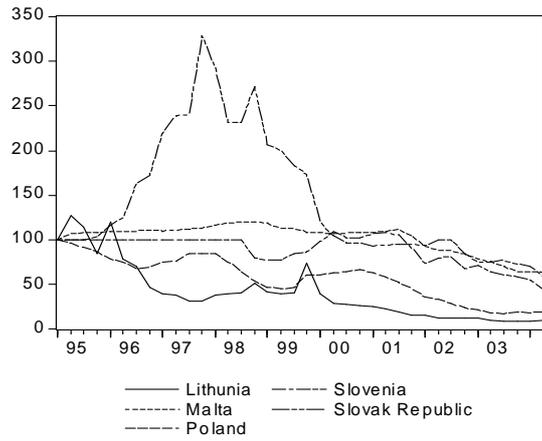Figure 1: Broad money, GDP, interest and exchange rates in the new EU member countries (1995.1)
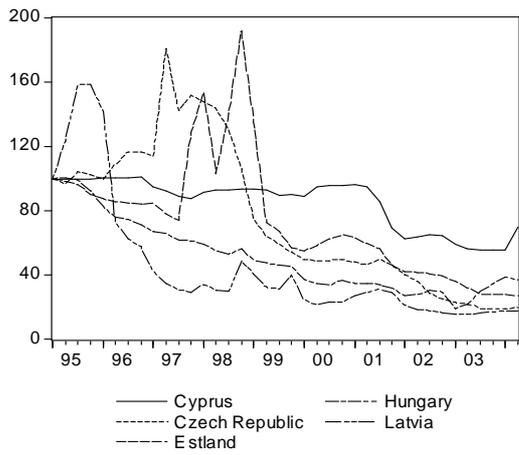
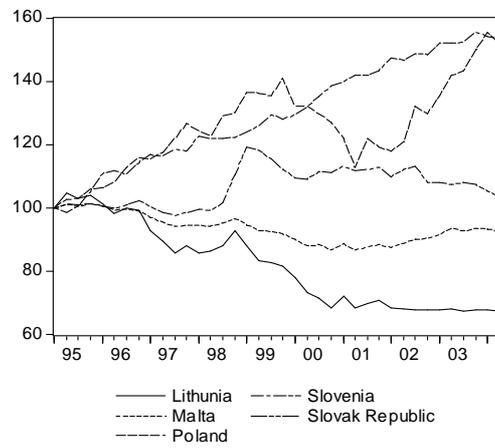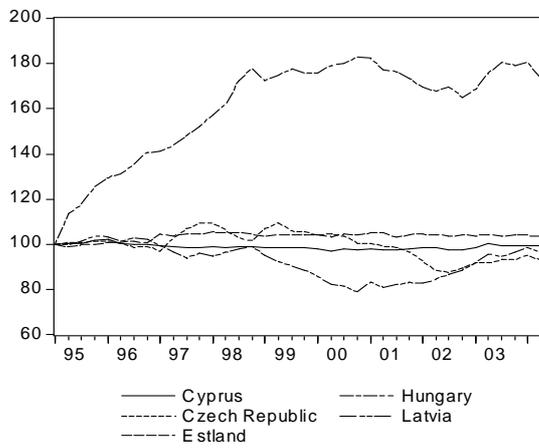A   Real broad money (deflated by CPI)

B   Real GDP

B   Nominal interest rates

Cyprus — Hungary
Czech Republic — Latvia
Estland

Lithunia — Slovenia
Malta — Slovak Republic
Poland

D   Exchange rate: domestic currency units per euro

Cyprus — Hungary
Czech Republic — Latvia
Estland

Lithunia — Slovenia
Malta — Slovak Republic
Poland

E   Exchange rate: domestic currency units per US-Dollar

Cyprus — Hungary
Czech Republic — Latvia
Estland

Lithunia — Slovenia
Malta — Slovak Republic
Poland
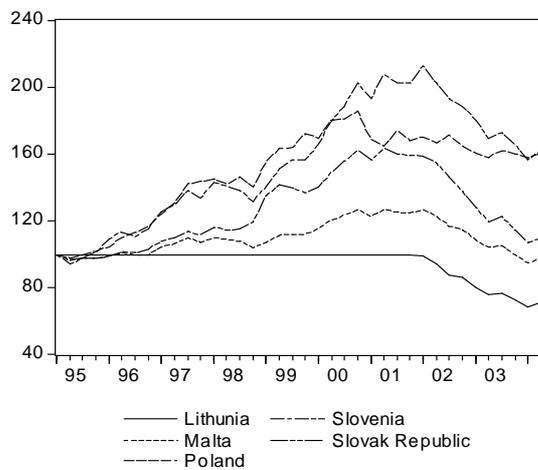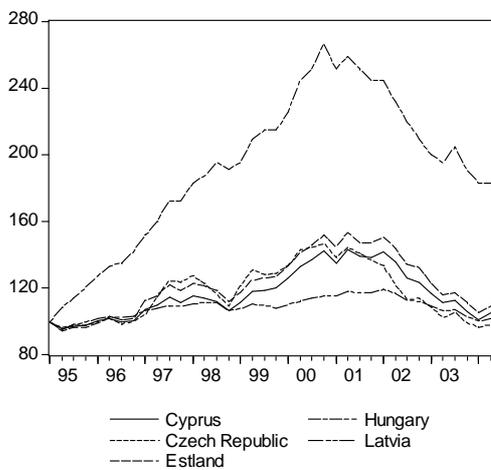
Table 1: Panel unit root test of the variables in the money demand function

A: Levels				
	LLC	Breitung	IPS	Hadri
Broad Money	-0.931	0.322	-1.491	6.769*
Income	0.591	0.739	1.026	5.834*
Interest rate	1.508	2.248	3.269	9.748*
Exchange rate against euro	-5.088*	2.064	-1.998	9.431*
Exchange rate against US-dollar	-2.767*	6.763	-0.149	7.151*
B: First differences				
Broad Money	-9.980*	-5.096*	-11.021*	2.026*
Income	-9.863*	-5.856*	-13.664*	-0.141
Interest rate	-8.894*	-5.070*	-9.273*	0.207
Exchange rate against euro	-15.901*	-6.336*	-15.578*	2.622*
Exchange rate against US-dollar	-10.693*	-5.530*	-9.588*	6.632*

LLC=Levin, Lin, Chu (2002), IPS=Im, Pesaran, Shin (2003). The other statistics are described in detail in Breitung (2000) and Hadri (2000). The statistics are asymptotically distributed as standard normal with a left hand side rejection area, except of the Hadri test, which is right sided. A * indicates the rejection of the null hypothesis of nonstationarity (LLC, Breitung, IPS) or stationarity (Hadri) at least on the 0.05 level of significance.

Table 2: Panel cointegration tests

Model without exchange rate		
	Pedroni (1999)	
Method:	Panel Statistics	Group Statistics
Variance ratio	0.881	
Rho statistic	0.078	0.327
PP statistic	-0.725	-1.005
ADF statistic	0.247	-0.315
	Kao and McCoskey (1998)	
LM statistic	FM: 0.202	DOLS: 1.548

Models including the exchange rate				
	Pedroni (1999)			
	USDollar exchange rate		Euro exchange rate	
	Panel Statistics	Group Statistics	Panel Statistics	Group Statistics
Variance ratio	2.542*		1.402	
Rho statistic	-0.790	0.117	-0.153	0.402
PP statistic	-1.831*	-1.666*	-1.129	-1.460
ADF statistic	-2.246*	-2.400*	-1.483	-2.802*
	Kao and McCoskey (1998)			
LM statistic	FM: -1.966*	DOLS: -0.695	FM: -1.638	DOLS: -0.471

Statistics are asymptotically distributed as normal. The Pedroni statistics are described in detail in Pedroni (1999). The variance ratio test is right-sided, while the other Pedroni tests are left-sided. The LM test from Kao and McCoskey (1998) is right-sided and carried out using either FM or DOLS residuals. A * indicates the rejection of the null hypothesis of no cointegration (Pedroni) or cointegration (Kao and McCoskey) at least on the 0.05 level of significance.

Table 3: Panel estimation of the cointegration vector

Model including *m-p, y, R*		
	Income	Interest rate
FM (Pedroni, 1999)	1.67 (0.08)	-0.09 (0.02)
DOLS (Mark and Sul, 2002)	1.46 (0.13)	-0.12 (0.03)
2-Step (Breitung, 2002)	1.46 (0.14)	-0.17 (0.04)

Model including *the exchange rate*			
	Income	Interest rate	Euro
FM (Pedroni, 1999)	1.54 (0.08)	-0.10 (0.02)	0.13 (0.16)
DOLS (Mark and Sul, 2002)	1.38 (0.12)	-0.14 (0.03)	0.01 (0.09)
2-Step (Breitung, 2002)	1.48 (0.10)	-0.12 (0.03)	0.11 (0.08)
Model including *m-p, y, R, US-Dollar*			
	Income	Interest rate	US-Dollar
FM (Pedroni, 1999)	1.73 (0.08)	-0.09 (0.02)	-0.28 (0.04)
DOLS (Mark and Sul, 2002)	1.94 (0.13)	-0.07 (0.03)	-0.22 (0.06)
2-Step (Breitung, 2002)	1.78 (0.10)	-0.06 (0.02)	-0.16 (0.04)

Elasticities of real money demand with respect to real income, interest and exchange rates are reported. Standard errors in parantheses.

Table 4: Cointegration analysis of common factors

	Euro	US Dollar
Income	0.959 (0.106)	1.047 (0.051)
Interest rate	-0.279 (0.070)	-0.160 (0.036)
Exchange rate	-0.355 (0.061)	-0.278 (0.023)
ADF	-2.427	-4.797*

ADF-test for stationarity of residuals obtained by DOLS methods. Elasticities of money demand with respect to real GDP, interest and exchange rates Standard errors in parantheses.