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Dynamic modelling of the demand for money in Latvia

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Abstract

This study develops a parsimonious stable coefficient money demand model for Latvia for the period from 1996 till 2005. A single cointegrating vector between the real money balances, the gross domestic product, the long-term interest rate, and the rate of inflation is found. 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, Latvia

JEL code: C32, E41.

1 Introduction

In the last 15 years since Latvia gained its independence a substantial economic progress has been made. However, its path of economic development was not nice and easy all the way. The yearly 90-ties were characterised by sharp fall in output and by rampant inflation – as a consequence of abrupt breakdown of the traditional economic ties through which Latvia was connected to the rest of the Soviet Union. Under such difficult economic conditions it was necessary to implement a package of stabilisation and structural reform programs. The aim of these programs were price liberalisation, elimination of subsidies to state enterprises, privatisation, tightening of fiscal system, introduction of liberal trade regime, establishment of national central bank and credible national currency, etc.

The implemented reforms have set the Latvian economy on the path of recovery. Despite detrimental influence of the internal banking crisis in 1995 and of the consequences of the Russian financial crisis in 1998, the Latvian economy has been steadily developing. Since 2000 its pace of development has reached record levels with the average GDP growth rate of about 7.6% per annum. The GDP growth comes mainly from the strong domestic demand and substantial increase in exports.

However, such rapid economic development has been accompanied by steep rise in the inflation rate which in the past three years was about 7% on the year-to-year basis. Definitely, this is an alarming phenomenon given the fact that since the beginning of the 1990s the inflation rate was continuously declining from 958.7% per annum in 1992 to 23.1% in 1995 and to 1.8% in 2000. Amongst the factors that contributed to such surge in inflation are the internal factors like rather high expansion of credit brought about by the relatively low interest rates and wage and salary increases beyond of the productivity growth as well as the external factors related to the increase in energy prices observed in this period.

As an acknowledgement of the success in reforming its economy as well as its political system Latvia joined the European Union in May, 2004. A year later, in 2005, it also joined the Exchange Rate Mechanism (ERM II) with the ultimate goal of adopting the Euro currency. After the introduction of the Euro, the responsibility regarding the monetary policy in Latvia (as well as other new Member States of the European Union) will rest on the Governing Council of the European Central Bank (ECB). Hence, due to anticipation of the introduction of the Euro and due to the high relevance of money within the two-pillar monetary strategy of the ECB and its concern regarding the price stability in the Eurozone (European Central Bank, 2003), the analysis of the determinants of the money demand in the new Member States is important.

In this paper we analyse the demand for money in Latvia. The motivation for the paper is twofold. First, we would like to determine which factors shaped the money demand in this country

and whether the inflation rate played a significant role. Second, we would like to investigate whether the developed here model for money demand has been structurally stable in the recent period characterised by the rapid economic development and sharp surge in the inflation level, as described above.

In doing so, our study contributes to the rather small literature on the determinants of money demand in the transition countries of Eastern Europe largely represented by the following articles: 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.

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. Section 5 tests for super exogeneity of the conditioning variables in the money demand equation. The final section concludes.

2 Theoretical considerations

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

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

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 opportunity costs of holding money OC , approximated by the long-term interest rate I and by the inflation π , which largely reflect two main purposes for holding money as stipulated by the economic theory¹. The first motive for holding money is to perform transactions and in our analysis the magnitude of transactions is represented by the real GDP. The second motive for holding money is portfolio diversification, where the interest rate indicates the rate of return to financial assets not included in the monetary aggregates and the inflation rate represents the costs of holding money in spite of holding real assets (Ericsson, 1998).

¹Often a short-run interest rate is included in the long-run money demand function in order to account for the rate of return of the short term assets included in the monetary aggregate. However, in our application we could not reject the null hypothesis that the short-run interest rate could be omitted from the long-run money demand equation once we have accounted for the presence of the long-run interest rate as well as of the inflation rate. Therefore the short-run interest rate has been omitted from the further analysis

The long-run money demand, when log transformed, reads as follows²

$$\ln\left(\frac{M}{P}\right) = \alpha_Y * \ln(Y) + \alpha_I * \ln(I) + \alpha_\pi * \pi + ec \quad (1)$$

where the coefficients α_Y , α_I , and α_π denote the long-run elasticities of money demand with respect to income and to interest rate, and the long-run semi-elasticity of money demand with respect to inflation, respectively. The former coefficient is positive as demand for money increases with income, 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.) and from the International Financial Statistics, see Table 1. The quarterly data span years 1996(1) – 2006(4), such that the sample size is $T = 44$. The following transformations of the original data have been carried out: $(m - p) = \ln(M2/CPI)$ – the real money balances, $y = \ln(GDP/PGDP)$ – the real GDP, $i = \ln(I)$ – the long-term interest rate, and $\pi = 4\Delta p$ is the annualised inflation rate. Observe that we have taken the logarithmic transformation of the long-term interest rate that we use in our subsequent analysis³. We have done this in order to account for the fact that the variation as well as level of the interest rate 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 variable 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 Latvia, we follow the general-to-specific approach advocated in Hendry and Mizon (1993) and Hendry and Juselius (2000, 2001), inter alia. In

²Essentially, this specification of the money demand function is very similar to that in Dreger and Wolters (2006).

³This is not an unusual thing to do, see discussion in Ericsson (1998, p. 297).

particular, we start with an unrestricted VAR(n) model transformed into the error-correction form

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

where $x_t = ((m-p)_t, y_t, i_t, \pi_t)'$ is the $k \times 1$ vector of variables described above and μ_t denotes the deterministic terms such as a constant term, a deterministic trend, and seasonal dummies.

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 test for the long-run weak exogeneity of the system variables. We use the results of the weak exogeneity tests in order to build a parsimonious time series model for the real money balances that satisfactorily passes all diagnostic tests, displays constant coefficients, and possesses the ability to accurately forecast the dependent variable 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 $T = 44$. It seems that the VAR(1) model can adequately describe the data as the misspecification tests report no serious departures from the underlying model assumptions, see Table 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; F_{RESET} – Ramsey (1969) test for functional form misspecification. 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). The CUSUM tests were performed in Eviews 5.1.

Having found the adequate unrestricted model, the next step is to proceed 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 3 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 in order to avoid explosive roots in the unrestricted model we have restricted the deterministic trend to the cointegration space. The test results suggest the presence of one cointegrating relation in the system.

Thus, we impose the cointegration rank $r = 1$ 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 a cointegrating vector. This test has an asymptotic χ^2 distribution with the $r = 1$ degrees of freedom. Finally, the test for the long-run weak exogeneity investigates whether the dependent variables adjust to the equilibrium errors represented by a cointegrating relation.

Table 4 reports the results of the tests for (trend-)stationarity and long-run exclusion, performed on the matrix of the long-run coefficients, and the tests for long-run weak exogeneity, performed on the matrix of the adjustment coefficients. According to the stationarity test, the null hypothesis that each variable is $I(0)$ around a linear deterministic trend is decisively rejected. The tests for the long-run exclusion rejects the null hypothesis that any of the $(m - p)_t$, y_t , and i_t variables can be individually excluded from the cointegrating vector at the 1% significance level and the inflation variable π – only at the 10% significance level.

According to the univariate long-run weak exogeneity test results, we cannot reject the null hypothesis that y_t and π_t variables are weakly exogenous with respect to the long-run parameters at the usual significance levels. However, the null hypothesis in question is marginally rejected for the interest rate variable i_t at the 5% level. Nevertheless, according to the joint test for the long-run weak exogeneity of these three variables we are unable to reject the null hypothesis with the log likelihood ratio test statistic of 5.104[p=0.164]. In order to check, whether this result is robust to the change in the sample size, we report the value of the recursive test statistics of the joint null hypothesis, scaled by the 5% critical value, in the lower panel of Figure 2. Observe, that the restriction that these three variables y_t , i_t , and π_t are weakly exogenous with respect to the long-run parameters is accepted for all sample sizes. Hence, this restriction seems to be reasonable, and in our further analysis we treat these three variables as weakly exogenous with respect to the long-run parameters.

4.2 Long-run money demand function

Imposing the long-run weak exogeneity restrictions on the y_t , i_t , and π_t variables results in the following cointegrating vector with the corresponding standard errors reported in parentheses below

the coefficient estimates

$$(m - p)_t = \underset{(0.181)}{2.851} y_t - \underset{(0.029)}{0.087} i_t - \underset{(0.121)}{0.426} \pi_t + \underset{(0.003)}{0.012} trend + constant + ec_t \quad (3)$$

Observe that all the coefficient estimates have the expected signs and all estimates are significantly different from zero at the conventional significance levels. The estimated cointegrating vector is depicted in Figure 3. The recursively calculated coefficients of the cointegrating relationship are displayed in Figure 4 and these are remarkably stable.

The point estimate of the income elasticity is 2.851 and it is significantly larger than unitary income elasticity. This implies that money tends to increase proportionally more than the real GDP. This is in line with observation that money are held not only for transaction motives but also for portfolio decisions. Moreover, as it is pointed out in Fase and Winder (1999), estimates of the income elasticity above one may suggest importance of the the wealth effects in the demand for money. The point estimate of the interest rate elasticity is -0.087 and it is significantly different from zero. It's relatively small size may point out to difficulties in controlling money stock. Lastly, the estimate of the semi-elasticity of the money demand with respect to inflation is -0.426 , which is also significantly different from zero. It indicates that the money tend to decrease proportionally less than inflation.

At this point, it is instructive to compare our 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.⁴ As seen, our point estimate of income elasticity is rather large in comparison with that reported for the Euro-zone.

Next, it appears that our point estimate of long-run interest rate elasticity is significantly smaller than the reported mean of the estimated long-run elasticities from 440 and 367 money demand regressions reported in Fase (1993) (-0.25) and Knell and Stix (2003) (-0.34), respectively, see Knell and Stix (2004).

Finally, although Knell and Stix (2004, 2006) does not summarise the semi-elasticity of money demand with respect to inflation, we can compare our estimate with that reported in Dreger and Wolters (2006) that is about -4.5 . The inflation semi-elasticity estimate obtained for Latvia is also significantly lower.

⁴The 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.

To summarise, our estimates of the parameters of the long-run money demand function for Latvia seem to significantly differ from those reported for the Euro-zone countries. The estimate of the income elasticity tends to be significantly larger, whereas the estimates of the interest rate elasticity and the inflation semi-elasticity tend to be significantly smaller. 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 Latvia 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 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, where we condition on the current and past values of the weakly exogenous variables, and on the first lag of the error correction term. After removing the variables that have turned to be insignificant, the estimated conditional model for the real money balances $(m - p)_t$ looks as follows

$$\begin{aligned}
 \Delta(m - p)_t = & \quad 6.534 \quad + \quad 1.262 \Delta y_t \quad - \quad 0.480 \Delta y_{t-1} \quad - \quad 0.142 \Delta \pi_{t-1} \\
 & (0.598) \quad \quad (0.135) \quad \quad (0.160) \quad \quad (0.0639) \\
 & - \quad 0.747 \, ec_{t-1} \quad + \quad 0.074 \, D0201_t \\
 & (0.069) \quad \quad (0.019) \\
 & \hat{\sigma} = 0.018, \, R^2 = 0.819, \, T = 44, \, F_{AR(1-4)}(4, 34) = 1.006[0.418], \\
 & F_{ARCH(1-1)}(1, 36) = 0.014[0.906], \, \chi^2_{Norm}(2) = 0.939[0.625], \\
 & F_{Hetero}(9, 28) = 0.756[0.656], \, F_{Hetero-X}(15, 22) = 0.772[0.692], \\
 & F_{RESET}(1, 37) = 1.2673[0.2675]
 \end{aligned} \tag{4}$$

with the corresponding standard errors reported in parentheses below the coefficient estimates. The conditional model (4) is parsimonious but at the same time the diagnostic tests show no signs of misspecification. Observe that the error-correction term is highly significant and it has the expected sign. As expected, the signs of the short-run dynamics coefficients indicate that money balances tend to react positively to upward changes in the economic activity and negatively to the surge in inflation.

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 one-step residuals, the recursive Chow stability tests (see Figure 6), the CUSUM test and the CUSUM test

of squares (see Figures 7 and 8, respectively). The recursively estimated coefficients are displayed in Figure 9. Finally, the conditional model is able to accurately forecast demand for real balances over the recent period 2003(1)-2006(4) (see Figure 10 for the one-step ahead forecasts), and this fact is supported by the Chow parameter constancy forecast F -test statistic which takes the value of $F(16, 22) = 0.440$ [$p = 0.951$].

5 Testing exogeneity

In the previous section we have developed the conditional model for the real money balances where we found that such variables as income, long-run interest rate as well as inflation are weakly exogenous with respect to the long-run parameters. As the next step we analyse whether these variables could be considered as *super exogenous* with respect to the parameters of the money demand model. To do so, we need to check whether these parameters are *invariant* to the interventions observed during the investigation period (see Engle, Hendry, and Richard, 1983)⁵. If this is the case then policy analysis can be carried out by changing the processes driving these variables. In the following we apply the methods suggested in Hendry (1988) and Engle and Hendry (1993) for testing super exogeneity of the conditioning variables.

5.1 Parameter constancy of marginal models

The testing procedure advocated in Hendry (1988) compares the outcomes of parameter constancy tests of the conditional and the marginal models. If the parameters of the conditional model are constant, as shown above, but the marginal models have nonconstant parameters, then the parameters of the conditional model cannot depend on those of the marginal models.

The following marginal models were estimated using the fourth-order autoregressive processes (and seasonal dummies when necessary)

$$\begin{aligned}
 \Delta i_t = & - 0.226 \Delta i_{t-3} + 0.191 \Delta i_{t-4} - 0.005 - 0.400 D972_t \\
 & (0.111) \quad (0.112) \quad (0.015) \quad (0.096) \\
 \hat{\sigma} = & 0.094, R^2 = 0.420, T = 43, F_{AR(1-4)}(4, 35) = 0.875[0.489], \\
 F_{ARCH(1-1)}(1, 37) = & 0.009[0.924], \chi^2_{Norm}(2) = 7.912[0.020]*, \\
 F_{Hetero}(5, 33) = & 0.206[0.958], F_{Hetero-X}(6, 32) = 0.167[0.984], \\
 F_{RESET}(1, 38) = & 0.014[0.905]
 \end{aligned} \tag{5}$$

⁵These interventions are modelled by means of the impulse dummies $Dyyq$ which take value of 1 in the corresponding quarter q of 19yy or 20yy and zero otherwise.

$$\begin{aligned}
\Delta\pi_t = & - 0.003 - 0.024 Q_t - 0.058 Q_{t-1} - 0.089 Q_{t-2} + 0.109 \Delta D961_t \\
& (0.004) \quad (0.011) \quad (0.011) \quad (0.011) \quad (0.019) \\
\hat{\sigma} = & 0.025, R^2 = 0.767, T = 44, F_{AR(1-4)}(4, 35) = 0.743[0.569], \\
F_{ARCH(1-1)}(1, 37) = & 4.018[0.052], \chi^2_{Norm}(2) = 2.623[0.267], \\
F_{Hetero}(5, 33) = & 1.673[0.169], F_{Hetero-X}F(5, 33) = 1.673[0.169], \\
F_{RESET}(1, 38) = & 3.2318[0.080]
\end{aligned} \tag{6}$$

$$\begin{aligned}
\Delta y_t = & 0.336 \Delta y_{t-4} + 0.013 - 0.009 Q_t + 0.023 Q_{t-1} - 0.011 Q_{t-2} \\
& (0.146) \quad (0.003) \quad (0.006) \quad (0.007) \quad (0.007) \\
\hat{\sigma} = & 0.013, R^2 = 0.752, T = 43, F_{AR(1-4)}(4, 34) = 0.198[0.937], \\
F_{ARCH(1-1)}(1, 36) = & 1.052[0.312], \chi^2_{Norm}(2) = 2.462[0.292], \\
F_{Hetero}(5, 32) = & 0.851[0.524], F_{Hetero-X}F(8, 29) = 0.584[0.782], \\
F_{RESET}(1, 37) = & 0.545[0.465]
\end{aligned} \tag{7}$$

The one-step-ahead residuals as well as the Chow parameter constancy tests of equations (5)–(7) are displayed in Figures (11)–(13). Parameter constancy is rejected for long-run interest rate as well as for inflation rate whereas the marginal model for income seems to display constant parameters. Thus, the joint occurrence of the structural breaks in the conditioning variables (inflation rate and long-run interest rate) and parameter constancy of the coefficients of the conditional model indicate that these variables are super exogenous in the money demand model (4).

5.2 Testing invariance

In order to conduct a formal test we follow Engle and Hendry (1993) who suggest to proceed as follows. After having determined that the marginal processes display parameter instability one can test invariance by first augmenting the marginal models with additional variables that capture changes in the respective parameters such that the parameter constancy is restored and second by inserting these intervention variables into the conditional model and testing for their significance. The rationale behind this type of tests is that if the parameters of the conditional model are invariant to changes in the marginal processes then including these intervention variables should not improve fit of the conditional model.

We have shown above that the marginal models for long-run interest and inflation rates display parameter instability. Therefore those models need to be augmented with intervention variables in order to restore parameter stability. The following marginal models were obtained using a number

of dummy variables:

$$\begin{aligned}
\Delta i_t = & - 0.276 \Delta i_{t-3} + 0.207 \Delta i_{t-4} - 0.021 - 0.381 D972_t \\
& (0.088) \quad (0.088) \quad (0.012) \quad (0.074) \\
& + 0.336 D041_t + 0.156 (D054_t + D062_t) \\
& (0.074) \quad (0.054) \\
\hat{\sigma} = & 0.072, R^2 = 0.671, T = 43, F_{AR(1-4)}(4, 33) = 1.700[0.174], \\
F_{ARCH(1-1)}(1, 35) = & 0.671[0.418], \chi^2_{Norm}(2) = 0.232[0.890], \\
F_{Hetero}(7, 29) = & 0.654[0.708], F_{Hetero-X}(9, 27) = 0.505[0.857], \\
F_{RESET}(1, 36) = & 0.103[0.747]
\end{aligned} \tag{8}$$

$$\begin{aligned}
\Delta \pi_t = & - 0.006 - 0.017 Q_t - 0.081 Q_{t-1} - 0.089 Q_{t-2} + 0.095 \Delta D961_t \\
& (0.003) \quad (0.007) \quad (0.00858) \quad (0.007) \quad (0.0133) \\
& - 0.050 \Delta D061_t + 0.041 (D024_t - D044_t) + 0.059 (D012_t + D042_t + D052_t) \\
& (0.013) \quad (0.012) \quad (0.012) \\
\hat{\sigma} = & 0.017, R^2 = 0.895, T = 44, F_{AR(1-4)}(4, 32) = 0.588[0.673], \\
F_{ARCH(1-1)}(1, 34) = & 0.667[0.420], \chi^2_{Norm}(2) = 2.608[0.272], \\
F_{Hetero}F(10, 25) = & 0.817[0.616], F_{RESET}(1, 35) = 3.528[0.069]
\end{aligned} \tag{9}$$

The corresponding outcomes of the parameter stability tests are displayed in Figures 8 and 9. Indeed, the sets of dummy variables $(D041_t, (D054_t + D062_t))$ and $((D024_t - D044_t), (D012_t + D042_t + D052_t), \Delta D061_t)$ for equations (8) and (9), respectively, are important for achieving stability, and therefore they can be seen as the determinants of non-constancies in the inflation and the interest rate processes. Testing the significance of these sets of dummies in the conditional model (4) yields the following test statistic $F(5, 32) = 0.800[0.559]$. Moreover, none of the individual dummies were found to be significant. Hence the null hypothesis of invariance in the money demand function cannot be rejected.

At this point, we can check the conclusions of the Johansen test for weak exogeneity (see Table 4) by adding the lagged error correction term ec_{t-1} to the marginal models. This yields the following test statistics: $t_{\Delta i} = -1.38, t_{\Delta \pi} = -1.28, t_{\Delta y} = 0.645$ for the long-run interest rate, inflation, and income variables, respectively. As seen, none of them is significant at the usual levels. In addition to that, one can perform a Wu-Hausman test for weak exogeneity (suggested in Engle and Hendry (1993)) which is tantamount to testing for independence between the conditioning variables and the residuals from the marginal models. This test is carried out by adding the corresponding residuals from equations (7), (9), and (8) to the conditional model (4) and testing

their significance therein. The corresponding F -test statistic is $F(3, 34) = 0.189[0.903]$. Hence the results of these additional tests for long-run weak exogeneity of the long-run interest rate, inflation, and income variables are consistent with those obtained using the Johansen procedure.

6 Conclusion

In this study, we have developed a parsimonious error correction model of money demand in Latvia based on the single cointegrating vector between the real money balances, the gross domestic product, the long-term interest rate, and the inflation rate. The model, which exhibits remarkable coefficient stability over the period characterised by the high growth rates of the GDP as well as by sharp increase in the inflation rate in the recent three years, was estimated from 1996 till 2006. The fact that the long-run interest rate and inflation variables were found to be super exogenous with respect to the parameters of the money demand function implies that suitable policy analysis can be conducted by varying the determinants of these variables.

Our main findings are following. On the one hand, we find that the long-run income elasticity is larger than unity and also tend to be larger than that typically reported for the countries of the Euro-zone. The long-run interest rate elasticity and the long-run inflation semi-elasticity seem to be smaller than those observed for the Euro area, on the other hand. Our results concord with those of Dreger et al. (2006), where the long-run money demand functions for the ten new EU member states were estimated using the panel data cointegration techniques, and together they may point out that the specification of the long-run money demand function for these economies may differ from that of the members of the Euro-zone.

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Table 1: Data information

Variable	Abbreviation	Database code	Database
Money supply (M2) - mln of Latvian Lats	M2	OA9410010.Q	World Market Monitor
Consumer price index 2000=100	CPI	OA9410008.Q	World Market Monitor
Government bond yield (long-term)	I	IA9410133.Q	World Market Monitor
Nominal GDP - bln of Latvian Lats	Y	OA9410002.Q	World Market Monitor
GDP deflator 2000=100	PGDP	OA9410111.Q	World Market Monitor

Table 2: VAR model: diagnostic tests

Single equation tests					Vector tests	
	$(m-p)_t$	y_t	i_t	π_t		
$F_{AR(1-4)}(4,31)$	[0.682]	[0.717]	[0.606]	[0.859]	$F_{AR(1-4)}(64,64)$	[0.099]
$\chi^2_{Norm}(2)$	[0.804]	[0.198]	[0.894]	[0.206]	$\chi^2_{Norm}(8)$	[0.341]
$F_{Hetero}(10,24)$	[0.377]	[0.815]	[0.130]	[0.008]**	$F_{Hetero}(100,119)$	[0.511]
$F_{Hetero-X}(20,24)$	[0.781]	[0.775]	[0.428]	[0.018]*	$F_{Hetero-X}(200,67)$	[0.998]
$F_{ARCH(1)}(1,33)$	[0.602]	[0.839]	[0.885]	[0.726]		

Table 3: VAR model: cointegration tests

rank	Asymptotic critical values				Osterwald-Lenum (1992) correction			
	Trace test	[Prob]	Max test	[Prob]	Trace test	(T-nm)	Max test	(T-nm)
0	90.09	[0.000]**	49.56	[0.000]**	81.9	[0.001]**	45.06	[0.000]**
1	40.53	[0.084]	20.35	[0.232]	36.84	[0.179]	18.5	[0.354]
2	20.18	[0.221]	11.81	[0.446]	18.35	[0.328]	10.73	[0.552]
3	8.37	[0.229]	8.37	[0.229]	7.61	[0.294]	7.61	[0.294]

Table 4: VAR model: restriction testing

$(m - p)_t$	y_t	i_t	π	trend	$\chi^2(v)$	p-value
Stationarity tests						
.	0	0	0	.	33.798	[0.000]**
0	.	0	0	.	34.040	[0.000]**
0	0	.	0	.	31.063	[0.000]**
0	0	0	.	.	28.939	[0.000]**
Long-run exclusion tests						
0	24.637	[0.000]**
.	0	.	.	.	27.991	[0.000]**
.	.	0	.	.	10.473	[0.001]**
.	.	.	0	.	5.1927	[0.023]*
.	.	.	.	0	14.305	[0.000]**
Long-run weak exogeneity tests						
0	23.728	[0.000]**
.	0	.	.	.	0.393	[0.531]
.	.	0	.	.	5.380	[0.020]*
.	.	.	0	.	0.305	[0.581]
.	0	0	0	.	5.962	[0.114]

Notes: '0' denotes the zero restriction on the coefficient of the corresponding variable, '.' denotes unrestricted coefficient. The number of the degrees of freedom v of the χ^2 distribution corresponds to the number of zero restrictions imposed.

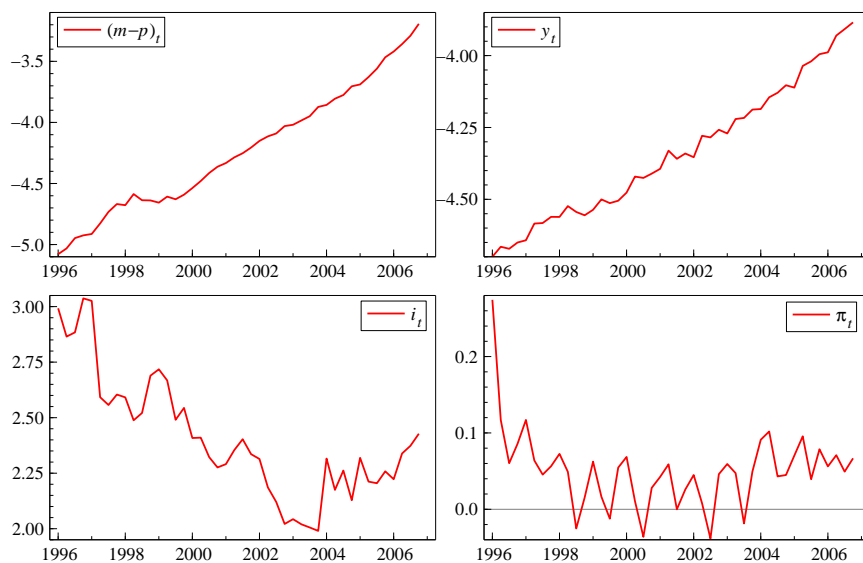
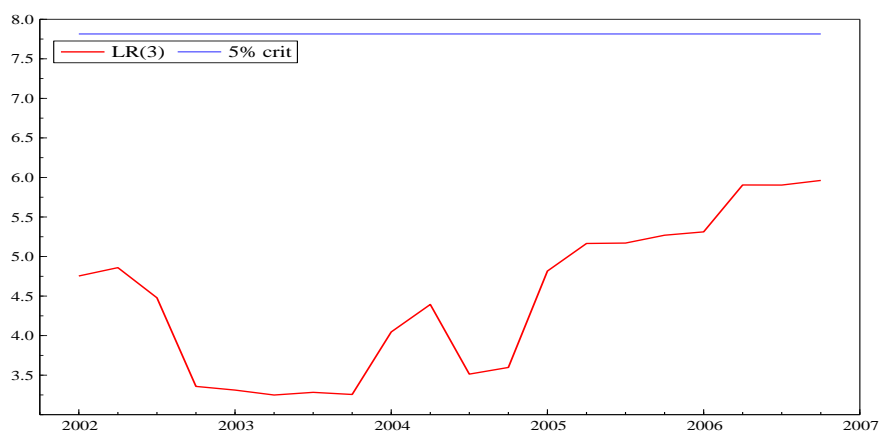


Figure 1: Data: 1996 - 2006

Figure 2: VAR model (2): Recursive LR test statistics for joint long-run weak exogeneity of y_t , i_t , and π variables, scaled by the 5% critical values

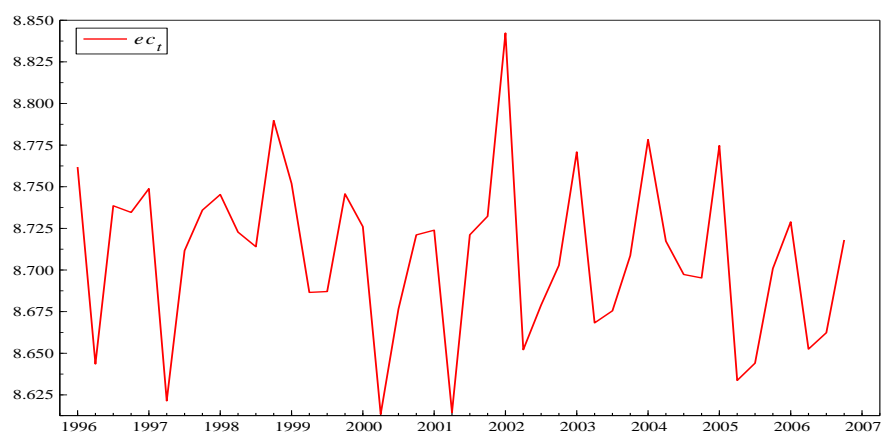


Figure 3: VAR model (2): Cointegrating vector

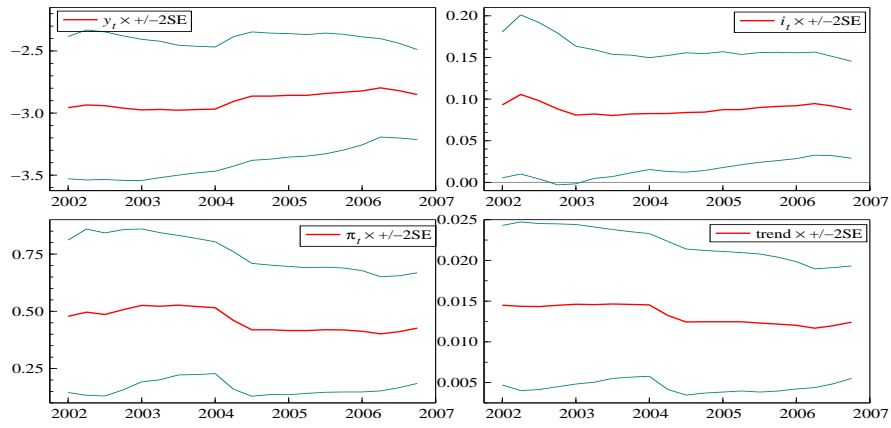


Figure 4: VAR model (2): Recursively estimated coefficients of the cointegrating vector

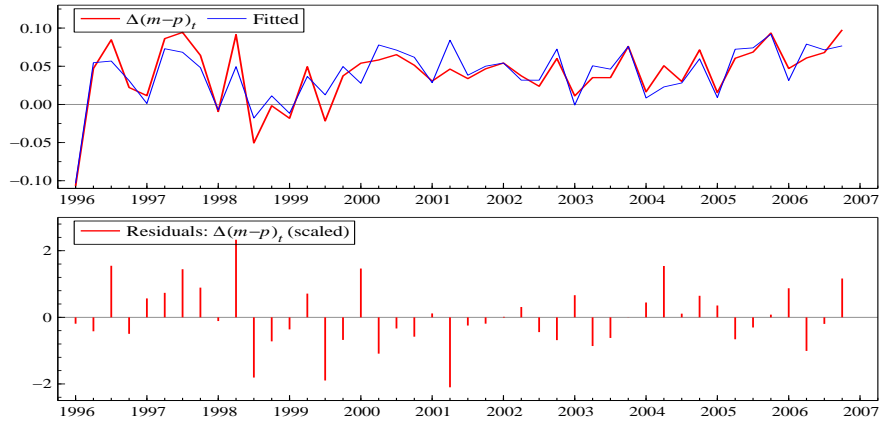


Figure 5: ECM model (4): Actual and fitted values (upper panel); Residuals (lower panel)

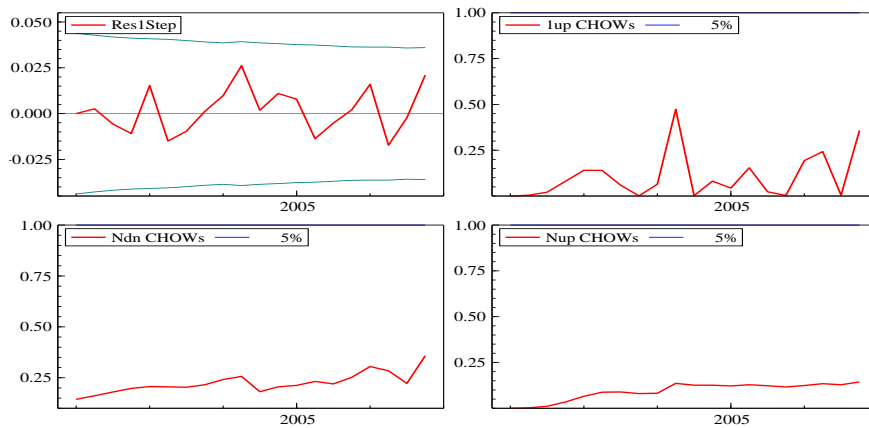


Figure 6: ECM model (4): 1-step residuals (Res1step) and Chow test statistics

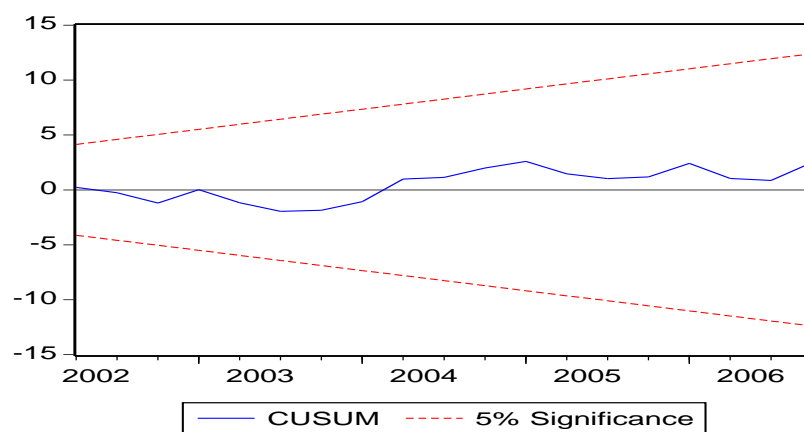


Figure 7: ECM model (4): CUSUM test

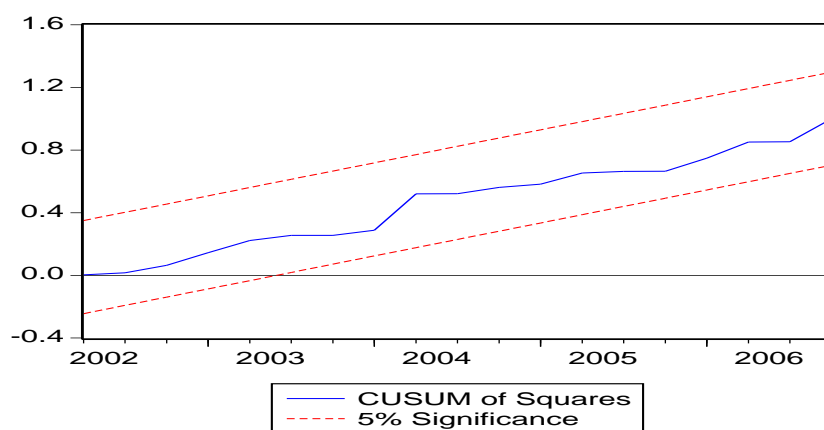


Figure 8: ECM model (4): CUSUM of squares test

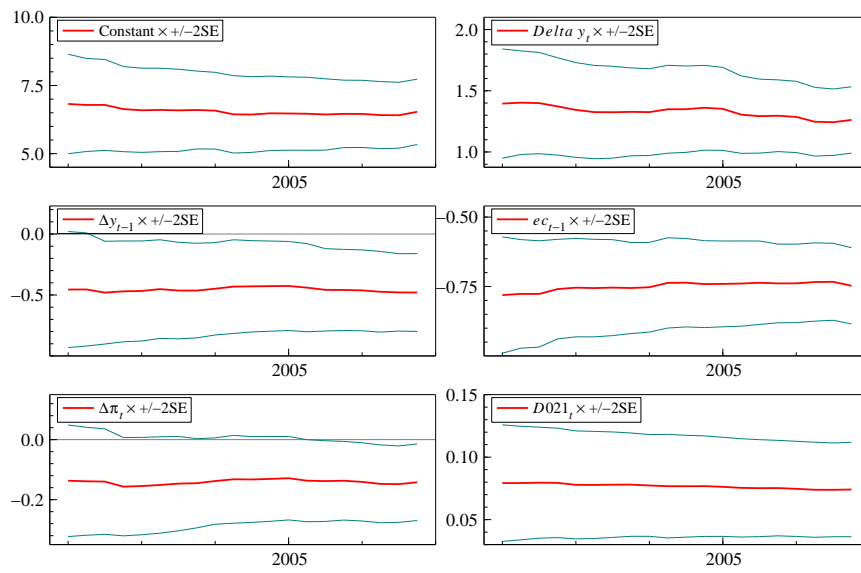


Figure 9: ECM model (4): Recursively estimated coefficients

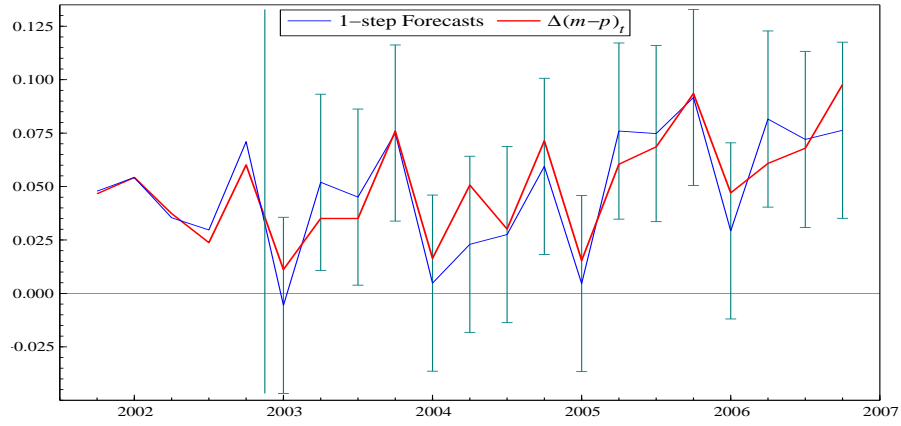


Figure 10: ECM model (4): 1-step ahead *ex post* forecasts

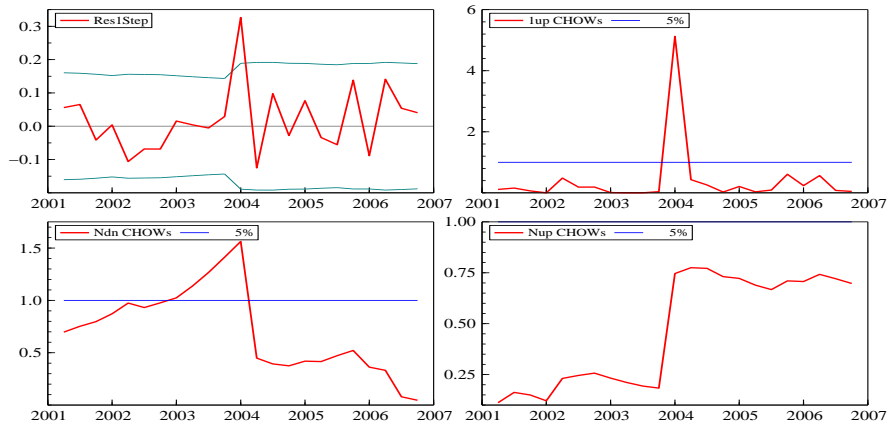


Figure 11: Marginal model for Δi_t (5): Parameter constancy tests

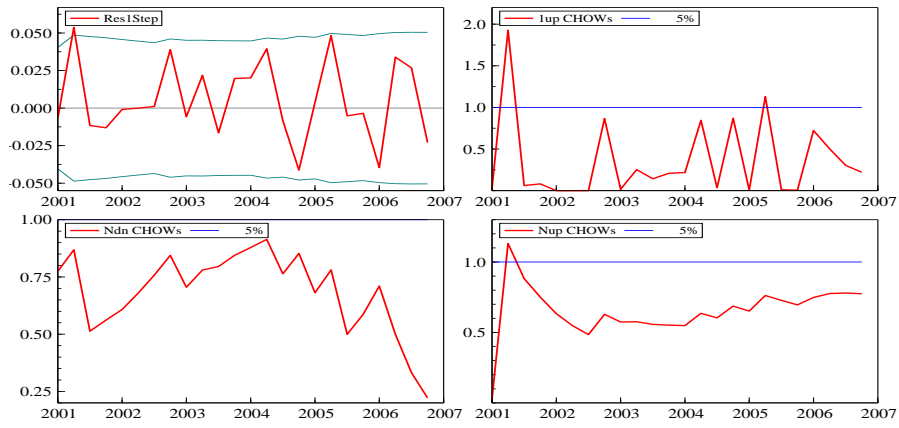


Figure 12: Marginal model for $\Delta\pi_t$ (6): Parameter constancy tests

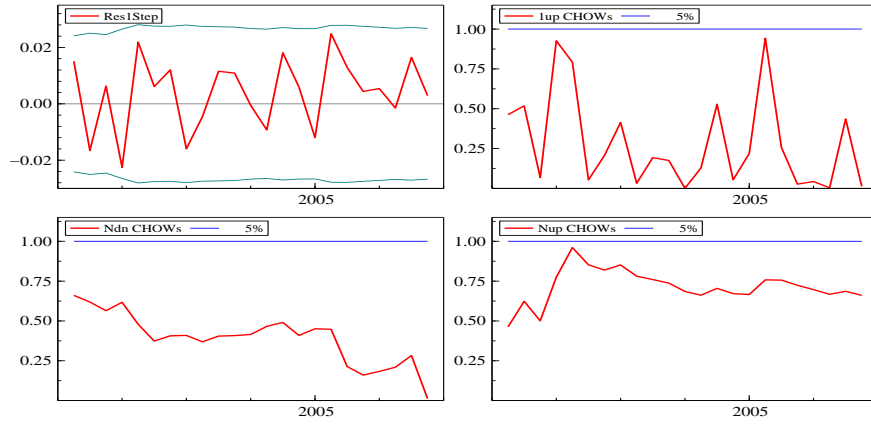


Figure 13: Marginal model for Δy_t (7): Parameter constancy tests

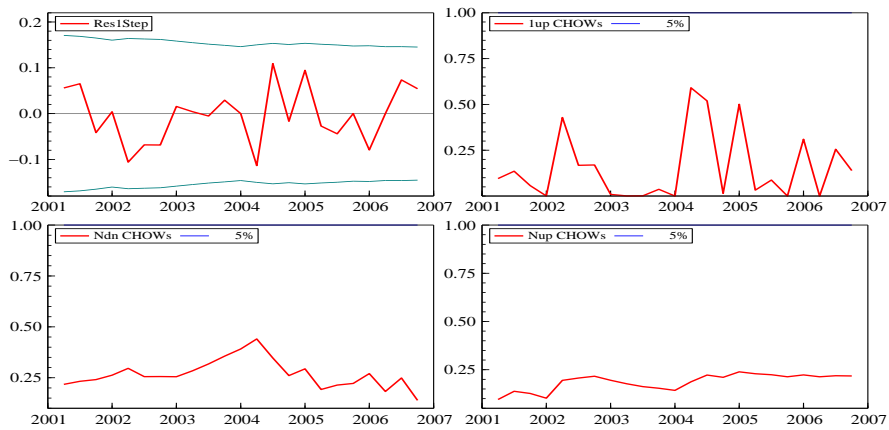


Figure 14: Marginal model for Δi_t (8): Parameter constancy tests

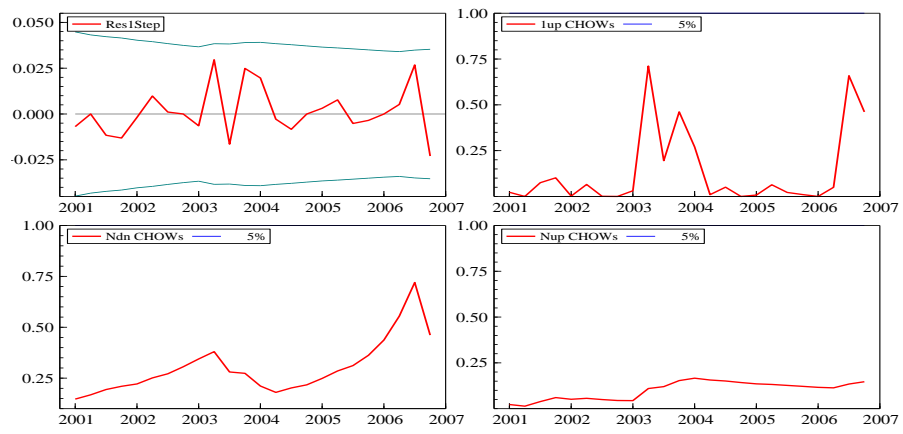


Figure 15: Marginal model for $\Delta\pi_t$ (9): Parameter constancy tests