Dynamic modelling of the demand for money in Latvia*

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Abstract

This study develops a money demand model for Latvia for the period from 1996 to 2006. The model is specified in the error-correction form based on a single co-integrating vector between real money balances, gross domestic product, long-term interest rate, and the rate of inflation. The model meets all three requirements for ‘stability’ put forward in Judd and Scadding (1982). The model is both well-specified and highly parsimonious. It demonstrates high explanatory power in sample as well as accurately forecasting real money balances out of sample.

Keywords: M2 money demand, stability, new EU Member States, Latvia
JEL code: C32, E41.

1 Introduction

“No proposition in macroeconomics has received more attention than that there exists, at the level of the aggregate economy, a stable demand for money function.”

Laidler (1982, p. 39)

Demand for money has been and still is one of the most researched topics in macroeconomics. Moreover, as emphasised in Laidler (1982, p. 39), attention is devoted to investigating whether a stable money demand function exists. The significance of the stability of money demand in economic theory is best illustrated by the fact that a stable money demand function can be found in a wide variety of theoretical models like monetarist models (Friedman, 1956), neoclassical models of monetary effects (Sargent and Wallace, 1975), some New Keynesian models (Mankiw, 1991) and empirical real business cycles models (e.g., King et al., 1991). The scope of the research is illustrated by several survey articles such as Fase (1993), Sriram (2001) and Knell and Stix (2003), in which money demand estimation results from 100 papers written between 1972 and 1992, from 28 papers published between 1990 and 1999, and from 68 papers from 1995 to 2002, respectively, are reviewed.

* The paper has benefited from comments made by Niklas Ahlgren and two anonymous referees.
From an empirical point of view, Judd and Scadding (1982, p. 993) outline three key elements characterising a stable demand function. The first condition is that the demand for money relation should be well-specified from a statistical point of view, i.e., the model should demonstrate satisfactory in-sample as well as out-of-sample predictive performance. Second, the model should be parsimonious. Third, the explanatory variables should reflect developments in the real sector of the economy. In sum, the stability of the money demand function implies that changes in monetary aggregates should be reliably related to developments in real economic variables.

This paper develops a stable demand function for money for Latvia. The resulting specification meets all three requirements for ‘stability’ put forward in Judd and Scadding (1982). The model is well-specified. It demonstrates high explanatory power in sample as well as accurately forecasting real money balances out of sample. Finally, the model involves only a small number of explanatory variables: income, the interest rate, and the inflation rate. The first variable is a scale variable, whereas the other two variables represent the opportunity costs of holding money.

Thus, by developing a stable money demand function for Latvia, the paper intends to close a gap in the literature, as it seems that so far factors determining money demand in Latvia and, especially, investigation of its stability, have received little attention. Arguably, Latvia represents an interesting case for studying stability of money demand in an open and small economy which underwent and is still undergoing structural adjustments on its path from a planned to a market economy. The period under investigation covering the years 1996-2006 is characterised by economic recovery after the banking crisis in 1995, by the detrimental consequences of the Russian financial crisis in 1998, and by a highly variable rate of inflation, which stood at about 20% at the beginning of 1996, dropping to 1.8% in 2000, and then rising to about 6-7% per annum in the years 2004-2006. At the same time it is worth mentioning that the Bank of Latvia has managed to maintain a stable exchange rate of the national currency in this volatile environment. The exchange rate of Lats has been stable since it was pegged to SDR at the rate of 0.7997 LVL per SDR in 1994 and subsequently it has been pegged to the Euro at the rate of 1 EUR = 0.702804 LVL in 2005.

Previous research on money demand in Latvia is represented by Tillers (2004) and more recently Dreger, Reimers, and Roffia (2007). The former study estimates the demand for money in Latvia using data from 1996 to 2003. Its main finding is that income elasticity is much higher than unity and that interest rate elasticity is not significantly different from zero. The latter finding is somewhat unsatisfactory, as it implies that money demand does not react at all to changes in interest rates and this in turn may point towards model misspecification. Furthermore, the coefficient estimates of the proposed money demand function are unstable. The main contribution of the latter study consists of use of panel co-integration techniques in order to estimate long-run money demand functions for the ten new EU Member States, including Latvia. In the approach of Dreger et al. (2007) common coefficients of long-run money demand functions are being imposed across all cross-sections, which may be questionable given the heterogeneity of these countries. In addition, the use of panel co-integration techniques relies on the assumption that cross-sections are independent. As acknowledged by the authors themselves, this assumption is most likely to be violated for the data used in their paper and hence their results may be distorted, e.g., by the presence of unmodelled cross-
sectional co-integration (Banerjee, Marcellino, and Osbat, 2001; Urbain, 2004). Moreover, the stability of country-specific long-run money demand functions is not addressed in Dreger et al. (2007).

An additional contribution of the present paper is that it allows us to compare factors determining the demand for money in the post-communist countries of Eastern Europe, which face more or less similar problems on their way to economic restructuring. This small but growing literature is largely represented by the following studies: Bolharyn and Babaian (1998) for the Ukraine, Karla (1999) for Albania, Buch (2001) for Hungary and Poland, Payne (2003) for Croatia, Slavova (2003) for Bulgaria, Andronescu et al. (2004) and Budina et al. (2006) for Rumania, Bahmani-Oskooee and Barry (2000), Kim and Pirttila (2004), and Oomes and Ohnsorge (2005) for Russia, and Pelipas (2006) for Belarus.

The rest of the paper is organised as follows. Section 2 briefly outlines the economic situation in Latvia to be followed by Section 3, which summarises the theoretical considerations behind empirical models for money demand. Section 4 describes data sources and data transformations. Section 5 contains a description of the modelling approach and presents estimation results. In Section 6 presents tests for super exogeneity of conditioning variables in the money demand equation. The final section sums up.

2 The economic situation in Latvia

During the last 15 years since Latvia regained its independence, substantial economic progress has been made. However, the path of economic development was not pleasant and easy all the way. The early 90s were characterised by a sharp fall in output as a consequence of the abrupt breakdown of traditional economic ties through which Latvia was connected to the rest of the Soviet Union. In addition, this period was characterised by rampant inflation brought about by excessive money creation in Russia and other CIS states. Under such difficult economic conditions it was necessary to implement a package of stabilisation and structural reform programmes. The aims of these programmes included price liberalisation, elimination of subsidies to state firms, privatisation, tightening of fiscal policy, introduction of a liberal trade regime, establishment of a national central bank and a credible national currency.

The Bank of Latvia has since its establishment pursued an independent monetary policy aimed at harnessing inflation and maintaining stability of the exchange rate. Its success in fighting inflation is reflected in the continuous fall of annual inflation from 958.7% in 1992 to 23.1% in 1995. The annual inflation rate was 1.8% in 2000. The central bank has also succeeded in maintaining a stable exchange rate. After an initial period of appreciation of the national currency, at the beginning of 1994 this was pegged to the SDR at the rate of 0.7997 LVL per SDR. Since then the exchange rate policy of the Bank of Latvia has been similar to that of a currency board, with the monetary base being backed by gold and foreign currency reserves. In 2005 the Lat was pegged to the Euro at the rate of 1 EUR = 0.702804 LVL. It is worth mentioning that the Bank of Latvia managed to maintain exchange rate stability during the banking crisis of 1995 and the financial crisis in Russia in 1998, which adversely affected the economic situation in Latvia.
Since 2000 the Latvian economy has expanded rapidly, with an average growth rate of about 7.6% per annum. GDP growth comes mainly from strong domestic demand, which further promotes development of services, especially in trade and construction, and a significant increase in exports as a result of a high investment share in GDP. Between 2001 and 2004 the share of investment in GDP was 24.5% on average and investment in fixed assets increased by 40%. Without doubt, such rapid investment growth would not have been possible without significant contributions from foreign direct investment, which Latvia managed to attract. Since the early 1990s, accumulated FDI stock doubled every 3–4 years, so that at the end of 2004 total FDI stock was about 2.3 billion Lats or approximately 30% of annual GDP.

Record GDP growth rates have, however, been accompanied by a sharp increase in inflation, which in 2004-2006 was about 6-7% per annum. Both external factors such as increases in the cost of electricity, natural gas, and other regulated prices as well as fluctuations in the oil price and internal factors such as wage and salary increases beyond productivity growth as well as high expansion of credit brought about by relatively low interest rates have contributed to the rapid surge in inflation. Clearly, such high inflation rates are of great concern to Latvian policy makers, as these may erode gains in competitiveness and as a result undermine export performance. Combined with a stronger demand for imported goods it may adversely affect the current account deficit, which may drop below the hitherto undesirable level.

Summarising, since regaining independence Latvia has made significant progress in reforming its economy towards a market-oriented economy. As it stands, the main challenge for Latvian policy makers is to bring inflation down, which may prove to be not an easy task. Under present economic conditions it is pertinent to investigate the determinants of long-run money demand in Latvia. Our findings will point out whether factors and their contribution to money demand are similar in Latvia as in other transition countries facing a similar task of economic restructuring, as well as in industrially developed countries.

3 Theoretical considerations

In the empirical section below, long-term demand for money is specified in the following general form:

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

where the demand for real balances \(M/P\) is measured as a ratio of a selected monetary aggregate \(M\) in nominal terms and the price level \(P\). Estimation of real demand for money implies that money neutrality and price homogeneity are assumed to hold in the long run. Below, demand for real money is modelled as a function of two categories of variables: a scale variable that reflects the scope of economic activity, typically approximated by real GDP \(Y\), and the opportunity cost of holding money \(OC\), approximated by long-term interest rate \(I\) and

1 Often a short-run interest rate is included in long-run money demand function in order to account for the rate of return of 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 analysis.
by inflation, which largely reflect the two main purposes of holding money as stipulated by economic theory. The first motive for holding money is to perform transactions and in our analysis the magnitude of transactions is represented by real GDP. The second motive for holding money is portfolio diversification, where the interest rate indicates the rate of return on financial assets not included in monetary aggregates, and the inflation rate represents the costs of holding money rather than holding real assets (Ericsson, 1998).

Long-run money demand, when log transformed, reads as follows

\[
\ln \left( \frac{M}{P} \right) = \alpha_Y \times Y + \alpha_I \times I + \alpha_\pi \times \pi + ec
\]

(1)

where \( \alpha_Y, \alpha_I, \) and \( \alpha_\pi \) respectively denote long-run elasticities of money demand with respect to income and the interest rate, and long-run semi-elasticity of money demand with respect to inflation. The first coefficient is positive as demand for money increases with income, whereas the latter two coefficients are negative. An increase in long-term interest rate leads to shifts in the portfolio towards longer-term investment and henceforth reduces demand for money. Similarly, a rise in inflation reduces the value of monetary assets and hence tends to reduce demand for it. Finally, the term \( ec \) denotes the error-correction term measuring deviations from the long-run equilibrium given in equation (1). As mentioned in the survey article of Sriram (2001), this is the standard specification for money demand that is common to most studies, even though each study may differ from the rest in choice of either dependent or independent variables, or both.

4 Data

Data were taken from the World Market Monitor (Global Insight, Inc.): see Table 1. Quarterly data span the years 1996(1) – 2006(4), so that the sample size is \( T = 44 \). The following transformations of original data have been carried out: \((m - p) = \ln(M2/CPI)\) – real money balances, \( y = \ln(GDP/PGDP)\) – real GDP, \( i = \ln(I)\) – long-term interest rate, and \( \pi = 4 \times \Delta \rho \) – annualised inflation rate. Observe that logarithmic transformation of the long-term interest rate has been carried out.

Table 1: Data information

<table>
<thead>
<tr>
<th>Variable</th>
<th>Abbreviation</th>
<th>Database code</th>
<th>Database</th>
</tr>
</thead>
<tbody>
<tr>
<td>Money supply (M2) - million of Latvian Lats</td>
<td>M2</td>
<td>OA9410010.Q</td>
<td>World Market Monitor</td>
</tr>
<tr>
<td>Consumer price index 2000=100</td>
<td>CPI</td>
<td>OA9410008.Q</td>
<td>World Market Monitor</td>
</tr>
<tr>
<td>Government bond yield (long-term)</td>
<td>I</td>
<td>IA9410133.Q</td>
<td>World Market Monitor</td>
</tr>
<tr>
<td>Nominal GDP - billion of Latvian Lats</td>
<td>Y</td>
<td>OA9410002.Q</td>
<td>World Market Monitor</td>
</tr>
<tr>
<td>GDP deflator 2000=100</td>
<td>PGDP</td>
<td>OA9410111.Q</td>
<td>World Market Monitor</td>
</tr>
</tbody>
</table>

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

3 This is not an unusual thing to do, see discussion in Ericsson (1998, p. 297).
well as the level of interest rate were much higher in the first half of our sample than in the second half. The advantage of such a transformation is, of course, that the estimated coefficient on the interest rate variable is to be read as elasticity. Transformed data are depicted in Figure 1.

5 Econometric model

5.1 Inference on co-integration

Modelling of the money demand function in Latvia proceeds according to the general-to-specific approach advocated in Hendry and Mizon (1993) and Hendry and Juselius (2000, 2001), _inter alia_. In particular, one starts with an unrestricted VAR(n) model transformed into error-correction form

\[ \Delta x_t = \Pi x_{t-1} + \sum_{i=1}^{n-1} \Gamma_i \Delta x_{t-i} + \mu_t + \epsilon_t, \quad \epsilon_t \sim N(0, \Sigma) \]  

(2)

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

The remainder of the section unveils as follows. After selecting the lag length of the unrestricted VAR model, a test for co-integration rank is performed and subsequently implied reduced rank restrictions are imposed on the unrestricted VAR model. Then follows a test for long-run weak exogeneity of system variables. The results of weak exogeneity tests are used
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in order to build a parsimonious time series model for real money balances that satisfactorily passes all diagnostic tests, displays constant coefficients, and possesses the ability to accurately forecast the dependent variable during the last four years of our sample period.

First, the lag length order of the unrestricted VAR($n$) model is determined. At this stage, the most parsimonious model should be preferred given the small 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. Univariate as well as multivariate model diagnostic tests comprise: $F_{AR}$ – test of no residual autocorrelation (see Godfrey (1978)); $\chi^2_{Norm}$ – test for normally distributed errors (see Doornik and Hansen (1994); $F_{Hetero}$ and $F_{Hetero-X}$ – White (1980) tests for heteroscedasticity based on original and squared regressors, and on original, squared regressors, and their cross-products; the $F_{ARCH}$ – Engle (1982) test of no residual AutoRegressive Conditional Heteroscedasticity; the $F_{RESET}$ – Ramsey (1969) test for functional form misspecification. 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.

Table 2: VAR model: diagnostic tests

<table>
<thead>
<tr>
<th>Test</th>
<th>Single equation tests</th>
<th>Vector tests</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>($m-p$)</td>
<td>$\gamma_t$</td>
</tr>
<tr>
<td>$F_{AR}$ (1-4)(4,31)</td>
<td>[0.682]</td>
<td>[0.717]</td>
</tr>
<tr>
<td>$\chi^2_{Norm}$ (2)</td>
<td>[0.804]</td>
<td>[0.198]</td>
</tr>
<tr>
<td>$F_{Hetero}$ (10,24)</td>
<td>[0.377]</td>
<td>[0.815]</td>
</tr>
<tr>
<td>$F_{Hetero-X}$ (20,24)</td>
<td>[0.781]</td>
<td>[0.775]</td>
</tr>
<tr>
<td>$F_{ARCH}$ (1)(1,33)</td>
<td>[0.602]</td>
<td>[0.839]</td>
</tr>
</tbody>
</table>

Notes: The symbols * and ** denote significance at 0.05 and 0.01 levels, respectively.

Having found an adequate unrestricted model, the next step is to proceed with imposing restrictions on that model. The co-integration rank of the estimated system is addressed by means of the Johansen Full Information Maximum Likelihood (FIML) procedure. Table 3 reports results of the trace and $\lambda$ - max tests, both using asymptotic critical values and critical values based on finite sample correction (see Osterwald-Lenum, 1992). Observe that in order to avoid a quadratic trend in the unrestricted model, the deterministic trend is restricted to the co-integration space. The test results suggest the presence of one co-integrating relation in the system. Thus, co-integration rank $r = 1$ is imposed on the system (2), which allows testing for (trend-)stationarity, long-run exclusion, and long-run weak exogeneity of the variables in our model conditional on $r = 1$. The test of stationarity of the variables in a VAR model was 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 co-integration space, it investigates whether any of the variables alone belong to the co-integration space. This test has an asymptotic $\chi^2$ distribution with $(k - r) = 3$ degrees of freedom.
The test for long-run exclusion (Johansen and Juselius, 1992) investigates whether any of the variables can be excluded from a co-integrating vector. This test has an asymptotic $\chi^2$ distribution with $r = 1$ degrees of freedom. Finally, the test for long-run weak exogeneity investigates whether the variables adjust to the equilibrium errors represented by the co-integrating relation.

Table 4 reports the results of tests for (trend-)stationarity and long-run exclusion, performed on the matrix of long-run coefficients, and tests for long-run weak exogeneity, performed on the matrix of adjustment coefficients. According to stationarity tests, the null hypothesis that each variable is I(0) around a linear deterministic trend is decisively rejected. The tests for

Table 3: VAR model: cointegration tests

<table>
<thead>
<tr>
<th></th>
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</tr>
</thead>
<tbody>
<tr>
<td>0</td>
<td>90.09 [0.000]**</td>
<td>49.56 [0.000]**</td>
<td>81.9 [0.001]**</td>
<td>45.06 [0.000]**</td>
</tr>
<tr>
<td>1</td>
<td>40.53 [0.084]</td>
<td>20.35 [0.232]</td>
<td>36.84 [0.179]</td>
<td>18.5 [0.354]</td>
</tr>
<tr>
<td>2</td>
<td>20.18 [0.221]</td>
<td>11.81 [0.446]</td>
<td>18.35 [0.328]</td>
<td>10.73 [0.552]</td>
</tr>
<tr>
<td>3</td>
<td>8.37 [0.229]</td>
<td>8.37 [0.229]</td>
<td>7.61 [0.294]</td>
<td>7.61 [0.294]</td>
</tr>
</tbody>
</table>

Notes: The symbols * and ** denote significance at 0.05 and 0.01 levels, respectively.

The test for long-run exclusion (Johansen and Juselius, 1992) investigates whether any of the variables can be excluded from a co-integrating vector. This test has an asymptotic $\chi^2$ distribution with $r = 1$ degrees of freedom. Finally, the test for long-run weak exogeneity investigates whether the variables adjust to the equilibrium errors represented by the co-integrating relation.

Table 4: VAR model: testing restrictions

<table>
<thead>
<tr>
<th>(m-p)_i</th>
<th>y_i</th>
<th>i_i</th>
<th>$\pi_i$</th>
<th>trend</th>
<th>$\chi^2(\nu)$</th>
<th>p-value</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
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<td></td>
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<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Stationarity tests</td>
<td></td>
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<td></td>
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<td></td>
<td></td>
</tr>
<tr>
<td>.</td>
<td>0</td>
<td>0</td>
<td>0</td>
<td>0</td>
<td>33.798</td>
<td>[0.000]**</td>
</tr>
<tr>
<td>0</td>
<td></td>
<td>0</td>
<td>0</td>
<td>0</td>
<td>34.040</td>
<td>[0.000]**</td>
</tr>
<tr>
<td>0</td>
<td>0</td>
<td></td>
<td>0</td>
<td>0</td>
<td>31.063</td>
<td>[0.000]**</td>
</tr>
<tr>
<td>0</td>
<td>0</td>
<td>0</td>
<td></td>
<td>0</td>
<td>28.939</td>
<td>[0.000]**</td>
</tr>
<tr>
<td>Long-run exclusion tests</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>0</td>
<td>.</td>
<td>.</td>
<td>.</td>
<td>.</td>
<td>24.637</td>
<td>[0.000]**</td>
</tr>
<tr>
<td>.</td>
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<td>.</td>
<td>.</td>
<td>.</td>
<td>27.991</td>
<td>[0.000]**</td>
</tr>
<tr>
<td>.</td>
<td>.</td>
<td>0</td>
<td>.</td>
<td>.</td>
<td>10.473</td>
<td>[0.001]**</td>
</tr>
<tr>
<td>.</td>
<td>.</td>
<td>.</td>
<td>0</td>
<td>.</td>
<td>5.1927</td>
<td>[0.023]**</td>
</tr>
<tr>
<td>.</td>
<td>.</td>
<td>.</td>
<td></td>
<td>0</td>
<td>14.305</td>
<td>[0.000]**</td>
</tr>
<tr>
<td>Long-run weak exogeneity tests</td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>0</td>
<td>.</td>
<td>.</td>
<td>.</td>
<td></td>
<td>23.728</td>
<td>[0.000]**</td>
</tr>
<tr>
<td>.</td>
<td>0</td>
<td>.</td>
<td>.</td>
<td></td>
<td>0.393</td>
<td>[0.531]</td>
</tr>
<tr>
<td>.</td>
<td>.</td>
<td>0</td>
<td>.</td>
<td></td>
<td>5.380</td>
<td>[0.020]**</td>
</tr>
<tr>
<td>.</td>
<td>.</td>
<td>.</td>
<td>0</td>
<td></td>
<td>0.305</td>
<td>[0.581]</td>
</tr>
<tr>
<td>.</td>
<td>0</td>
<td>0</td>
<td>0</td>
<td></td>
<td>5.962</td>
<td>[0.114]</td>
</tr>
</tbody>
</table>

Notes: ‘0’ denotes zero restriction on the coefficient of the corresponding variable, ‘.’ denotes unrestricted coefficient. The degrees of freedom parameter $\nu$ of the $\chi^2$ distribution corresponds to the number of zero restrictions imposed. The symbols * and ** denote significance at 0.05 and 0.01 levels, respectively.
long-run exclusion reject the null hypothesis that any of the variables \((m-p)_t, y_t\) and \(i_t\) can be individually excluded from the co-integrating vector at the 1% significance level and the inflation variable \(\pi_t\) – at the 10% significance level.

According to the univariate long-run weak exogeneity test results, we cannot reject the null hypothesis that the variables \(y_t\) and \(\pi_t\) are weakly exogenous with respect to long-run parameters at usual significance levels. However, the null hypothesis in question is marginally rejected for the interest rate variable at the 5% level. Nevertheless, according to the joint test for 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 changes in sample size, values of recursive test statistics of the joint null hypothesis, scaled by the 5% critical value, are reported in Figure 2. Observe that the restriction that these three variables \(y_t, i_t\) and \(\pi_t\) are weakly exogenous with respect to the long-run parameters is accepted for all sample sizes. Hence, in the further analysis these three variables will be treated as weakly exogenous with respect to long-run parameters.

![Figure 2: VAR model (2): Recursive LR test statistics for joint long-run weak exogeneity of \(y_t, i_t\) and \(\pi_t\) scaled by the 5% critical values](image)

### 5.2 Long-run money demand function

Imposing long-run weak exogeneity restrictions on the variables \(y_t, i_t\) and \(\pi_t\) results in the following cointegrating vector with the corresponding standard errors reported in parentheses below the coefficient estimates

\[
(m - p)_t = 2.851 y_t - 0.087 i_t - 0.426\pi_t + 0.012 \text{ trend} + \text{constant} + ec_t \tag{3}
\]

Observe that all coefficient estimates have the expected signs and all estimates are significantly different from zero. The estimated co-integrating relation is depicted in Figure 3. Recursively calculated coefficients of the co-integrating relationship are displayed in Figure 4 and these are remarkably stable.
The point estimate of income elasticity is 2.851, which is significantly larger than unit income elasticity. This implies that money demand increases proportionally more than real GDP. This, at the first glance puzzling, outcome does not appear to contradict results reported in other studies on transitional economies. It reconciles with the conclusion reached in Dreger et al. (2007), where the panel co-integration method was used in order to estimate the parameters of the money demand function for the ten new EU Member States. They find that income elasticity of money demand is significantly larger than unity, in fact around 1.7. Furthermore, the result reported here is also consistent with those reported in an earlier study.

Figure 3: VAR model (2): Cointegrating relation

Figure 4: VAR model (2): Recursively estimated coefficients of the cointegrating vector
on money demand in Latvia (Tillers, 2004) with income elasticity of about 2.2. Karla (1999) reports income elasticity of 3.75 using data for Albania. It is pointed out that such a high value of elasticity ‘is consistent with estimates for developing and transitional economies, where velocity shows a downward trend on account of many factors, including financial deepening’ (p. 93). Additionally, rather high income elasticity is consistent with the fact that a rather broad measure of the monetary base is used in this paper. As stated in Laidler (1993, p. 169), ‘Broader definitions of money … produce higher estimates of the income or wealth elasticity of the demand for money than narrower ones’. Furthermore, a high value of income elasticity could also be explained by the fact that as wages, which measure the opportunity cost of time, increase with growth in income, economic agents tend to increase money holding for a planned volume of transactions (Fried, 1973). Moreover, as pointed out in Fase and Winder (1999), estimates of income elasticity above one may suggest the importance of the wealth effect in the demand for money.

The point estimate of interest rate elasticity is $-0.087$, which is significantly different from zero. This contrasts with the finding of Tillers (2004) of insignificant interest rate effect on money balances. The difference between the results reported in equation (3) and those of Tillers (2004) could be explained by several factors. First, the sample size used in the present study is longer than that employed in Tillers (2004). Second, in contrast to Tillers (2004), logarithmic transformation of interest rate is used here in order to take into account the fact that the variation as well as the level of interest rate was much higher in the first half of the sample than in the second half. Third, inclusion of inflation as an additional explanatory variable may also have contributed to the result. Lastly, the estimate of semi-elasticity of money demand with respect to inflation is $-0.426$, which is also significantly different from zero. This indicates that money demand tends to decrease proportionally less than inflation.

At this point, it is instructive to compare our elasticity estimates with those obtained by 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 income elasticity estimates lie around unity, but they show a large dispersion. Moreover, they report that Eurozone countries’ income elasticity is about 1.28 or 1.42, depending on the way the results are summarised. As seen, the reported point estimate of income elasticity is rather large in comparison with that found for the Eurozone.

Next, it appears that the reported point estimate of long-run interest rate elasticity is substantially smaller that the reported mean of estimated long-run elasticities from the 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) do not summarise the semi-elasticity of money demand with respect to inflation, one can compare the reported estimate with that reported in Dreger and Wolters (2006), which is about $-4.5$. The inflation semi-elasticity estimate obtained for Latvia is also significantly lower.

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4 The former figure was obtained by weighted-averaging all broad money income elasticity estimates for individual Eurozone countries, whereas the latter figure was obtained by taking the average of income elasticity estimates reported in studies that estimated joint money demand for several European countries, i.e., data aggregation was done before estimation.
To summarise, estimates of the parameters of long-run money demand function for Latvia are significantly different from those reported for Eurozone countries. The estimate of income elasticity is significantly larger, whereas estimates of interest rate elasticity and inflation semi-elasticity are significantly smaller. Thence, we can conclude that the long-run money demand function for a transitional economy may differ from that of members of the Eurozone. To this end, the estimation results obtained for Latvia support a similar conclusion reached in Dreger et al. (2007), where the long-run money demand function was estimated for the new EU member states using panel co-integration techniques.

5.3 Short-run error correction model

As shown in Johansen (1992), the status of long-run weak exogeneity of some variables allows us to reformulate model (2) in terms of a conditional model, where one conditions on the current and past values of weakly exogenous variables and on the first lag of the error correction term. After removing the variables that have turned out to be insignificant, the estimated conditional model for real money balances \((m–p)_t\) looks as follows:

\[
\Delta(m–p)_t = 6.534 + 1.262\Delta y_t – 0.480\Delta y_{t-1} – 0.142\Delta \pi_{t-1} – 0.012 \text{ec}_{t-1} + 0.074D021_t
\]

\(\sigma = 0.018, \ R^2 = 0.819, \ T = 44, \ F_{AR(1-4)}(4, 34) = 1.006[0.418], \ F_{ARCH(1-1)}(1, 36) = 0.014[0.906], \)

\(= \chi^2(2) = 0.939[0.625], \ F_{Hetero}(9, 28) = 0.756[0.656], \ F_{Hetero-X}(15, 22) = 0.772[0.692], \)

\(F_{RESET}(1, 37) = 1.267[0.268], \)

with corresponding standard errors reported in parentheses below the coefficient estimates. The variable \(D021_t\) denotes the impulse dummy that takes the value of one in 2002:1 and zero otherwise. Conditional model (4) is parsimonious but at the same time the diagnostic tests show no signs of misspecification.

Figure 5: ECM model (4): Actual and fitted values (upper panel); Residuals (lower panel)
Observe that the error-correction term is highly significant and has the expected sign. Signs of short-run dynamics coefficients indicate that money balances tend to react positively to upward changes in economic activity and negatively to surges in inflation.

The conditional model has good explanatory power, as can be assessed by looking at the actual values and fitted values as well as the regression residuals (displayed in Figure 5). Coefficient estimates are well determined and exhibit remarkable stability according to one-step residuals, recursive Chow stability tests (see Figure 6), the CUSUM test and the CUSUM test of squares (see Figures 7 and 8, respectively). Recursively estimated coefficients are displayed in Figure 9. Finally, the conditional model is able to accurately forecast demand for real balances over the period 2003(1)-2006(4) (see Figure 10 for one-step ahead forecasts). This fact is supported by the Chow parameter constancy forecast F-test statistic, which takes the value $F(16, 22) = 0.440 [p = 0.951]$.

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
6 Testing exogeneity

In the previous section the conditional model for real money balances was estimated and it was found that variables such as income, long-run interest rate, and inflation are weakly exogenous with respect to long-run parameters. As the next step one can 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 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 conditioning variables.

6.1 Parameter constancy of marginal models

The testing procedure advocated in Hendry (1988) compares the outcomes of parameter constancy tests of the conditional and marginal models. If the parameters of the conditional model are constant, as shown above, but the marginal models have non-constant parameters, then the parameters of the conditional model cannot depend on those of marginal models. The following univariate marginal models were estimated using fourth-order autoregressive processes (and seasonal dummies when necessary):

\[
\Delta_i = -0.226 \Delta i_{t-3} + 0.191 \Delta i_{t-4} - 0.005 - 0.400 D972_t, \quad (5)
\]

\[
\begin{align*}
\sigma &= 0.094, \quad R^2 = 0.420, \quad T = 43, \quad F_{\text{AR}(1-4)}(4, 35) = 0.875[0.489], \quad F_{\text{ARCH}(1-1)}(1, 37) = 0.009[0.924], \\
\chi^2_{\text{Norm}}(2) &= 7.912[0.020], \quad F_{\text{Hetero}}(5, 33) = 0.206[0.958], \quad F_{\text{Hetero-}\lambda}(6, 32) = 0.167[0.984], \\
F_{\text{RESET}}(1, 38) &= 0.014[0.905],
\end{align*}
\]

\[
\Delta \pi_i = -0.003 - 0.024 Q_t - 0.058 Q_{t-1} - 0.089 Q_{t-2} + 0.109 D961_t, \quad (6)
\]

\[
\begin{align*}
\sigma &= 0.025, \quad R^2 = 0.767, \quad T = 44, \quad F_{\text{AR}(1-4)}(4, 35) = 0.743[0.569], \quad F_{\text{ARCH}(1-1)}(1, 37) = 4.018[0.052], \\
\chi^2_{\text{Norm}}(2) &= 2.623[0.267], \quad F_{\text{Hetero}}(5, 33) = 1.673[0.169], \quad F_{\text{Hetero-}\lambda}(5, 33) = 1.673[0.169], \\
F_{\text{RESET}}(1, 38) &= 3.232[0.080],
\end{align*}
\]

\[
\Delta y_t = 0.336 \Delta y_{t-4} + 0.013 - 0.009 Q_t + 0.023 Q_{t-1} - 0.011 Q_{t-2}, \quad (7)
\]

\[
\begin{align*}
\sigma &= 0.013, \quad R^2 = 0.752, \quad T = 43, \quad F_{\text{AR}(1-4)}(4, 34) = 0.198[0.937], \quad F_{\text{ARCH}(1-1)}(1, 36) = 1.052[0.312], \\
\chi^2_{\text{Norm}} &= 2.462[0.292], \quad F_{\text{Hetero}}(5, 32) = 0.851[0.524], \quad F_{\text{Hetero-}\lambda}(8, 29) = 0.584[0.782], \\
F_{\text{RESET}}(1, 37) &= 0.545[0.465].
\end{align*}
\]

One-step-ahead residuals as well as Chow parameter constancy tests of equations (5)–(7) are displayed in Figures 11–13. Parameter constancy is rejected for the long-run interest rate as well as for the inflation rate. The marginal model for income seems to display constant parameters. The fact that marginal equations for conditioned variables such as long-run interest

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\(^5\) These interventions are modelled by means of impulse dummies \(D_{yyq}\), which take the value of 1 in the corresponding quarter \(q\) of 19yy or 20yy and zero otherwise.
rate and inflation rate are found unstable whereas the parameters of the conditional model for real money balances exhibit stability indicates that both the interest rate and the inflation rate are super exogenous in the money demand model (4). The interpretation of this fact is that structural breaks observed in conditioned variables do not affect stability of the model for real money balances that uses the interest rate and inflation as explanatory variables.

Figure 11: Marginal model for $\Delta i_t$ (5): Parameter constancy tests

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

In order to conduct a formal test, the approach of Engle and Hendry (1993) is chosen. Engle and Hendry (1993) suggest the following sequence of testing procedure. After having determined that marginal processes display parameter instability, one can test invariance by first augmenting marginal models with additional variables that capture changes in the respective parameters so that 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 test is that if the parameters of the conditional model are invariant to changes in marginal processes, then including these intervention variables should not improve the fit of the conditional model.

It has been shown above that marginal models for long-run interest rate and inflation rate 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:

\[
\Delta i_t = -0.276 \Delta i_{t-3} + 0.207 \Delta i_{t-4} - 0.021 - 0.381 D972 + 0.336 D041 + 0.156(D054 + D062) \tag{8}
\]

\[
\sigma = 0.072, R^2 = 0.671, T = 43, F_{AR(1-4)}(4, 33) = 1.700[0.174], F_{ARCH(1-1)}(1, 35) = 0.671[0.418],
\]

\[
\chi^2_{Norm} = 0.232[0.890], F_{Hetero}(7, 29) = 0.654[0.708], F_{Hetero-X}(9, 27) = 0.505[0.857],
\]

\[
F_{RESET}(1, 36) = 0.103[0.747]
\]

and
\[
\Delta \pi_t = -0.006 - 0.017Q_{it-1} - 0.081Q_{it-2} - 0.089\Delta D961_t - 0.050\Delta D061_t + 0.041(D024_t - D044_t) + 0.059(D012_t + D042_t + D052_t)
\]

\[
\sigma = 0.017, \quad R^2 = 0.895, \quad T = 44, \quad F_{ARB(1-4)}(4, 32) = 0.588[0.673], \quad F_{ARCH(1-1)}(1, 34) = 0.667[0.420], \quad \chi^2_{\text{Norm}} = 2.608[0.272], \quad F_{\text{Hetero}}(10, 25) = 0.817[0.616], \quad F_{\text{RESET}}(1, 35) = 3.528[0.069].
\]

The corresponding outcomes of parameter stability tests are displayed in Figures 14 and 15. Indeed, the sets of dummy variables \((D041_t, (D054_t + D062_t))\) and \(((D024_t - D044_t), (D012_t + D042_t + D052_t), D061_t)\) for equations (8) and (9), respectively, are important for achieving stability. Therefore they can be seen as determinants of non-constancies in the inflation and interest rate processes. Testing the significance of these sets of dummies in conditional model (4) yields the 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 money demand function cannot be rejected.

![Figure 14: Marginal model for \(\Delta i_t\) (8): Parameter constancy tests](image1)

Summarising, the results of tests conducted in this section suggest that long-run interest rate and inflation rate could be considered super exogenous for parameters of the conditional money demand model. It was shown that marginal processes for these variables display instability whereas the money demand function which uses these variables as explanatory variables exhibits constant parameters. Furthermore, the parameters of conditional models were found to be invariant to changes in marginal processes, as the determinants of their instability once added to the conditional model turned out to be insignificant. This means that policy analysis could be performed by suitably modifying the processes driving these variables.
6.3 Additional tests for long-run weak exogeneity

At this point, one 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 marginal models. This yields the following test statistics: $t_{\delta_i} = -1.38$, $t_{\delta \pi} = -1.28$ and $t_{\delta y} = 0.645$ for long-run interest rate, inflation, and income, respectively. As seen, none of these is significant at the usual levels. In addition, one can perform a Wu-Hausman test for weak exogeneity (suggested in Engle and Hendry (1993)), which is tantamount to testing for independence between conditioning variables and residuals from marginal models. This test is carried out by adding corresponding residuals from equations (7), (8) and (9) to conditional model (4) and testing their significance. 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 long-run interest rate, inflation, and income are consistent with those obtained using the Johansen procedure.

7 Conclusion

This study suggests a parsimonious error correction model of money demand in Latvia. This is based on a single co-integrating vector between real money balances, gross domestic product, long-term interest rate, and inflation rate. The model exhibits remarkable coefficient stability over the period covering the years between 1996 and 2006, when most important economic restructuring of the Latvian economy took place, and which is characterised by both internal (banking crisis and its consequences) and external (the Russian financial crisis) economic shocks as well as a highly varying rate of inflation, which was above 20% per annum at the beginning of 1996, then lowered to a meagre 1.8% in 2000 and then further increased to about 6-7% at the end of the estimation sample period.
The results of exogeneity tests suggest that the three explanatory variables such as income, long-run interest rate, and inflation could be considered as weakly exogenous with respect to long-run parameters. Furthermore, long-run interest rate and inflation rate were found to be super exogenous with respect to the parameters of the money demand function. This implies that suitable policy analysis can be conducted by varying the processes driving these two variables.

With respect to the factors determining money demand for Latvia, the findings are as follows: On the one hand, long-run income elasticity of money demand is larger than unity and also tends to be larger than that typically reported for the countries of the Eurozone. Long-run interest rate elasticity and 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. (2007), where long-run money demand functions for the ten new EU Member States were estimated using panel data co-integration 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 Eurozone.

One of the limitations of the present study is that the effects of the credit boom witnessed in Latvia in the last decade were not directly incorporated into the formal model. This is largely due to data considerations as the relevant variable measuring the credit supply is only available since 2003, which is rather short for modelling purposes within the co-integration framework adopted in this paper. Nevertheless, assessing the effects of the credit boom on money balances and, possibly, on other economic variables such as consumption, is an interesting and highly relevant extension of the current paper that is worthwhile pursuing in future research.

References


