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Money Demand and the Role of Monetary Indicators in Forecasting Euro Area Inflation

Christian Dreger and Jürgen Wolters[1]

Abstract

This paper examines the forecasting performance of a broad monetary aggregate (M3) in predicting euro area inflation. Excess liquidity is measured as the difference between the actual money stock and its fundamental value, the latter determined by a money demand function. The out-of sample forecasting performance is compared to widely used alternatives, such as the term structure of interest rates. The results indicate that the evolution of M3 is still in line with money demand even in the period of the financial and economic crisis. Monetary indicators are useful to predict inflation at the longer horizons, especially if the forecasting equations are based on measures of excess liquidity. Due to the stable link between money and inflation, central banks should implement exit strategies from the current policy path, as soon as the financial conditions are expected to return to normality.

Keywords: Money demand, excess liquidity, money and inflation

JEL Classification: C22, C52, E41

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1 Introduction

Due to the recent recession caused by the financial crisis, the interbank market collapsed and the ECB as well as other central banks had to inject a huge amount of liquidity at low interest rates. Despite the massive increase in the monetary base, however, the money stock in the euro area did not show any noticeable increase. As the interbank market did not allow for a redistribution of liquidity between banks, central banks had to design unconventional policy measures (Freixas, 2009). While the interventions have been rather successful in avoiding a sudden meltdown of the financial system, many analysts have argued that these policies have laid the foundation to destabilise inflation expectations and generate inflation pressures in the future. If financial intermediation returns to normality, the precautionary demand for liquidity will decline, implying that the huge accumulation of reserve balances can result in a rapid increase in the money stock. For example, the euro area monetary multiplier defined as the ratio between M3 and M0 is currently 20 percent below its level at the outbreak of the crisis. Therefore, the ECB is expected to implement exit strategies from the expansionary policy path in order to avoid higher inflation pressure (Minegishi and Cournède, 2010).

In particular, achieving and maintaining the stability of the price level is a main goal for central banks all over the world. Especially in the medium and long run, inflation is inherently a monetary phenomenon (Benati, 2009). However, while the monetary conditions became abnormally loose well before the crisis, inflation did not accelerate at all. Hence, the link between money growth and inflation needs to be revisited. If the relationship has become fragile, money growth might not be well-suited neither for predicting future inflation prospects nor for supporting policy decisions. To contribute to this debate, the analysis investigates the forecasting power of indicators derived from the M3 monetary aggregate with respect to consumer price inflation by taking the period of the financial and economic crisis into account.

Monetary growth does not indicate future inflation per se, as money balances are also driven by real economic activity. For that reason, money demand is crucial for monitoring the inflation process, at least as a long run reference (ECB, 2004). The money demand function links the monetary development to its fundamental determinants, such as real income and the opportunity costs of holding money. By comparing the actual money stock with the long run equilibrium according to money demand, measures of excess liquidity can be derived and might be used to forecast inflation (Dreger and Wolters, 2010b).

Excess liquidity measures are based on the assumption of a stable money demand function. However, recent evidence has cast serious doubts on the robustness of money demand, especially if data after the introduction of the euro as the common currency are used. See for example Gerlach (2004) and Carstensen (2006). However, as Dreger and Wolters (2010a and 2010b) have demonstrated, the instability problem can be resolved by an appropriate specification of the opportunity costs. Specifically, an almost stable money demand function for M3 is obtained if inflation is included. There is still a minor permanent shift in the income elasticity from 2002 onwards (Dreger and Wolters, 2010b). It might reflect better income expectations in the monetary union. Alternatively, wealth variables have become more important to explain the evolution of real money balances since then. Some authors like Greiber and Setzer (2007) and Beyer (2009) included house prices as a proxy for financial wealth. Dreger and Wolters (2009) provided evidence on the impact of wealth effects on money velocity, i.e. they restrict the income elasticity to unity.

However, little is known when the most recent development is taken into account. As an exception, Beyer (2009) has reported evidence for a stable money demand function for M3 using preliminary data until the end of 2008. Similar to Dreger and Wolters (2010a and 2010b), the inclusion of inflation is decisive to achieve this result.

The first contribution of our paper is to examine whether money demand has remained stable over an extended period until 2010Q2 covering the financial and economic crisis. As a main finding, the demand for real money balances appears to be very robust, especially, if real house prices are included as a proxy of wealth. The long run parameters show only minor variation over time.

The second contribution is to explore the forecasting properties of M3 indicators with respect to inflation over the recent period. The importance of money growth and/or excess money measures for inflation has been discussed in various papers. For example, Gerlach and Svensson (2003) found that both the output gap and the real money gap, i.e. the difference between the actual real M3 money stock and its equilibrium value derived from a long run money demand relation contains useful information with respect to one- and two-year ahead HICP inflation rates. However, the nominal M3 annual growth rate provides almost no information regarding the future inflation process. Trecroci and Vega (2002) reported similar results for GDP inflation. Following Nicoletti-Altimari (2001) the real M3 gap or real M3 overhang are important complements to monetary aggregates when inflation is predicted for a two years period. Kaufmann and Kugler (2008) detected a robust cointegration relationship between money growth and inflation. According to their results, shocks in M3 growth account for a substantial part of the inflation forecast error variance. In contrast, the effects of output gap and interest rate shocks on inflation are only transitory and their forecasting variance shares are negligible at the medium term horizons. Carstensen et al (2009) reported evidence that an aggregated monetary overhang can predict country-specific inflation in huge euro area countries, but it does not encompass measures of the country-specific monetary overhang. Most of these papers are based on data ending at the very beginning of the ECB period. In contrast, our analysis produces inflation forecasts for the period from 2003Q4 to 2010Q2, thus including the period of the financial and economic crisis. This period is chosen as the ECB changed its policy strategy in 2003 giving less weight to the monetary pillar. The results indicate that models based on monetary indicators can outperform the benchmark as well as standard alternatives such as the term structure of interest rates, especially at longer forecasting horizons. The increase in the predictive accuracy is markedly when compared to the benchmark. Compared to the term structure, the excess liquidity model reveals a similar performance.

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 are discussed. The specification and estimation of money demand functions in error correction form has been the customary approach to capture the nonstationary behavior of the relevant data. Evidence regarding the cointegration properties and error correction is provided in section 4. The forecasting exercise is performed in section 5. Finally, section 6 concludes.

2 Specification of money demand

A widely used specification of money demand is chosen as the starting point of the analysis, see Ericsson (1998) and Beyer (2009). This specification of money demand leads to a long run relationship of the form

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

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. Opportunity costs of holding money are proxied by nominal long (*R*) and short (*r*) term interest rates and the annualized inflation rate, i.e. $\pi=4\Delta p$, in case of quarterly data. The index *t* denotes time.

Price homogeneity is imposed as a long-run restriction to map the money demand analysis into a system of I(1) variables; see Holtemöller (2004). The income variable exerts a positive effect on nominal and real money balances. Often, its impact is restricted to unity on theoretical grounds, see Dreger and Wolters (2009) for a discussion. Money holdings are also related to portfolio allocation decision. 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 money 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. For the inclusion of the inflation rate see also Dreger and Wolters (2010a). The inflation rate is part of the opportunity costs as it represents the costs of holding money in spite of holding real assets. Its inclusion provides a convenient way to generalize the short run homogeneity restriction imposed between money and prices. In addition, adjustment processes in nominal or real terms can be distinguished (Hwang, 1985).

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 *r* 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 on theoretical grounds.

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 to draw robust conclusions, they have to be extented by artificial data. Euro area series prior to 1999 are obtained by aggregating national time series (Artis and Beyer 2004). By comparing aggregation methods, Bosker (2006) and Beyer and Juselius (2010) have 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. See Juselius (1998) for evidence on a change in the monetary transmission mechanism in European countries in March 1983. Therefore, 1983Q1-2010Q2 is chosen as the observation period. Quarterly seasonally adjusted series are used.

Nominal money balances for M3 are taken from the ECB monthly bulletin database and quarterly data refer to end-of-period values. The short and long term interest rates r and R come also from this source and are defined by the end-of-period 3-month Euribor and ten-year government bond rates, respectively. Nominal and real GDP, as a proxy for income, are taken from Eurostat, the latter defined as chain-linked volumes with 2000 as the reference year. The GDP deflator (2000=1) is constructed as the ratio of nominal to real GDP. The Brand and Cassola (2004) GDP data are used for periods prior to 1991, as these data yield stable and economically interpretable results. 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. The nominal series is taken from the Bank of International Settlement (Borio and Lowe, 2002) and interpolated to the quarterly frequency. HICP inflation is obtained from the ECB monthly bulletin and defined on a year-on-year basis, as this measure is relevant for central banks to monitor inflation. Figure 1 shows the evolution of the time series in levels (A) and first differences (B) during the period from 1983Q1 until 2010Q2. Our data set already starts in 1981Q1 to cover initial values.

-Figure 1 about here-

Except of GDP and HICP inflation, unit root tests indicate that all other variables are integrated of order 1, I(1), implying that they are nonstationary in levels, but stationary in first differences. For the annualized GDP inflation rate and the annual HICP inflation rate ADF tests reject the unit root hypothesis, while KPSS tests reject the null of stationarity. Contrary to the data generating process for unit root tests, where high positive autocorrelation is assumed, the starting point for stationarity tests is a sum of two components, one is stationary and the other one a random walk (see e.g. Kirchgässner and Wolters, 2007). This can lead to different results in testing stationarity versus nonstationarity. The cointegration analysis presented below provides indirect evidence for the nonstationarity of the GDP inflation rate, as this variable is important to get a cointegration vector that can be interpreted as a money demand relationship.

Outliers are detected in the real money balances. 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). 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 due to the financial crisis. Despite the fact that monetary developments have been largely favourable, massive production losses occurred.

-Figure 2 about here-

In the subsequent analysis, outliers are acknowledged by two impulse dummies, which are equal to 1 in the respective period and 0 otherwise ($d902$ and $d011$). The break in the income elasticity is captured by an additional income variable y^*, defined as the product of y and a step dummy $s021$ equal to 1 from 2002.1 until the end of the sample and 0 in the period before. See Lütkepohl, Teräsvirta and Wolters (1999) for this strategy, that has been applied in Dreger and Wolters (2010b). As an alternative to capture the break, real house prices (w^*) are included as the product of w and the step dummy $s021$ (Dreger and Wolters, 2009). As the aim is to measure excess liquidity, dummies according to the financial crisis are not introduced at all.

4 Cointegration analysis and error correction

In systems including real money balances, real income, nominal interest rates and inflation, at least one cointegrating relationship should represent a long run money demand function. The cointegration properties are explored using the Johansen (1995) trace test. The lag length of the VAR in levels is determined by the Schwarz criterion and equal to one throughout the analysis. Moreover, an unrestricted constant term and the two impulse dummies are added.

Under these settings, Dreger and Wolters (2010b) have provided evidence in favour of a stable and well behaved money demand function for the 1983Q1-2006 Q4 period, when inflation is considered as an opportunity cost measure and y^* is used to capture the break in the income elasticity. Investigating money velocity, Dreger and Wolters (2009) present results that the introduction of financial wealth is able to model this break. Therefore, we use y^* and w^* as alternatives to estimate money demand equations. In addition, it has been demonstrated that the cointegration parameters can be estimated with higher precision, if the term structure, i.e. the difference between the long and short term interest rate is included[2].

Table 1 reveals the trace test for the null hypothesis of no cointegration and the parameters of the long run money demand equation for a varying sample period. Results for the model with a break in the income (house price) elasticity are reported in the upper (lower) part of the table.

[2] As the term structure of interest rates is a stationary series, the cointegration rank of the system is larger than one. Since the focus of the paper is on money demand, the respective results are not reported in the analysis. They can be obtained from the authors upon request.

For both specifications at least one highly significant cointegration relationship in line with a long run money demand function can be detected. The model with a break in the income elasticity (Table 1, panel A) shows the theoretically expected signs of the coefficients. Nonetheless, the parameters fluctuate strongly over varying sample endpoints, especially for opportunity cost measures. For example, the variation between the largest and smallest coefficient is 56 percent for the semielasticity of money demand with respect to inflation (=(10.61-4.714)/10.61). The variables in the model with real house prices (Table 1, panel B) also reveal the correct signs. For wealth, the scale effect dominates the substitution effect. The long run parameters appear to be quite stable if the sample is successively enlarged from 2006Q4 to 2010Q2. For example, the maximum variation of the inflation coefficient is only 18 percent. In addition, the estimates for inflation and the term structure are more precise than in the model with a break in the income elasticity. Overall, the analysis points to a stable relation between the variables entering the money demand function even during the period of the financial and economic crisis. In fact, the cointegration relationship can be interpreted in terms of a long run money demand equation, as the null of weak exogeneity can be rejected for real money balances in each case.

-Table 1 about here-

As a further check of robustness, conditional single equation models are evaluated. In general, a conditional model may lead to constant coefficients even if a shift occurs in the reduced form (Johansen, 1995). Given the identification problems in full systems with multiple cointegration vectors, a structural model for an individual variable might be easier to specify and to test in a single equation context. The error correction model for money demand is estimated in one step, where the long run parameters are obtained jointly with the short run dynamics (Stock, 1987). At the initial stage, the contemporaneous and the first four lags of the changes of all variables, a constant and the two impulse dummies are included in addition to the one period lagged levels of the variables included in the cointegration vector. Then, the variables with the lowest and insignificant *t*-values are eliminated subsequently, where a 0.1 level is used. The final equations confirm the findings of the multivariate analysis, as the implied cointegration vectors are very similar, where the correspondence for the house price model is better than for the income model (Table 2). This is also confirmed in Figure 3, where the error correction terms for both approaches are plotted.

-Table 2 and Figure 3 about here-

In both cases, the residuals are well behaved, as they are normal, homoscedastic and do not show autocorrelation patterns. Moreover, the tests on the functional form do not reveal problems. However, the house price error correction model turns out to be superior: first, comparing the coefficients of the one period lagged level of real money and its *t*-value, the evidence for cointegration and the adjustment towards equilibrium is much stronger for this specification. Second, with a smaller number of explanatory variables, the fit in the *w** model is better than in the *y** alternative. Third, according to Chow forecast tests, the *w** model is stable even in the period of the financial and economic crisis, in contrast to the *y** specification. This evidence is also confirmed by the recursive residuals and the cusums-of-squares tests in Figure 4. Therefore, the real house price variant is selected for the forecasting exercise.

-Figure 4 about here-

5 Monetary models and inflation forecasts

The evolution of monetary aggregates provides information on future inflation pressures when they can improve inflation forecasts. The different inflation rates are defined as follows:

$$\pi_{c,t}^{k} = \frac{4}{k}\log(pc_t / pc_{t-k}) \quad , \quad k = 1,4,8,12$$

In the out-of sample forecast experiment, the annual change of the consumer price index (*pc*), *k*=4 is used, as well as average cumulative inflation rates over the two and three years horizon (*k*=8, 12). They are also relevant for the monetary authorities, as they reveal information on the inflation potential in the long run. Temporary changes in high volatile prices are removed if these measures are selected. The forecasting equations are given by

$$\text{(2)} \qquad \pi^{k}_{c,t+k} = \alpha(L)\pi^{1}_{c,t} + \beta x_t + u_{t+k} \quad , \quad k = 4,8,12$$

where $\alpha(L)$ is a lag polynom, ensuring that the equations are balanced. Future inflation for k=4, 8 or 12 quarters ahead is predicted by current and lagged inflation up to order 3 and additional variables known at the time the forecast is made (x). Since the forecast error u follows a moving average process of order k-1, the autocorrelation and heteroscedasticity consistent covariance estimator proposed by Newey and West (1987) is used to evaluate the significance of the regression parameters.

To constitute a benchmark, future inflation is predicted by current and lagged inflation. Alternative models arise if further variables are added to the benchmark. Several specifications are explored. While the first alternative is based on annual M3 growth, the second one includes the error correction term at period t, i.e. the deviation from the long run money demand function. This accounts for the fact that money is not an indicator for inflation per se. Instead excess liquidity matters. The forecasting properties of the term structure of interest rates serve as a further competitor. If real interest rates are almost constant and risk premia for government bonds with different periods to maturity do not fluctuate much, future inflation differentials can be predicted by the nominal interest rate spread, see e.g. Fama (1990), Mishkin (1990).

The forecasting performance of the different models is evaluated in an out-of-sample exercise. This mimics the actual situation the forecaster is confronted with. Naturally we do not perform a real time analysis due to data availability and since revisions in monetary aggregates and interest rates are neglibile. Moreover, as the long run money demand equation from the wealth model is very stable it makes no difference using the error correction term from the full sample or corresponding subsamples.

In particular, the forecasts are obtained in a recursive manner. The first estimation subsample is 1983Q1-2002Q4 and the forecast subsample is 2003Q4-2010Q2 in case of annual inflation rates. After producing the forecast for 2003Q4, the estimation period is extended by one quarter (1983Q1-2003Q1) and the forecast for 2004Q1 is made. This process is repeated until the end of the sample is reached (2010Q2). For the multiyear forecasting horizons, the first estimation subsample is again 1983Q1-2002Q4. Hence, 27 annual, 23 biennial and 19 triennial forecasts are derived. The forecast accuracy is evaluated by the root mean square forecast error, expressed relative to the benchmark model. For robustness, the relative mean absolute forecast error is also considered. The results are exhibited in Table 3.

-Table 3 about here-

The average root mean square forecast error exceeds the mean absolute forecast error at the shorter horizons due to possible outliers. In general, the average forecast errors decline with the forecast horizon, as idiosyncratic shocks are smoothed out at the longer intervals. There are no huge differences in the forecast accuracy between the models at the annual horizon. However, the predictive accuracy can be improved for longer periods if the forecasting equation is extended by additional variables. This means that fundamental information becomes more important. In particular, the model based on excess liquidity leads to a root mean square forecast error which is 15 percent below the one of the benchmark at the 3-year horizon. Compared to money growth, excess liquidity is superior to predict inflation. However, the models using excess liquidity or the term structure show a similar forecasting performance at the two- and three-years horizon.

6 Conclusions

This paper examines the forecasting performance of a broad monetary aggregate (M3) in predicting euro area inflation. Excess liquidity is measured as the difference between the actual money stock and its fundamental value, the latter derived from a money demand function. The out-of sample forecasting performance is compared to widely used alternatives, such as the term structure of interest rates. The results indicate that the evolution of M3 is still in line with money demand even in the period of the financial and economic crisis, especially if real house prices are included as a proxy for financial wealth. The long run parameters appear to be very stable, and the error correction model passes all standard specification tests.

Compared to the benchmark of an autoregressive process for inflation a payoff can be realized if additional variables are used as predictors. The relative accuracy is raised with the forecasting horizon. In particular, the model based on excess liquidity leads to a forecast error which is 15 percent below the one of the benchmark at the 3-year horizon. Compared to money growth, excess liquidity is superior to predict inflation. While these gains seem to be substantial, excess liquidity does not outperform the term structure and vice versa. According to these results, higher inflation is expected to materialize, if the monetary multipliers increase to the pre-crisis levels. Hence, the ECB should implement exit strategies from the expansionary policy path, as soon as financial conditions tend to return to normality.

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Figure 1A: Variables in level

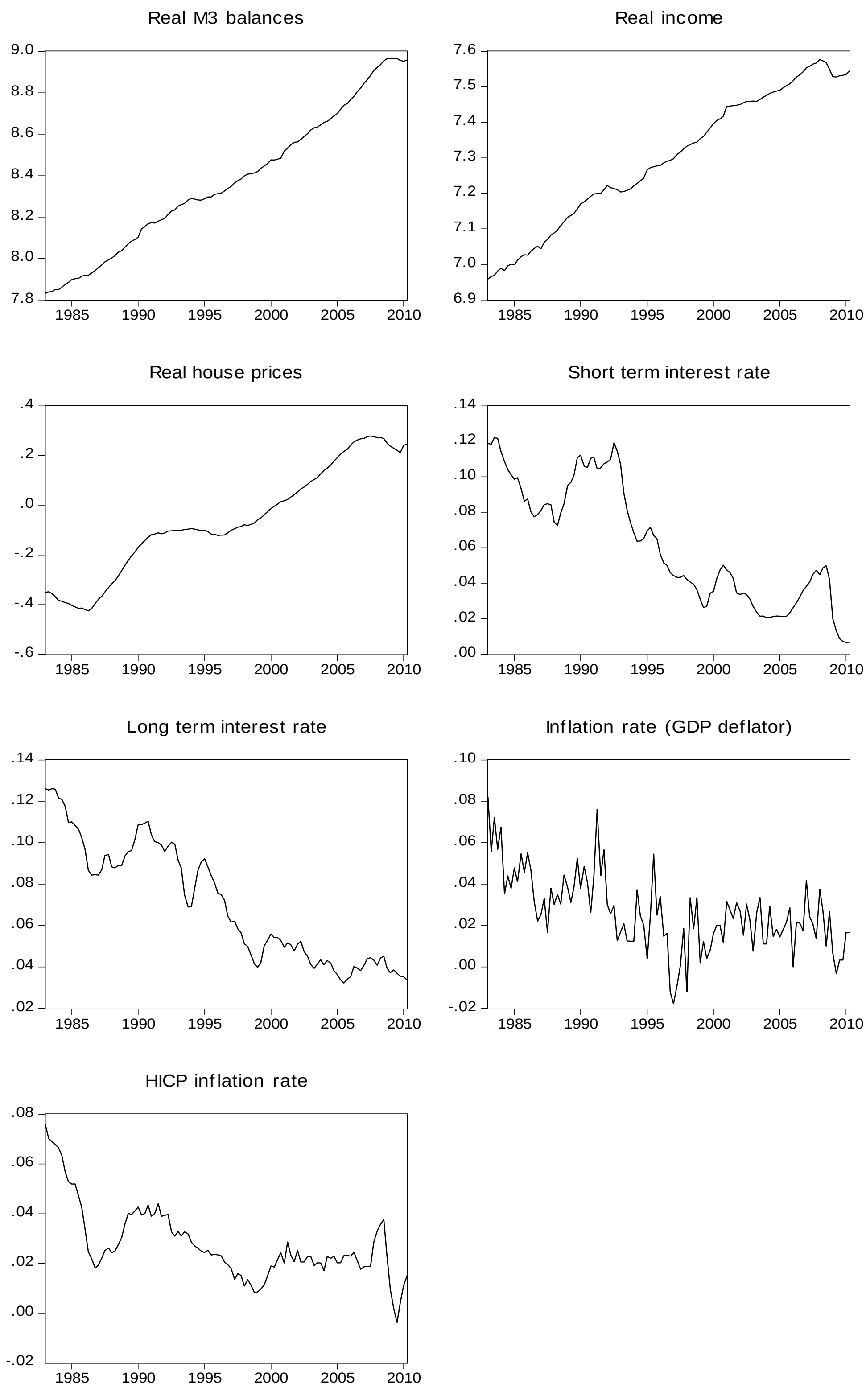
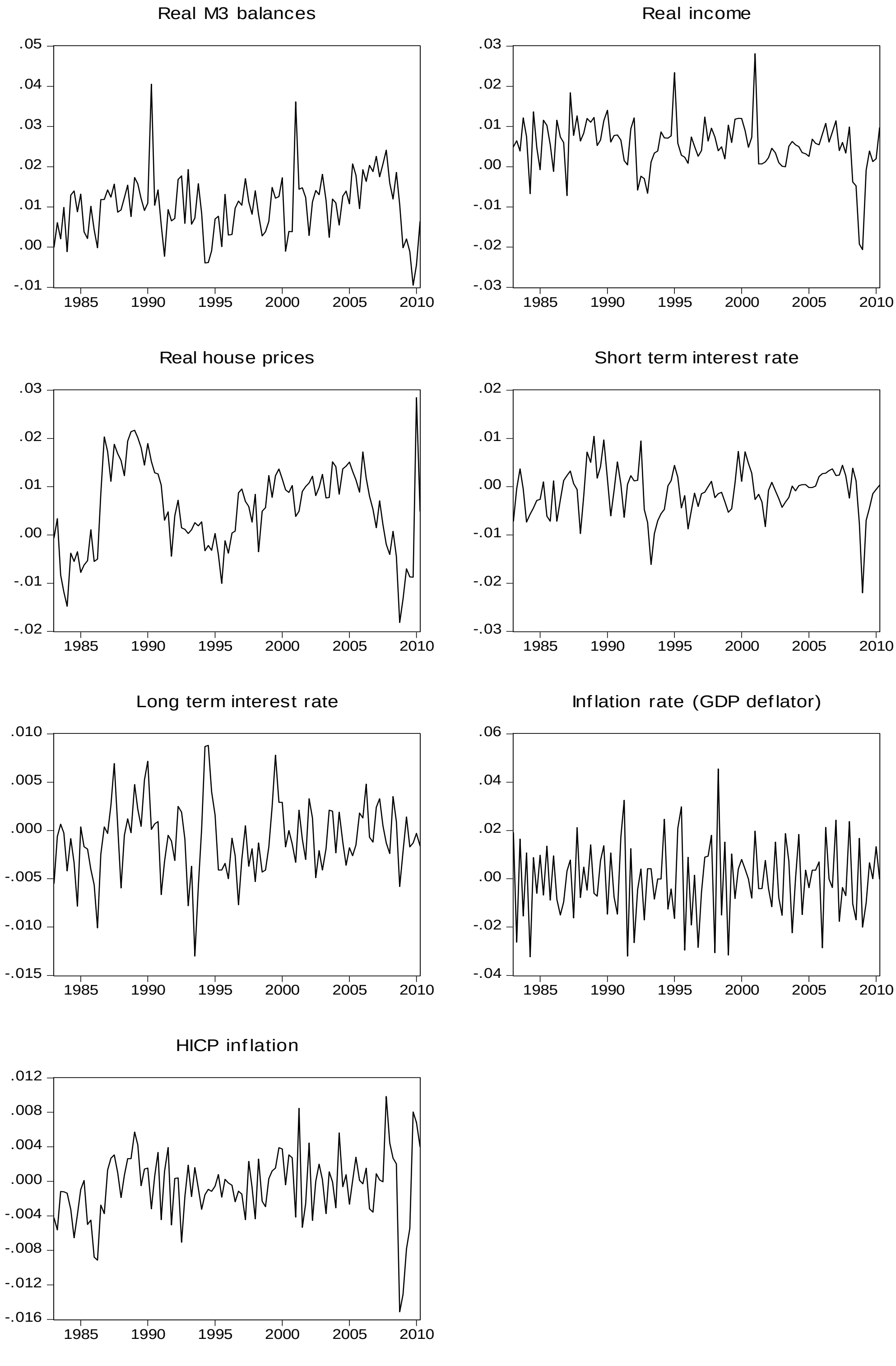
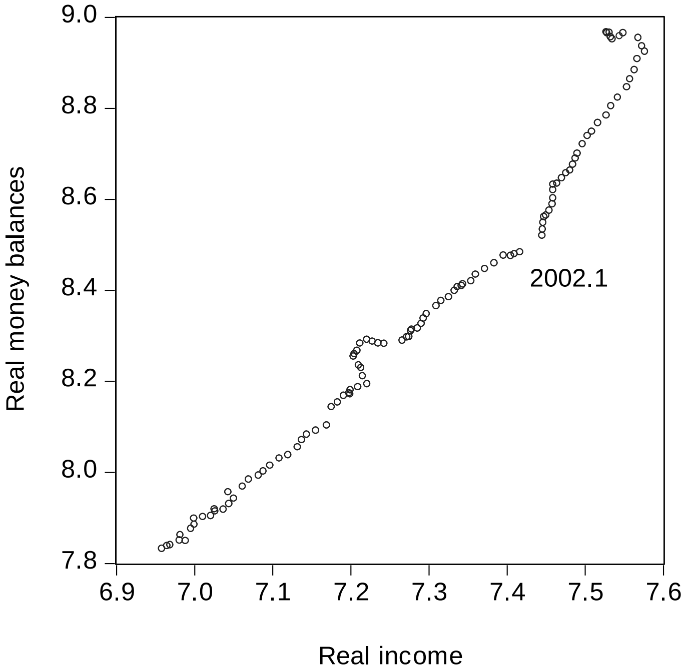


Figure 1B: Variables in first differences



Note: Sample period 1983.1-2010.2. Real money, real GDP and real house prices in logs. Inflation q-o-q change in the GDP deflator (2000=1) or y-o-y change in the HICP.

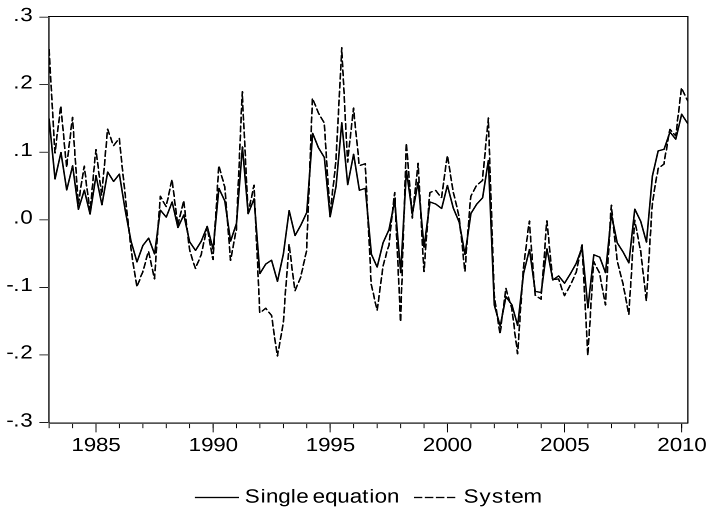
Figure 2 Structural break in income elasticity



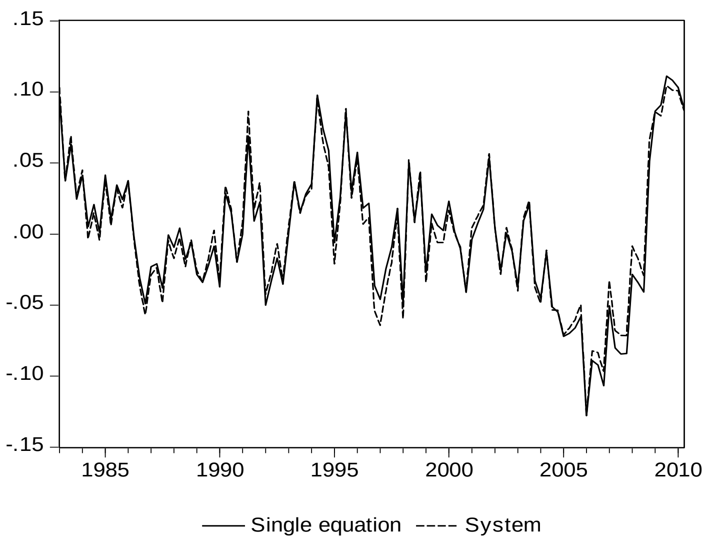
Note: Sample period 1983.1-2010.2.

Figure 3 Error correction terms for system and single equation approach

A Break in the income elasticity



B Break in the house price elasticity

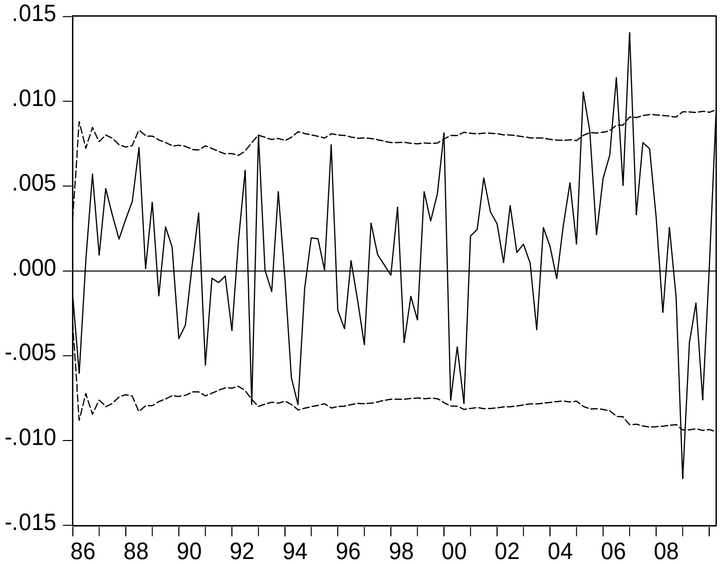


Note: Sample period 1983.1-2010.2. Error correction terms (mean adjusted) obtained from the Johansen (1995) system and the Stock (1987) single equation approach.

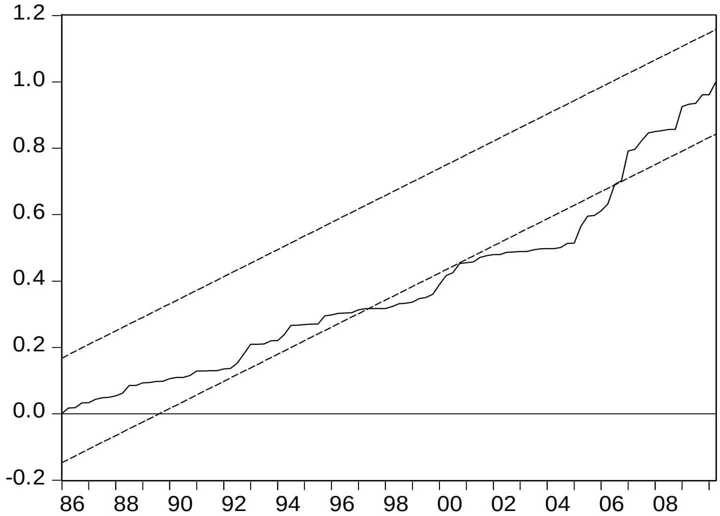
Figure 4 Parameter stability of the error correction equations

A Income model

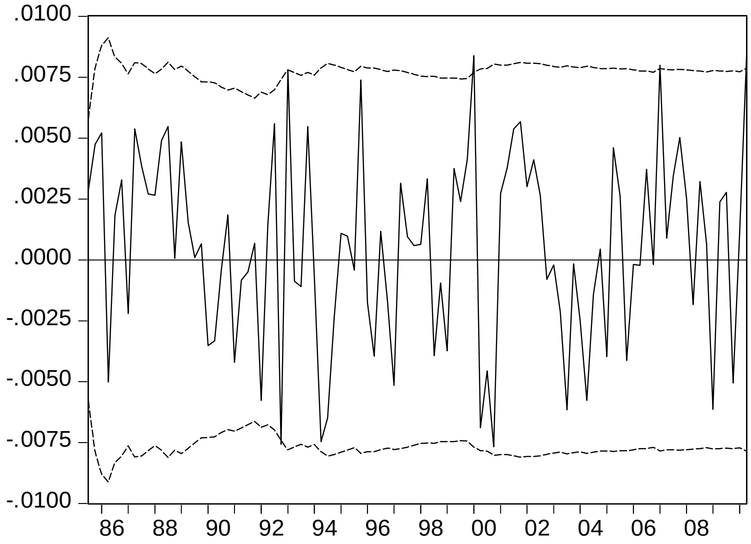
Recursive residuals

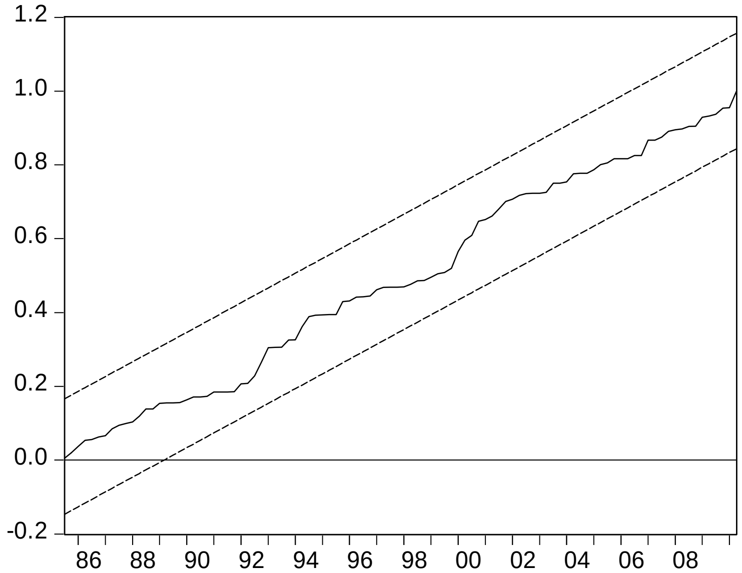


Cusum of squares



B House price model





Sample period 1983.1-2010.2. Dashed lines represent 0.05 significance levels.

Table 1 Cointegration analysis of money demand

A Break in the income elasticity

Sample	Trace	y	y^*	π	ts
83Q1-06Q4	113.8**	1.184 (0.067)	0.026 (0.003)	-4.714 (0.471)	-3.647 (0.759)
83Q1-07Q2	112.4**	1.114 (0.095)	0.035 (0.005)	-6.044 (0.665)	-5.593 (1.068)
83Q1-07Q4	114.8**	1.015 (0.125)	0.044 (0.006)	-7.742 (0.877)	-7.662 (1.384)
83Q1-08Q2	117.7**	0.936 (0.146)	0.051 (0.007)	-9.079 (1.023)	-9.290 (1.582)
83Q1-08Q4	123.9**	0.849 (0.171)	0.059 (0.008)	-10.61 (1.198)	-11.31 (1.810)
83Q1-09Q2	141.1**	1.059 (0.120)	0.044 (0.006)	-7.503 (0.841)	-8.535 (1.264)
83Q1-09Q4	148.8**	1.200 (0.087)	0.034 (0.004)	-5.426 (0.607)	-6.675 (0.902)
83Q1-10Q2	151.5**	1.190 (0.095)	0.036 (0.004)	-5.707 (0.658)	-7.190 (0.958)

B Break in the house price elasticity

	Trace	y	w^*	π	ts
83Q1-06Q4	185.6**	1.371 (0.027)	0.797 (0.055)	-2.185 (0.192)	-1.145 (0.310)
83Q1-07Q2	189.2**	1.377 (0.027)	0.835 (0.052)	-2.143 (0.194)	-1.200 (0.310)
83Q1-07Q4	192.6**	1.379 (0.027)	0.853 (0.049)	-2.139 (0.195)	-1.247 (0.308)
83Q1-08Q2	194.2**	1.376 (0.027)	0.856 (0.047)	-2.171 (0.192)	-1.286 (0.299)
83Q1-08Q4	201.9**	1.375 (0.027)	0.852 (0.046)	-2.176 (0.188)	-1.331 (0.290)

83Q1-09Q2	228.3**	1.386 (0.024)	0.807 (0.041)	-1.988 (0.172)	-1.114 (0.263)
83Q1-09Q4	239.6**	1.390 (0.024)	0.813 (0.040)	-1.968 (0.173)	-1.085 (0.260)
83Q1-10Q2	225.2**	1.353 (0.030)	0.899 (0.050)	-2.393 (0.216)	-1.406 (0.316)

Note: All models estimated with unrestricted constant terms and two impulse dummies for the German unification (1990Q2) and stock market turbulences (2001.Q1). The trace test is for the null hypothesis of no cointegration. Critical values are from MacKinnon, Haug and Michelis (1999). A ** denotes significance at the 0.01 level. Lag order of 1 in underlying VAR models (level specification), due to the Schwarz criterion. Numbers below the (semi)elasticities denote standard errors.

Table 2 One step estimation of the ECM

A Break in the income elasticity

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

Con	$d902$	$d011$	$(m\text{-}p)_{t\text{-}1}$	$y_{t\text{-}1}$	$y^*_{t\text{-}1}$	$\pi_{t\text{-}1}$	$(R\text{-}r)_{t\text{-}1}$
-0.055 (1.195)	0.031 (6.382)	0.032 (6.105)	-0.051 (5.075)	0.067 (4.186)	0.001 (5.360)	-0.165 (4.016)	-0.191 (3.294)

$\Delta\pi_t$	$\Delta(m\text{-}p)_{t\text{-}1}$	$\Delta(m\text{-}p)_{t\text{-}3}$	$\Delta(m\text{-}p)_{t\text{-}4}$
-0.170 (4.621)	0.232 (3.124)	0.151 (2.087)	-0.166 (2.287)

Long run: $m-p=1.323y+0.029y^*-3.246\pi_t-3.767(R-r)$

R2=0.613, SE=0.005

JB=0.39 (0.82)	ARCH(1)=0.22 (0.64)	ARCH(2)=1.32 (0.27)	LM(1)=0.24 (0.62)
LM(2)=1.98 (0.14)	LM(4)=1.86 (0.12)	LM(8)=1.81 (0.09)	RESET(1)=2.41 (0.12)
RESET(2)=4.06 (0.13)	CF(07.1)=2.54 (0.00)	CF(08.1)=1.60 (0.12)	CF(09.1)=2.55 (0.03)

B Break in the real house price elasticity

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

Con	$d902$	$d011$	$(m\text{-}p)_{t\text{-}1}$	$y_{t\text{-}1}$	$w^*_{t\text{-}1}$	$\pi_{t\text{-}1}$	$(R\text{-}r)_{t\text{-}1}$
-0.188 (5.356)	0.029 (7.258)	0.026 (6.389)	-0.118 (11.82)	0.163 (11.15)	0.108 (11.21)	-0.246 (7.288)	-0.218 (5.404)

$\Delta\pi_t$	$\Delta(m\text{-}p)_{t\text{-}4}$
-0.176 (5.847)	-0.148 (2.505)

Long run: $m-p=1.381y+0.914w^*-2.086\pi_t-1.849(R-r)$

R2=0.735, SE=0.004

JB=1.19 (0.55)	ARCH(1)=0.01 (0.95)	ARCH(2)=0.02 (0.98)	LM(1)=0.01 (0.99)
LM(2)=0.22 (0.80)	LM(4)=1.09 (0.37)	LM(8)=1.18 (0.32)	RESET(1)=1.30 (0.26)
RESET(2)=0.70 (0.50)	CF(07.1)=1.30 (0.23)	CF(08.1)=1.10 (0.37)	CF(09.1)=1.66 (0.14)

Note: Sample period 1983.1-2010.2. One-step estimation of the error correction model in the upper part, standard specification tests in the lower part of each subtable. 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 conditional heteroscedasticity up to order k, RESET=Ramsey specification test, CF=Chow forecast test. Upper (lower) part: t-values (p-values) in parantheses.

Table 3 Out-of-sample forecasting performance of different models

Horizon	RMSFE	MAFE	Money growth	Excess liquidity	Term structure
4	1.308	0.933	0.981 | 0.918	1.019 | 1.012	0.961 | 0.941
8	0.896	0.665	1.085 | 0.989	0.937 | 0.892	0.923 | 0.923
12	0.636	0.518	0.950 | 0.858	0.842| 0.833	0.856 | 0.812

Note: The root mean squared forecast error (RMSFE) and mean absolute forecast error (MAFE) are taken from the benchmark and expressed in percent. The three colums on the right report the RMSFE (left) or MAFE (right) relative to that of the benchmark. Excess liquidity is derived as error correction term from the model with a break in the real house price elasticity.