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On the Stability of Euro Area Money Demand and its Implications for Monetary Policy

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Abstract

We employ a recent time-varying cointegration test to revisit the usefulness of long-run money demand equations for the ECB, addressing the issue of their instability by means of a model evaluation exercise. Building on the results, we make a twofold contribution. First, we propose a novel stable money demand equation relying on two crucial factors: a speculative motive, represented by domestic and foreign price-earnings ratios, and a precautionary motive, measured by changes in unemployment. Second, we use the model to derive relevant policy implications for the ECB, since excess liquidity looks more useful for forecasting stock market busts than future inflation. Overall, this evidence points to (*i*) a possible evolution of the monetary pillar in the direction of pursuing financial stability and (*ii*) the exclusion of a sudden liquidity–driven inflationary burst after the exit from the prolonged period of unconventional monetary measures.

JEL Classification: E41, E52, C32

Keywords: money demand; time-varying cointegration; monetary policy; financial stability; price stability.

The views here expressed are those of the authors and do not necessary reflect those of the Banca d'Italia or the Eurosystem.

1 Introduction

Monetary aggregates have been largely neglected by the prevailing macroeconomic consensus until the Global Financial Crisis of 2008–09 triggered a renewed interest for money's role in the business cycle and the financial cycle. Following the rising and diffusion of the New Keynesian model as a tool for policy analysis and forecasting, money has been indeed gradually disappearing from monetary analysis (Nelson, 2003).[1] Such theoretical developments, together with empirical evidence on the instability of the relationship between inflation and money growth, had a substantial impact on the conduct of monetary policy. For instance, the relative weights of the two pillars on which the strategy of the European Central Bank (ECB) is based – one resting on monetary analysis the other one on economic analysis – has changed over time, with the latter taking center stage in 2003 (European Central Bank, 2003). Accordingly, the great emphasis put on the estimation of money demand equations decreased, also in relation to their instability over time (see Papademos and Stark, 2010, for a recent survey). On the other hand, a number of contributions has recommended not to neglect money's role in macroeconomics, both on theoretical and empirical grounds, because of its impact on economic activity and inflation (Reynard, 2007; Nelson, 2008; Beck and Wieland, 2008; Favara and Giordani, 2009; Canova and Menz, 2011).

Given this debate, in this paper we re-evaluate the relevance of monetary aggregates for price and financial stability. In particular, by means of a newly developed econometric method we propose a new stable money demand equation from which we derive a measure of excess liquidity (with respect to its long-run equilibrium) that provides a valuable signal for both future financial imbalances and inflationary pressures. Financial intermediaries' balance sheets are under the constant monitoring of policy makers and the literature has suggested that monetary aggregates, which are the counterpart of bank lending, convey information on the stage of the financial cycle (Shin and Shin, 2011). Indeed, one can look at the demand for money, i.e. banks' liabilities, as supply of funding to banks. Therefore, a novel empirical analysis of the determinants of the long-run stock of money is required for studying the role of

[1] In contrast, Ireland (2004) has challenged this view and a number of more recent studies have included money in models aimed to monetary policy analysis (Zanetti, 2012; Benati *et al.*, 2017).

monetary aggregates in shaping policies oriented to financial stability (European Central Bank, 2012; Kim *et al.*, 2012; Allen *et al.*, 2014).

In order to accomplish this task we first have to address the well-known issue of the instability of money demand (De Santis *et al.*, 2013), which in turn requires comparing different existing models. While the emergence of unstable money demand behaviour in the Euro Area (EA) seems to originate in the early 2000s, the exact date of such break remains a matter of discussion. Hence, we investigate the existence of a stable long-run money demand equation in the EA by using a time-varying cointegration test proposed by Bierens and Martins (2010), a generalization of the standard methodology by Johansen (1996), which avoids the need for specifying the exact timing of the structural break.

Our contribution is twofold. First, we run a model evaluation exercise of different specifications of EA money demand proposed by the literature (Carstensen, 2006; Dreger and Wolters, 2010a, and many others; see Table 1 for a complete list), in order to determine which ones, if any, are stable and therefore effective for policy analysis. We consider a quarterly dataset spanning the sample from 1980:Q1 to 2016:Q3 and we find that most of the considered models are indeed unstable after 2001, although improvements are observed when including either (*i*) the spread between EA and US price-earnings ratios (De Santis *et al.*, 2013), or (*ii*) the changes in EA unemployment rate (De Bondt, 2010). Second, we develop further this result by proposing, estimating, and identifying a new stable model for money demand relying on *both* those motives. Finally, we derive useful implications for monetary policy. Specifically, we find that: (*i*) the monetary overhang computed from our model is a leading indicator of the probability of stock market busts and (*ii*) it provides some incremental predictive content for forecasting inflation, although this tends to vanish when excess liquidity is considered together with real GDP growth, proxying the economic pillar of the ECB.

To our knowledge, this is the first contribution proving that excess liquidity measures derived from estimated long-run money demand equations may be helpful for predicting financial crises. Moreover, this finding is still valid when controlling for the slope of the yield curve, the credit-to-GDP ratio and non-core banks liabilities, considered among the most relevant predictors of financial stress (Bank of International Settlements, 2010; Hahm *et al.*, 2013). Hence, we add to the increasing debate on central banks financial stability mandate (Kim *et al.*, 2013; Hahm *et al.*, 2013) suggesting that information from the monetary pillar, in particular from excess liquidity measures, should mainly be used to assess the risk of financial crises in the EA.

Results are robust to the estimation of our model on different time spans, i.e. both in "normal times" (1980:Q1–2008:Q4) and when extending the sample to most recent data (1980:Q1–2016:Q3), covering the Global Financial Crisis, the Sovereign Debt Crisis and the unconventional measures implemented by the ECB. Furthermore, our model helps in predicting M3 growth in the EA both in "normal times" and in times of financial stress.

Overall, our findings support the usefulness of money demand equations as tools for monetary policy, provided they are properly specified, and suggest that they matter more for financial stability than for controlling inflation.

The rest of the paper is organized as follows. Section 2 reviews the literature on modeling long-run money demand in the EA. Section 3 presents the data employed, while further details are given in Appendix A. Section 4 sketches the methodology by Bierens and Martins (2010) and reports the results of the model evaluation exercise, before moving in Section 5 to the estimation of a new stable dynamic model of money, unemployment and stock markets. Diagnostics of the estimated model are in Appendix B. Section 6 draws the policy implications for the ECB. Finally, in Section 7 we conclude. Supplementary results referred to in the paper are available in an online appendix on the authors' personal webpages.

2 Modeling Euro Area money demand

We define M_t as the nominal monetary aggregate, usually M3 in empirical studies on the EA, P_t as the aggregate price level, and Y_t as real GDP. Typically, both the quantity of real money holdings, or real balances, $\log(M_t/P_t) =: (m_t - p_t)$ and income $\log Y_t =: y_t$, display a unit root, i.e. are $I(1)$ processes. Therefore, the literature on money demand usually focuses on the estimation of a long-run equilibrium, i.e. cointegration, relation as

$$(m_t - p_t) = \beta_0 + \beta^y y_t + \boldsymbol{\beta}^{\mathbf{X}} \mathbf{X}_t + \eta_{1,t}, \quad \eta_{1,t} \overset{iid}{\sim} (0, \sigma^2) \tag{1}$$

Table 1: Models for Euro Area money demand. Specifications considered.

Ref.	Acronym	$\mathbf{X}_t$	r	Sample	k
Coenen and Vega (2001)	CV	ℓ_t, s_t, π_t	3	1980:Q1-1998:Q4	5
Calza *et al.* (2001)	CGL1	$(s_t - o_t), (\ell_t - o_t)$	1	1980:Q1-1999:Q4	4
Calza *et al.* (2001)	CGL2	$(s_t - o_t)$	1	1980:Q1-1999:Q4	3
Gerlach and Svensson (2003)	GS	$(\ell_t - s_t)$	2	1980:Q1-2001:Q4	3
Vlaar (2004)	V	$(\ell_t - s_t), \pi_t$	3	1980:Q1-2000:Q4	4
Carstensen (2006)	C	$(s_t - o_t), (\Delta e_t - o_t), v_t$	1	1980:Q1-2003:Q4	5
Beyer (2009)	B1	$o_t, s_t, \pi_t, \Delta h_t$	2	1980:Q1-2008:Q4	6
Beyer (2009)	B2	Δh_t	2	1980:Q1-2008:Q4	3
Beyer (2009)	B3	$\Delta h_t, \pi_t$	2	1980:Q1-2008:Q4	4
Dreger and Wolters (2010a)	DW1	π_t	1	1983:Q1-2004:Q4	3
Dreger and Wolters (2010b)	DW2	s_t, l_t, π_t, f_t	2	1983:Q1-2010:Q2	6
Dreger and Wolters (2010b)	DW3	f_t, π_t	1	1983:Q1-2010:Q2	4
Dreger and Wolters (2010b)	DW4	$f_t, \pi_t, (l_t - s_t)$	2	1983:Q1-2010:Q2	5
De Bondt (2010)	D1	$w_t, o_t, \Delta e_t, \Delta u_t$	1	1983:Q1-2007:Q2	6
De Bondt (2010)	D2	$w_t, o_t, \Delta e_t$	1	1983:Q1-2007:Q2	5
De Bondt (2010)	D3	$o_t, \Delta u_t$	2	1983:Q1-2007:Q2	4
De Santis *et al.* (2013)	DFR1	$\ell_t, \ell_t^*, q_t, q_t^*, o_t$	3	1980:Q1-2007:Q3	7
De Santis *et al.* (2013)	DFR2	$\ell_t, \ell_t^*, q_t, q_t^*$	3	1980:Q1-2007:Q3	6
De Santis *et al.* (2013)	DFR3	$(\ell_t - \ell_t^*), (q_t - q_t^*)$	3	1980:Q1-2007:Q3	4
Barigozzi and Conti	BC	$\Delta u_t, (q_t - q_t^*), (\ell_t - o_t), \pi_t$	3	1980:Q1-2008:Q4	6
Barigozzi and Conti	BC	$\Delta u_t, (q_t - q_t^*), (\ell_t - o_t), \pi_t$	3	1980:Q1-2014:Q2	6

The time-varying long run money demand equation considered is $(m_t - p_t) = \beta_0 + \beta_t^y y_t + \boldsymbol{\beta}_t^X \mathbf{X}_t$, where: $(m_t - p_t)$ = log-real balances, y_t = log-real income, ℓ_t = long-term interest rate, s_t = short-term interest rate, o_t = own rate, π_t = q-o-q inflation rate (in Beyer (2009) y-o-y inflation rate), Δe_t = q-o-q equity returns, v_t = log-volatility of equity returns, Δh_t = y-o-y housing wealth growth rate, f_t = log-real financial wealth, w_t = log-real wealth, Δu_t = y-o-y differences of unemployment rate, ℓ_t^* = US long-term interest rate, q_t = log-price to earnings ratio, q_t^* = US log-price to earnings ratio. The last three columns report the cointegration rank r, the sample, and the number of variables k, used in the original paper.

where $\mathbf{X}_t$ is a vector of variables representing either the opportunity cost of holding money or other possible motives as, for example, the own-rate of M3, i.e. the return of the less liquid assets contained in M3, the short-term interest rate, i.e. the policy instrument, and inflation. We denote as $\eta_{1,t}$ the cointegration residual, thus allowing for the possibility of more than one cointegration relation in the system.

In this Section we review the possible specifications proposed in the literature which are summarized in Table 1, together with the respective choice of $\mathbf{X}_t$, the cointegration rank r, and the time sample considered. In the benchmark model for EA money demand proposed by Calza *et al.* (2001), the vector $\mathbf{X}_t$ contains the spread between the long-term and short-term interest rates and the long-term and own rate of M3. This model has been used in the Quarterly Monetary Assessment by ECB during the period 2001-2006 (see Fischer *et al.*, 2009). Similar specifications are also considered by Coenen and Vega (2001), Gerlach and Svensson

(2003), and Vlaar (2004). In all these works, to which we refer as to *classical specifications*, real balances depend on real GDP and different combinations of the interest rates or their spreads, while Dreger and Wolters (2010a) add also inflation.[2]

In general, the classical specifications provide a good fit of the demand for EA M3 from 1980 to 2001. However, when extending the sample over 2001 all these models fail in delivering a stable relation, and, in particular, the estimated coefficient of output $\widehat{\beta}^y$ becomes unstable (see e.g. Dreger and Wolters, 2010b), with no additional benefit from allowing for a linear trend to capture velocity shift (see Beyer, 2009).[3] More precisely, instability is usually assessed in two ways: first, by estimating recursively the model and using the related tests for stability of the estimated parameters (see Juselius, 2006, for more details), and, second, by studying the estimated cointegration residual $\widehat{\eta}_{1,t} = (m_t - p_t) - \widehat{\beta}_0 - \widehat{\beta}^y y_t - \widehat{\boldsymbol{\beta}}^{\mathbf{X}} \mathbf{X}_t$. For classical specifications $\widehat{\eta}_{1,t}$ is found to be no more mean reverting after 2001 (see Figure 1.a and De Santis *et al.*, 2013).

Such unstable behavior of the estimates may suggest either the omission of some key explanatory variable in $\mathbf{X}_t$ or some structural change in the parameters of the model.[4]

More recent works claim to solve the problem of instability by adding to the vector $\mathbf{X}_t$ new variables related to financial, housing, or labor markets. In particular, Carstensen (2006) proposes to add equity returns and stock market volatility to account for the stock market impact on real balances, while Beyer (2009) stresses the role of housing wealth in capturing the trending behaviour of money in the first ten years of ECB existence. Dreger and Wolters (2010b, 2011) consider also real house prices as a measure of real financial wealth and then further develop their argument by analyzing the relation between money and inflation in the Global Financial Crisis.

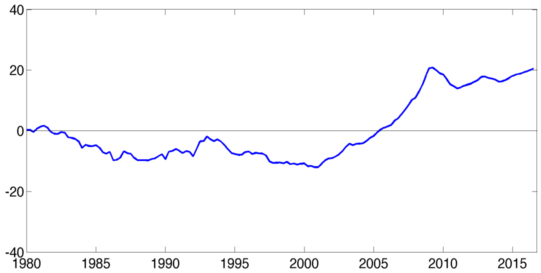
Finally, as we show in the rest of this paper, two specifications have a particular role in studying money demand. First, De Bondt (2010) focuses on the role of equity and labor market dynamics. The latter motive is measured by changes in unemployment which represent a precautionary saving motive: money is hoarded when labor market weakens, i.e. when

[2]Other similar specifications may be found in Funke (2001) and Brand and Cassola (2004).

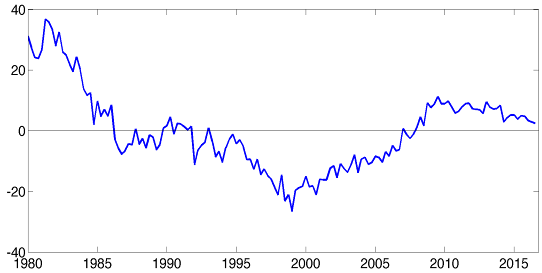
[3]The same is often true also for the estimated coefficient of the the short-term spread.

[4]It may in turn reflect also changes in the deep parameters of the model, of which the coefficients are convolutions (Galí, 2008).

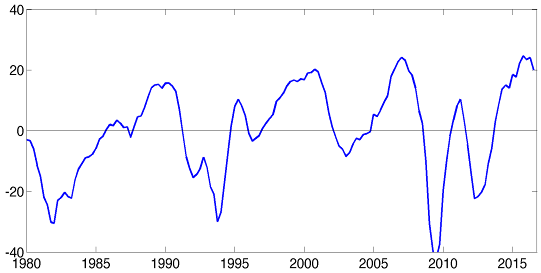
Figure 1: Residuals of money demand equations.



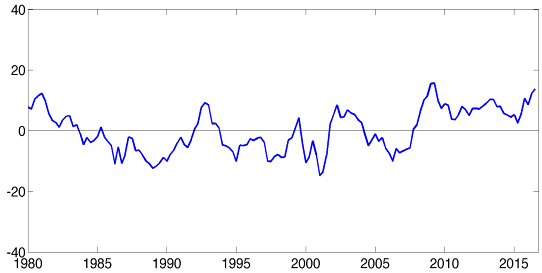
(a) Calza *et al.* (2001)



(b) Dreger and Wolters (2010b)



(c) De Bondt (2010)



(d) De Santis *et al.* (2013)

Residuals of money demand equations (percentages) computed according to the indicated references and rescaled to have zero mean.

the unemployment rate rises. Moreover, De Bondt (2010) shows that the annual change in unemployment is strongly correlated with consumer confidence sentiment. An improvement in labor markets conditions, normally associated to a higher GDP and, thus, to a higher demand for real balances, also lowers long-run precautionary demand for money. Second, De Santis *et al.* (2013) adopt an international portfolio allocation approach based on the assumption of wealth diversification among money, domestic and foreign risky assets. They augment the classical set of explanatory variables with price-earnings ratios, as proxies of Sharpe ratios in the EA and US markets, and EA and US long term interest rates.

When considering the money demand cointegration residual for those recent specifications, a more stable behavior is found (see Figures 1.b, 1.c, and 1.d). Furthermore, the recursive tests computed on the estimated coefficients can not reject the stability of the estimated long-run coefficients. However, since each model proposes a different money demand equation, i.e. different explanatory variables, it remains unclear which are the key variables of interest for the ECB in terms of policy. Thus, a comparison among different specifications is needed in order to investigate this issue more in depth. This is the first contribution of the paper and is based on a time-varying cointegration test, see Section 4 for details.

3 Data

Before presenting the empirical findings, we give a brief description of the data used in the analysis. Appendix A contains a detailed description of the sources and the construction of the variables used in this paper. Plots of the variables and of their first differences are available in the online supplementary appendix. For additional details we refer to the cited works.

Throughout the paper we use quarterly time series covering the period 1980:Q1–2016:Q3. The chosen sample allows for a careful evaluation of both the period in which money demand instability starts to emerge, before the Global Financial Crisis - that is "normal times" - and the aftermath of the Crisis, characterised by higher financial stress and low inflation, two factors which could exacerbate the unstable behaviour. Moreover, we are also enabled to draw some policy conclusions on the effects of the program of monthly asset purchases of private and public sector securities (Asset Purchases Program, APP) launched by the ECB at the end of 2014 to contrast risks of de-anchoring in inflation expectations. In particular, we evaluate the role of excess liquidity for predicting stock market busts and inflation.[5]

Monetary aggregates M3 and M2, loans to non financial corporations, housing and financial wealth are taken from the ECB Statistical Data Warehouse or ECB Monthly Bulletin. The own rate of M3 comes from ECB calculations. Output, total wealth, GDP deflator, unemployment rate, short and long-term interest rates are taken from the Area Wide Model Database (AWM) available from the Euro Area Business Cycle Network website. This allows us to avoid comparability problems due to revisions, extensions, and unavailability of some of the original data after 2001, and, furthermore, AWM is the standard source for macroeconomic studies on the EA (see, for example, Smets and Wouters, 2003). Finally, price-earnings ratios for EA and US economy, Dow Jones EuroStoxx50 and the German DAX30, used to backdate the latter until 1980:Q1 (Carstensen, 2006), are taken from Datastream, while the

[5] It has to be however noted that for the APP we only have two years of quarterly observations.

US long-term interest rate is downloaded by FRED St. Louis website.

A possible issue relates to EA data before 1998. Although these are artificial data, the literature has pointed out that EA aggregated data may result in poor approximations only before 1980 (see e.g. Bosker, 2006). Moreover, Beyer and Juselius (2010) show that cointegration estimates are robust to the choice of the aggregation method.

Regarding the stationary properties of the data used, we transform each time series in the same way as in the original models proposed by the literature in order to correctly replicate the analysis therein and perform the model evaluation exercise. When we move to estimate a new specification, we run the usual stationarity tests, in order to confirm the proper choice of the order of integration (see Appendix B for details).

4 Testing for time-varying cointegration

In this Section, we first present the theory and the empirical set-up for the Bierens and Martins (2010) methodology, and we then report the results of the test for time-varying cointegration applied to the specifications listed in Table 1.

4.1 Methodology

Using the same notation as in model (1), we define the vector $\mathbf{Z}_t := ((m_t - p_t), y_t, \mathbf{X}_t')'$ and we denote by k its dimension, while we indicate with T the sample size. Our starting point is the time-varying Vector Error Correction Model (VECM) of order p

$$\Delta \mathbf{Z}_t = \boldsymbol{\gamma}_0 + \boldsymbol{\alpha}\boldsymbol{\beta}_t'\mathbf{Z}_{t-1} + \sum_{j=1}^{p-1} \boldsymbol{\Gamma}_j \Delta \mathbf{Z}_{t-j} + \boldsymbol{\varepsilon}_t, \qquad \boldsymbol{\varepsilon}_t \overset{iid}{\sim} \mathrm{N}(\mathbf{0}, \boldsymbol{\Omega}), \qquad t = 1, \ldots, T, \tag{2}$$

where $\boldsymbol{\gamma}_0$ is $k \times 1$, $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}_t$ are $k \times r$, with $r \leq k$ being the cointegration rank, $\boldsymbol{\Gamma}_j$ and $\boldsymbol{\Omega}$ are $k \times k$. Notice that contrary to the standard VECM framework, in (2) the long-run coefficients $\boldsymbol{\beta}_t$ are allowed to change with time.

In Lemma 1 of their paper, Bierens and Martins (2010) prove that $\boldsymbol{\beta}_t$ can be approximated by a finite sum of Chebyshev polynomials $P_{h,T}(t)$ of decreasing smoothness:

$$\boldsymbol{\beta}_t = \sum_{h=0}^{m} \boldsymbol{\xi}_h P_{h,T}(t), \qquad P_{h,T}(t) = \sqrt{2}\cos\left(h\pi\frac{t-0.5}{T}\right), \quad h \geq 0, \quad t = 1,\dots T, \tag{3}$$

where $m = 1,\dots,T-1$ and $\boldsymbol{\xi}_h$ are $k \times r$ matrices. By substituting (3) into (2), we get

$$\Delta\mathbf{Z}_t = \boldsymbol{\gamma}_0 + \boldsymbol{\alpha}\left(\sum_{h=0}^{m} \boldsymbol{\xi}_h P_{h,T}(t)\right)' \mathbf{Z}_{t-1} + \sum_{j=1}^{p-1} \boldsymbol{\Gamma}_j \Delta\mathbf{Z}_{t-j} + \boldsymbol{\varepsilon}_t. \tag{4}$$

Therefore, since $P_{0,T}(t) = 1$, when $m = 0$, the time-varying VECM (4) becomes the usual time-invariant VECM as in Johansen (1996). Testing for time-invariant cointegration is equivalent to testing the null-hypothesis $m = 0$, against the alternative hypothesis $m > 0$, which implies time-varying cointegration. The two cases are therefore nested and can be estimated by maximum likelihood. Consequently, a likelihood ratio test for the null-hypothesis of time-invariant cointegration emerges naturally as

$$LR_T = 2\left[\hat{\ell}_T(r, m > 0) - \hat{\ell}_T(r, m = 0)\right], \tag{5}$$

where $\hat{\ell}_T(r,\cdot)$ are the sample Gaussian log-likelihoods computed in the estimated values of the parameters. In Theorem 1 of Bierens and Martins (2010) it is proved that, as $T \to \infty$, the statistics LR_T is distributed as $\chi^2_{(rmk)}$. Empirical critical values computed via Monte Carlo simulations are shown to be very close to the asymptotic critical values already for $T = 100$. Since in this paper we consider quarterly data from 1980:Q1 to 2016:Q3, i.e. $T = 147$, we compute p-values of the test based on the asymptotic distribution and its quantiles.[6] The methodology presented above is just one of the possible options for evaluating the time invariance of cointegration vectors. Other approaches are for example the classical recursive tests surveyed, among others, in Juselius (2006).[7] However, the test proposed by Bierens and Martins (2010) (*i*) extends to the multivariate case of VECM the similar test proposed by Park and Hahn (1999) and applied in Park and Park (2013), while (*ii*) it is particularly useful

[6]The minimum number of observations used is $T = 116$, when considering the "normal times" sample.

[7]For a Bayesian perspective on time-varying cointegration, which we do not consider here, see e.g. Koop *et al.* (2011).

for comparing different specifications in terms of stability, as it nests the time-invariant model under the null hypothesis. In this sense it allows us to perform a model evaluation and selection exercise similar to the Encompassing test principle by Mizon and Richard (1986) rather than just a test of structural change in the parameters. On the other hand, it is not available a procedure for determining r under model (2), hence we have to run the test conditional on different values of the cointegration rank r.

4.2 Results of the test

Before presenting the results of the test, few remarks are necessary. First, we compute the likelihood ratio test just for the case $m = 1$, i.e. when allowing for the first Chebyshev polynomial only as source of time variation. From (3), we notice that $m = 1$ corresponds to the slowest time-varying effect and moreover to the only monotonic change in the level of the coefficients. Both features are supported by previous studies where a slow change in the long-run money demand equation is observed, a behavior which in turn is mainly due to an upward shift in income elasticity as shown in Dreger and Wolters (2010b). Second, we report results for two different samples, both starting in 1980:Q1 and ending in (*i*) 2008:Q4, i.e. right before the recent Financial Crisis, and in (*ii*) 2016:Q3, i.e. including the most recent available data. Results on the samples originally considered by the works cited are available in the online appendix. Third, we show here only results for $p = 2$ lags in equation (2). Results for other values of p are qualitatively similar and are available in the online appendix. Finally, since no cointegration rank test is available in the time-varying framework, the whole procedure is repeated for $r = 1, 2$, and 3, thus including all the values of r considered in the cited literature (see Table 1).

Results of the test are in Tables 2 and 3. We report the likelihood-ratio statistics LR_T together with their corresponding p-value for the null-hypothesis of no time variation, jointly with the reference to the original work considered as defined in Table 1, and for each sample and cointegration rank considered.

Table 2 shows results for the classical specifications (Calza *et al.*, 2001, Coenen and Vega, 2001, Gerlach and Svensson, 2003, and Vlaar, 2004). Although those models were found to be

Table 2: LIKELIHOOD-RATIO TEST, LR_T, FOR STABILITY. PART I.

	1980:Q1-2008:Q4			1980:Q1-2016:Q3		
	$r=1$	$r=2$	$r=3$	$r=1$	$r=2$	$r=3$
CV	24.61	47.61	73.37	37.62	57.66	68.13
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
CGL	21.65	39.67	56.13	39.37	45.60	51.18
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
CGL	18.08	37.45	39.24	37.03	40.81	43.17
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
GS	17.43	34.77	35.34	34.30	41.71	45.19
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
V	15.79	32.81	52.50	33.80	49.26	56.78
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)

Values for the Bierens and Martins (2010) likelihood ratio test LR_T for the null-hypothesis of time-invariant cointegration, p-values in parenthesis. Results refer to a model with $p = 2$ lags. See Table 1 for details on each specification.

stable on pre-euro samples, when considering more recent samples as in this paper we reject the time-invariant null hypothesis for all specifications.

The first part of Table 3 presents results for those specifications claiming to find a stable relation once extending the set of explanatory variables to new motives related to financial and housing markets (Carstensen, 2006, Beyer, 2009, Dreger and Wolters, 2010a, and Dreger and Wolters, 2010b). Instability is always observed when extending the sample to the most recent years. This finding suggests the need for looking for other explanatory variables which might stabilize the long-run money demand equation.

Indeed, we gain more insights on the determinants of money demand when we consider those models based on labor markets (De Bondt, 2010) and international financial markets (De Santis *et al.*, 2013). Results are in the second part of Table 3. In the majority of the considered cases for these two specifications, the hypothesis of time-invariant cointegration is rejected. However, stable specifications are found when considering as explanatory variables: (i) the own rate of M3 and annual changes in the unemployment rate (labeled D3 in Table 3) and (ii) the spread between US and EA long-term rates and price-earnings ratios (DFR2 and DFR3 in Table 3). However, it has to be noted that the evidence of a stable money demand on *both* the pre-crisis (1980:Q1–2008:Q4) *and* the post-crisis (1980:Q1–2014:Q2) samples is very limited. Furthermore, although interesting, those cases in which we could not reject the null

Table 3: Likelihood-ratio test, LR_T, for stability. Part II.

	1980:Q1-2008:Q4			1980:Q1-2016:Q3		
	$r=1$	$r=2$	$r=3$	$r=1$	$r=2$	$r=3$
C	26.13	49.05	64.22	36.16	42.80	53.39
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
B	25.26	49.77	64.51	27.12	60.77	72.14
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
B	21.47	27.90	31.13	23.28	34.50	39.03
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
B	20.57	33.49	49.27	17.35	35.97	47.89
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
DW1	20.31	33.33	38.94	59.96	107.72	130.60
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
DW2	30.43	65.45	88.68	20.17	29.58	33.71
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
DW3	25.15	44.60	58.99	18.76	33.72	39.03
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
DW4	16.64	43.37	58.98	36.54	51.90	66.99
	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
D1	21.00	39.25	67.29	18.66	60.27	85.99
	(0.03)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
D2	17.22	43.30	73.22	36.02	54.94	83.16
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
D3	13.45	53.58	68.90	3.04	34.76	42.53
	(0.01)	(0.00)	(0.00)	(0.55)	(0.00)	(0.00)
DFR1	16.15	45.88	79.75	28.94	48.01	73.20
	(0.02)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
DFR2	8.32	34.92	54.41	20.10	44.19	65.55
	(0.22)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
DFR3	11.70	38.37	53.88	15.41	27.58	39.93
	(0.02)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)

Values for the Bierens and Martins (2010) likelihood ratio test LR_T for the null-hypothesis of time-invariant cointegration, p-values in parenthesis. Results refer to a model with $p = 2$ lags. See Table 1 for details on each specification.

hypothesis of time-invariant long-run coefficients of money demand occur only when $r = 1$, a fact which poses problems of interpretation as the specifications considered have usually more than one cointegrating relation, namely $r = 1$ or 2 in De Bondt (2010), and $r = 3$ in De Santis *et al.* (2013).

The results described help us in discriminating among the possible explanatory variables of money demand used in the literature, but do not provide a definitive answer to our main question: which are the crucial determinants (if any) for observing a time-invariant long-run money demand equation in the EA?

5 A new model: Money, stock markets and unemployment

In this Section we build on the evidence provided by the model comparison exercise to propose and estimate a new stable specification for EA money demand, which is then used in Section 6 to draw policy implications for the ECB.

5.1 Stability analysis

Starting from the results in Table 3, we propose a novel model accounting for different motives for holding money. In particular, our specification includes income (y_t), price-earnings ratios of both EA and US (q_t and q_t^* respectively), the annual change in the unemployment rate (Δu_t), the quarterly inflation rate (π_t), the long-term rate (ℓ_t), and the M3 own rate (o_t).[8]

We thus test model (2) for time-invariant cointegration when $\mathbf{Z}_t = ((m_t - p_t), y_t, \Delta u_t, (q_t - q_t^*), (\ell_t - o_t), \pi_t)'$. The null hypothesis can not be rejected with rank $r = 1, 2$, and 3, thus validating the time-invariant approach (see the first row of Table 4). Similar stability results hold for possible alternative novel specifications, based on variables such as housing wealth combined to price-earnings ratios, inflation and interest rates (see the other rows of Table 4).[9]

It has to be noticed that when extending the sample until 2016:Q3, we have a small p-value of 0.01 for $r = 3$, a finding probably due to very recent data. Indeed, when running the test over different possible end dates subsequent to 2008:Q4, we can not reject the stability of our model up to 2014:Q2 for $r = 3$ with a larger p-value equal to 0.06. This is not surprising since during the summer of 2014 risks of deflation increased, as a consequence of a stagnant economy and the collapse of oil prices (Ciccarelli and Osbat, 2017; Conti *et al.*, 2017). When the economy hit the Zero Lower Bound (ZLB) of interest rates, in September 2014, the ECB was then forced to undertake a broad set of unconventional measures, such as pushing its deposit facility rate into negative territory (-0.20%) and launching a program of monthly asset purchases of private and public sector securities (APP). Therefore, it is likely that this new regime may alter the dynamic relations between money, interest rates, financial variables

[8] We use annual changes in unemployment rather than quarterly for consistency with the choice by De Bondt (2010). Notice that stationarity tests suggest indeed the presence of a unit root (see Table B-1 for details).

[9] Using the short-term rate instead of the own rate of money leaves results unaffected because of the high correlation between the two (De Santis *et al.*, 2013; Giannone *et al.*, 2012).

Table 4: LIKELIHOOD-RATIO TEST, LR_T, FOR STABILITY. PART III.

	1980:Q1-2008:Q4			1980:Q1-2016:Q3		
	$r=1$	$r=2$	$r=3$	$r=1$	$r=2$	$r=3$
$\Delta u_t, (q_t - q_t^*), (\ell_t - o_t), \pi_t$	12.28	16.34	36.68	9.83	19.68	35.45
	(0.11)	(0.41)	(0.11)	(0.13)	(0.07)	(0.01)
$\Delta u_t, (q_t - q_t^*), (\ell_t - o_t), \pi_t, h_t$	11.54	38.33	53.96	9.10	23.46	49.18
	(0.12)	(0.00)	(0.00)	(0.25)	(0.05)	(0.00)
$(q_t - q_t^*), (\ell_t - o_t), \pi_t$	7.02	23.70	43.71	7.41	13.96	23.19
	(0.22)	(0.01)	(0.00)	(0.19)	(0.17)	(0.08)
$\Delta u_t, (\ell_t - o_t), \pi_t$	8.15	19.49	37.02	9.96	21.13	35.61
	(0.15)	(0.03)	(0.00)	(0.08)	(0.02)	(0.00)
$(q_t - q_t^*), (\ell_t - o_t), \pi_t, h_t$	17.52	32.26	47.16	12.28	36.03	47.38
	(0.01)	(0.00)	(0.00)	(0.06)	(0.00)	(0.00)
$\Delta u_t, (\ell_t - o_t), \pi_t, h_t$	9.60	45.36	56.07	11.12	25.39	39.63
	(0.14)	(0.00)	(0.00)	(0.08)	(0.01)	(0.00)

Values for the Bierens and Martins (2010) likelihood ratio test LR_T for the null-hypothesis of time-invariant cointegration, p-values in parenthesis. Results refer to a model with $p = 2$ lags. See Table 1 for details on each specification.

and the real economy in such a way that a *linear, time-invariant* model can not capture. Consistently with these results and given the peculiar behaviour of the economy in the very recent years, we then choose to estimate our time-invariant VECM until 2014:Q2, leaving the last observations as the out-of-sample period used in one of our forecasting exercises.

Let us stress that our model is the only one robust to the beginning of the crisis, as for all the other models we evaluated the hypothesis of time-invariant cointegration is rejected. Beyond the new test developed by Bierens and Martins (2010), also the other usual stability tests based on the recursive estimation of the parameters (Juselius, 2006) support the stability of our proposed model (see below).

5.2 Estimation settings and diagnostics

Once assessed the stability of our model, we can move to estimate the time-invariant version of (2), when the long-run coefficients $\boldsymbol{\beta}$ do not depend on time, i.e.

$$\Delta \mathbf{Z}_t = \boldsymbol{\gamma}_0 + \boldsymbol{\alpha}\boldsymbol{\beta}'\mathbf{Z}_{t-1} + \sum_{j=1}^{p-1} \boldsymbol{\Gamma}_j \Delta \mathbf{Z}_{t-j} + \boldsymbol{\varepsilon}_t, \qquad \boldsymbol{\varepsilon}_t \overset{iid}{\sim} \mathrm{N}_k(\mathbf{0}, \boldsymbol{\Omega}), \qquad t = 1, \ldots, T, \tag{6}$$

$$\mathbf{Z}_t = ((m_t - p_t),\ y_t,\ \Delta u_t,\ (q_t - q_t^*),\ (\ell_t - o_t),\ \pi_t)'.$$

Table 5: TESTS FOR DETERMINING THE COINTEGRATION RANK r.

$k-r$	r	Eig	Trace	$c_{95\%}$	p-val	p-val*
			1980:Q1–2008:Q4			
6	0	0.422	180.842	95.514	0.00	0.00
5	1	0.412	118.380	69.611	0.00	0.00
4	2	0.235	57.781	47.707	0.00	0.03
3	3	0.134	27.206	29.804	0.10	0.24
2	4	0.090	10.856	15.408	0.22	0.56
1	5	0.001	0.164	3.841	0.69	0.78
			1980:Q1–2014:Q2			
6	0	0.433	191.885	95.514	0.00	0.00
5	1	0.310	114.610	69.611	0.00	0.00
4	2	0.224	64.100	47.707	0.00	0.01
3	3	0.130	29.241	29.804	0.06	0.14
2	4	0.056	9.831	15.408	0.29	0.46
1	5	0.020	2.060	3.841	0.15	0.21

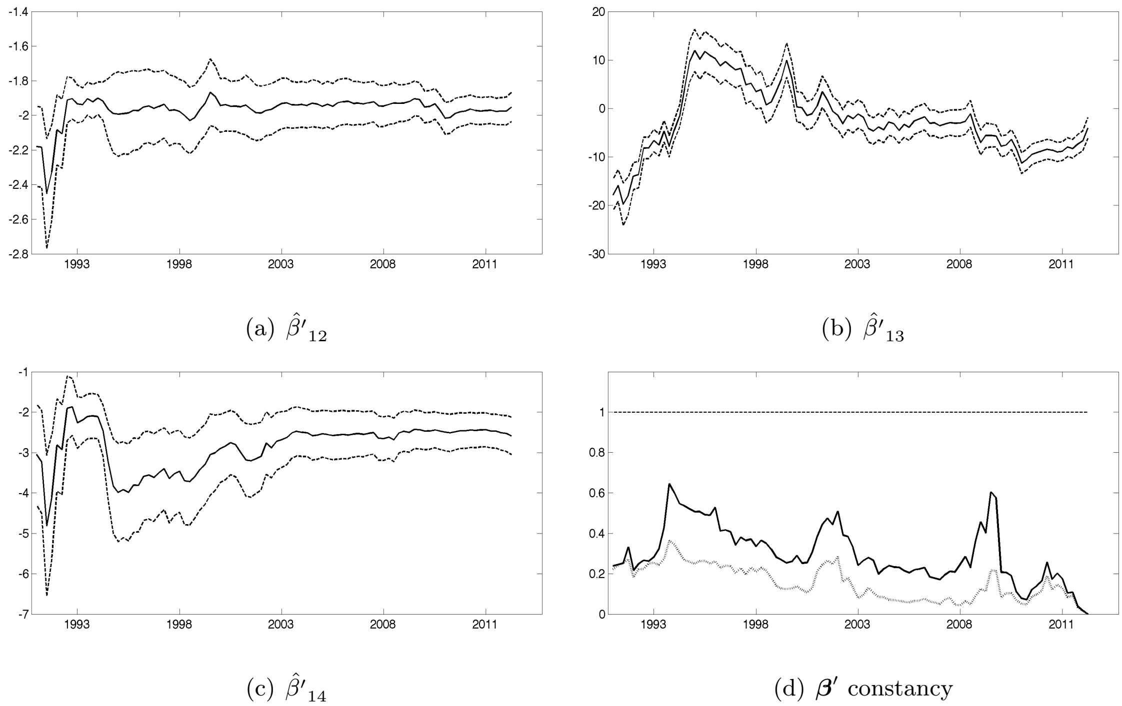
Values of Johansen largest eigenvalue and trace test computed for model (6); $c_{95\%}$: 95% critical value. p-val* denotes the small-sample corrected p-value.

Estimation of (6) can be carried out as in Stock and Watson (1993) or Johansen (1996). Here we adopt the latter approach, based on maximum likelihood, for consistency with the Bierens and Martins (2010) approach and for comparability with other studies of EA money demand. When estimating the model, we consider an unrestricted constant, thus allowing for an intercept in the cointegration space and for trending variables, but not for a trend in the cointegrating vectors.[10] We set the spread between price-earnings ratios and between long interest rate and the return on M3 to have unit coefficients since both restrictions are strongly supported by data. The size of the model is $k = 6$, and given the discussion above, we comment here only the results for the sample 1980:Q1–2008:Q4, i.e. $T = 116$, and for the sample 1980:Q1–2014:Q2, i.e. $T = 138$; the number of lags is $p = 2$ in both cases. In particular, according to the Johansen trace test, the hypothesis of three cointegrating vectors can not be rejected at the 10% level and we thus select $r = 3$ (see Table 5).

As an additional check for the stability of our specification, we also run the usual recursive tests on stability of the parameters of model (6). We show in Figure 2 the recursive estimates of

[10] Presently the LR_T test by Bierens and Martins (2010) is available for the drift case only. Moreover, when estimating the time-invariant model allowing for the trend in cointegration space, results are unaffected, as for the sample 1980:Q1–2008:Q4 the exclusion of the trend from the system can not be rejected by the test $\chi^2_{(7)} = 6.177$, with a corresponding p-value of 0.52. The same statistics for the longer sample 1980:Q1-2014:Q2 is $\chi^2_{(7)} = 8.896$, with a p-value of 0.26.

Figure 2: Recursive estimated money demand coefficients and Nyblom test.



Panels (a)-(c): Recursively estimated coefficients of the money demand equation (7) (initial sample: 1980:Q1–1989:Q4). Solid line: recursive estimation of the identified long-run coefficients; dashed line: 95% confidence interval. Panel (d): Nyblom test of constant long-run parameters. Values greater than 1 imply rejection of the null-hypothesis of constant log-likelihood at 5% level. Black solid line: test for the full model; light grey line: test for for the model conditional on the short-run dynamics, i.e. on the $\boldsymbol{\Gamma}_j$ matrices.

the estimated and identified cointegration vectors of the money demand equation (see equation (7) below) and the Nyblom test of constant long-run parameters. All these plots are built using as initial estimation sample the period 1980:Q1–1989:Q4.

Finally, the standard battery of diagnostics tests are reported in Appendix B. A small degree of heteroskedasticity and non-normality is detected in the residuals, the first being related to price-earnings ratios, while the latter essentially due to a large residual in M3 in 1993:Q2 and corresponding to the European Monetary System crisis. However, Cavaliere *et al.* (2010) show that cointegration estimates are robust to conditional heteroskedasticity. Furthermore, we show that introducing dummy variables to account for institutional events responsible for these fluctuations (see e.g. Dreger and Wolters, 2010b) removes the problem of non-normality leaving unaffected the whole set of results here presented (see Appendix B

Table 6: Estimates of the identified long run structure - $\widehat{\boldsymbol{\beta}}'$.

	$(m_t - p_t)$	y_t	Δu_t	$(q_t - q_t^*)$	$(\ell_t - o_t)$	π_t
		1980:Q1–2008:Q4				
$\widehat{\boldsymbol{\beta}}_1$	1.000 [n.a.]	−1.893 [−38.345]	−14.478 [−16.613]	−2.052 [−11.477]	0.000 [n.a.]	0.000 [n.a.]
$\widehat{\boldsymbol{\beta}}_2$	0.000 [n.a.]	0.000 [n.a.]	1.000 [n.a.]	0.000 [n.a.]	−0.099 [−3.192]	0.000 [n.a.]
$\widehat{\boldsymbol{\beta}}_3$	0.000 [n.a.]	0.000 [n.a.]	0.000 [n.a.]	0.000 [n.a.]	1.000 [n.a.]	−0.560 [−10.758]
		1980:Q1–2014:Q2				
$\widehat{\boldsymbol{\beta}}_1$	1.000 [n.a.]	−1.953 [−45.468]	−4.036 [−3.682]	−2.588 [−10.931]	0.000 [n.a.]	0.000 [n.a.]
$\widehat{\boldsymbol{\beta}}_2$	0.000 [n.a.]	0.000 [n.a.]	1.000 [n.a.]	0.000 [n.a.]	−0.145 [−4.143]	0.000 [n.a.]
$\widehat{\boldsymbol{\beta}}_3$	0.000 [n.a.]	0.000 [n.a.]	0.000 [n.a.]	0.000 [n.a.]	1.000 [n.a.]	−0.546 [−12.069]

Estimated long run coefficients $\widehat{\boldsymbol{\beta}}$, t-statistics in parenthesis.

for details).[11] Overall, then, the results confirm a reasonable fit for our specification and in the following we stick to the case with no dummies.

5.3 Long-run structure

We identify the three cointegrating relations as: (i) a long-run money demand equation, (ii) an aggregate demand equation, and (iii) a relation between the yield curve and inflation. Table 6 presents the restricted estimates of the long-run coefficients, $\widehat{\boldsymbol{\beta}}' = (\widehat{\boldsymbol{\beta}}_1, \widehat{\boldsymbol{\beta}}_2, \widehat{\boldsymbol{\beta}}_3)'$.

The likelihood ratio test for the restricted model is distributed as $\chi^2_{(4)}$, and is equal to 3.737 and 1.627, giving a p-value of 0.44 and 0.80 for the samples 1980:Q1–2008:Q4 and 1980:Q1–2014:Q2, respectively, thus finding no evidence against the imposed restrictions.

Our focus is in particular on the long-run money demand equation, which on the two

[11] When extending the sample up to 2014:Q2 it is more difficult to accept the hypothesis of multivariate normality, especially because of kurtosis of output residuals. We have then to allow for a Global Financial Crisis dummy variable (2008:Q3–2009:Q4). However, Gonzalo (1994) shows that the maximum likelihood estimates of cointegrating vectors are robust to non-normality. Furthermore, we have also re-estimated our money demand equation by Fully Modified OLS (Phillips and Hansen, 1990)and by Dynamic OLS (Stock and Watson, 1993), and we achieve very similar estimates for the long-run coefficients of interest.

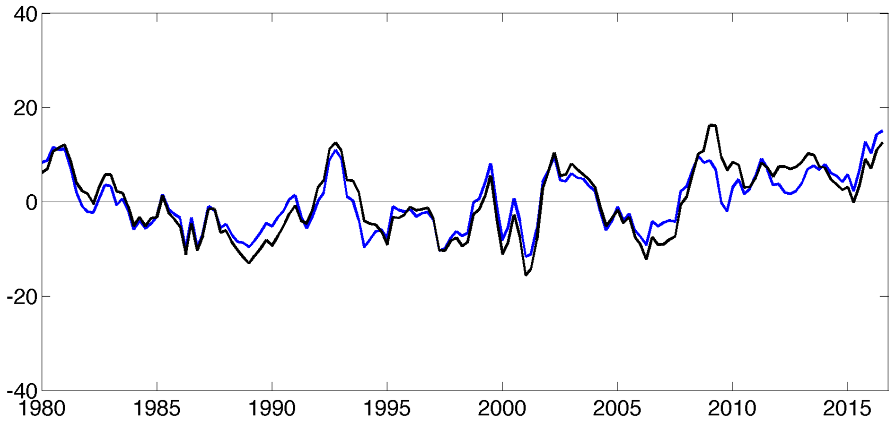
samples considered is estimated as (t-statistics in parenthesis):

$$\begin{array}{lcllllllll}
(m_t - p_t) & = & \widehat{\beta}'_{12} y_t & + & \widehat{\beta}'_{13} \Delta u_t & + & \widehat{\beta}'_{14}(q_t - q^*_t) & & & \\
(m_t - p_t) & = & 1.893\ y_t & + & 14.478\ \Delta u_t & + & 2.052\ q_t & - & 2.052 q^*_t. & (1980{:}Q1\text{-}2008{:}Q4) \\
 & & [38.345] & & [16.613] & & [11.477] & & [-11.477] & \\
(m_t - p_t) & = & 1.953\ y_t & + & 4.052\ \Delta u_t & + & 2.588\ q_t & - & 2.588 q^*_t. & (1980{:}Q1\text{-}2014{:}Q2) \\
 & & [45.419] & & [3.682] & & [10.931] & & [-10.931] &
\end{array} \tag{7}$$

All the estimated parameters are strongly significant and display the expected sign. Real money holdings are a positive function of income, with a coefficient slightly lower than 2, consistently with the *wealth* effect found by the literature (De Santis *et al.*, 2013; Dreger and Wolters, 2010b). Also, they increase with the annual change of unemployment, which represents the *precautionary* motive and reflecting the uncertainty about the state of the economy (De Bondt, 2010). Finally, real balances increase with the spread between EA and US price-earnings ratios, i.e. the *speculative* motive (De Santis *et al.*, 2013). Thus, an increase in the relative convenience of EA assets produces a higher demand of euro from abroad, a finding consistent with the idea of a consolidation of the euro as reserve currency, which would explain the emergence of the speculative motive only after 2001. It has to be noticed that, when estimating the VECM up to 2014:Q2 the long-run structure is broadly consistent with the one estimated until 2008:Q4. The only relevant change is related to the lower size of the precautionary motive coefficient, about one third with respect to 2008:Q4 estimation. One possible interpretation is that the ZLB and the launch of the APP by the ECB, driving to an expansion of economic activity and financial assets, de-potentiated the insurance role of hoarding money. However, this requires further analysis, possibly with a non-linear setup, which is beyond the scope of this paper.

Few remarks regarding the other two cointegrating vectors are necessary. Firstly, $\widehat{\boldsymbol{\beta}}_2$ displays a relation between the interest rates spread and the change in unemployment. This is a long-run aggregate demand equation in which unemployment is linked to the slope of the yield curve, i.e. to $(\ell_t - o_t)$. Secondly, $\widehat{\boldsymbol{\beta}}_3$ represents the link between the yield curve and inflation. The latter somewhat differs from the standard Fisher Interest Parity observed,

Figure 3: Residual of money demand equation and excess money.



Blue line: residual of money demand equation $\widehat{\eta}_{1,t}$. Black line: excess money $\widehat{\eta}_t$ obtained by solving the system of cointegration equations in Table 6 with zero residuals. The estimated coefficients are obtained over the sample 1980:Q1–2014:Q2, then used for illustrative purposes also in 2014:Q3–2016:Q3.

among others, in Coenen and Vega (2001) and Vlaar (2004) on samples ending before the launch of the euro. The rationale for such a long-run equation is that the yield curve is a relevant predictor of inflation. Among others, we refer to Estrella (2005), who shows that if monetary policy is essentially reactive to deviations of inflation from target and of output from potential, the predictive relationships for prices depend primarily on the magnitudes of the reaction parameters. Furthermore, similar relations for US economy are found for example in Beyer and Farmer (2007). Although beyond the scope of this paper, which focuses on money demand, a deeper analysis of the last two cointegrating relations could give some useful insights on the sacrifice ratio and the natural rate of unemployment in the EA, additionally helping for a careful inclusion of the yield curve in a macroeconomic model.

To conclude the analysis of the long-run, we show in blue in Figure 3 the cointegration residual of the money demand equation (7), defined as $\widehat{\eta}_{1,t} = \widehat{\boldsymbol{\beta}}_1'\mathbf{Z}_t$. We observe the expected mean reverting behavior, which qualitatively confirms the goodness of fit of our model and its stability. However, when considering a system with more than one cointegration relations, equilibrium money should be obtained by solving the full system (with zero residuals) and the excess money or excess liquidity is computed as deviation from this measure. In our policy applications in Section 6 we adopt this definition, which we denote by $\widehat{\eta}_t$, as opposed to $\widehat{\eta}_{1,t}$,

Table 7: ESTIMATES OF THE IDENTIFIED SHORT RUN DYNAMICS - $\widehat{\boldsymbol{\alpha}}$.

	$\widehat{\boldsymbol{\alpha}}_1$	$\widehat{\boldsymbol{\alpha}}_2$	$\widehat{\boldsymbol{\alpha}}_3$	$\widehat{\boldsymbol{\alpha}}_1$	$\widehat{\boldsymbol{\alpha}}_2$	$\widehat{\boldsymbol{\alpha}}_3$
		1980:Q1–2008:Q4				
$\Delta(m_t - p_t)$	-0.002	-0.441	-0.265	-0.001	-0.447	-0.246
	[-0.157]	[-2.257]	[-3.261]	[-5.107]	[-4.033]	[-3.292]
Δy_t	-0.047	-0.588	0.068	-0.042	-0.607	0.075
	[-4.025]	[-4.073]	[1.135]	[-4.247]	[-5.160]	[1.261]
$\Delta\Delta u_t$	0.013	0.000	-0.002	0.013	-0.003	0.004
	[4.653]	[0.003]	[-0.131]	[4.451]	[-0.086]	[-0.317]
$\Delta(q_t - q_t^*)$	0.220	2.811	0.469	0.199	2.625	0.366
	[6.029]	[6.205]	[2.492]	[5.600]	[5.857]	[2.114]
$\Delta(\ell_t - o_t)$	0.007	0.166	-0.082	-0.000	0.039	-0.085
	[0.896]	[1.654]	[-1.970]	[-2.023]	[0.741]	[-2.459]
$\Delta\pi_t$	0.055	0.404	0.403	0.038	0.141	0.345
	[2.167]	[1.285]	[3.084]	[1.860]	[0.563]	[2.721]
		1980:Q1–2014:Q2				
$\Delta(m_t - p_t)$	-0.017	-0.396	-0.242	-0.001	-0.439	-0.225
	[-1.603]	[-4.306]	[-3.661]	[-4.860]	[-5.408]	[-3.845]
Δy_t	-0.029	-0.0053	0.105	-0.033	-0.163	0.144
	[-3.233]	[-0.689]	[1.891]	[-3.029]	[-2.908]	[1.713]
$\Delta\Delta u_t$	0.011	-0.167	-0.013	0.011	-0.156	-0.011
	[4.341]	[-7.865]	[-0.872]	[4.508]	[-9.583]	[-0.960]
$\Delta(q_t - q_t^*)$	0.104	0.182	0.131	0.101	0.122	0.042
	[4.889]	[1.011]	[1.011]	[4.407]	[0.833]	[0.385]
$\Delta(\ell_t - o_t)$	-0.006	0.090	-0.116	-0.003	0.063	-0.112
	[-0.997]	[1.817]	[-3.256]	[-0.492]	[1.276]	[-5.042]
$\Delta\pi_t$	0.037	-0.015	0.358	0.001	-0.001	0.356
	[2.129]	[-0.100]	[3.390]	[3.866]	[-0.011]	[3.568]

Estimated short-run coefficients $\widehat{\boldsymbol{\alpha}}$, t-statistics in parenthesis. Left panel: estimates of unrestricted model; right panel: same estimates when eliminating the non-significant coefficients $\widehat{\boldsymbol{\Gamma}}_j$.

to highlight that it comes from the full system. A plot of $\widehat{\eta}_t$ is in black in Figure 3.

5.4 Short-run structure

Moving to the short-run dynamics, the left panel of Table 7 shows that most variables respond to the three cointegration residuals $\widehat{\eta}_{1,t}$, $\widehat{\eta}_{2,t}$, and $\widehat{\eta}_{3,t}$, for both the considered periods. In particular, money is strongly error correcting with respect to the second and the third cointegration relation ($\widehat{\alpha}_{12}$ and $\widehat{\alpha}_{13}$), whereas the feedback coefficient obtained for the disequilibrium in money holdings ($\widehat{\alpha}_{11}$) has the expected negative sign, but is not statistically different from zero. This result is also found by Beyer (2009), and requires a deeper analysis.

In order to study the role of money in the short-run, we first constrain to zero the coefficient of $(m_t - p_t)$, thus imposing the normalization on y_t. The $\chi^2_{(7)}$ test of exclusion of money has a value of 43.893 and 23.360 on the periods 1980:Q1–2008:Q4 and 1980:Q1–2014:Q2, respectively, and money redundancy is strongly rejected at 1% level. Second, we test the hypothesis $\alpha_{11} = \alpha_{12} = \alpha_{13} = 0$ to assess the possibility of real balances not reacting to any of the estimated cointegration equations, hence being weakly exogenous: again, the test rejects the hypothesis at 1% level, with $\chi^2_{(7)} = 19.746$ for the period 1980:Q1–2008:Q4 and $\chi^2_{(7)} = 29.663$ when extending the analysis up to 2014:Q2. Moreover, as in Beyer (2009), we re-estimate the model with fixed values for $\widehat{\boldsymbol{\beta}}$, as given in (7), and by iteratively erasing non–significant variables, i.e. fixing at zero the correspondent elements of $\widehat{\boldsymbol{\Gamma}}_j$ in (6). Results are reported in the right panel of Table 7 and we find a small but highly significant coefficient for the error correction term of money.

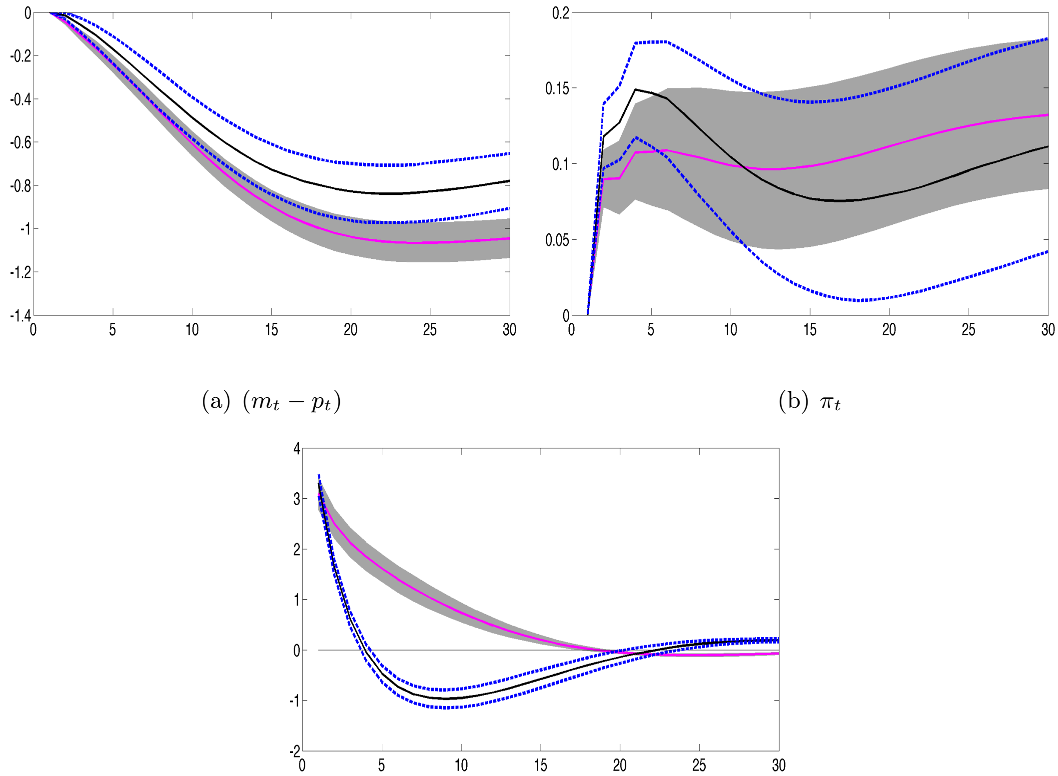
When considering other variables, income is strongly error-correcting with respect to both the first and the second cointegration vector ($\widehat{\alpha}_{21}$ and $\widehat{\alpha}_{22}$), i.e. the money demand equation and the aggregate demand relation. Unemployment is pulled away from its equilibrium with respect to changes in money demand ($\widehat{\alpha}_{31}$), while the spread between the long-term rate and the own rate of M3 is adjusting to disequilibrium in the third relation ($\widehat{\alpha}_{53}$), the one between interest rates and inflation. Last, inflation seems not error-correcting in neither of the above relations. Notice, however, that $\widehat{\alpha}_{61}$ is positive and significant, meaning that inflation is pushed up by deviation of real balances from their long-run equilibrium.

Finally, we can compute the dynamic adjustment of variables of interest at different horizons (see Fanelli and Paruolo, 2010). Following Pesaran and Shin (1998) we estimate the generalized impulse response (GIRF) of real balances to $\widehat{\eta}_t$. The result, obtained from a VAR(2) using as variables $\mathbf{Z}_t$ and $\widehat{\eta}_t$, is in Figure 4.a and confirms a significant negative reaction at all considered lags. It is worth mentioning that the generalized impulse responses to the money disequilibrium $\widehat{\eta}_t$ look fairly stable across the two considered samples, i.e. before and after 2008:Q4. The exception is related to the reaction of $\widehat{\eta}_t$ to itself: it is more persistent when considering data up to 2014:Q2, possibly reflecting the sustained increase in liquidity due to unconventional ECB policies.[12]

Notice also that following a shock to excess liquidity inflation only modestly rises, as one can infer from considering the generalized impulse response of inflation to money disequilibrium, shown in Figure 4.b.

[12]When considering the GIRFs from a VECM (which includes all three cointegration residuals $\widehat{\eta}_{1t}$, $\widehat{\eta}_{2t}$, and $\widehat{\eta}_{3t}$) instead of using $\widehat{\eta}_t$ results are very similar.

Figure 4: GENERALIZED IMPULSE RESPONSES TO THE MONEY DISEQUILIBRIUM.



(a) $(m_t - p_t)$ (b) π_t

(c) $\widehat{\eta}_t$

Black lines: generalized impulse responses for the period 1980:Q1–2008:Q4 to a 3% shock in excess liquidity, cumulated effects, together with their 68% confidence intervals obtained with 1000 Monte Carlo replications (dashed blue lines). Magenta lines: generalized impulse responses for the period 1980:Q1–2014:Q2 to a 3% shock in excess liquidity, cumulated effects, together with their 68% confidence intervals obtained with 1000 Monte Carlo replications (light grey shaded area).

5.5 Prediction

To further validate our model we run two forecasting exercises considering in-sample and out-of-sample predicitons, analogously to De Santis *et al.* (2013) and Giannone *et al.* (2012). All predictions are computed conditional on two observed samples 1980:Q1–2008:Q4 and 1980:Q1–2014:Q4. In particular, we compute one-step-ahead conditional expectations obtained from estimating a VAR(2) using all the variables $\mathbf{Z}_t$ in (6) together with our estimated measure of excess money $\widehat{\eta}_t$, which is computed by solving the system of long-run equations and putting

Table 8: PREDICTIONS OF REAL M3 AND GDP Y-O-Y GROWTH RATES.

Estimation Sample	Prediction Sample		BC	DB	DFR
1980:Q1–2008:Q4	1980:Q1–2008:Q4	$\Delta_4(m_t - p_t)$	1.000	1.115	1.019
		$\Delta_4 y_t$	1.000	1.180	1.066
1980:Q1–2014:Q2	1980:Q1–2014:Q2	$\Delta_4(m_t - p_t)$	1.000	1.043	1.035
		$\Delta_4 y_t$	1.000	1.058	1.063
1980:Q1–2008:Q4	2009:Q1–2010:Q4	$\Delta_4(m_t - p_t)$	1.000	2.016	3.612
		$\Delta_4 y_t$	1.000	2.674	2.103
1980:Q1–2004:Q4	2005:Q1–2010:Q4	$\Delta_4(m_t - p_t)$	1.000	1.299	0.849
		$\Delta_4 y_t$	1.000	2.535	0.630
1980:Q1–2008:Q4	2009:Q1–2016:Q3	$\Delta_4(m_t - p_t)$	1.000	1.065	1.045
		$\Delta_4 y_t$	1.000	0.950	0.982
1980:Q1–2004:Q4	2005:Q1–2016:Q3	$\Delta_4(m_t - p_t)$	1.000	0.970	1.042
		$\Delta_4 y_t$	1.000	1.069	1.038
1980:Q1–2014:Q2	2014:Q3–2016:Q3	$\Delta_4(m_t - p_t)$	1.000	0.795	0.879
		$\Delta_4 y_t$	1.000	0.677	0.817

RMSEs of De Bondt (2010) (DB) and De Santis *et al.* (2013) (DFR) specifications relative to our model (BC). Top panel: in-sample results; bottom panel: one-step-ahead conditional predictions.

the residuals to zero.[13]

In Table 8 we show the root-mean-squared-errors (RMSE) to compare our predictions with those obtained by estimating the models by De Santis *et al.* (2013) and De Bondt (2010). The estimated or predicted paths of real M3 and GDP growth are shown in Figure 5.

First, we compare the actual annual growth rates with those predicted by our model for the periods 1980:Q1–2008:Q4 and 1980:Q1–2014:Q2, i.e. the full sample estimation. Results confirm the ability of our model in describing the dynamics not only of M3 but also of GDP growth rates. Second, we compute one-step-ahead forecasts for different sub-samples of interest: (*i*) a period of deep crisis, i.e. the one immediately subsequent the Great Recession, 2009:Q1–2010:Q4 and (*ii*) 2009:Q1–2016:Q3, a period also including the subsequent events, such as the Sovereign Debt Crisis, the unconventional monetary policies undertaken by the ECB, and the recent low inflation period.

[13] Recall that $\widehat{\eta}_t$ by construction implies $\widehat{\eta}_{2t} = \widehat{\eta}_{3t} = 0$. Similar results to those presented are achieved when forecasting by means of the VECM (6), which instead of $\widehat{\eta}_t$, includes all three cointegration residuals $\widehat{\eta}_{1t}$, $\widehat{\eta}_{2t}$, and $\widehat{\eta}_{3t}$.

Even in a period of crisis as the one observed in recent years, our model mimics fairly well the behaviour of the observed variables (see Figure 5). In particular, our model almost always outperforms those of De Bondt (2010) and De Santis *et al.* (2013), especially in predicting $m_t - p_t$. These findings also hold when estimating the model from 1980:Q1 to 2004:Q4 and computing one-step-ahead predictions for the period 2005:Q1–2010:Q4 and 2005:Q1–2016:Q3 (see the bottom panel of Figure 5). This confirms the goodness of our model, since it is able to forecast quite well the evolution of M3 and real GDP even when stopping the estimation sample well before the insurgence of the crisis, i.e. not exploiting dynamic correlations already incorporating financial markets turmoils. By contrast, on the very last part of the sample – 9 quarters, from 2014:Q3 to 2016:Q3 – the performance of our model deteriorates relatively to those by De Bondt (2010) and De Santis *et al.* (2013). This is probably related to the inclusion in the latter models of variables linked to US monetary policy. At the ZLB monetary policy is *globally* conducted by means of unconventional actions which mainly affect the term structure and stock prices. Hence, having in the model also US monetary policy variables could bear a superior predictive content for EA real GDP – and thus M3 – in the most recent period However, while acknowledging that our model is outperformed in the prediction exercise ran over the last 9 available quarters, we also stress that the models by De Bondt (2010) and De Santis *et al.* (2013) are not stable over this time span.

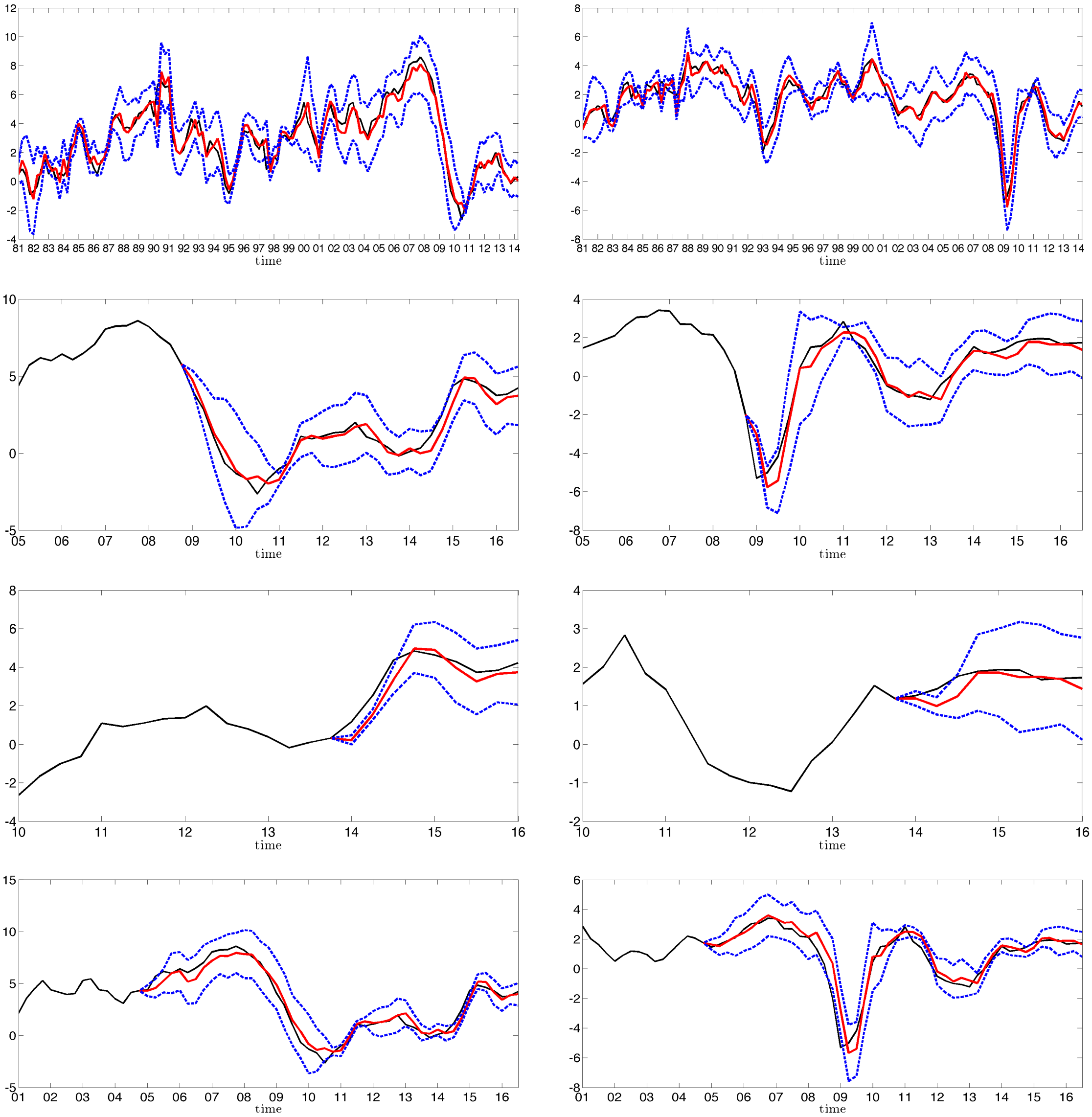
6 Money demand and monetary policy

In this last section, we evaluate the usefulness of our estimates for policy makers. We consider two related topics: (*i*) the role of long-run money demand in policies oriented to financial stability and (*ii*) the impact of excess liquidity on inflation.

6.1 Financial stability

The recent Global Financial Crisis has triggered an increasing debate on the central bank mandate with respect to the pursue of financial stability jointly with price stability (see, e.g. Borio, 2006; Buiter, 2012, and references therein). In this first exercise, we evaluate the

Figure 5: Predictions of real M3 and GDP growth rates.



(a) Real M3 growth $100\Delta_4(m_t - p_t)$

(b) GDP growth $100\Delta_4 Y_t$

Predictions of our model. Black line: percentage annualized growth rates of observed series; red line: in-sample conditional expectations for the period 1980:Q1-2016:Q3 (top panel) and out-sample one-step-ahead predictions for the period 2009:Q1-2016:Q3 (second panel), 2014:Q3-2016Q3 (third panel) and 2005:Q1-2016:Q43 (bottom panel); 95% confidence intervals are also shown (blue dashed lines).

performance of our measure of excess liquidity $\widehat{\eta}_t$ in predicting financial crises. To this aim, we build a binary indicator of stock prices busts, B_t, and we compare different explanatory variables for the probability of a bust at time t. The indicator B_t is based on the quarterly growth rates of equity Δe_t and of house prices Δf_t.[14] The binary variable indicating stock market busts is defined as follows

$$B_t = \begin{cases} 1 & \text{if } (\Delta e_t < \mu_t^e - \delta\, \sigma_t^e) \text{ or } (\Delta f_t < \mu_t^f - \delta\, \sigma_t^f), \\ 0 & \text{otherwise,} \end{cases} \tag{8}$$

where μ_t^e, μ_t^f, σ_t^e, and σ_t^f are the mean and standard deviations of Δe_t and Δf_t computed recursively, and we set $\delta = 1.25$. Results for $\delta = 1.00$ are qualitatively the same and are available in the online appendix. Similar indicators are proposed by Gerdesmeier *et al.* (2010) and Alessi and Detken (2011). Due to the recursive calculations required to build B_t, few observations at the beginning of the sample are lost and in what follows our analysis covers the period 1985:Q2–2014:Q2.

We then define a bust if $B_t = 1$ for at least two consecutive quarters. Using the described procedure, we find four periods of crisis: the first one in 1990:Q2–1990:Q3, corresponding to a period of banking and currency crises, the second one in 2002:Q2–2002:Q4, related to the burst of the dot-com bubble, the third one in 2007:Q4–2009:Q1, capturing the recent Global Financial Crisis, while the last one runs from 2011:Q2–2012:Q1, corresponding to the Sovereign Debt Crisis.

We then consider different possible leading indicators of stock market busts: (i) the monetary overhang $\widehat{\eta}_t$ implied by our model for a given sample, (ii) the spread between the long-term interest rate and the own rate of money, $(\ell_t - o_t)$, (see e.g. Estrella and Mishkin, 1998), (iii) the credit-to-GDP ratio, cy_t, (see, among others, Angelini *et al.*, 2014; Schularick and Taylor, 2012), and (iv) a measure of banks' non-core liabilities, ncl_t, which are represented by sources of funding different from retail deposits and then defined as the difference between the broad monetary aggregate M3 and the narrower M2, divided by M2 (Hahm *et al.*, 2013).

We estimate probit regressions for the probability of a bust, i.e. for Prob$(B_t = 1)$ on the samples 1985:Q2–2008:Q4 and 1985:Q2–2014:Q2, and using different combinations of the four considered explanatory variables computed two quarters before a bust.[15] Results are in Table

[14] Details on the series used in this Section are in Appendix A.
[15] Using different lags does not alter results.

9. The monetary overhang is always significant and has the expected positive sign, thus it increases the probability of a stock market busts. The best fit is observed in the last column (i.e. model V), which shows that only the credit-to-GDP ratio has an additional predictive content with respect to the monetary overhang, while the yield curve is barely significant and non-core liabilities are not significant.

Finally, as a measure of particular interest for policy makers, we can consider the noise-to-signal ratio (NtS), i.e. the ratio between the percentage of bad signals over the percentage of good signals as predicted by the estimated probit model.

This measure depends on how we define a *signal*. Indeed, a threshold level ζ is usually chosen, such that the estimated probability of a bust corresponds to a signal only if it is larger than ζ. By increasing ζ, NtS can be made smaller. However, it has to be understood that minimizing NtS is not necessarily the optimal choice. Instead it depends on the weights assigned to Type I errors (missed busts as percentage of periods in which a bust occurred) and to Type II errors (false alarms as percentage of periods in which no bust occurred). This is a policy problem studied in the literature, which has proposed different methods to choose ζ from the data (see e.g. Alessi and Detken, 2011). Although important, this is an issue beyond the scope of this paper and therefore, following the literature, we fix exogenously $\zeta = 0.25$ as in Gerdesmeier *et al.* (2010). With these choices we see that using only the monetary overhang as explanatory variable can lead to no predictive power (model I) while combining it with credit-to-GDP ratio improves predictability. In particular for model V a percentage of good signals (given by the sum of correctly predicted busts and quiet periods) of 91% and 90% and a noise-to-signal ratio of 0.10 and 0.10 on the samples 1985:Q2–2008:Q4 and 1985:Q2–2014:Q2, respectively.

Overall, our findings suggest that our monetary overhang, $\widehat{\eta}_t$, is a significant leading indicator of stock markets busts complementary to loans-to-GDP ratio, which is among the most relevant predictors of crises. This result may help in the implementation of macroprudential policies (De Grauwe and Gros, 2009; Bank of International Settlements, 2010).

Table 9: IMPLICATIONS FOR FINANCIAL STABILITY. PREDICTING CRISES.

	1985:Q2–2008:Q4					1985:Q2–2014:Q2				
			Model					Model		
Prob($B_t = 1$)	(I)	(II)	(III)	(IV)	(V)	(I)	(II)	(III)	(IV)	(V)
$\widehat{\eta}_{t-2}$	0.469	0.579	0.392	0.392	0.492	0.338	0.475	0.267	0.332	0.407
	[2.344]	[2.680]	[1.889]	[1.839]	[2.115]	[2.160]	[2.797]	[1.654]	[2.114]	[2.361]
cy_{t-2}		0.074			0.080		0.062			0.068
		[2.441]			[2.549]		[2.557]			[2.666]
$(\ell_{t-2} - o_{t-2})$			-0.180		-0.189			-0.173		-0.229
			[-1.254]		[-0.853]			[-1.357]		[-1.483]
ncl_{t-2}				0.599	0.220				0.460	-0.182
				[0.923]	[0.201]				[1.001]	[-0.306]
$\bar{R}^2$	0.10	0.21	0.12	0.11	0.25	0.06	0.14	0.08	0.07	0.17
R^2	0.06	0.13	0.08	0.07	0.16	0.04	0.10	0.06	0.05	0.12
Good	0.90	0.91	0.91	0.89	0.91	0.86	0.86	0.86	0.86	0.90
NtS	0.14	0.10	0.10	0.19	0.10	1.00	0.29	0.63	0.38	0.10

Probit regressions with dependent variable the binary bust indicator, B_t, with explanatory variables: $\widehat{\eta}_t$ = monetary overhang from our model estimated over a given sample; cy_t = first differences of loans-to-income ratio; $(\ell_t - o_t)$ = spread between long and own rates ; ncl_t = non-core liabilities. For each estimated model we report the maximum likelihood value of the estimated coefficients, t-statistics in parenthesis, the McFadden $\bar{R}^2$ and the R^2 as computed in Estrella and Mishkin (1998). The fraction of good signals (*Good*) and the noise-to-signal ratio (*NtS*) are computed with a threshold $\zeta = 0.25$ (the fraction of bad signals is equal to $(1 - Good)$).

6.2 Price stability

Beyond the traditional strand of literature aimed at assessing risks for price stability due to excess liquidity (see e.g. Woodford, 2008; McCallum and Nelson, 2010; De Santis *et al.*, 2013), a growing interest is devoted to understanding the effects of liquidity injections by the ECB and other central banks in terms of future inflation (see e.g. Dreger and Wolters, 2011) and the link between money and inflation during and after financial crises (Reynard, 2012). To address these issues, in our second policy exercise we evaluate the relevance of our monetary overhang measure in explaining inflation. As an initial exercise, we estimate full sample and sub-sample regressions including our money overhang measure which turns out to be not significant. This evidence is consistent with De Santis *et al.* (2013). However, when instead looking at first differences of our excess liquidity measure, $\Delta\widehat{\eta}_t$, which basically represent the distance of real money growth from its equilibrium, we find evidence of a statistically significant relation between inflation and excess liquidity.[16]

[16] In order to save space, we do not report here the tables for these in-sample exercises. They are available upon request from the authors.

We then move to investigate the predictive content of $\Delta\widehat{\eta}_t$ by running an out-of-sample exercise. In order to do so, we adopt the Stock and Watson (1999) approach to forecasting inflation and we estimate the following regressions:

$$\pi_{t+k}^{(k)} = a^{(k)} + \sum_{h=1}^{4} b_h^{(k)} \pi_{t-h} + \sum_{j=1}^{4} \mathbf{\Lambda}_j^{(k)\prime} \mathbf{Y}_{t-j} + u_{t+k}, \quad k = 4, 8, \tag{9}$$

where $\pi_{t+k}^{(k)}$ represents k-step ahead (cumulated) inflation, π_t represents q-o-q inflation, $\mathbf{Y}_t$ denotes a vector of explanatory variables. In particular we consider four specifications for $\mathbf{Y}_t$: (I) the benchmark case in which $\mathbf{Y}_t = \mathbf{0}$, (II) the case in which $\mathbf{Y}_t$ contains only changes in excess liquidity, $\Delta\widehat{\eta}_t$, (III) the case in which $\mathbf{Y}_t$ contains only q-o-q changes in real GDP Δy_t, and (IV) the case in which $\mathbf{Y}_t = (\Delta\widehat{\eta}_t,\ \Delta y_t)'$, thus mimicking a two pillars Phillips Curve (Gerlach, 2004). We choose $k = 4, 8$ since these are the horizons relevant short and medium-run predictions for the ECB.

We estimate each model on a fixed sample terminating in 2008:Q4 or in 2014:Q2. Estimation starts either in 1980:Q1, which is the first period of data available, or in 1999:Q1, marking the date in which the ECB effectively started to conduct monetary policy. We then recursively forecast inflation at 4 or 8 steps ahead until the end of the sample in 2016:Q3. RMSEs, relative to the benchmark model I, are reported for each model and sample considered are in Table 10.

Some findings are worth mentioning. First, for the pre-crisis periods ending in 2008:Q4, excess liquidity measures are useful for forecasting inflation, however their additional predictive content is basically outperformed or in line with the one provided by considering real GDP dynamics only (model III). This is particularly true when restricting the focus on samples staring in 1999:Q1. Second, in the post-crisis estimation sample ending in 2014:Q2 and thus considering the most recent data, it is very difficult to outperform models based on pure inflation at shorter horizons (model I), whereas in the medium-run specifications including economic activity developments dominate those based on excess liquidity (models III and IV). Notice however that results on this last sample are based on a small number of observations. Overall, this evidence suggest a modest gain in using excess liquidity based measures for

Table 10: Implications for price stability. Predicting inflation.

Estimation Sample	Prediction Sample	Model (I)	(II)	(III)	(IV)
	$k = 4$				
1980:Q1–2008:Q4	2009:Q4–2016:Q3	1.000	0.979	0.874	0.909
1999:Q1–2008:Q4	2009:Q4–2016:Q3	1.000	0.877	0.802	0.683
1980:Q1–2014:Q2	2015:Q2–2016:Q3	1.000	1.063	1.067	1.093
1999:Q1–2014:Q2	2015:Q2–2016:Q3	1.000	1.034	1.023	1.024
	$k = 8$				
1980:Q1–2008:Q4	2010:Q4–2016:Q3	1.000	0.954	0.988	1.052
1999:Q1–2008:Q4	2010:Q4–2016:Q3	1.000	0.922	0.631	0.694
1980:Q1–2014:Q2	2016:Q2–2016:Q3	1.000	1.052	0.394	0.373
1999:Q1–2014:Q2	2016:Q2–2016:Q3	1.000	0.907	0.530	0.537

RMSEs relative to the benchmark model based only on lags of q-o-q inflation only ($\mathbf{Y}_t = \mathbf{0}$).

forecasting inflation, in line with the weakening of the quantitative relation between money and inflation in the EA, highlighted for example by Hofmann (2009).

7 Concluding remarks

In this paper we make two contributions to the literature on EA money demand. First, by means of the new time-varying cointegration test by Bierens and Martins (2010), we perform a model comparison among different specifications claiming to find a solution to the problem of instability of long-run relations for M3. We find that the key drivers for the dynamics of the EA broad monetary aggregate after 2001 are the spread in price-earnings ratios with US assets, representing a speculative motive for holding money, and changes in the unemployment rate, capturing a precautionary motive.

Second, based on these findings, we estimate a new stable VECM of money, stock markets, and unemployment. Identification of the long-run run structure has a meaningful economic interpretation similar to the one found by Beyer and Farmer (2007) for the US. Indeed, we identify a money demand equation, an aggregate demand equation, and a relation between the yield curve and inflation.

Furthermore, this model has useful implications for the ECB monetary policy. In particular, the derived monetary overhang measure is a leading indicator of periods of financial stress and has also a statistically significant although weak incremental predictive content for forecasting inflation. Overall, these results suggest that the monetary pillar may evolve into policies more focused on financial rather than price stability (Galí, 2012).

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Appendix

A Data description and transformations

The following is the list of variables used in the empirical analysis and a brief description of the way they were transformed. For additional details see the cited works.

1. $m_t = 100\log(M_t)$, where M_t is the monetary aggregate M3, outstanding amounts at the end of the period (source: ECB, series ID: BSI.M.U2.Y.V.M30.X.1.U2.2300.Z01.E);
2. $p_t = 100\log(P_t)$, where P_t is GDP deflator, see Fagan *et al.* (2001) for details (source: Area Wide Model Database, series ID: YED);
3. $y_t = 100\log(Y_t)$, where Y_t is real GDP, see Fagan *et al.* (2001) for details (source: Area Wide Model Database, series ID: YER);
4. s_t is short-term interest rate, see Fagan *et al.* (2001) for details (source: Area Wide Model Database, series ID: STR);
5. ℓ_t is long-term interest rate, see Fagan *et al.* (2001) for details (source: Area Wide Model Database, series ID: LTR);
6. o_t is own rate of M3, see e.g. Calza *et al.* (2001) and Giannone *et al.* (2012) for details (source: ECB calculations);
7. $\pi_t = 400\Delta\log P_t$ or $\pi_t = 100\Delta_4\log P_t$ is the q-o-q (or y-o-y) inflation rate, where P_t is defined above;
8. Δe_t in Section 5 is the 3 years average of $400\Delta(\log E_t)$, where E_t is the equity index, based on the German DAX30 for the period 1980-1986 and on the Dow Jones Euro Stoxx50 for 1987-2007, see Carstensen (2006) for details, (source: Datastream), in Section 6 $\Delta e_t = 400\Delta(\log E_t)$;
9. v_t is the 2 years average log-volatility of equity returns, estimated with a leverage GARCH model, see Carstensen (2006) for details;
10. $\Delta h_t = 400\Delta(\log H_t)$, where H_t is net household housing wealth at current replacement costs, see Beyer (2009) for details (source: ECB, series ID: IEAQ.Q.I6.N.V.LE.TU.S1M.A1.S.1.X.E.Z);
11. $f_t = 100\log F_t$, where F_t is real financial wealth, see Dreger and Wolters (2010b) for details (source: ECB Monthly Bulletin);
12. $w_t = 100\log(W_t/P_t)$, where P_t is defined above and W_t is total wealth, see Fagan *et al.* (2001) for details (source: Area Wide Model Database, series ID: WLN);
13. $\Delta u_t = \Delta_4 U_t$, where U_t is the unemployment rate, see Fagan *et al.* (2001) for details (source: Area Wide Model Database, series ID: URX);
14. ℓ_t^* is 10-year US Treasury notes and bonds yields end of month (source: Federal Reserve Bank);
15. q_t is the log of EA price to earnings ratio, see De Santis *et al.* (2013) for details (source: Datastream);
16. q_t^* is the log of US price to earnings ratio, see De Santis *et al.* (2013) for details (source: Datastream);
17. $cy_t = 100(\log(L_t/Y_t))$, where Y_t is defined above and L_t is loans, outstanding amounts of loans to non-financial corporations at the end of the period (source: ECB, series ID: BSI.Q.U2.N.A.A20.A.1.U2.2240.Z01.E);
18. $ncl_t = 100(m_t - \widetilde{m}_t)/\widetilde{m}_t$, where m_t is defined above and $\widetilde{m}_t = 100\log(\widetilde{M_t})$, where $\widetilde{M}_t$ is the monetary aggregate M2, outstanding amounts at the end of the period (source: ECB, series ID: BSI.M.U2.Y.V.M20.X.1.U2.2300.Z01.E);
19. $\Delta f_t = 400\Delta(\log F_t)$, where F_t is defined above.

B Goodness-of-fit tests for the time-invariant VECM

Table B-1: Tests of exclusion, stationarity, and weak exogeneity: 1980Q1–2008Q4.

Test of exclusion						
r	m_t	y_t	Δu_t	$(q_t - q_t^*)$	$(\ell_t - o_t)$	π_t
1	0.995 (0.32)	1.110 (0.29)	0.040 (0.84)	1.197 (0.27)	1.095 (0.30)	0.069 (0.79)
2	26.535 (0.00)	27.360 (0.00)	29.858 (0.00)	30.328 (0.00)	1.911 (0.39)	0.107 (0.95)
3	39.396 (0.00)	35.803 (0.00)	39.441 (0.00)	42.452 (0.00)	13.124 (0.00)	14.211 (0.00)
4	41.483 (0.00)	39.429 (0.00)	45.088 (0.00)	48.070 (0.00)	17.379 (0.00)	15.725 (0.00)
5	43.203 (0.00)	44.677 (0.00)	55.418 (0.00)	55.871 (0.00)	27.906 (0.00)	25.281 (0.00)
Test of stationarity						
r	m_t	y_t	Δu_t	$(q_t - q_t^*)$	$(\ell_t - o_t)$	π_t
1	41.890 (0.00)	43.026 (0.00)	7.264 (0.20)	45.857 (0.00)	45.898 (0.00)	49.648 (0.00)
2	40.051 (0.00)	41.170 (0.00)	6.117 (0.19)	44.945 (0.00)	44.036 (0.00)	48.077 (0.00)
3	14.056 (0.00)	14.934 (0.00)	5.599 (0.09)	22.962 (0.00)	15.691 (0.00)	19.280 (0.00)
4	1.301 (0.52)	1.773 (0.41)	5.975 (0.05)	8.943 (0.01)	4.289 (0.12)	5.225 (0.07)
5	0.013 (0.91)	0.421 (0.52)	0.817 (0.37)	6.573 (0.01)	0.234 (0.63)	0.667 (0.41)
Test of weak exogeneity						
r	m_t	y_t	Δu_t	$(q_t - q_t^*)$	$(\ell_t - o_t)$	π_t
1	0.420 (0.52)	1.464 (0.23)	0.002 (0.97)	1.663 (0.20)	1.665 (0.20)	0.108 (0.74)
2	8.490 (0.01)	15.791 (0.00)	29.818 (0.00)	22.823 (0.00)	4.906 (0.09)	1.918 (0.38)
3	14.896 (0.00)	17.570 (0.00)	38.836 (0.00)	33.868 (0.00)	9.277 (0.03)	9.728 (0.02)
4	18.611 (0.00)	21.019 (0.00)	41.165 (0.00)	34.973 (0.00)	11.995 (0.02)	9.854 (0.04)
5	26.352 (0.00)	26.761 (0.00)	51.362 (0.00)	36.153 (0.00)	20.845 (0.00)	16.764 (0.01)

Each line reports the $\chi^2_{(r)}$ test statistics and in parenthesis the corresponding p-value for the null hypothesis of excluding a given variable (top panel), stationarity (middle panel), and weak exogeneity (bottom panel) of a given variable, conditioning on the cointegration rank r.

Table B-2: Tests of exclusion, stationarity, and weak exogeneity: 1980Q1–2014Q2.

Test of exclusion						
r	m_t	y_t	Δu_t	$(q_t - q_t^*)$	$(\ell_t - o_t)$	π_t
1	0.227	0.282	1.312	14.581	0.219	0.053
	(0.634)	(0.595)	(0.252)	(0.000)	(0.640)	(0.817)
2	14.378	12.142	6.003	31.655	4.923	15.025
	(0.001)	(0.002)	(0.050)	(0.000)	(0.085)	(0.001)
3	23.172	18.549	13.499	40.947	14.823	25.913
	(0.000)	(0.000)	(0.004)	(0.000)	(0.002)	(0.000)
4	32.692	28.446	19.755	50.39	24.435	34.808
	(0.000)	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)
5	34.591	32.544	26.271	52.676	31.49	40.969
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Test of stationarity						
r	m_t	y_t	Δu_t	$(q_t - q_t^*)$	$(\ell_t - o_t)$	π_t
1	59.656	60.417	11.697	58.925	60.009	63.57
	(0.000)	(0.000)	(0.039)	(0.000)	(0.000)	(0.000)
2	38.249	37.565	11.69	35.701	37.278	41.563
	(0.000)	(0.000)	(0.020)	(0.000)	(0.000)	(0.000)
3	18.704	17.91	11.565	18.625	18.887	21.9
	(0.000)	(0.000)	(0.009)	(0.000)	(0.000)	(0.000)
4	7.459	6.561	10.501	8.883	10.288	10.434
	(0.024)	(0.038)	(0.005)	(0.012)	(0.006)	(0.005)
5	3.077	1.569	1.591	1.259	1.241	1.174
	(0.079)	(0.210)	(0.207)	(0.262)	(0.265)	(0.279)
Test of weak exogeneity						
r	m_t	y_t	Δu_t	$(q_t - q_t^*)$	$(\ell_t - o_t)$	π_t
1	13.819	1.017	14.47	2.492	1.575	0.28
	(0.000)	(0.313)	(0.000)	(0.114)	(0.209)	(0.597)
2	21.875	3.279	30.015	14.636	4.981	7.241
	(0.000)	(0.194)	(0.000)	(0.001)	(0.083)	(0.027)
3	25.471	7.169	41.731	23.495	11.607	10.829
	(0.000)	(0.067)	(0.000)	(0.000)	(0.009)	(0.013)
4	29.297	16.484	48.139	24.259	20.173	11.094
	(0.000)	(0.002)	(0.000)	(0.000)	(0.000)	(0.026)
5	34.595	19.689	53.908	25.829	24.199	16.293
	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)	(0.006)

Each line reports the $\chi^2_{(r)}$ test statistics and in parenthesis the corresponding p-value for the null hypothesis of excluding a given variable (top panel), stationarity (middle panel), and weak exogeneity (bottom panel) of a given variable, conditioning on the cointegration rank r.

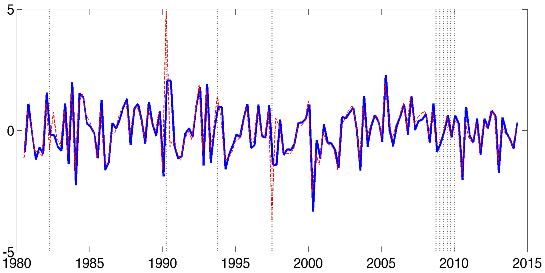
Table B-3: MULTIVARIATE AND UNIVARIATE TESTS ON THE RESIDUALS.

Multivariate test of no autocorrelation				
		1980:Q1–2008:Q4		1980:Q1–2014:Q2
LM(1)	$\chi^2_{(36)}$	48.772 (0.08)	$\chi^2_{(36)}$	74.340 (0.00)
LM(2)	$\chi^2_{(36)}$	46.549 (0.11)	$\chi^2_{(36)}$	57.653 (0.01)
Multivariate test of conditional heteroschedasticity				
No dummies				
ARCH-LM(1)	$\chi^2_{(441)}$	489.343 (0.06)	$\chi^2_{(441)}$	587.976 (0.00)
ARCH-LM(2)	$\chi^2_{(882)}$	1022.839 (0.00)	$\chi^2_{(882)}$	1051.957 (0.00)
Dummies				
ARCH-LM(1)	$\chi^2_{(441)}$	489.343 (0.06)	$\chi^2_{(441)}$	569.060 (0.00)
ARCH-LM(2)	$\chi^2_{(882)}$	1022.839 (0.00)	$\chi^2_{(882)}$	1059.881 (0.00)
Multivariate test of normality				
No dummies	$\chi^2_{(12)}$	87.754 (0.00)	$\chi^2_{(12)}$	94.415 (0.00)
Dummies	$\chi^2_{(12)}$	20.006 (0.07)	$\chi^2_{(12)}$	20.970 (0.06)

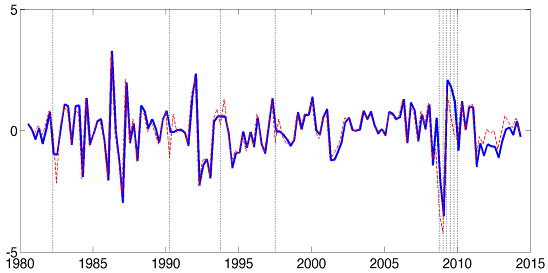
Univariate Test	Ljung-Box(4)	ARCH(2)	Normality No dummies	Normality Dummies
		1980:Q1–2008:Q4		
$\Delta(m_t - p_t)$	4.186 (0.38)	0.277 (0.87)	49.394 (0.00)	0.624 (0.73)
Δy_t	5.580 (0.23)	5.747 (0.26)	7.552 (0.02)	6.919 (0.03)
$\Delta\Delta u_t$	12.909 (0.01)	0.467 (0.79)	14.219 (0.00)	1.203 (0.55)
$\Delta(q_t - q_t^*)$	2.657 (0.62)	21.718 (0.00)	2.859 (0.24)	3.707 (0.16)
$\Delta(\ell_t - o_t)$	2.622 (0.62)	0.017 (0.99)	4.998 (0.08)	5.460 (0.07)
$\Delta\pi_t$	0.983 (0.91)	0.790 (0.67)	3.159 (0.21)	5.143 (0.08)
		1980:Q1–2014:Q2		
$\Delta(m_t - p_t)$	5.738 (0.21)	0.555 (0.76)	53.707 (0.00)	1.467 (0.48)
Δy_t	1.790 (0.77)	29.018 (0.00)	20.522 (0.00)	14.211 (0.01)
$\Delta\Delta u_t$	25.863 (0.00)	4.252 (0.12)	12.443 (0.00)	2.145 (0.10)
$\Delta(q_t - q_t^*)$	1.083 (0.90)	9.251 (0.01)	1.590 (0.45)	3.740 (0.15)
$\Delta(\ell_t - o_t)$	3.930 (0.42)	0.144 (0.93)	0.818 (0.66)	1.941 (0.38)
$\Delta\pi_t$	0.796 (0.94)	0.507 (0.78)	6.146 (0.05)	2.145 (0.34)

Top panel: Joint tests on residuals for the null-hypothesis of no autocorrelation, no conditional heteroschedasticity (LM tests with 1and 2 lags), and normality (Doornik-Hansen test); p-values are in parenthesis. Bottom panel: Tests on univariate residuals for the null-hypothesis of no autocorrelation (Ljung-Box test with 4 lags), no conditional heteroschedasticity (ARCH test with 2 lags), and normality (Doornik-Hansen test); p-values are in parenthesis; estimation with dummies is obtained when controlling for institutional changes in 1982:Q2, 1990:Q2, 1993:Q4, and 1997:Q3 and for a dummy crisis 2008:Q3-2009:Q4.

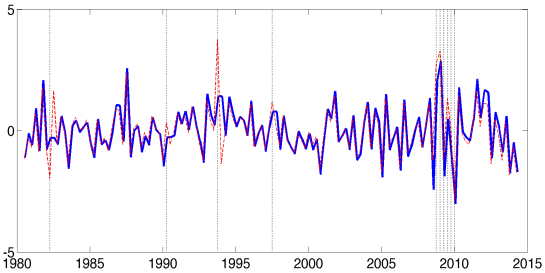
Figure B-1: MODEL RESIDUALS.



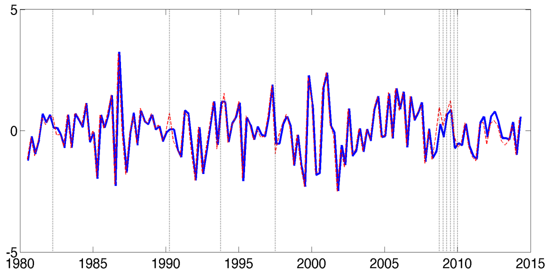
(a) $\Delta(m_t - p_t)$



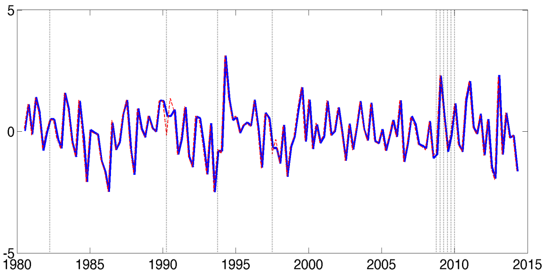
(b) Δy_t



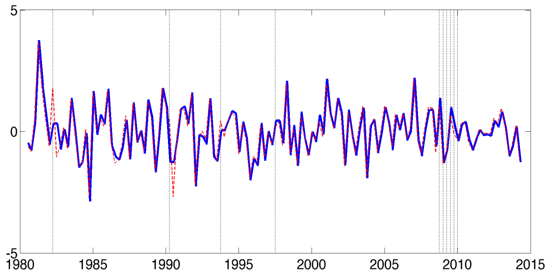
(c) $\Delta\Delta u_t$



(d) $\Delta(q_t - q_t^*)$



(e) $\Delta(\ell_t - o_t)$



(f) $\Delta\pi_t$

Red line: residuals of model (6). Blue line: residuals of the same model with dummies controlling for institutional changes in 1982:Q2, 1990:Q2, 1993:Q4, and 1997:Q3 and for a dummy crisis 2008:Q3-2009:Q4(indicated by dashed black lines). The estimation sample is 1980:Q1–2014:Q2.