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Money Demand in Estonia

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Abstract

This study develops a parsimonious stable coefficient money demand model for Estonia for the period from 1995 till 2006. Using the Johansen Full Information Maximum Likelihood framework the two cointegrating vectors are found among the system variables including the real money balances, the gross domestic product, the long- and short-term interest rates, and the rate of inflation. The first cointegrating vector is identified as the money demand function whereas the second as the interest rate parity. Our study contributes to better understanding of the factors shaping the demand for money in the new Member States of the European Union that committed themselves to adopting of the Euro currency in the near future.

Keywords: M2 money demand, stability, new EU member states, Estonia

JEL code: C32, E41.

*The paper has benefited from comments of Jiří Slačálek.

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1 Introduction

Demand for money has been and still is one of the most researched topics in macroeconomics. This is reflected in a steady stream of theoretical and empirical research that has been undertaken over the past several decades. The scope of the research is illustrated by several survey articles such as Fase (1993), Sriram (2001), and Knell and Stix (2003) where money demand estimation results are reviewed from 100 papers written from 1972 to 1992, from 28 papers published between 1990 and 1999, and from 68 papers from 1995 to 2002, respectively. While most of the research has traditionally focused on the money demand in the developed industrialised countries, there exist only few studies that address money demand issue in the transition economies of the Eastern Europe. These include Bolharyn and Babaian (1998) – for Ukraine, Karla (1999) – for Albania, Buch (2001) – for Hungary and Poland, Payne (2003) – for Croatia, Slavova (2003) – for Bulgaria, Andronescu et al. (2004) – for Rumania, Bahmani-Oskooee and Barry (2000) and Oomes and Ohnsorge (2005) – for Russia.

Nevertheless, the importance of understanding of the determinants of money demand in the Eastern European countries, especially in those that are already EU Member states, cannot be understated due to the anticipation of introduction of Euro in these countries in the near future and due to high relevance of money within the two-pillar monetary strategy of the European Central Bank and its strong concern regarding price stability in the Eurozone (European Central Bank, 2003). As after introduction of the Euro in the new Member States of the European Union, the responsibility regarding the monetary policy in those countries will ultimately rest on the Governing Council of the European Central Bank.

In this paper we intend to contribute to the literature on money demand in the transitional economies of Eastern Europe by focusing on Estonia. To the best of our knowledge, there are only two references in the literature that attempt to estimate the long-run money demand function for Estonia: Dabusinskas (2005) and Dreger et al. (2006). In our opinion, each of these two studies have left certain issues unresolved and hence our intention is first to identify these problematic issues and then address them in this paper.

The former study finds rather weak statistical evidence on the existence of the cointegrating relationship in Estonian data that would correspond to the money demand function. Clearly, such finding could be largely attributed to the rather small sample size used in Dabusinskas (2005), i.e., from the first quarter of 1997 till the third quarter of 2003 which comprises only 29 quarterly observations. Another finding which is somewhat problematic from the point of view of application of cointegration technique to estimation of money demand in Estonia is that the real money and the real GDP variables were found of different order of integration – trend-stationarity of the real money and I(1) non-stationarity of the real GDP – an artifact that also may well be attributed to the size of the sample.

The main contribution of the latter study constitutes use of panel cointegration techniques in order to estimate the long-run demand function for the ten new EU member states, including Estonia. According to this approach, the common coefficients of the long-run money demand function are being imposed across all individual cross-sections, which may be questionable given heterogeneity of these countries. In addition, use of panel cointegration techniques relies on the assumption that the cross-sections are independent. As acknowledged by the authors themselves, this assumption is most likely violated for the data used in their paper and hence their results may be distorted, e.g., by the presence of unmodelled cross sectional cointegration (Banerjee et al., 2001; Urbain, 2004). Moreover, the stability of the country-specific long-run money demand functions is not addressed at all in Dreger et al. (2006). Thence, additional research that develops a long-run money demand function for each country individually is needed in order to verify the plausibility of reported results therein.

In sequel, we try to overcome these problematic issues of these two studies as follows. First, in contrast to Dreger et al. (2006), we address estimation of the money demand function within the unified framework of the full information maximum likelihood method of Johansen (1995). This methodology allows us to make inference on the integration order of the variables, on the existence and on the number of cointegrating relationships in the data, and to develop a parsimonious model for demand for real money balances in Estonia over the examination period by imposing statistically acceptable zero restrictions on the system. Lastly, we subject our model to the battery of the diagnostic tests including those for structural stability.

Second, in comparison to Dabusinskas (2005), we use longer time series data spanning the period from 1995 till 2006. We have chosen to focus on the estimation period that starts after four years since Estonia gained its independence for the following reason. During the early 90-ties the Estonian economy initiated transformation process from the planned to the market economy by enforcing several important economic reforms including introduction of the national currency and establishment of the currency board arrangement in 1992, rapid implementation of the external liberalisation by removing restrictions on trade, investment and financial flows, initiation of privatisation and other measures aimed at the complete restructuring and stabilisation of the national economy. In this period the Estonian economy underwent severe economic crisis that manifested itself in severe slump in the economic activity and by the rampant inflation (around 1000 per cent in 1992). However, starting from 1995 the earlier implemented reforms started to yield positive results bringing about stabilisation of Estonian economy and reversing the downward trend in the economic activity.

Our main findings are following. We have found very strong statistical evidence for existence of a long-run equilibrium relationship that may well be interpreted as the long-run money demand function in Estonia. Both the long-run coefficients and the short-run parameters of the conditional vector error correction model exhibit stability over the observation period. Moreover, our results confirm the conclusion of Dabusinskas (2005) and Dreger et al. (2006) that the income elasticity of demand for real money balances in Estonia is significantly larger than unity.

The rest of the paper is organised as follows. Section 2 briefly outlines the theoretical considerations behind the empirical models for money demand. Section 3 describes the data sources and the data transformations. Section 4 contains description of the modelling approach and presents the estimation results. The final section concludes.

2 Theoretical considerations

In the empirical section below we specify the long-term demand for money in the following form:

$$\frac{M}{P} = f(Y, R)$$

where the demand for real balances M/P is measured as a ratio of a selected money aggregate M in nominal term and the price level P. Estimation of the real demand for money implicitly implies that money neutrality and price homogeneity hold in the long run.

Below the demand for real money is modelled as a function of the two categories of variables: the scale variable that reflects the scope of economic activity, typically approximated by the real GDP Y, and the vector of returns on different assets $R = (I^s, I^l, \pi)'$, where I^s denotes money "own" return, approximated by the short-term deposit rate, I^l is the return on the assets outside of money, approximated by the government bond yield, and π is the rate of inflation that measures the return to holding goods. This set of the explanatory variables largely reflect three main purposes for holding money as stipulated by the economic theory: transactions, precautionary, and speculative motives (Keynes, 1936).

The long-run money demand, when log transformed, reads as follows

$$\ln\left(\frac{M}{P}\right) = \gamma_1 * \ln(Y) + \gamma_2 * \ln(I^s) + \gamma_3 * \ln(I^l) + \gamma_4 * \pi + ec \qquad (1)$$

where the coefficients γ_1, γ_2, γ_3, and γ_4 denote the long-run elasticities of money demand with respect to income, to short- and long-run interest rates, and the long-run semi-elasticity of money demand with respect to inflation, respectively. The former two coefficients are expected to have a positive sign as demand for money increases with income and with own interest rate, whereas the latter two coefficients are negative. An increase in the long-term interest rate leads to shifts in portfolio towards the longer-term investment and henceforth reduces demand for money. Similarly, rise in inflation reduces the value of monetary assets and henceforth tends to reduce demand for it. Finally, the term ec denotes the error-correction term that measures deviations from the long-run equilibrium given in equation (1).

As mentioned in the survey article of Sriram (2001), this is the ultimate specification structure for the money demand that is common to the most of the studies even though each study may be different from the rest in choice of either of the dependent or independent variables and/or both.

3 Data

The data were taken from the World Market Monitor (Global Insight, Inc.), see Table 1. The quarterly data span years 1995(1) – 2006(2), such that the sample size is $T = 46$. The following transformations of the original data have been carried out: $(m - p) = ln(M2/CPI)$ – the real money balances, $y = ln(GDPR)$ – the real GDP, $\{i^s, i^l\} = \{ln(I^s), ln(I^l)\}$ are the short- and the long-term interest rates, and $\pi = 4\Delta ln(CPI)$ is the annualised inflation rate. Observe that we have taken the logarithmic transformation of the interest rates that we use in our subsequent analysis. We have done this in order to account for the fact that the variation as well as level of the interest rates were much higher in the first half of our sample than in the second one. The advantage of such transformation is, of course, that the estimated coefficient values of the interest rate variables are to be read as elasticities. The transformed data are depicted in Figure 1.

4 Econometric model

4.1 Inference on cointegration

In our modelling of money demand function in Estonia, we follow the general-to-specific approach advocated in Hendry and Mizon (1993) and Hendry and Juselius (2000, 2001), inter alia. In particular, we start with an unrestricted VAR(n) model transformed into the error-correction form

$$\Delta x_t = \Pi x_{t-1} + \sum_{i=1}^{n-1} \Gamma_i \Delta x_{t-i} + \mu_t + \Psi D_t + \varepsilon_t, \varepsilon_t \sim N_k(0, \Sigma) \quad (2)$$

where $x_t = ((m-p)_t, y_t, i^s_t, i^l_t, \pi_t)'$ is the $k \times 1$ vector of variables described above. The matrix $\Pi = \alpha\beta'$ is the long-run matrix which in the presence of r cointegrating relations among the k I(0) variables has reduced rank r, where β is a $k \times r$ matrix with r cointegrating vectors and α is a $k \times r$ matrix with loading coefficients. μ_t denotes the deterministic terms such as a constant term and the seasonal dummies. Lastly, in order to control for the large outliers in the empirical model we use the following intervention dummies $D_t = (D9704, D9804, D0302)'$. The first two dummy variables account for the effects of the Asian and the Russian crises, whereas the latter accounts for the sharp fall in the long-run interest rate in the second quarter of 2003. These intervention dummies $Dyy0q$ take value of 1 in the corresponding quarter $0q$ of $19yy$ or $20yy$ and zero otherwise.

In the remainder of the section we proceed as follows. After selecting the lag length of the unrestricted VAR model, we test for the cointegration rank and subsequently impose the implied reduced rank restrictions on the unrestricted VAR model. Then we impose the (over-)identifying restrictions on the space spanned by the columns of estimated matrix of the long-run coefficients β and address the long-run weak exogeneity of the system variables. We use the results of the weak exogeneity tests in order to build a parsimonious representation of the system in the form of the conditional vector error correction model (VECM) that satisfactorily passes diagnostic tests, displays constant coefficients, and possesses the ability to accurately forecast the dependent variables in the recent time period.

First, we determine the lag length order of an unrestricted VAR(n) model. At this stage, we would like to get the parsimonious model given rather moderate number of observations $T = 46$. Table 2 contains results of the formal lag order selection procedures.[1] As seen, the optimal lag length varies between one (according to the Schwarz information criterion, SC) and four (according to the Akaike information criterion, AIC). At the same time, the sequential likelihood ratio test statistic (LR), the Final Prediction Error(FPE), and the Hannan-Quinn (HQ) information criterion select $n = 2$. Hence, it seems that the VAR(2) model can adequately describe the data at hand. This decision on the lag length of the unrestricted VAR model is further reinforced by the battery of the misspecification tests which report no serious departures from the underlying

[1] The lag order selection was conducted using Eviews 5.1.

model assumptions, see Table 3.[2] The univariate as well as multivariate model diagnostic tests comprise: F_{AR} – test of no residual autocorrelation (see Godfrey (1978)); χ^2_{Norm} – test for the normally distributed residuals (see Doornik and Hansen (1994); F_{Hetero} and $F_{Hetero-X}$ – White (1980) tests for heteroscedasticity based on the original and squared regressors, and on the original, squared regressors, and their cross-products; F_{ARCH} – Engle (1982) test of no residual AutoRegressive Conditional Heteroscedasticity. The graphics, regression output, and residual diagnostic tests were calculated using GiveWin 2.2 and Pc-Give 10.2 (see Doornik and Hendry, 2001a,b).

Having found an adequate unrestricted model, the next step is to proceed to imposing restrictions on that model. Hence, we address the cointegration rank of the estimated system. We use the Johansen Full Information Maximum Likelihood (FIML) procedure for this purpose. Table 4 reports the results of the trace and λ-max tests both using the asymptotic critical values and the critical values based on the finite sample correction (see Osterwald-Lenum, 1992). Observe that regardless of what kind of critical values are used, the test results strongly suggest the presence of two cointegrating relations in the system.

Thus, we impose the cointegration rank $r = 2$ on the system (2) and proceed with testing for (trend-)stationarity, long-run exclusion, and long-run weak exogeneity of the variables in our model. The test of stationarity of the variables in a VAR model has been suggested in Johansen and Juselius (1992). This is a multivariate version of the Augmented Dickey-Fuller test with the null hypothesis of stationarity rather than non-stationarity. Since a linear combination of I(1) variables that is I(0), or I(0) variables themselves, could only belong to the cointegration space, it investigates whether any of the variables alone belong to the cointegration space. This test has an asymptotic χ^2 distribution with the $(k - r) = 3$ degrees of freedom.

The test for the long-run exclusion (Johansen and Juselius, 1992) investigates whether any of the variables can be excluded from the cointegration space. This test has an asymptotic χ^2 distribution with the $r = 2$ degrees of freedom.

Table 5 reports the results of the tests for (trend-)stationarity and long-run exclusion, performed on the matrix of the long-run coefficients. According to the stationarity test, the null hypothesis

[2] At the same time, the misspecification tests indicate remaining residual autocorrelation in the VAR(1) model.

that each variable is either I(0) or I(0) around a linear deterministic trend is decisively rejected. The tests for the long-run exclusion reject the null hypothesis that any of the $(m-p)_t$, y_t, i_t^s, i_t^l, and π_t variables can be individually excluded from the cointegration space at the 1% significance level.

4.2 Long-run money demand function

After finding that there are two cointegrating relationships in the system, it remains to determine whether they are unique and whether they contain information regarding the structural economic relationships underlying the long-run model. Having imposed the (over-)identifying restrictions on the cointegration space for this purpose, we proceed with testing for the long-run weak exogeneity of the variables in the system. Table 6 summarises the results of imposing the over-identifying restrictions on the cointegration space (step 1) and of imposing the long-run weak exogeneity restrictions (step 2).

As seen, the imposed restrictions on the matrix of the long-run coefficients satisfy conditions for identification, as each cointegrating vector has a variable that is unique to it. The over-identifying restrictions which are imposed in step 1 are accepted according to the likelihood ratio test statistic $\chi^2(3) = 7.277[0.064]$. Furthermore, imposing several weak exogeneity restrictions in step 2 yields the likelihood ratio test statistic $\chi^2(8) = 16.544[0.035]$, which is marginally significant at the 5% level. Nevertheless, we have chosen to maintain it as these restrictions seem to hold for different sample sizes which is evident from Figure 2, where the corresponding recursively estimated test statistic (scaled by the 1% critical value) is reported.

The detected long-run relationships are following

$$(m-p) = 2.177y + 0.346i_t^s - 1.533\pi \tag{3}$$

$$i_t^l = 825i_t^s. \tag{4}$$

The first of them (3) may well be interpreted as a money demand relationship in line with the theoretical considerations outlined in Section 2. The income elasticity is significantly larger than unity (2.177) being consistent with a trend decline of the velocity of circulation $V2$, see Figure 1. This result may also be in line with observation that money is held not only for transaction motives but also for portfolio decisions. The short-run interest rate elasticity is 0.346, and, thus, it indicates a tendency of money holdings to increase when assets inside the broad monetary aggregate promise higher returns. The estimate of the semi-elasticity of the money demand with respect to inflation is -1.533 implying more than proportionate decrease in money holding as inflation rises.

The second cointegrating relationship can be interpreted as stationary interest spread as proposed by the expectation theory of the term structure, albeit the hypothesis of the second cointegrating vector being $\beta_2 = (1, -1)'$ could not be marginally rejected at the 1% significance level. The corresponding likelihood ratio statistic is $\chi^2(1) = 6.593[0.0102]$. The likely reason is rather moderate sample size available for the analysis and turbulence in the interest rates caused by the Asian and the Russian financial crises, see Figure 1.

The estimated cointegrating vectors are depicted in Figure 3. The recursively calculated coefficients of the cointegrating relationship are displayed in Figure 4 and these are rather stable in time.

At this point, it is instructive to compare our income elasticity estimates with those obtained from other studies. Knell and Stix (2004, 2006), where the results of more than 500 studies of money demand are analysed, report that the mean and the median of all income elasticity estimates taken together lies around unity but nevertheless they shows a large dispersion. Moreover, they report that the Euro-zone countries income elasticity of about 1.28 and 1.42, depending on the way the results are summarised.[3] As seen, our point estimate of income elasticity is rather large in comparison with that reported for the Euro-zone. At this point, it is interesting to note that our estimate of the long-run income elasticity very closely matches that reported in Dabusinskas (2005, Table 5), estimated with a different methodology for the shorter time period (only 28 observations) and with the seasonally adjusted data. However, it must be noted that the reported

[3] The former figure was obtained by weighted-averaging all broad money income elasticity estimates for individual Euro-zone countries, whereas the latter figure – by taking average of income elasticity estimates reported in the studies that estimated a joint money demand for (several) European countries, i.e. data aggregation was done before the estimation.

income elasticity estimates therein seems to be sensitive both to the specification of the estimated equilibrium relationship and to a method applied.

To summarise, our estimate of long-run money elasticity for Estonia significantly differs from those reported for the Euro-zone countries. The estimate of the income elasticity tends to be significantly larger. Thence, we can tentatively conclude that the long-run money demand function for a transitional economy may differ from that of the members of the Euro-zone. To this end, our estimation results obtained for Estonia seem to support the similar conclusion reached in Dreger et al. (2006) where the long-run money demand function was estimated for the new EU member states using the panel cointegration techniques.

4.3 Short-run error correction model

As discussed above, Table 6 contains results for the long-run weak exogeneity testing. We found that for the variables y_t and π we cannot reject the null hypothesis of the long-run weak exogeneity as the corresponding adjustment coefficients for both long-run equilibrium relationships were not significant for these two variables. As shown in Johansen (1992), the status of long-run weak exogeneity of some variables allows us to reformulate the model (2) in terms of a conditional model vector error correction model (VECM), where we condition on the current and past values of the weakly exogenous variables, and on the first lag of the error correction term as follows

$$\Delta y_t = \Gamma_0 z_t + \Gamma_1 \Delta x_{t-1} + \alpha_1(\beta' x_{t-1}) + \mu_t + \Psi D_t + \varepsilon_t, \varepsilon_t \sim N_{k_y}(0, \Sigma) \tag{5}$$

where $y_t = ((m-p)_t, i_t^l, i_t^s)'$ and $z_t = (y_t, \pi_t)'$. α_1 is the 3×2 loading matrix of the cointegrating relations in this conditional VECM.

In sequel, we first estimate the unrestricted VECM with the full information likelihood method and then perform the valid system reduction in order to achieve the most parsimonious representation of the model subject to the condition of no serious departures from model assumption. The estimated conditional model for the real money balances $(m-p)_t$ and the two interest rate variables i_t^l and i_t^s is presented in Table 7, where the corresponding standard errors reported in parentheses below the coefficient estimates. The imposed zero restrictions are accepted by the likelihood ratio test of over-identifying restrictions $\chi^2(16) = 23.309[0.106]$. The resulting conditional VECM is parsimonious but at the same time the diagnostic tests show no signs of misspecification. Observe that the adjustment coefficients of the error-correction terms are significantly different from zero and their magnitude is comparable to that reported in Table 5.

The conditional model has good explanatory power as it can be assessed by looking at the actual values and the regression fitted values as well as the regression residuals (see Figure 5). The coefficient estimates are well determined and exhibit remarkable stability according to the recursive Chow system stability tests and the one-step residuals, (see Figures 6 and 7). Finally, the conditional model is able to accurately forecast demand for real balances over the recent period of three years 2003(3)-2006(2) (see Figure 8 for the one-step ahead forecasts), and this fact is supported by the parameter constancy forecast F-test statistics based on Ω and on $V[e]$ which take the value of $F(36, 24) = 1.605[0.113]$ and $F(36, 24) = 1.198[0.325]$.[4]

5 Conclusion

In this study, we have developed a parsimonious error correction model of money demand in Estonia based on two cointegrating vectors among the system variables including the real money balances, the gross domestic product, the long- and short-term interest rates, and the rate of inflation. The first cointegrating vector is identified as the money demand function whereas the second as the interest rate parity. The model, which exhibits remarkable coefficient stability, was estimated from 1995(1) till 2006(2).

Our main finding is that the long-run income elasticity of money demand in Estonia is larger than unity and moreover it also tends to be larger than that typically reported for the countries of the Euro-zone. In this respect, our results concord with those of Dabusinskas (2005) despite the fact that he estimates the money demand income elasticity in Estonia using much shorter sample period and the different estimation method. Our results also confirm those of Dreger et al. (2006),

[4]For description of these tests see Doornik and Hendry (2001b).

where the long-run money demand functions for the ten new EU member states were estimated using the panel data cointegration techniques. Thus, when taken together the results of our study and the results of the other two studies may point out that the specification of the long-run money demand function in Estonia may differ from that in the countries of the Western Europe that are the members of the Euro-zone.

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Appendix

Table 1: Data information

Variable	Abbreviation	Database code	Database
Money supply (M2) - bln of Estonian Kroons	M2	OA9390010.Q	Global Market Monitor
Consumer price index, 2000=100	CPI	OA9390008.Q	Global Market Monitor
Government bond yield (long-term)	I^l	IA9390133.Q	Global Market Monitor
Deposit rate	I^s	OA9390130.Q	Global Market Monitor
Real GDP - bln of Estonian Kroons, in 2000 prices	Y	OA9390003.Q	Global Market Monitor

Table 2: VAR model: Lag order selection criteria

Lag	LogLik	LR	FPE	AIC	SC	HQ
0	184.266	NA	1.06e-09	-6.4898	-5.0984	-5.9686
1	414.143	339.81	1.50e-13	-15.397	-13.012*	-14.504
2	460.160	58.022*	6.81e-14*	-16.311	-12.932	-15.045*
3	484.597	25.498	8.89e-14	-16.286	-11.914	-14.648
4	525.636	33.901	6.84e-14	-16.984*	-11.617	-14.973

Notes: '*' indicates lag order selected by the criterion; LR – sequential modified LR test statistic (each test at 5% level); FPE – Final prediction error; AIC – Akaike information criterion; SC – Schwarz information criterion; HQ – Hannan-Quinn information criterion

Table 3: VAR model: Diagnostic tests

	Single equation tests					Vector tests	
	$(m-p)_t$	y_t	i_t^l	i_t^s	π_t		
$F_{AR(1-4)}(4,25)$	[0.138]	[0.223]	[0.914]	[0.385]	[0.618]	$F_{AR(1-4)}(100,29)$	[0.749]
$\chi^2_{Norm}(2)$	[0.145]	[0.127]	[0.206]	[0.416]	[0.608]	$\chi^2_{Norm}(10)$	[0.052]*
$F_{Hetero}(22,6)$	[0.995]	[0.998]	[0.993]	[0.909]	[0.998]	$\chi^2_{Hetero}(330)$	[0.853]
$F_{ARCH(1)}(1,27)$	[0.855]	[0.636]	[0.566]	[0.344]	[0.651]		

Notes: ***, **, * indicate significance at the 1%, 5%, and 10% levels, respectively.

Table 4: VAR model: Cointegration tests

	Asymptotic critical values				Osterwald-Lenum (1992) correction			
rank	Trace test	[Prob]	Max test	[Prob]	Trace test	[Prob]	Max test	[Prob]
0	152.18	[0.000]***	71.57	[0.000]***	119.10	[0.000]***	56.01	[0.000]***
1	80.61	[0.000]***	54.57	[0.000]***	63.09	[0.001]***	42.71	[0.000]***
2	26.04	[0.131]	18.85	[0.103]	20.38	[0.408]	14.75	[0.319]
3	7.19	[0.563]	5.30	[0.706]	5.62	[0.741]	4.14	[0.838]
4	1.89	[0.169]	1.89	[0.169]	1.48	[0.224]	1.48	[0.224]

Notes: ***, **, * indicate significance at the 1%, 5%, and 10% levels, respectively.

Table 5: VAR model: Restriction testing

Variable	Test for stationarity, $\chi^2(3)$		Test for trend-stationarity, $\chi^2(3)$		Test for long-run exclusion, $\chi^2(2)$	
$(m-p)_t$	33.175	[0.000]***	44.764	[0.000]***	18.168	[0.000]***
y_t	33.302	[0.000]***	42.148	[0.000]***	20.835	[0.000]***
i_t^l	31.801	[0.000]***	40.209	[0.000]***	38.990	[0.000]***
i_t^s	30.368	[0.000]***	18.509	[0.000]***	48.799	[0.000]***
π	32.001	[0.000]***	50.534	[0.000]***	26.209	[0.000]***

Notes: ***, **, * indicate significance at the 1%, 5%, and 10% levels, respectively.

Table 6: VAR model: Identification of cointegration relationships

step		$(m-p)_t$	y_t	i^l_t	i^s_t	π
1	β_1	1.000	-2.116	0	-0.306	1.555
			(0.14)		(0.06)	(0.23)
	α'_1	-0.131	0.035	0.495	0.734	-0.048
		(0.04)	(0.03)	(0.11)	(0.13)	(0.05)
	β_2	0	0	1.000	-0.819	0
					(0.04)	
	α'_2	0.113	-0.003	-0.451	0.154	0.002
		(0.04)	(0.02)	(0.09)	(0.11)	(0.04)
2	β_1	1.000	-2.177	0	-0.346	1.533
			(0.15)		(0.06)	(0.24)
	α'_1	-0.143	0	0.440	0.782	0
		(0.03)		(0.09)	(0.07)	
	β_2	0	0	1.000	-0.825	0
					(0.04)	
	α'_2	0.118	0	-0.487	0	0
		(0.03)		(0.08)		

Table 7: VECM: Estimation results

	$\Delta(m-p)_t$	Δi^l_t	Δi^s_t
Δy_t			
Δy_{t-1}		1.302**	1.147
		(0.62)	(0.75)
Δi^l_{t-1}	-0.069		0.315**
	(0.04)		(0.15)
Δi^s_{t-1}	0.064*		
	(0.03)		
Δi^s_{t-5}		-0.235***	-0.163*
		(0.06)	(0.08)
$\Delta \pi_t$	-0.344***	-0.714**	
	(0.11)	(0.34)	
$\Delta \pi_{t-1}$		-0.768**	
		(0.30)	
$\beta_{1,t-1}$	-0.123**	0.256**	
	(0.05)	(0.11)	
$\beta_{2,t-1}$	0.120***	-0.190*	0.395***
	(0.03)	(0.09)	(0.10)
$\widehat{\sigma}$	0.028	0.085	0.109

Notes: The deterministic terms (constant, seasonal dummies, and intervention dummies) are not shown. ***, **, * indicate significance at the 1%, 5%, and 10% levels, respectively.

Table 8: VECM: Diagnostic tests

	Single equation tests			Vector tests	
	$(m-p)_t$	i^l_t	i^s_t		
$F_{AR}(1-4)(4{,}26)$	[0.136]	[0.100]	[0.048]**	$F_{AR}(1-4)(36{,}65)$	[0.999]
$\chi^2_{Norm}(2)$	[0.183]	[0.554]	[0.852]	$\chi^2_{Norm}(6)$	[0.396]
$F_{Hetero}(24{,}11)$	[0.997]	[0.999]	[0.176]	$F_{Hetero}(142{,}42)$	[0.998]
$F_{ARCH(1)}(1{,}34)$	[0.646]	[0.138]	[0.917]		

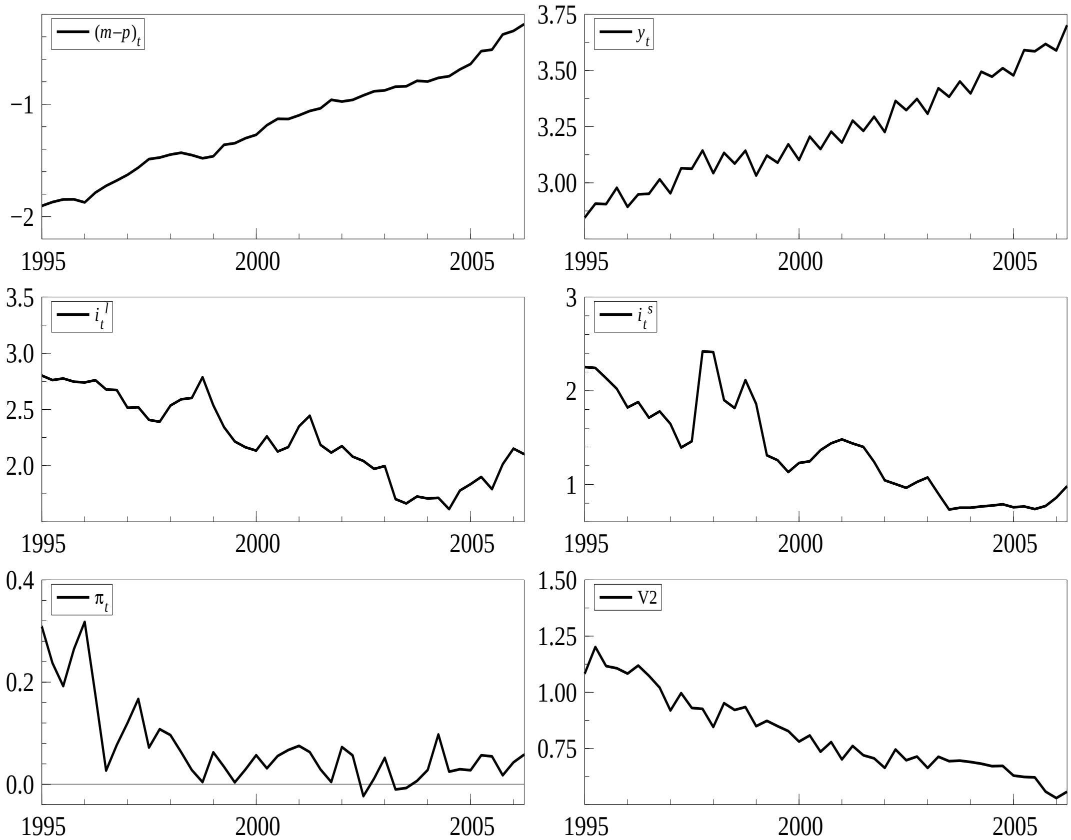

Figure 1: Data: 1995:I - 2006:II

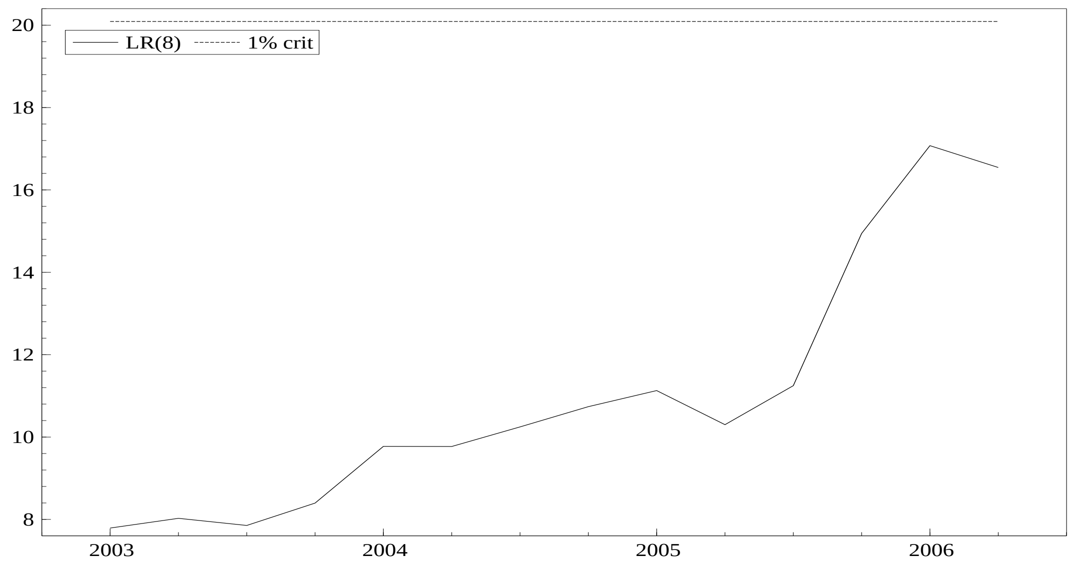

Figure 2: VAR model: Recursively estimated likelihood ratio test statistic of over-identifying restrictions, scaled by the 1% critical value

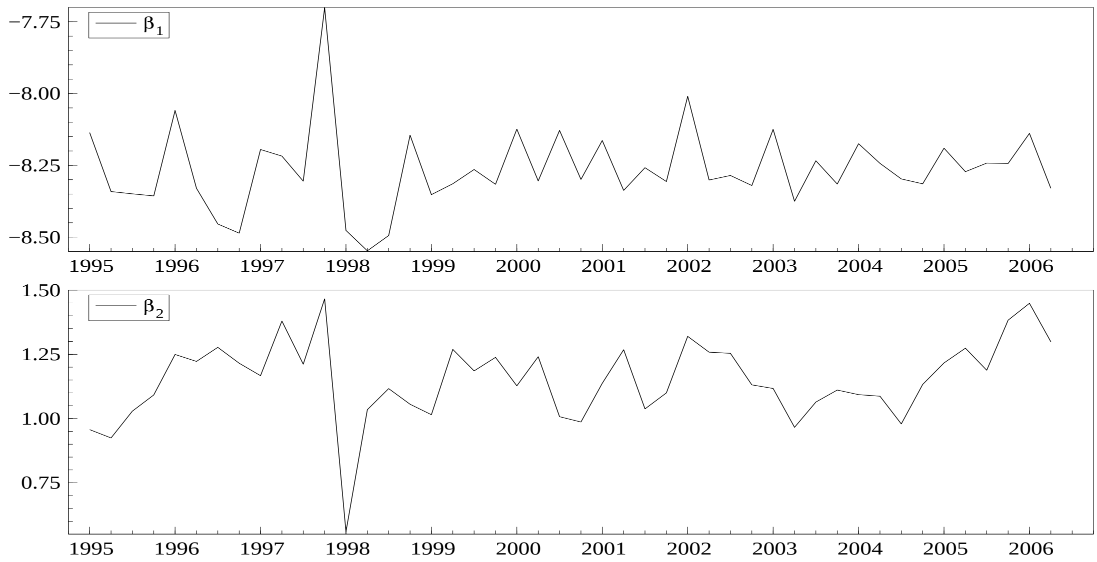

Figure 3: VAR model: Cointegrating vectors

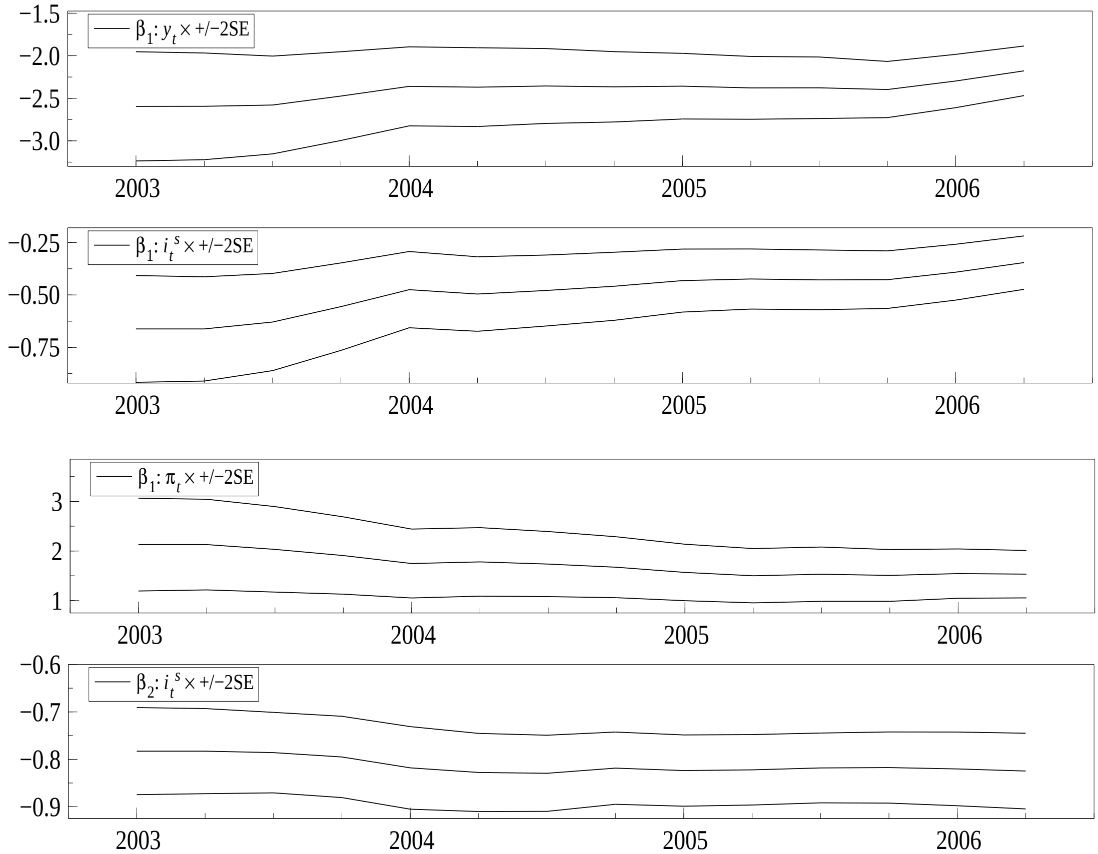

Figure 4: VAR model: Recursively estimated coefficients of the identified cointegrating vectors β_1 and β_2

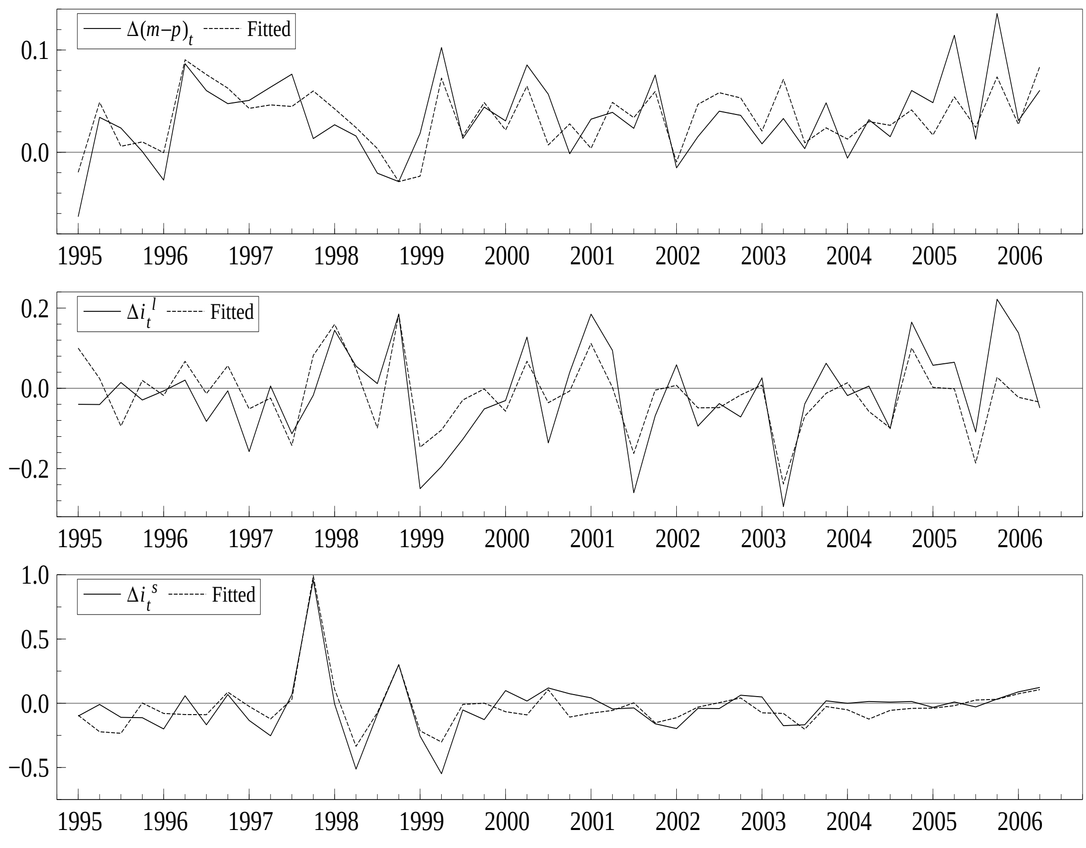

Figure 5: VECM: Actual and fitted values

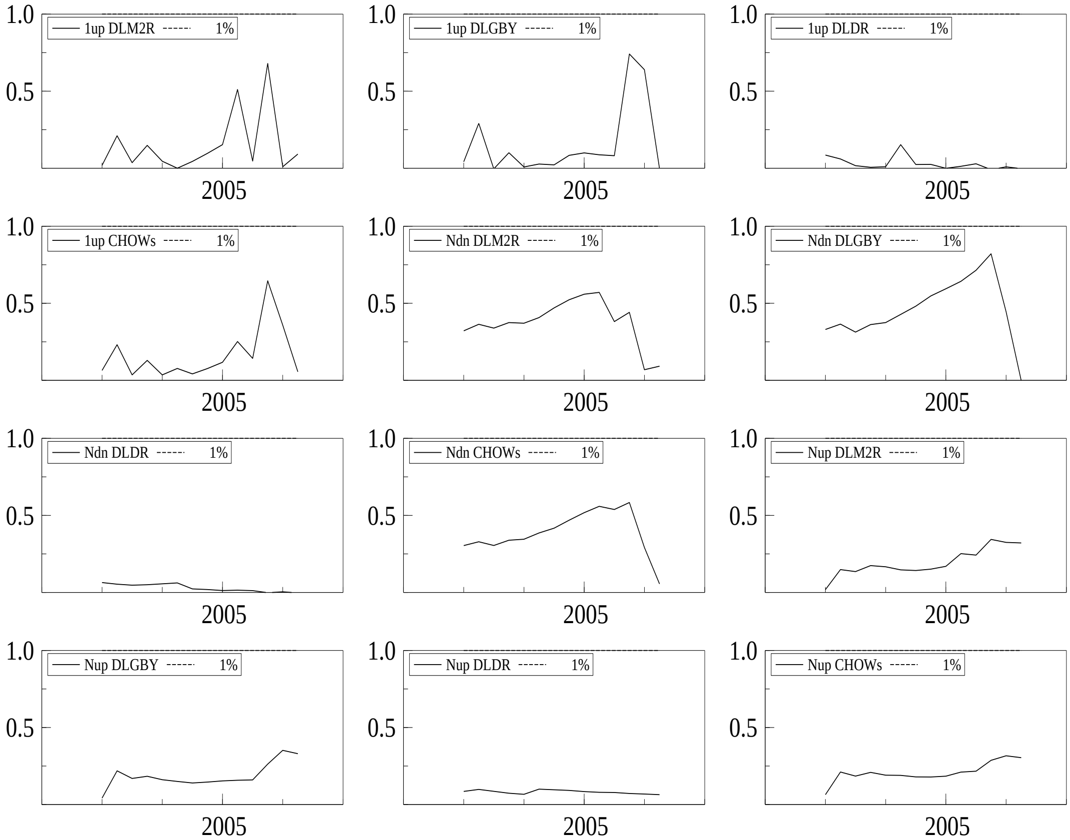

Figure 6: VECM: Chow test statistics, scaled by the 1% critical values

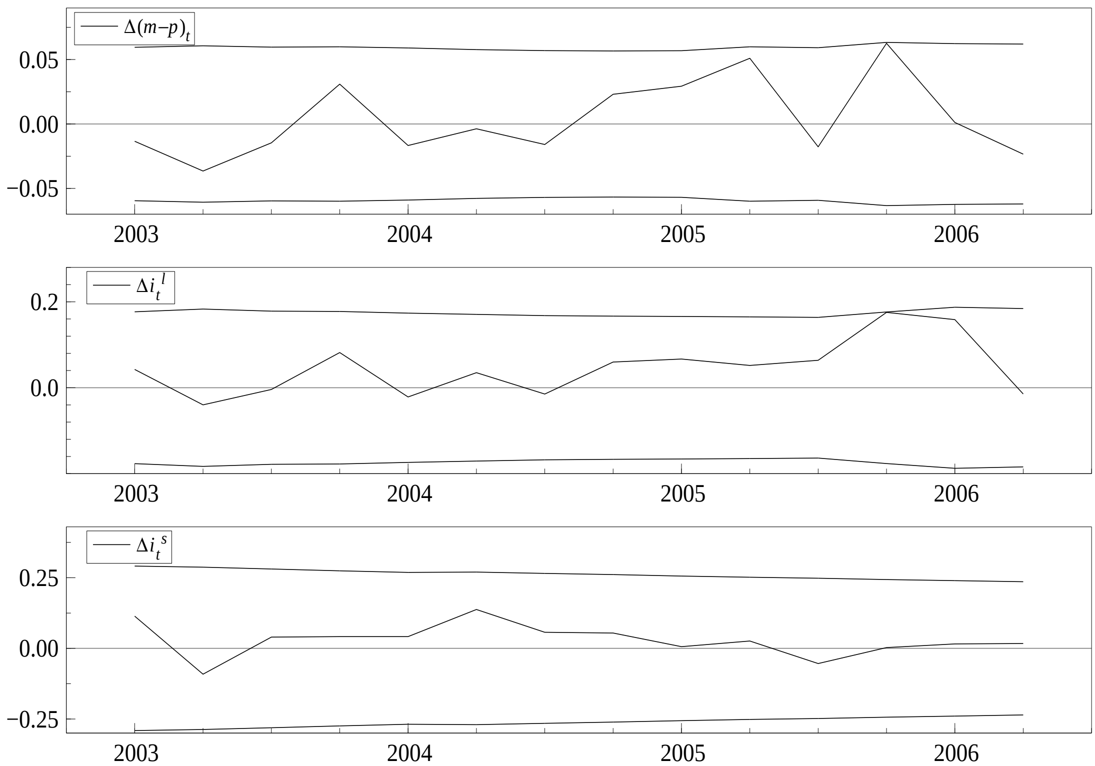

Figure 7: VECM: One-step ahead residuals

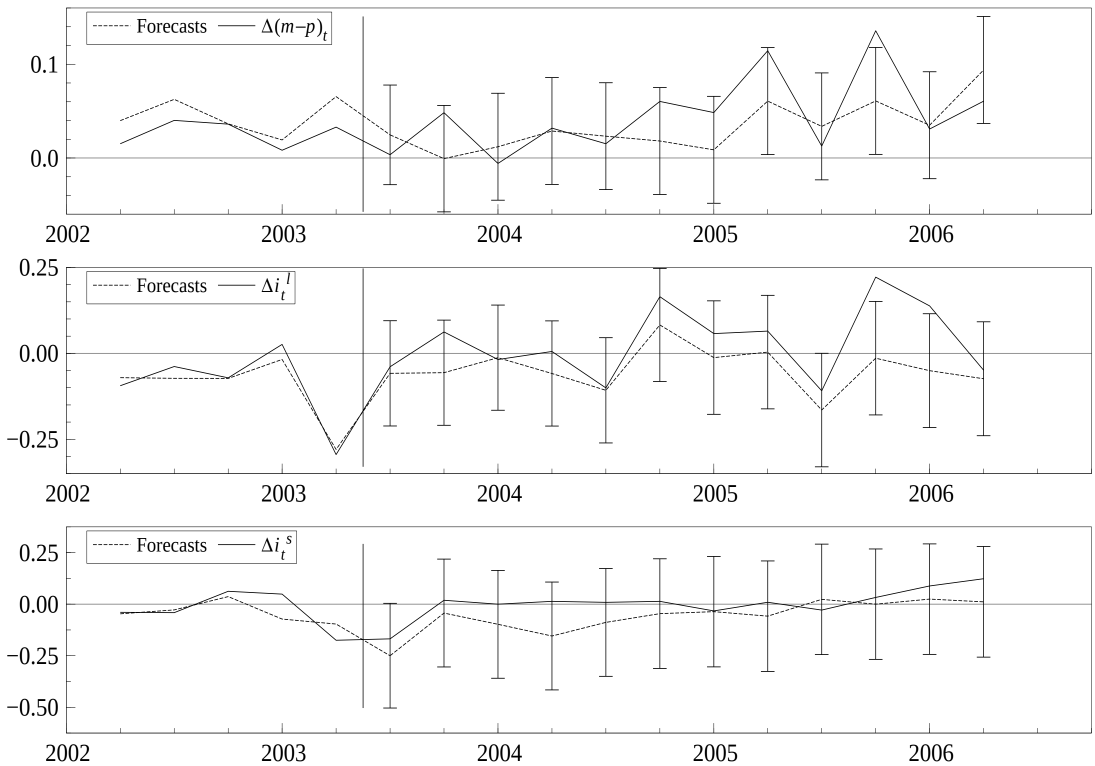

Figure 8: VECM: 1-step ahead *ex post* forecasts