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On the empirical relevance of the Lucas critique: The case of euro area money demand

Christian Dreger and Jürgen Wolters[1]

Abstract: This paper examines the relevance of the Lucas critique for euro area money demand. Based on the money in the utility function approach, a vector error correction model is specified to investigate the relationship between money and inflation in times of policy shifts. A well defined equation for money demand is obtained. The results indicate that the evolution of M3 is still in line with money demand. In the long run, inflation is affected by asset prices and detrended output. Our results show that the Lucas critique can be refuted in case of euro area money demand for the period of quantitative easing. Thus, the estimated money demand equation provides reliable information for the conduct of future monetary policy.

Keywords: Lucas critique, money demand, superexogeneity

JEL Classification: C22, C52, E41

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1 Introduction

According to Lucas (1976) private agents base their decisions in part upon their forecasts of future actions of policymakers. Changes in policy rules have an immediate impact on the decisions of agents. Any policy evaluation based on backward looking macroeconomic models is misleading whenever such shifts occur. The Lucas critique refers to the fact that model parameters might not be constant under different policy regimes. Instead, they are expected to vary, as rational agents are forward looking and adapt to the expected effects of the new regime. This would generally alters the model structure: Since the model estimates are based on historical time series, they are suitable for the evaluation of historical policy measures but they may not be appropriate to forecast the future policy impact. Previous evidence suggests that the Lucas critique is particularly relevant for monetary policy and financial markets (Clarida, Gali, and Gertler, 2000). On the other hand, many studies have found that the quantitative impact is only modest (Estrella and Fuhrer, 2003, Rudebusch, 2005). Forward looking models appear to be less stable than their backward looking counterparts. One interpretation of this finding is that policy changes are not large enough to affect the behaviour of agents in a critical way (Leeper and Zha, 2003). Whether the Lucas critique is a crucial assumption is an empirical question. If the Lucas critique is not relevant, backward looking models might be still valuable for policy advice.

The contribution of this paper is an assessment of the Lucas critique for euro area money demand. The recent period is predestinated for such an analysis. In the financial crisis, interbank markets dried up and hampered the redistribution of liquidity between banks. As short-term interest rates reached lower bounds, monetary authorities turned to unconventional policy measures to provide further stimulus. This fundamental policy

shift might have affected the relationship between money and its fundamental determinants.

For the role of money demand in a world of interest rate based operating procedures see especially Duca and VanHoose (2004). In the recent period with interest rates fixed near zero, they are no longer a powerful instrument for monetary policy. In fact, if interest rates are at the zero bound, the opportunity costs of holding money are zero and agents should be indifferent between holding money or bonds. As a consequence, money demand might break down, see e.g. Reis (2013). However, the existence of a well defined money demand function cannot be ruled out in advance, if the opportunity costs are captured by the interest rate spread and/or inflation rates. To analyze the role of money in the transmission process from the monetary to the real sector the existence and stability of money demand is a necessary condition.

Especially at longer horizons, inflation is inherently a monetary phenomenon (Benati, 2009). Since money defines the unit of account, monetary developments are integral to the determination of prices and inflation. Excessive liquidity can also provide early signals for the emergence of speculative bubbles in asset prices with potential risks to inflation and the real economy, see Borio (2007) and Adalid and Detken (2007). Consequently, monetary developments play a key role in the two pillar strategy of the ECB. While one pillar is based on the economic analysis of price risks in the short term, the other one is built on the monetary analysis of risks to price stability in the medium and long run. The explicit reliance on money as a guide for monetary policy is a distinguishing feature of the ECB compared to other central banks, see Hall, Swamy and Tavlas (2012).

The concept of money demand is crucial for an appropriate interpretation of monetary conditions in an economy, see Fischer, Lenza, Pill and Reichlin (2008). The money demand function links the monetary development to its fundamental determinants, such as the price level, real income and the opportunity costs of holding money. By comparing the money stock with its long run equilibrium according to money demand, measures of excess liquidity are derived. Because of increasing financial uncertainty and portfolio shifts, money demand stability has been doubted even long before the financial crisis, see Greiber and Lemke (2005) and Carstensen (2006). Due to repeated surges of M3 growth beyond its reference value, financial markets have put less and less weight on the signals stemming from the monetary analysis. Likewise, monetary aggregates play no role in state-of-the-art macroeconomic models, see Smets and Wouters (2007) and Woodford (2008). This might be inefficient, as information on potential inflation risk is not properly addressed.

Dreger and Wolters (2010a) demonstrated that a long run money demand relation can be restored if inflation is allowed to enter as part of the opportunity costs. In addition, financial wealth is increasingly important for the development of real money balances (Greiber and Setzer, 2007, Beyer, 2009, Hall, Swamy and Tavlas, 2012). The wealth effect is often proxied by house prices, as this specification tends to be superior over alternatives like stock prices. Based on a panel analysis, Arnold and Roelands (2010) concluded that the inclusion of house prices is necessary to obtain an income elasticity of one, and that country specific developments in house prices can explain a significant part of the deviations from money demand in the period before the crisis. Beyer (2009) found a stable money demand function for M3 up to the crisis. Similar to Dreger and

Wolters (2010a, 2010b and 2014), the opportunity costs of money holdings include the inflation rate.

This paper explores whether the money demand relationship has remained stable over the recent financial crisis, despite the marked policy shift towards unconventional measures. Evidence is based on a cointegration analysis, where inflation and real house prices are allowed to enter the long run money demand relationship. It is shown that the cointegration rank is not unique. By restricting the cointegration space, money demand and inflation equations can be identified to explore the joint dynamics of money and inflation. The results indicate that both relationships are well defined and stable over time. In particular, the evolution of M3 is still in line with money demand. While the weak exogeneity hypothesis is rejected for real money balances and inflation, real income, house prices and the interest rate spread do not respond to long run deviations. Furthermore the short run dynamics are driven by excess liquidity and excess inflation. Finally, it is shown that a conditional money demand equation provides reliable information for monetary policy. As the Lucas critique can be rejected, this equation is sufficient to monitor the monetary development.

The rest of the paper is organized as follows. Section 2 reviews the specification of the money demand function. In section 3 the time series used in the analysis and their stochastic properties are discussed. Evidence regarding the cointegration space, long run relations and weak exogeneity are provided in section 4. Section 5 presents the conditional error correction money demand equation and the test of the Lucas critique. Finally section 6 concludes.

2 Specification of money demand

The conventional specification of money demand postulates a long run relationship of the form

$$(1) \qquad (m-p)_t = \delta_0 + \delta_1 y_t + \delta_2 w_t + \delta_3 rl_t + \delta_4 rs_t + \delta_5 \pi_t$$

where m denotes nominal money balances taken in logs, p is the log of the price level, y is log of real income, representing the transaction volume in the economy, and w is log of real financial wealth (Goldfeld and Sichel, 1990). Opportunity costs of holding money are proxied by nominal long (rl) and short (rs) term interest rates and the annualized inflation rate, i.e. $\pi=4\Delta p$, in case of quarterly data. The index t denotes time.

Although equation (1) is an empirical specification, it does not lack a microeconomic foundation (Walsh, 2010). The theoretical rationale is provided by the money in the utility model. Representative households receive utility both from consumption and real money. The marginal utility of money is positive, since money allows to shift purchasing power between the periods. Utility is maximized subject to an intertemporal budget constraint, where the ressources of the household include income, interest bearing bonds and other financial assets. Equation (1) can be considered as a logarithmic approximation of the first order conditions.

The income variable exerts a positive effect on nominal and real money balances. In addition, money holdings are related to a portfolio allocation decision of private households. For example, a surge in asset prices may trigger a rise in demand for liquidity due to an increase in net household wealth. While the scale effect points to a positive impact of wealth, the substitution effect works in the opposite direction, as higher asset prices make assets more attractive relative to money holdings. If the opportunity costs of mon-

ey holdings refer to earnings on alternative financial assets, possibly relative to the own yield of money balances, their coefficients should enter with a negative sign. The inflation rate is part of the opportunity costs because it represents the costs of holding money in spite of holding real assets. Furthermore, its inclusion provides a convenient way to generalize the short run homogeneity restriction imposed between money and prices that might not hold on empirical grounds. In addition, adjustment processes in nominal or real terms can be distinguished (Hwang, 1985), compare also Dreger and Wolters (2010a).

The parameters δ_1>0and δ_2 denote the elasticities of money demand with respect to the scale variables, income and wealth. The impact of the return of other financial assets and inflation is captured by the semielasticities δ_3<0, δ_4 and δ_5, respectively. The parameter δ_4 should be positive when *rs* is mainly a proxy for the own rate of interest of holding money balances, but negative otherwise. Due to the ambuigity in the interpretation of the wealth and inflation variables, the signs of their impact cannot be specified a priori.

3 Data and preliminary analysis

Since the introduction of the euro in 1999, the ECB is responsible for the monetary policy in the euro area. As the time series under the new institutional framework are too short for robust conclusions, they have to be extented by artificial data. Euro area series prior to 1999 are obtained by aggregating national series (Artis and Beyer 2004). By comparing different aggregation methods, Bosker (2006) and Beyer and Juselius (2010) stressed that differences are substantial prior to 1983, especially for interest rates and

inflation, but they are almost negligible from 1983 onwards. In addition, the European Monetary System started working in 1983 and financial markets have become more integrated since then. Juselius (1998) reports evidence in favour of a change in the monetary transmission mechanism in March 1983 for some European countries. Therefore, 1983Q1-2011Q4 is chosen as the observation period. To cover initial values, the data already start in 1981Q1. Quarterly seasonally adjusted series are used.

Nominal money balances for M3 are from the ECB monthly bulletin database. The short and long term interest rates *rs* and *rl* are also from this source and are defined by the 3-month Euribor and ten-year government bond rates, respectively. Real GDP, as a proxy for income, is taken from Eurostat, defined as chain-linked volumes with 2000 as the reference year. The GDP deflator (2000=1) is also from Eurostat. GDP volume and deflator information prior to 1995 is based on Brand and Cassola (2004). To derive real money balances, nominal money stocks are deflated with the GDP deflator. Real financial wealth is approximated by nominal house prices deflated by the GDP deflator. Nominal house prices are taken from the Bank of International Settlement and interpolated to the quarterly frequency. Figure 1 shows the series in levels (A) and first differences (B).

-Figure 1 and Table 1 about here-

According to the DF-GLS unit root test, all level variables appear to be integrated of order 1, I(1), implying that they are nonstationary in levels, but stationary in first differ-

ences, see Table 1[2]. For the interest rate spread, i.e. the difference between the long and short term nominal interest rate, the nonstationarity hypothesis is rejected at the 0.01 level.

Outliers are detected in the real money balances, see Figure 1B. The first one (1990Q2) is due to the German unification, while the other one (2001Q1) reflects the stock market turbulences in the aftermath of the new economy bubble, see Kontolemis (2002). In the subsequent analysis, these outliers are acknowledged by two impulse dummies. They are equal to 1 in the respective period and 0 otherwise (*d*902 and *d*011). Breaks are also relevant in the income elasticity, see Figure 2. In particular, the parameter has risen after the introduction of the euro to the public (2002Q1), see Dreger and Wolters (2010b). There has been also a sharp increase because of the financial crisis. Despite the fact that monetary developments have been largely favourable, massive production losses occurred.

-Figure 2 about here-

To measure the change in the income elasticity from 2002Q1 onwards, wealth is included. This approach leads to a stable relationship that might reflect the rising presence of wealth in money demand. A variable *w** is introduced, which is defined as the product of *w* and a step dummy *s*021 equal to 1 from 2002Q1 onwards and 0 before.

[2] The DF-GLS test is more efficient than the standard ADF test in estimating the deterministic terms. See Elliott, Rothenberg and Stock (1996). The lag lengths for the unit root and cointegration tests is determined by the Akaike criterion. For unit root and cointegration testing see Kirchgässner, Wolters and Hassler (2013).

4 Cointegration analysis in the system of variables

The empirical analysis is based on a vector autoregressive (VAR) model in levels which is equivalently rewritten as a vector error correction model

$$(2) \qquad \Delta \boldsymbol{Y}_t = \boldsymbol{\alpha}\boldsymbol{\beta}' \boldsymbol{Y}_{t-1} + \sum_{j=1}^{p-1} \boldsymbol{\Gamma}_j \Delta \boldsymbol{Y}_{t-j} + \boldsymbol{D}_t + \boldsymbol{\varepsilon}_t$$

where $\boldsymbol{\alpha}$ denotes the matrix of feedback parameters, $\boldsymbol{\beta}$ the matrix of the cointegrating relations and $\boldsymbol{D}_t$ the deterministic terms. The cointegration properties of different sets of variables are explored using the Johansen (1995) trace test. The lag length p of the VAR in levels is determined by the Akaike criterion and set equal to two throughout the analysis. The specification investigated by Dreger and Wolters (2010a) serves as a starting point, implying $\boldsymbol{Y}=(m\text{-}p, y, \pi)'$. It comprises real money balances, real income and the inflation rate. An unrestricted constant and a linear time trend restricted to the cointegrating space are embedded and might allow for possible changes in the long run relations over the sample period.

-Table 2 about here-

The three variables are cointegrated, where the cointegrating rank is equal to one, see Table 2. As the sample period is not very large both the standard trace test (Johansen 1995) and the small sample Bartlett corrected trace statistics (Johansen, 2000) are presented. Although the cointegration property can be established even in small systems of variables, the long run parameters are not very convincing, according to economic theo-

ry. Because of the permanent break in the income elasticity from 2002Q1 onwards, they are quite unstable.

Therefore, as discussed above the system is enlarged by the w^* variable to account for wealth effects in money demand. In the extended specification, $\mathbf{Y}=(m-p,\ y,\ w^*,\ \pi)$, the cointegration rank increased by one, i.e. it is equal to two. To estimate the long run parameters with higher precision, the model can be further enlarged by the interest rate spread, $\mathbf{Y}=(m-p,\ y,\ w^*,\ \pi,\ rl-rs)$. The interest rate spread can be interpreted as the net opportunity costs of holding money; i.e. the interest rate adjusted for the own rate of money holdings. Its inclusion leads to an increase in the cointegration rank using the standard trace test, but the result is no longer significant with the Bartlett corrected version (p=0.27). Such a further cointegrating relation would exist only because of the stationarity of the interest rate spread. Hence the following analysis is restricted to the more interesting first two long run vectors shown in the left part of Table 3. The normalization is on the coefficients of real money balances and inflation, respectively.

-Table 3 about here-

The first vector may be a money demand function, while the second one can be consistent with an inflation equation. For identification, zero restrictions are put upon the respective long run parameters. In particular, the trend coefficient is set equal to zero in the first relationship. In the second vector, real money balances and the interest rate spread are excluded. The corresponding likelihood ratio test is distributed as chi-squared with one degree of freedom. The test statistic is 1.062, with a *p*-value of 0.303.

Thus, the identifying restrictions cannot be rejected. The restricted cointegrating vectors are reported in the right part of Table 3[3].

The variables in the money demand function bear the theoretically expected signs, and the long run coefficients are plausible. For housing wealth, the scale effect dominates the substitution effect. The linear time trend in the inflation equation accounts for the secular decline of inflation, but it might be also interpreted as a rough indicator of potential output. In that case, inflation would depend on detrended output and real house prices. Thus, monetary policy should closely monitor the price evolution in asset markets, as it could have implications for the overall price development. The fact that real money balances is excluded from a restricted inflation equation does not imply that money is irrelevant for inflation. This can be inferred from the feedback coefficients, see Table 4.

-Tables 4 and 5 about here-

Both real money balances and the inflation rate respond significantly to deviations from the two cointegrating relationships, and the adjustment coefficients are all well signed. In contrast, real income, real house prices and the interest rate spread do not adjust to equilibrium errors and can be considered as weakly exogeneous. The reaction of real house prices is in line with the findings of Dreger and Wolters (2011a): money reacts to changes in house prices, but the reversed direction is not relevant for the euro area. In fact, a joint test for weak exogeneity of real income, real house prices and the interest

[3] Interestingly Hendry and Doornik (1994) have found similar long run relationships in a system for UK M1.

rate spread cannot be rejected. The corresponding likelihood ratio test is distributed as chi-squared with 7 degrees of freedom. The test statistic is 6.901, with a *p*-value of 0.439. To obtain more efficient estimates, the bivariate error correction model for money and inflation is evaluated in the following. The other variables (income, wealth and the interest rate spread) are treated as exogenous. The identified cointegrating vectors and adjustment parameters are displayed in Table 5. As expected, excess liquidity has a significant positive effect on inflation. According to our empirical evidence, a joint relationship between money and inflation can be established with a direct link from excess liquidity to inflation. Thus, models neglecting money balances in the inflation equation may be misspecified.

To assess the relevance of the Lucas critique for money demand in the long run, forward recursive tests are applied to determine whether the parameters in the restricted model can be assumed to be constant. The importance of parameter constancy in money demand equations is stressed by Ericsson (1992), see also Favero and Hendry (1992). Since the main issue is to test whether the relations are stable after the identified problems of financial uncertainty and portfolio shifts, the baseline sample runs from 1983Q1 to 2001Q4. Moreover, as the focus is on the stability of the long run parameters, the results for the so called *R*-form are preferred (Juselius, 2006). This is the concentrated model version, where the short run dynamics estimated from the full sample have been removed.

Figure 3 displays the eigenvalues of the system. The idea is to investigate whether the more recent observations are also generated by the baseline model. The test provides general information with regard to the constancy and non constancy of the cointegration space, because the eigenvalues can be expressed as quadratic functions of the feedback

and cointegration parameters (Juselius, 2006). If both are constant, the eigenvalues will share this property. In fact, they show only minor variation, and the fluctuations are not significant. Figure 4 displays the Nyblom test for constancy of the cointegrating space, as suggested by Hansen and Johansen (1999). The test statistic is divided by the 0.95 quantile of the distribution under the null hypothesis of a constant parameter regime. Two variants are considered, i.e. a test either based on the full model or the concentrated version. In both cases, the test statistic is well below the rejection line of 1.0. Therefore, the full sample estimates of the cointegration vectors are in the space spanned by the long run relationships in each subsample.

Figures 5 and 6 reveal the recursively estimated individual coefficients related to the cointegrating vectors and the feedback parameters. All parameters are highly significant and the size of the variation can be neglected. Overall, the results are robust, even in the financial crisis period. Hence, the joint dynamics of money and inflation during the operation of the ECB have been very stable and therefore, the system can serve as a starting point for monetary policy analysis. Furthermore, the parameters have remained constant despite of different policy regimes and estimation methods.

-Figures 3 to 6 about here-

5 Conditional error correction modelling

Since monetary policy has to be conducted in real time, a model based on a system of variables might not be the optimal strategy for policymakers. Given the identification problems in full systems with multiple cointegrating vectors, a conditional single equa-

tion analysis can produce similar insights, but is much easier to handle. Thus, conditional on the two cointegrating relations (*ec_liquidity* and *ec_inflation*, $\boldsymbol{\beta}_1$ and $\boldsymbol{\beta}_2$ from Table 5) and the contemporaneous change of real house prices, an error correction model for money demand is derived (Table 6). All other weakly exogeneous variables (ΔY, Δrl, Δrs) do not have significant impacts on the change of real money balances, neither in contemporaneous nor lagged specification. The coefficients are well signed. The residuals are well behaved, because they are normally distributed, homoscedastic and do not show autocorrelation. Tests of the functional form do not reveal any problems, and Chow forecast tests for the period of the financial crisis do not indicate instabilities, as well as the Chow breakpoint test for the beginning of the financial crisis in 2007Q4 ($F(6,103)$-value 0.33, p-value 0.92). The parameter constancy can be also inferred from the cusum of squares (Figure 7) and the recursive estimated coefficents available from the authors upon request.

-Table 6 and Figure 7 about here-

The Lucas critique that the parameters depend on the policy regime can be refuted if the parameters of the conditional model remain constant even if the parameters in the marginal process describing the evolution of the weakly exogeneous regressors might be subject to policy regime shifts (Favero and Hendry, 1992, Ericsson and Irons, 1995). In our case, this implies that house prices should be superexogenous to be able to reject the Lucas critique.

-Table 7 and Figure 8 about here-

Superexogeneity encompasses both weak exogeneity and structural invariance. It can be seen as a pre-condition for meaningful policy analysis (Fischer and Peytrignet, 1991, Belke, 2000). The testing procedure requires a constant parameter model for house prices[4]. After a specification search among the weakly exogeneous variables of the system the estimated equation presented in Table 7 is obtained. For capturing the period of quantitative easing a step dummy equal to 1 from 2007Q4 (Lehman crash) onwards and 0 elsewhere (*s*074) is introduced. This equation is well specified. The residuals do not show any kind of autocorrelation and heteroscedasticity. The parameters of the marginal process are constant, as indicated by the cusum of squares (Figure 8), and the recursively estimated coefficients which can be obtained from the authors upon request.

By extending the conditional model with the explanatory variables from the marginal model, the condition of superexogeneity is explored (Ericsson and Irons, 1995). Since the additional variables do not improve the conditional model, the null hypothesis of superexogeneity is not rejected ($F(4,105)$-value 0.445, p-value 0.776) for real house prices. As a result, the Lucas critique is not relevant for euro area money demand.

6 Conclusion

This paper examines the relevance of the Lucas critique for a euro area money demand equation. Based on a money in the utility function approach, a vector error correction model is specified to investigate the relation between money and its fundamental deter-

[4] The marginal model is estimated for the 2002Q1-2011Q4 sample, as house prices did not affect money demand in earlier periods.

minants in times of policy regime shifts. A well defined equation for money demand in the definition of M3 is obtained. While the weak exogeneity hypothesis is rejected for real money balances and inflation, the other variables do not respond to deviations from the long run. A single equation analysis still provides reliable information for the conduct of monetary policy. The instantaneous regressor fulfills the condition of superexogeneity. It is weakly exogeneous, and its marginal process is not relevant in the conditional model. Hence, policy regime shifts will not affect the conditional model structure. Overall the empirical findings suggest that the monetary policy shift towards a regime of quantitative easing did not generate instability in the relationship. Overall we find that the Lucas critique does not apply to euro area money demand. In addition to the standard set of variables, house price developments should be closely monitored by the ECB, since they are crucial to interpret the monetary evolution in an appropriate way.

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Figure 1A: Variables in levels

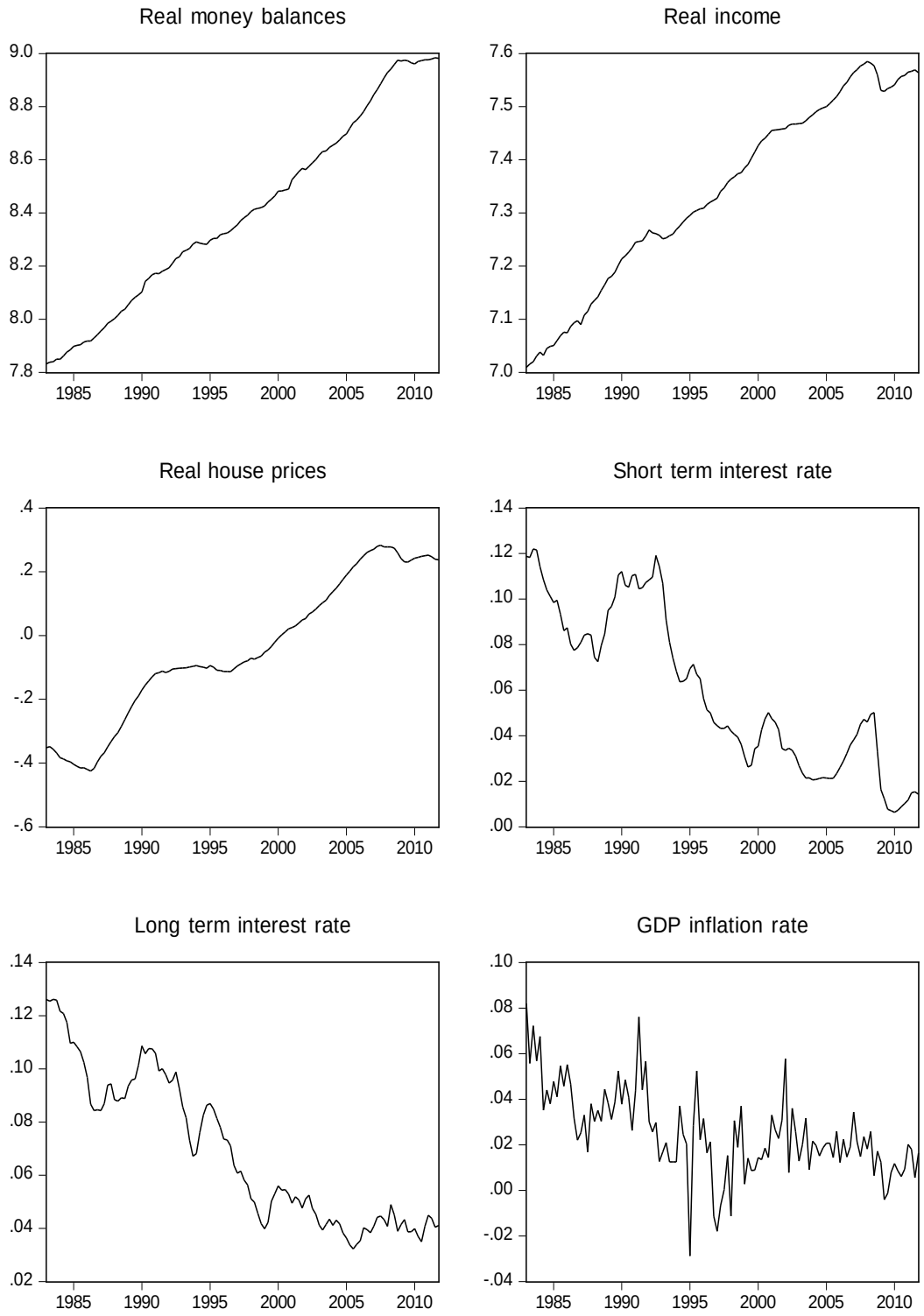
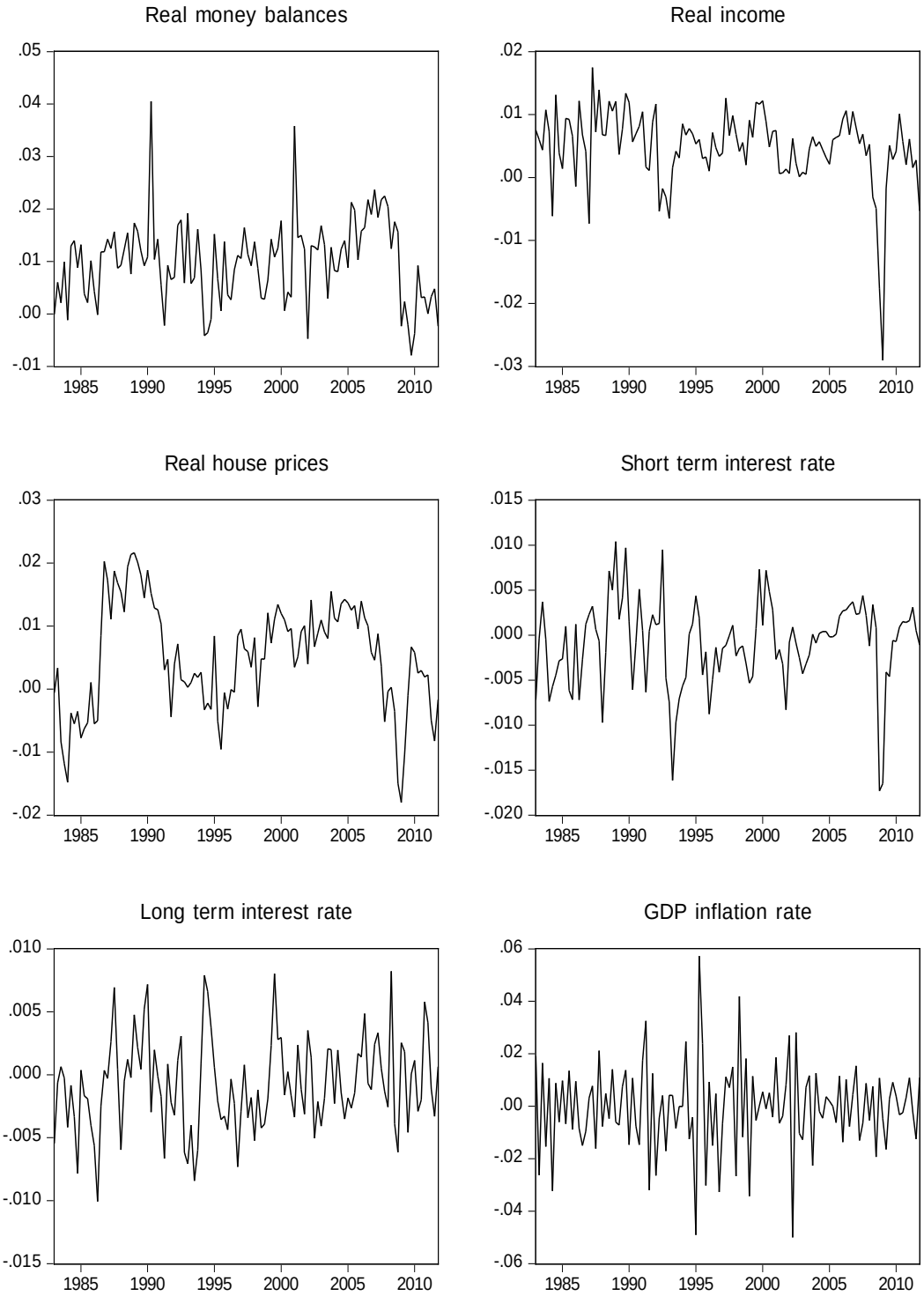
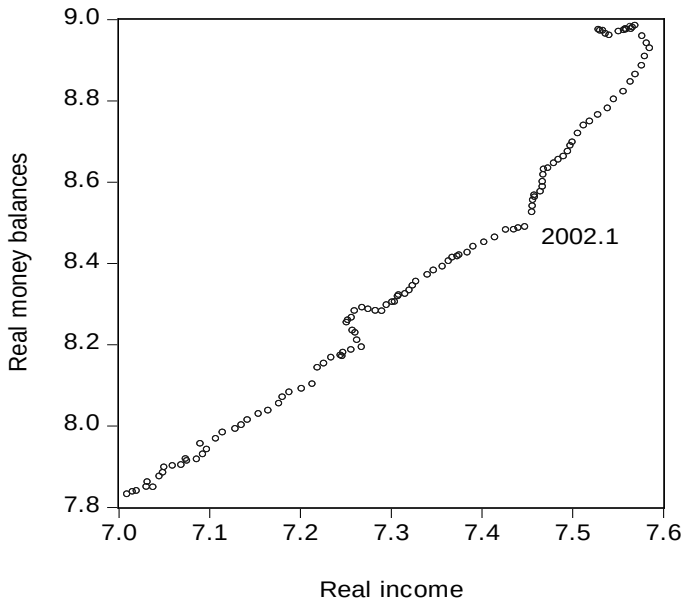


Figure 1B: Variables in first differences



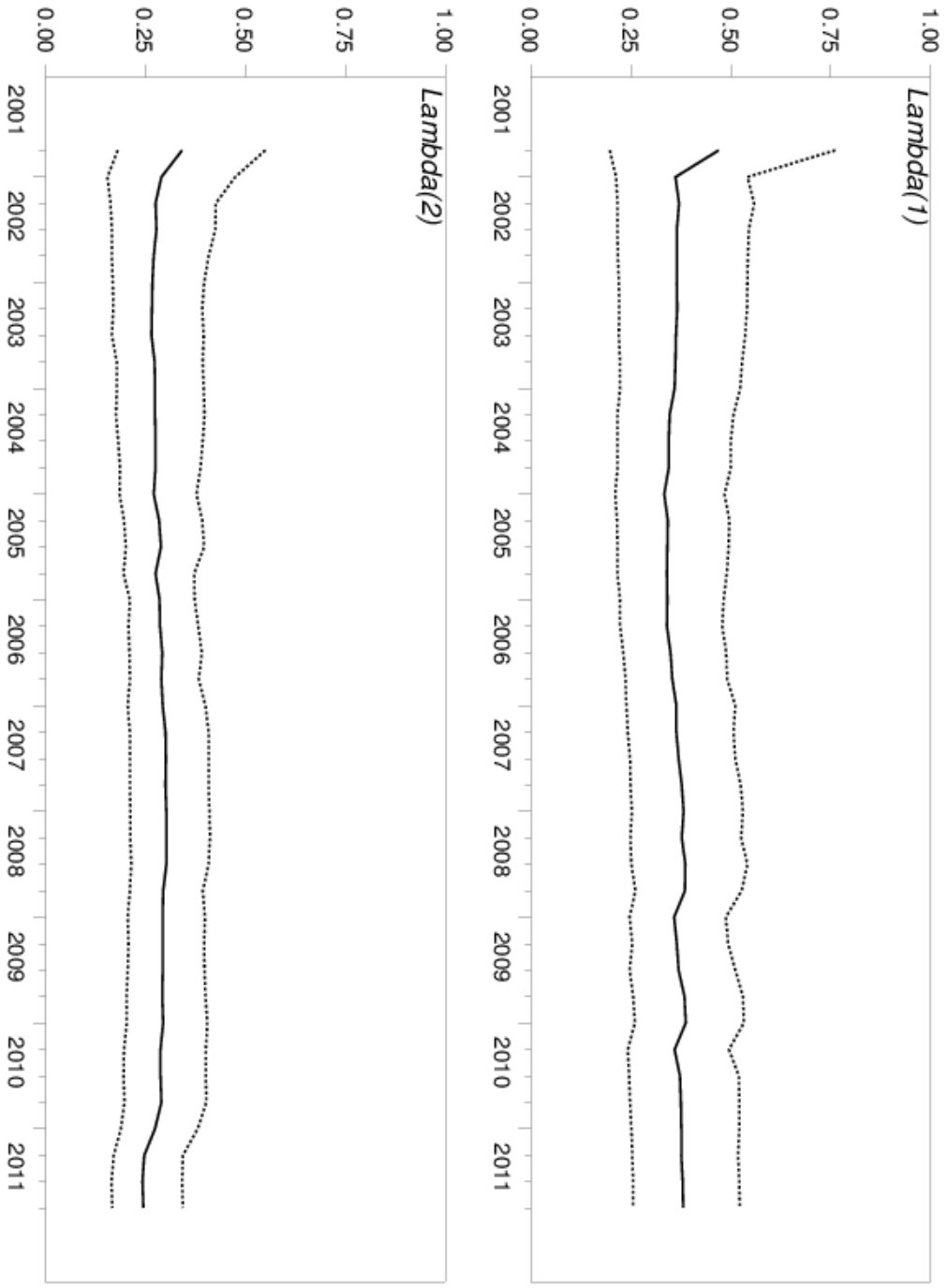
Note: Sample period 1983.1-2011.4. Real money, real GDP and real house prices in logs. Inflation q-o-q change in the GDP deflator (2000=1), annualized.

Figure 2 Structural break in income elasticity



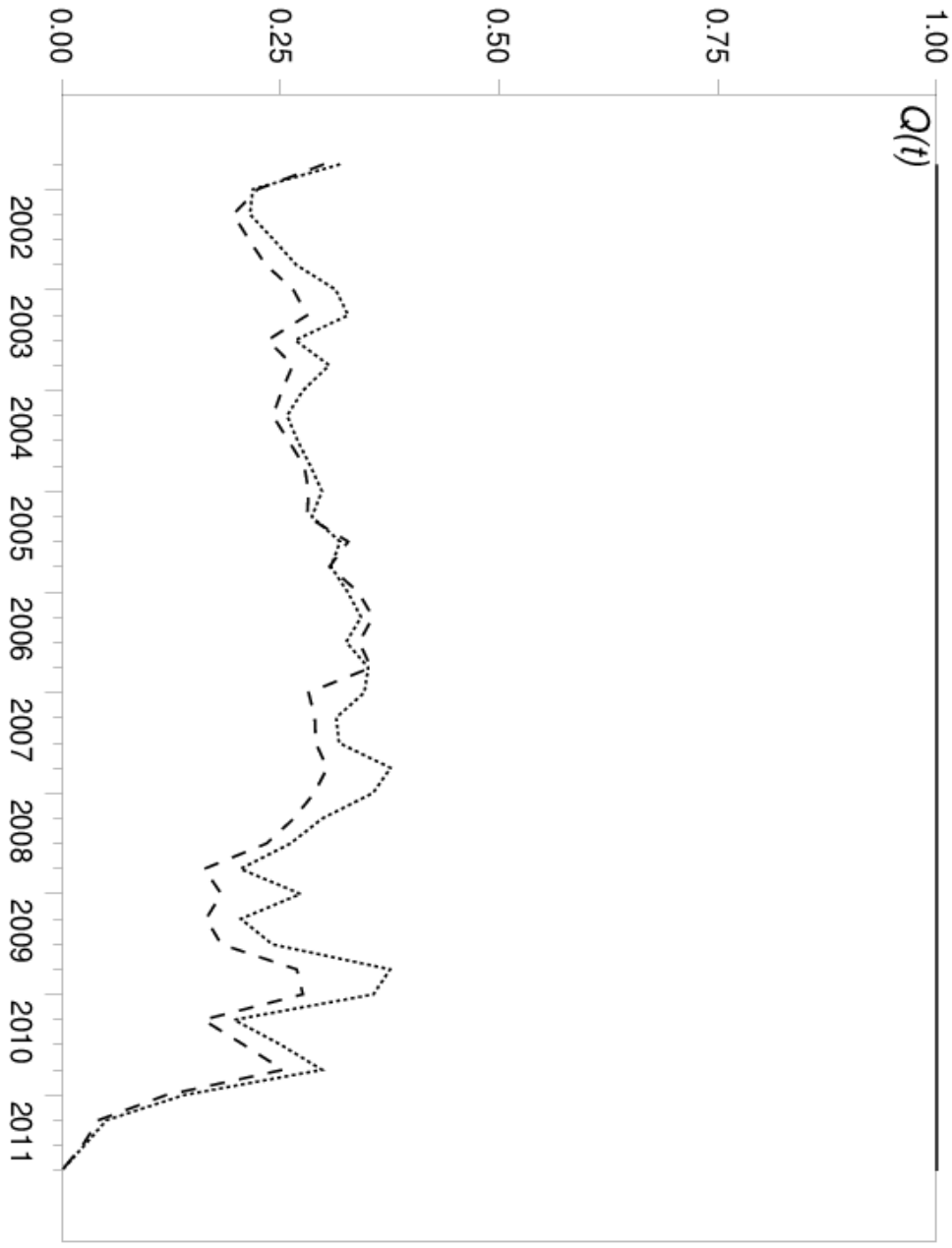
Note: Sample period 1983.1-2011.4.

Figure 3: Eigenvalues of the system



Note: Recursively calculated first two eigenvalues of the restricted cointegration system. Dotted lines are 0.95 confidence bands.

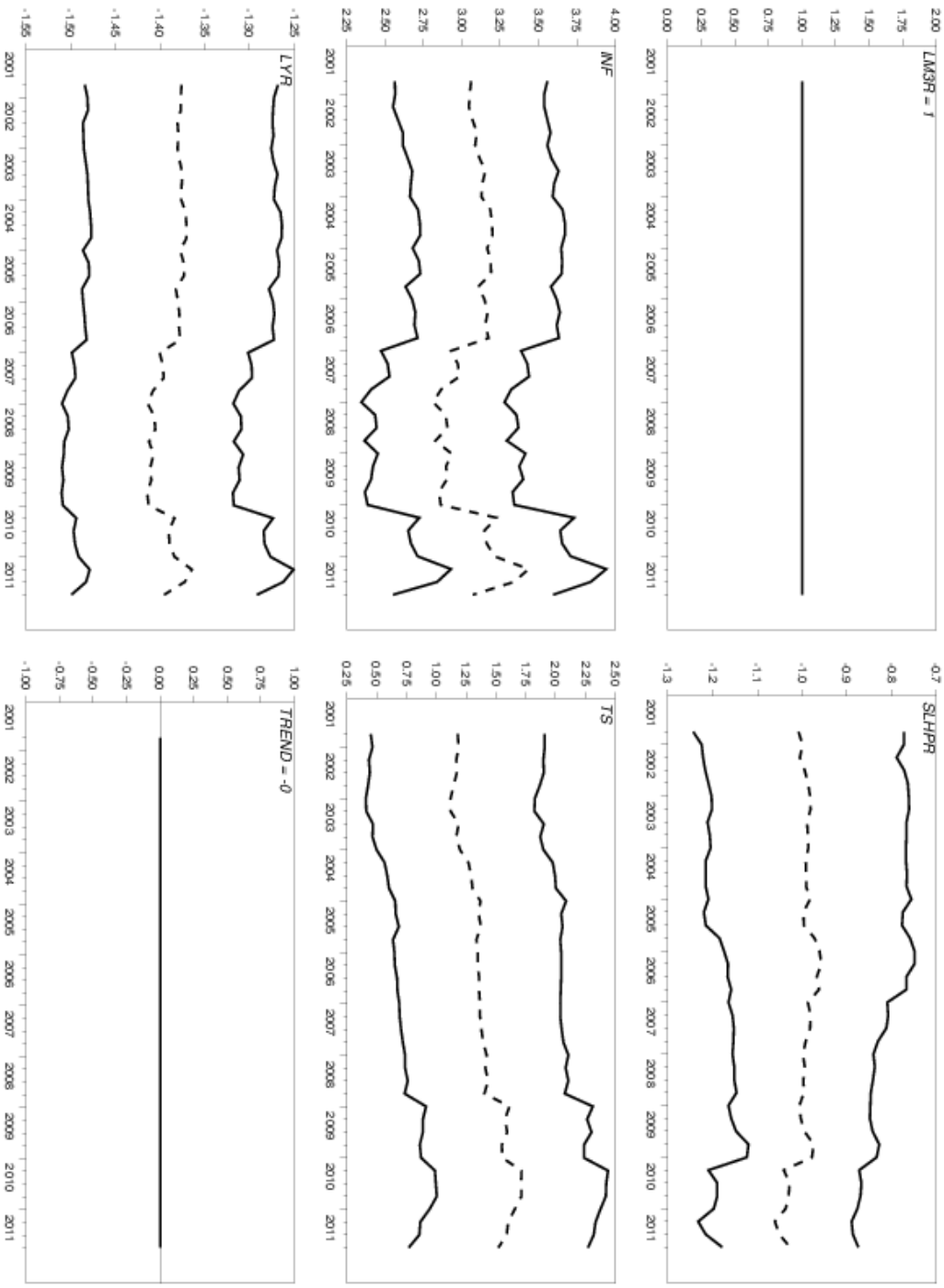
Figure 4: Nyblom test on beta constancy



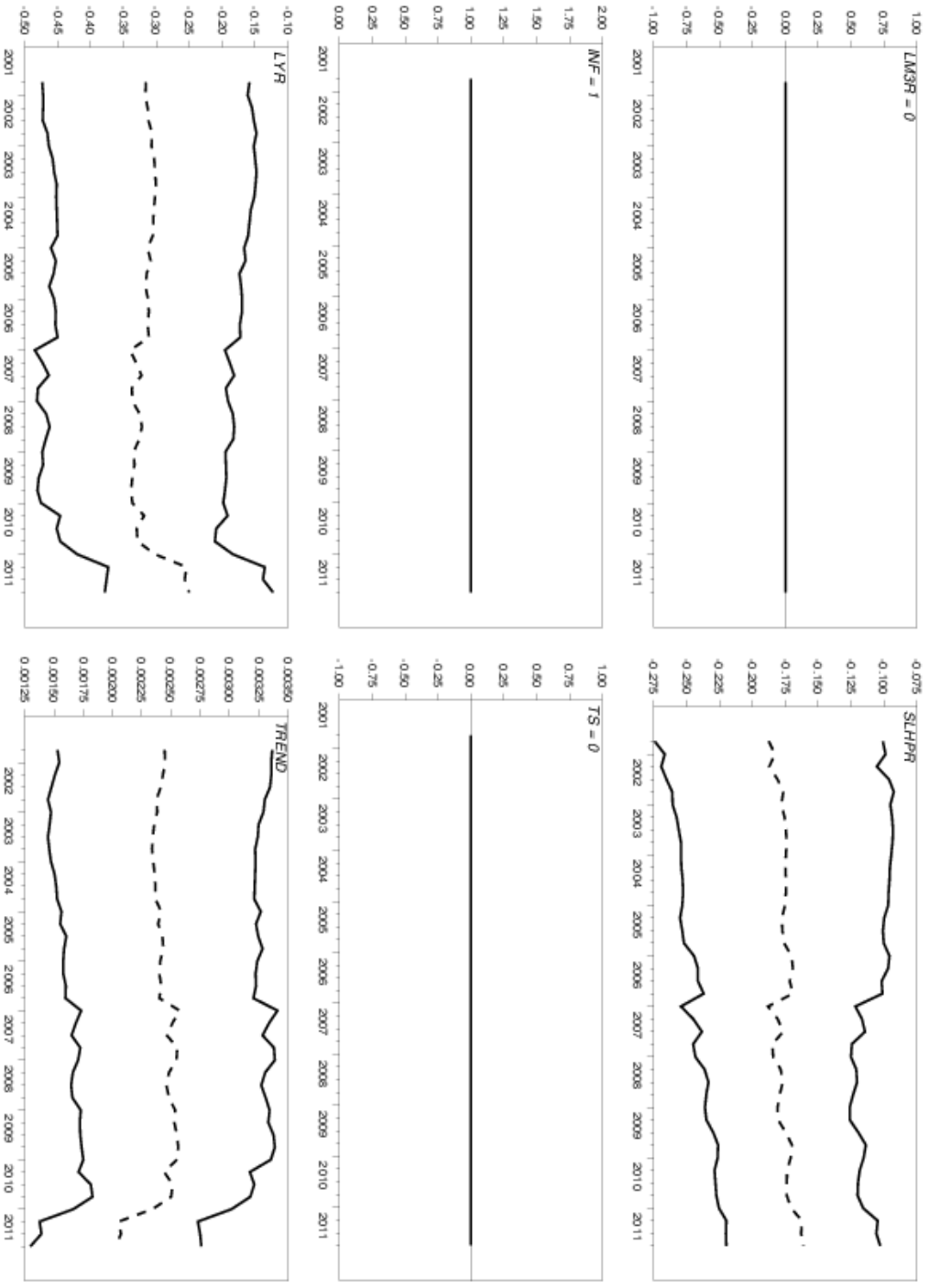
Note: Test statistics above the 1.0 line imply a rejection of the null hypothesis of constancy of the cointegration vectors. The dotted line shows the test statistic obtained for the full model, while the dashed line is calculated on the base of a concentrated model version, where the short run dynamics have been regressed out.

Figure 5: Constancy of individual cointegration parameters

A First cointegration vector (money demand equation)



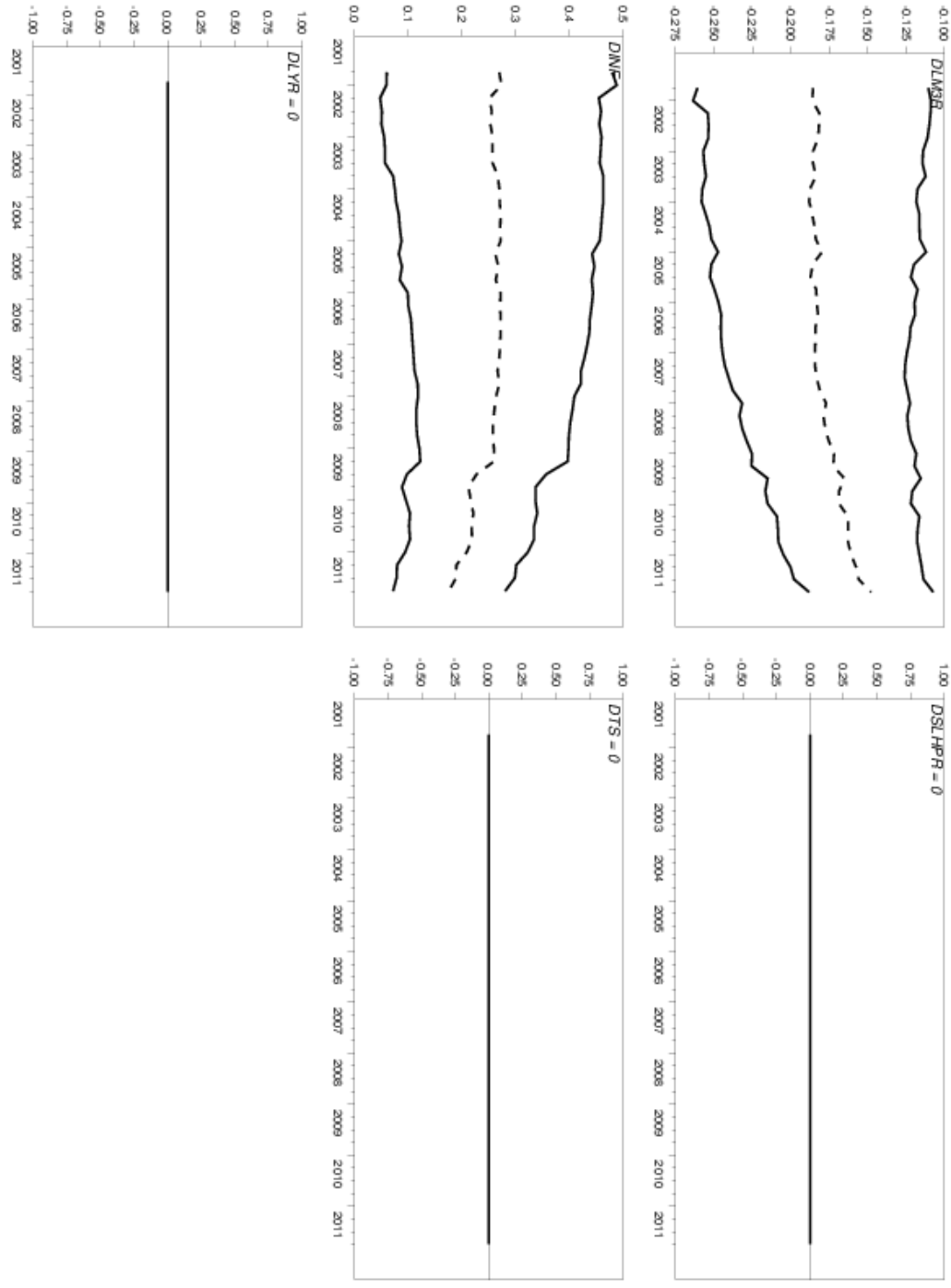
B Second cointegration vector (inflation equation)



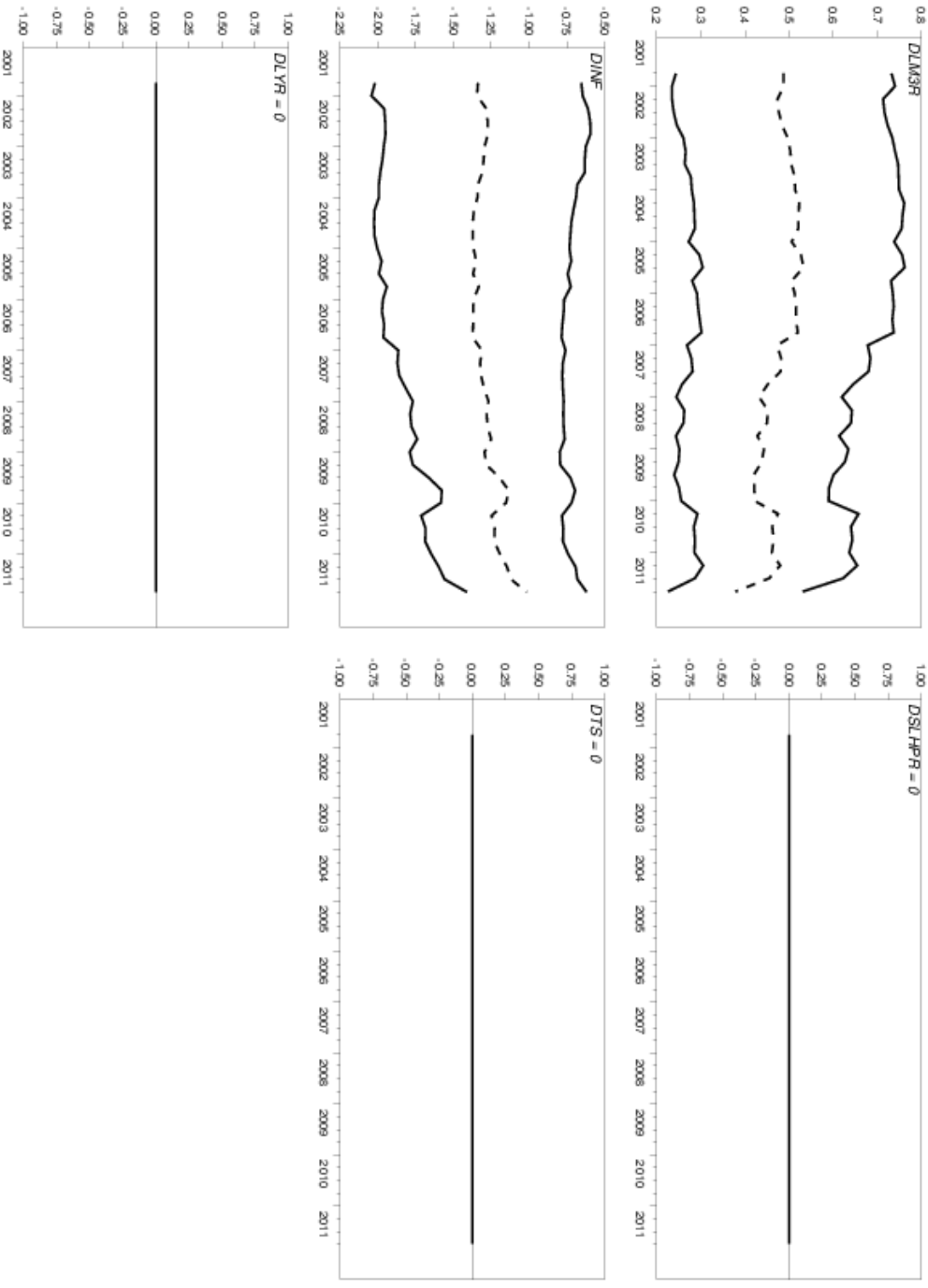
Note: Recursively estimated coefficients based on a concentrated model version. First vector (5A) normalized on real money balances, second vector (5B) on inflation. Solid lines represent 0.95 confidence bands.

Figure 6: Constancy of feedback parameters

A First feedback mechanism

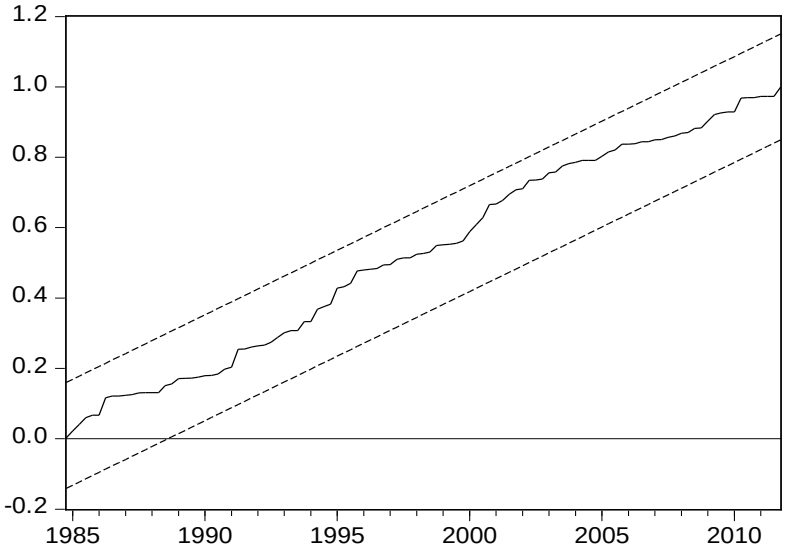


B Second feedback mechanism



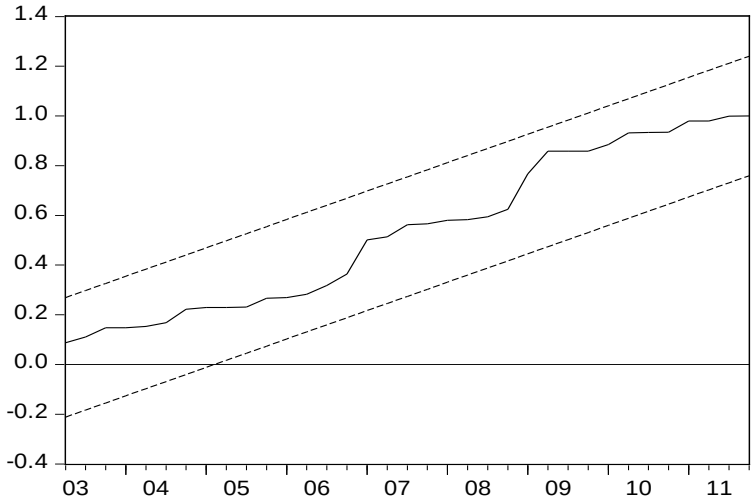
Note: Recursively estimated coefficients based on a concentrated model version. First vector (6A) shows feedback parameters to deviations from the long run money demand equation, while the second vector (6B) displays the responses to deviations from the long run inflation relationship. Solid lines represent 0.95 confidence bands.

Figure 7: Cusum of squares for conditional error correction model



Note: Dashed lines represent 0.05 significance levels.

Figure 8: Cusum of squares for the marginal house price model



Note: Dashed lines represent 0.05 significance levels.

Table 1 DF-GLS unit root test

Variables	Lag length	Test statistic
m-p	3 | 2	-2.24 | -3.98***
y	1 | 1	-0.92 | -3.03***
π	3 | 2	-0.34 | -9.67***
rl	6 | 5	0.48 | -4.04***
rs	1 | 0	-0.02 | -5.98***
rl-rs	1	-2.64***
*w**	1 | 0	-1.10 | -3.55***

Note: Sample period 1983Q1-2011Q4. For *w**, results refer to the 2002Q1-2011Q4 period. Left (right) entry corresponds to level (first difference) of the respective variable. The lag order is determined by the Akaike criterion. Real money balances, income and housing prices have a constant and linear time trend in the level specification and a constant in the first differences. The interest rates, the interest rate spread and inflation include a constant in their levels, while the first differences do not have any deterministic terms. Critical values are -3.57 (0.01), -3.02 (0.05), -2.73 (0.10) for the model with constant and linear time trend and -2.59 (0.01), -1.94 (0.05), -1.61 (0.10) for the model with a constant. A ***, **, * indicates rejection of the unit root hypothesis at the 0.01, 0.05 and 0.10 level, respectively.

Table 2 Cointegration rank of subsets of variables

Variables	Rank null hypothesis	Trace test	Trace test (Bartlett correction)	Rank decision
m-p, y, π	0	44.06 (0.036)	41.66 (0.065)	1
	1	22.18 (0.135)	20.87 (0.188)	
	2	7.12 (0.342)	5.83 (0.492)	
m-p, y, w, π*	0	102.29 (0.000)	93.27 (0.000)	2
	1	47.04 (0.017)	40.73 (0.080)	
	2	19.31 (0.268)	16.21 (0.484)	
	3	6.01 (0.470)	5.57 (0.526)	
m-p, y, w, π, rl-rs*	0	129.08 (0.000)	115.79 (0.000)	2
	1	73.48 (0.005)	61.97 (0.070)	
	2	40.83 (0.079)	34.49 (0.271)	
	3	18.29 (0.332)	15.70 (0.525)	
	4	5.41 (0.548)	4.91 (0.616)	

Note: Sample period 1983Q1-2011Q4, *p* values in parantheses. All models include the impulse dummies *d*902 and *d*011, an unrestricted constant and a linear time trend which is restricted to the cointegration relationship. The lag order of the VAR in levels is determined by the Akaike criterion and equal to 2 throughout the analysis.

Table 3 Long run vectors in the cointegration space

	Unrestricted model		Restricted model	
	$\boldsymbol{\beta}_1$	$\boldsymbol{\beta}_2$	$\boldsymbol{\beta}_1$	$\boldsymbol{\beta}_2$
m-p	1.000	-0.271	1.000	0.000
y	-1.158	-0.167	-1.372 (0.052)	-0.187 (0.066)
*w**	-0.808	-0.037	-1.029 (0.075)	-0.152 (0.029)
π	1.454	1.000	3.034 (0.274)	1.000
rl-rs	1.169	-0.965	1.370 (0.391)	0.000
trend	-0.002	0.004	0.000	0.002 (0.001)

Note: Sample period 1983Q1-2011Q4. Estimated cointegration vectors in the model including real money balances, real income, real house prices, the inflation rate, the interest rate spread and a linear time trend. Standard errors in parantheses.

Table 4: Feedback parameters in the restricted cointegration space

	$\boldsymbol{\alpha}_1$	$\boldsymbol{\alpha}_2$
$\Delta(m\text{-}p)$	-0.135 (0.019)	0.312 (0.070)
Δy	0.029 (0.021)	-0.152 (0.078)
Δw^*	-0.016 (0.023)	0.130 (0.085)
$\Delta\pi$	0.131 (0.051)	-0.803 (0.187)
$\Delta(rl\text{-}rs)$	0.008 (0.016)	-0.016 (0.058)

Note: Sample period 1983Q1-2011Q4. Estimated feedback vectors in the model including real money balances, real income, real house prices, the inflation rate, the interest rate spread and a linear time trend. Standard errors in parantheses.

Table 5: Restricted feedback parameters and cointegrating vectors (bivariate model)

	$\boldsymbol{\alpha}_1$	$\boldsymbol{\alpha}_2$
$\Delta(m\text{-}p)$	-0.148 (0.021)	0.379 (0.078)
$\Delta\pi$	0.176 (0.054)	-1.015 (0.202)

	$\boldsymbol{\beta}_1$	$\boldsymbol{\beta}_2$
m-p	1.000	0.000
y	-1.395 (0.053)	-0.250 (0.065)
*w**	-1.026 (0.078)	-0.161 (0.030)
π	3.072 (0.267)	1.000
rl-rs	1.521 (0.384)	0.000
trend	0.000	0.002 (0.001)

Note: Sample period 1983Q1-2011Q4. Feedback and cointegrating vectors in the bivariate error correction model for real money balances and inflation, treating real income, real house prices, and the interest rate spread and a linear time trend as exogenous. Standard errors in parantheses.

Table 6 Conditional error correction model

Dependent variable $\Delta(m\text{-}p)$

Con	*d*902	*d*011	ec_liquidity$_{t\text{-}1}$	*ec_inflation*$_{t\text{-}1}$	Δw^*
0.371	0.027	0.020	-0.149	0.374	-0.113
(4.497)	(44.110)	(25.289)	(10.903)	(6.146)	(2.611)

R2=0.606, SE=0.005

JB=2.74 (0.25)	ARCH(1)=0.36 (0.55)	ARCH(2)=0.70 (0.50)	LM(1)=0.01 (0.92)
LM(4)=1.10 (0.36)	RESET(1)=1.09 (0.28)	RESET(2)=0.61 (0.55)	CF(07.1)=0.82 (0.68)
CF(08.1)=0.94 (0.53)	CF(09.1)=1.07 (0.40)	CF(10.1)=0.97 (0.47)	CF(11.1)=0.82 (0.51)

Note: Sample period 1983Q2-2011Q4. Estimation of the error correction model in the upper part with HAC *t*-values in parentheses. Specification tests in the lower part. R2=R squared adjusted, SE=standard error of regression, JB=Jarque-Bera test, LM(*k*)=Lagrange multiplier test for no autocorrelation in the residuals up to order *k*, ARCH(*k*)= LM test for no conditional heteroscedasticity up to order *k*, RE-SET=Ramsey specification test, CF=Chow forecast test. *p*-values in parantheses.

Table 7: Marginal process for real house prices

Dependent variable Δw^*

Con	Δy_{t-1}	Δy_{t-3}	Δrl_{t-2}	*s*074
0.008	0.648	-0.258	-0.918	-0.010
(8.367)	(8.491)	(5.756)	(8.539)	(9.457)

R2=0.837, SE=0.003

JB=3.71 (0.16)	ARCH(1)=0.00 (0.96)	LM(1)=0.12 (0.73)	LM(4)=0.49 (0.74)
RESET(1)=1.23 (0.28)			

Note: Sample period 2002Q1-2011Q4. R2=R squared adjusted, SE=standard error of regression, JB= Jarque-Bera test, LM(*k*)=Lagrange multiplier test for no autocorrelation in the residuals up to order *k*, ARCH(*k*)= LM test for no conditional heteroscedasticity up to order *k*. RESET=Ramsey specification test. Upper (lower) part: HAC *t*-values (*p*-values) in parantheses.