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On the Source of Euro Area Money Demand Stability. A Time-Varying Cointegration Analysis

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ON THE SOURCES OF EURO AREA MONEY DEMAND STABILITY A TIME-VARYING COINTEGRATION ANALYSIS

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Abstract

We adopt a time-varying cointegration test to discriminate among different empirical studies claiming to find a stable Euro Area money demand equation. A time-invariant relation explaining real balances is rejected by data, even when accounting for housing, financial and labour markets. Conversely, an international portfolio allocation approach provides stabilization. In particular, international financial markets, rather than monetary policy, are the key determinant of the observed diverging path of money growth. In terms of policy, we provide empirical support for a New Two Pillars Strategy aimed to achieve financial stability through money and credit and price stability through interest rates.

Keywords Euro Area Money Demand; Time-Varying Vector Error Correction Model; International Portfolio Allocation; Financial Stability.

JEL Classification: E41; E44; C32.

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1 Introduction

Since 1999 the European Central Bank (ECB) Governing Council assigned a prominent role to monetary aggregates, therefore starting monitoring a reference value for the growth rate of M3, and establishing that an annual increase of 4.5% should have been consistent with an inflation target close or below 2%. This was known as the *monetary pillar* aimed to achieve the statutory objective of price stability jointly with the *economic analysis pillar*, i.e. the Two Pillars Strategy of ECB.¹ For this scope, a stable relation between money demand and income was felt to be a necessary prerequisite (Galí, 2008b). However, since the end of 2001, the monetary aggregate diverged from the reference value of 4.5%, while inflation remained stable around its target at 2% (De Grauwe and Gros, 2009, and Figure 1). This sharp rise in M3 growth associated with a period of relatively low inflation has casted doubts on the linkage between money and inflation and on the stability of Euro Area (EA) money demand. Consistently, since the end of 2001, signs of instability started to appear in estimating reduced-form money demand equations (Kontolemis, 2002), becoming stronger the more data became available. Investigating the causes of such instability can be extremely useful both for economic modeling and from a policy perspective. Indeed, despite the fact that, regardless of diverging money growth, price stability is under control, monitoring money aggregates could help in achieving financial stability by adopting a macro-prudential control of banks, i.e. by constraining credit supply through countercyclical reserve accumulation (De Grauwe and Gros, 2009).

In this paper we argue that, contrary to the existing literature, adopting a smooth time-varying framework represents the proper way to understand the (in)stability nature of EA money demand. Previous results on an unstable relation between money demand and income starting from the Euro launch suggest the presence of a structural break taking place around 2001. Not recognizing it can lead to estimation errors and unreliability of the model in general. If this is the case, the traditional time-invariant approach is not suitable for studying the recent dynamics of money demand. A desirable way to model economic change, that avoids the need of specifying the exact timing of the structural break, consists in assuming a smooth evolution of the parameters. For this reason we use the smooth time-varying cointegration likelihood-ratio test for parameter stability developed in Bierens and Martins (2010b) who in turn generalize the framework by Park and Hahn (1999). The test we employ is needed to verify if expanding the set of variables included in the cointegrating vector provides a solution to the apparent disequilibrium path in long-run money demand. This approach enables us to disentangle between changes in parameters and new motives for holding money. In this framework the usual time-invariant Vector Error Correction model is nothing else but a particular case of a most general model with time-varying cointegrating vectors modeled by means of Chebyshev polynomials.

First, we show that the classical model by Calza *et al.* (2001) is stable until 2001, while it becomes unstable afterwards, consistently with the analysis in De Santis *et al.* (2008), Beyer (2009) and Fischer *et al.* (2009). Second, when considering data spanning 1980:Q1-2007:Q2, we strongly reject the null hypothesis of a time-invariant money demand equation against the alternative of a time-varying one for all the specifications considered and proposed by the literature as possible solutions to the instability problem, with one exception. More precisely,

¹The existence of a separate monetary pillar has been widely discussed and investigated by Rudebusch and Svensson (2002); Gerlach and Svensson (2003); Gerlach (2004) among others, and often criticized (Woodford, 2007, 2008).

in order to obtain robust evidence, we examine eight of the most influential specifications proposed in the literature, namely the classical models by Calza *et al.* (2001); Coenen and Vega (2001), and Vlaar (2004) and the more recent works by Carstensen (2006); De Santis *et al.* (2008); Dreger and Wolters (2009); Beyer (2009), and De Bondt (2010) (see Section 3 for a detailed review).² In the second group of papers the classical set of explanatory variables (e.g. income and interest rates) is extended on the base of a number of interesting underlying economic reasons related to financial, labour, and housing markets. In our framework none of these motives is found to be a significant explanatory variable of the observed instability. However, when considering the model by De Santis *et al.* (2008), based on international portfolio allocation, we accept the null-hypothesis of a time-invariant long-run relation between money demand and income. This result suggests that international financial markets variables are the key determinants of the diverging path observed in M3 growth rate, while we show that monetary policy has only an indirect role. Additional results on a close relation between money demand, loans supplied by banks, and international asset prices, allow us to suggest policy implications related to the prevention of financial crises.

Summing up, we contribute to the literature on Euro Area money demand in four different aspects. First, we make use of the new methodology by Bierens and Martins (2010b) to properly test for the stability of cointegrating relations over time. Second, we discriminate among different contributions who claim to find a stable specification, only validating the one in De Santis *et al.* (2008). Third, we go one step forward with respect to De Santis *et al.* (2008) in highlighting the role of international financial markets, rather than monetary policy instruments, as the direct determinants of M3 growth rate. Fourth, and most significantly for ECB policy, we provide empirical support for a New Two Pillars Strategy (De Grauwe and Gros, 2009) aimed to achieve financial stability through monetary aggregates and price stability through interest rates.

Section 2 presents the standard economic framework for modeling long-run money demand. Section 3 reviews the contributions dealing with recent years money demand instability in the Euro Area and considered in this paper. Section 4 outlines the econometric methodology by Bierens and Martins (2010b). Section 5 presents the results which provide empirical evidence on money demand instability. Section 6 analyzes the key role played by international markets in explaining money demand instability, while Section 7 suggests policy implications for achieving financial stability. Finally, Section 8 concludes.

2 Money demand specification

Economic theory assumes the quantity of real money hold to be a function f of the amount of transactions in the economy, i.e. income Y_t which represents the so called *transactional* motive, and of a set of explanatory variables \mathbf{X}_t representing either the opportunity cost of

²Other contributions not considered in this paper and presenting a stable money demand equation are, among others, Bruggemann *et al.* (2003); Artis and Beyer (2004); Brand and Cassola (2004). Finally, Setzer and Wolff (2009) specify individual member states money demand in deviations from the EA average, and show that country specific income elasticity and interest rate (semi)-elasticity have remained stable. Nevertheless, they adopt a panel-cointegrated approach, while we focus here to pure time series models.

holding money or other possible motives:

$$\frac{M_t}{P_t} = f(Y_t, \mathbf{X}_t). \quad (1)$$

M_t denotes nominal money and P_t is the price level. \mathbf{X}_t , for example, may comprehend the own rate of money, the short-term interest rate (the policy instrument), inflation, wealth, etc. (see Golinelli and Pastorello, 2002; De Bondt, 2010, for a comprehensive survey on different specifications in the classical literature and for a survey on more recent studies, respectively). In this paper we examine eight possible specifications considered by the literature as the most influential (see next Section).

Empirical literature on money demand is mainly based on the estimation of a long-run relation by means of cointegration techniques in the spirit of Johansen (1995). Price homogeneity is assumed to hold in the long-run: M_t and P_t are assumed to be integrated of order 2, $I(2)$, but, if they are cointegrated - with vector $[1, -1]'$, then real balances, defined as $(\log M_t - \log P_t)$, are $I(1)$. Correctly specifying the vector of explanatory variables \mathbf{X}_t is a key point for estimating equation (1). The most general specification in the empirical literature on money demand takes as a good approximation a (semi) log-linear form of equation (1)

$$(m_t - p_t) = \beta_0 + \beta^y y_t + \beta^{s-o}(s_t - o_t) + \beta^{\ell-o}(\ell_t - o_t) + \beta^\pi \pi_t \quad (2)$$

where m_t denotes the log of nominal monetary aggregate M3, p_t is the log of GDP deflator, y_t is the log of real GDP, s_t and ℓ_t represent the short and long term interest rates respectively, and o_t stands for the own rate of M3 (see e.g. Calza et al., 2001, for its definition). Finally, when using quarterly data, $\pi_t = \Delta p_t \equiv (p_t - p_{t-1})$ is quarter over quarter inflation rate. As for the sign of coefficients in equation (2), one should expect - building on theory and empirical evidence - $\beta^y > 0$, $\beta^{\ell-o} < 0$, $\beta^\pi < 0$, whereas the impact of β^{s-o} is a more controversial: raising the policy rate should lower demand for money, but, depending on the monetary aggregate chosen in the estimation, one can find both signs. Vlaar (2004) finds a positive correlation between M3 and the short term interest rate in the EA, posing issues for the controllability of money stock by monetary authorities, as a contractionary monetary policy raises demand for money. Obviously, this would be a big concern since M3 is the aggregate monitored by ECB as reference value for price stability, the monetary pillar in the two pillars strategy. However, most of empirical studies on EA find a negative semielasticity of real balances to the short term interest rate (see e.g. De Bondt, 2010).

When analysing long-run money demand, the multivariate analysis is based on the specification of the following VAR in levels

$$\mathbf{Z}_t = \sum_{h=1}^p \mathbf{A}_h \mathbf{Z}_{t-h} + \boldsymbol{\varepsilon}_t, \quad \boldsymbol{\varepsilon}_t \stackrel{\text{i.i.d.}}{\sim} N_k(\mathbf{0}, \boldsymbol{\Omega}), \quad t = 1, \dots, T, \quad (3)$$

where T is the number of observations and

$$\mathbf{Z}_t = [(m_t - p_t) \quad y_t \quad (s_t - o_t) \quad (\ell_t - o_t) \quad \pi_t]'. \quad (4)$$

Equation (3) implies the time-invariant Vector Error Correction model (TI-VECM) of order p

$$\Delta \mathbf{Z}_t = \boldsymbol{\gamma}_0 + \boldsymbol{\Pi} \mathbf{Z}_{t-1} + \sum_{j=1}^{p-1} \boldsymbol{\Gamma}_j \Delta \mathbf{Z}_{t-j} + \boldsymbol{\varepsilon}_t, \quad \boldsymbol{\varepsilon}_t \stackrel{\text{i.i.d.}}{\sim} N_k(\mathbf{0}, \boldsymbol{\Omega}), \quad t = 1, \dots, T. \quad (5)$$

$\boldsymbol{\Omega}$ and $\boldsymbol{\Gamma}_j$ are $k \times k$ matrices where k is the size of \mathbf{Z}_t , i.e. $k = 5$ in (2). If cointegration between variables exists, one can decompose $\boldsymbol{\Pi} = \boldsymbol{\alpha} \boldsymbol{\beta}'$ where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are fixed $k \times r$ matrices with reduced rank $r < k$ representing the number of cointegration vectors. In principle, we cannot exclude the presence of more than one cointegrating relation, like, say, the Fisher Interest Parity and the expectations theory (Coenen and Vega, 2001; Vlaar, 2004). If $r = 1$ only a normalization with respect to the coefficient of interest is required in order to identify the cointegrating relation. In this case, $\boldsymbol{\beta} = [1 \quad \beta^y \quad \beta^{s-o} \quad \beta^{\ell-o} \quad \beta^\pi]'$, and $\boldsymbol{\gamma}_0 = \boldsymbol{\alpha} \beta_0$. On the contrary, if $r > 1$, (5) has to be restricted to achieve identification of the long-run structure. We can use predictions coming from theory on long-run income elasticity, (e.g. $\beta^y = 1$ according the quantity theory or $\beta^y = 0.5$ in the Baumol-Tobin model), or arguing on the proper inclusion of the interest rates, imposing a common stochastic trend between s_t and ℓ_t , which implies $\beta^s = 1, \beta^\ell = -1$, i.e. imposing and testing for a cointegrating vector $[1, -1]'$ between the variables (Coenen and Vega, 2001). Alternatively, especially to depict money demand, literature often employs the contemporaneous short-run homogeneity constraint (Coenen and Vega, 2001; Galí, 1992), meaning that $\beta^\pi = 0$.³ As noted by Coenen and Vega (2001), the inclusion (exclusion) of π_t in equation (2) is subject to some controversy. They note that including π_t allows for theoretically plausible and convenient reparametrizations in dynamic money demand models. They also interpret inflation as the opportunity cost of holding money instead of real assets, as showed in Ericsson (1998). Furthermore, Dreger and Wolters (2009) show that accounting for inflation can stabilize EA money demand.

3 (In)Stability of Euro Area money demand

3.1 Classical specifications

Calza *et al.* (2001) propose a model which is now considered as the workhorse M3 equation used in the Quarterly Monetary Assessment by ECB during the period 2001-2006 Fischer *et al.* (2009) and it is taken as starting point in the most recent contributions. The estimated equation relates real balances to real GDP and two different spreads ($s_t - o_t$) and ($\ell_t - o_t$), hence

$$\mathbf{X}_t = [(s_t - o_t) \quad (\ell_t - o_t)]'.$$

When estimating (5) with this choice of \mathbf{X}_t we obtain the cointegrating vector (t -statistics in parenthesis)

$$(m_t - p_t) = \underset{(33.50)}{1.34} y_t - \underset{(-2.97)}{0.86} (s_t - o_t) + \underset{(0.03)}{0.01} (\ell_t - o_t),$$

clearly $\beta^{\ell-o}$ is not significant and can be omitted from the analysis. The disequilibrium path defined as $(m_t - p_t) - \beta^y y_t$ is no more mean reverting after 2001 (see Figure 2). This result is interpreted in the literature as a sign of recent instability of money demand, also known as

³Galí (1992) uses this restriction to get *short-run* identification in a Structural VAR model on US data, noting that it has some theoretical foundations in the absence of costs of adjusting nominal money holdings.

money overhang (see Alves *et al.*, 2007; De Santis *et al.*, 2008; Beyer, 2009, among others). In particular, the estimated coefficients of output and the short-term spread become dramatically unstable from 2003 onwards (see Beyer, 2009), with no additional benefit stemming from allowing for a linear trend to capture velocity shift. Analogous results on instability are found when considering the specifications by Coenen and Vega (2001) and Vlaar (2004). While the former includes as explanatory variable only real GDP and the spread ($\ell_t - s_t$), the latter adds also inflation π_t . In the following, we briefly present some of the most recent and significant contributions to the empirical literature on money demand claiming to find a stable long-run relation. Each work takes into account some novel variables with respect to the classical specifications. We distinguish the contribution by Dreger and Wolters (2009) from the others, as they claim to achieve stability using only a combination of the classical set of explanatory variables. In Section 5 we apply a time-varying cointegration test to each of these eight specifications.

3.2 *Dreger and Wolters (2009)*

Dreger and Wolters (2009) claim that the classical set of explanatory variables is sufficient to achieve a stable relation once relaxing the short-run homogeneity constraint between real balances and inflation. They motivate their result arguing that in a low inflation environment the opportunity cost of holding money has decreased. Starting from the consideration that inflation represents the cost of holding money instead of other real assets, they provide a number of justifications for the inclusion of π_t . Incorporating inflation in the analysis can help to disentangle nominal and real terms of the cost adjustment process (Wolters and Lütkepohl, 1998) or to provide a convenient generalization of the short-run homogeneity restriction imposed between money and prices. Hence,

$$\mathbf{X}_t = [s_t \quad \ell_t \quad \pi_t]'$$

When investigating the cointegrating properties of \mathbf{Z}_t and its different subsystems on quarterly data spanning 1983:Q1-2004:Q4, the Johansen trace test always displays a unique cointegrating vector, which can be identified as a long-run money demand relationship.

3.3 *Carstensen (2006)*

Carstensen (2006) is the first author introducing new variables with respect to the classical set in specifying money demand. Specifically, he argues that high money growth observed since 2001 may indicate neglecting stock market impact when estimating unstable money demand equations. His view is empirically supported by shifts from risky assets to funds in M3 composing a large fraction of excess of money growth (European Central Bank, 2003). He thus proposes to augment the Calza *et al.* (2001) model to account for the effect of stock market variables on money holdings, which he proxies by means of equity returns e_t and market volatility v_t . Hence,

$$\mathbf{X}_t = [(s_t - o_t) \quad (e_t - o_t) \quad v_t]'$$

As for the cointegration analysis, conducted on the sample 1980:Q1-2003:Q2, the Johansen trace statistics reveals the existence of one cointegrating vector, with real balances demand negatively affected by equity returns and positively related to market volatility. After assessing

the stability of this specification through a battery of recursive tests, Carstensen uses his model to evaluate the concern that the excess of liquidity originated in the EA may fuel inflation. He computes deviations between actual M3 and the long-run equilibrium implied by money demand equations (i.e. money overhang) under the two different specifications, the Calza *et al.* (2001) and the stock market-augmented. It turns out that, after taking into account financial markets developments, there is no indication of excess liquidity and serious inflationary pressure.

3.4 *De Santis, Favero and Roffia (2008)*

De Santis *et al.* (2008) propose an international portfolio allocation approach to evaluate EA money demand. The aim of their paper is studying the causes of money and prices relationship break-down observed after 2001. They document instability of Calza *et al.* (2001) model, asking whether it is possible to reconstruct a stable money demand by including variables omitted in the classical specification. Assuming that accounting for international portfolio flows may explain recent monetary developments in the EA, they build up a simple Tobin portfolio model of asset choice in an open-economy environment, where shareholders may diversify their wealth among money, domestic and foreign risky assets. They evaluate the model by augmenting the Calza *et al.* (2001) set of explanatory variables with determinants of stocks and bonds Sharpe ratios in the EA and US markets and long term interest rates:

$$\mathbf{X}_t = [(q_t - d_t)^{EA} \quad (q_t - d_t)^{US} \quad \rho_t^{EA} \quad \rho_t^{US}]'$$

where $(q_t - d_t)$ stands for the price-earnings ratio on the two markets, as a proxy of Sharpe ratio, and ρ_t^{US} is the long-term US interest rate. The trace statistics detects three cointegrating vectors, which are identified as a stable portfolio augmented money demand and a relation between price-earnings and the domestic interest rate of each country. They conclude presenting decisive evidence in favour of incorporating asset prices into the framework analysis for money demand in the EA and, more generally for monetary developments, thus switching from a “simple” domestic environment to an international portfolio context, which they suggest as possible useful reference for the increasing literature on the two-directions linkages between money and asset prices (Baumeister *et al.*, 2008; Goodhart and Hofmann, 2008).

3.5 *Beyer (2009)*

Beyer (2009) emphasizes the role of housing wealth in capturing the trending behaviour of money in the first ten years of ECB existence. He finds that adding housing wealth as an explanatory variable for real balances produces a stable long run relation. His specification of the opportunity cost variables \mathbf{X}_t coincides with the original one in Calza *et al.* (2001), with the addition of housing wealth growth rate, Δh_t , as a further scale variable to real GDP. The cointegration analysis, conducted on the sample 1980:Q1-2007:Q2, reveals two cointegrating vectors which are identified as a money demand equation and a long-run wealth growth relationship.

3.6 *De Bondt (2010)*

De Bondt (2010) focuses on the role of equity and labour markets. He finds evidence of a stable money demand equation when extending the opportunity cost vector to wealth, equity

returns and the annual change in the unemployment rate, thus highlighting the importance of financial channel and the existence of a novel precautionary motive. The first effect can be decoupled in a positive one coming from wealth (a financial transaction motive) and a negative one deriving from expected equity returns (a speculative motive). The second effect is less standard and represents a precautionary saving motive: money is hoarded when labour market weakens, i.e. unemployment rate rises. An improvement in labour 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. He specifies the vector \mathbf{X}_t as

$$\mathbf{X}_t = [o_t \quad e_t \quad w_t \quad u_t]',$$

where w_t is wealth and u_t is the unemployment growth rate. As for the estimation, conducted on the sample 1983:Q1-2007:Q2, the cointegration rank is always found to be equal to one, pointing at the existence of one cointegrating vector which is identified as a money demand equation.⁴ Providing evidence of such a stable money demand, he suggests to closely monitor equity and signals from labour market developments in order to assess the money growth path.

4 Time-varying cointegration

For any of the specifications considered, all the variables are collected into a k -dimensional column vector \mathbf{Z}_t , defined as $\mathbf{Z}_t = [(m_t - p_t) \quad y_t \quad \mathbf{X}_t']'$, where \mathbf{X}_t depends on the money demand equation considered. Following Bierens and Martins (2010b), we assume that for some t there are fixed $r < k$ linearly independent columns of the time-varying $k \times r$ matrix $\boldsymbol{\beta}_t = (\beta_{1t} \dots \beta_{rt})$, i.e. they are a basis for the time-varying space of cointegrating vectors. We consider the time-varying Vector Error Correction model (TV-VECM) of order p

$$\Delta \mathbf{Z}_t = \boldsymbol{\gamma}_0 + \boldsymbol{\Pi}_t \mathbf{Z}_{t-1} + \sum_{j=1}^{p-1} \boldsymbol{\Gamma}_j \Delta \mathbf{Z}_{t-j} + \boldsymbol{\varepsilon}_t, \quad \boldsymbol{\varepsilon}_t \stackrel{\text{i.i.d.}}{\sim} N_k(\mathbf{0}, \boldsymbol{\Omega}), \quad t = 1, \dots, T, \quad (6)$$

where T is the number of observations and $\boldsymbol{\Omega}$ and $\boldsymbol{\Gamma}_j$ are $k \times k$ matrices. We want to estimate (6) and to test the null-hypothesis of time-invariant (TI) cointegration $\boldsymbol{\Pi}_t = \boldsymbol{\Pi}$, against time-varying (TV) cointegration of the type $\boldsymbol{\Pi}_t = \boldsymbol{\alpha} \boldsymbol{\beta}_t'$, where $\boldsymbol{\alpha}$ is a fixed $k \times r$ matrix with rank r and $\boldsymbol{\beta}_t$ is a time-varying $k \times r$ matrix also with rank r . We could also think of time-varying speed of adjustments $\boldsymbol{\alpha}_t$. In this way, we could model time differences in short-run convergence to equilibrium. However, it should be clear from the literature presented in previous Sections and from Figure 2 that money overhang mainly concerns the residual of cointegration $\boldsymbol{\beta}_t' \mathbf{Z}_{t-1}$.

Bierens and Martins (2010b) (Lemma 1) prove that under standard smoothness conditions the parameters of the TV cointegrating vector $\boldsymbol{\beta}_t$ can be approximated by a finite sum of Chebyshev polynomials $P_{h,T}(t)$ of decreasing smoothness:

$$\boldsymbol{\beta}_t = \boldsymbol{\beta}_{m,t} = \sum_{i=0}^m \boldsymbol{\xi}_i P_{i,T}(t), \quad t = 1, \dots, T, \quad (7)$$

⁴The only exception to this result is represented by the specification without stock market variables, in which only precautionary motive is added to the classical framework: here the trace test indicates $r = 2$.

where $1 \leq m < T - 1$ and $\boldsymbol{\xi}_i = \frac{1}{T} \sum_{t=1}^T \boldsymbol{\beta}_t P_{i,T}(t)$, for $i = 0, \dots, m$ are unknown $k \times r$ matrices. Therefore, $\boldsymbol{\beta}_t$ is a function of rescaled time t/T , i.e. is defined in $[0, 1]$. Chebyshev time polynomials are defined by (see Figure 3)

$$\begin{aligned} P_{0,T}(t) &= 1, \\ P_{h,T}(t) &= \sqrt{2} \cos\left(h\pi \frac{t-0.5}{T}\right), \quad t = 1, \dots, T, \quad h \geq 1, \end{aligned}$$

such that, for all couples of integers i, j , we have $\frac{1}{T} \sum_{t=1}^T P_{i,T}(t) P_{j,T}(t) = \mathbf{1}_{i=j}$, i.e. they are orthonormal polynomials.

Testing for TI cointegration corresponds to the null and alternative hypothesis:

$$\begin{aligned} \text{TI)} \quad H_0 &: \boldsymbol{\xi}_i = \mathbf{0}_{k \times r} \text{ for } i = 1, \dots, T-1 \\ \text{TV)} \quad H_1 &: \lim_{T \rightarrow \infty} \boldsymbol{\xi}_i \neq \mathbf{0}_{k \times r} \text{ for some } i = 1, \dots, m, \\ &\text{and } \boldsymbol{\xi}_i = \mathbf{0}_{k \times r} \text{ for } i > m. \end{aligned}$$

Substituting (7) in (6), we get

$$\Delta \mathbf{Z}_t = \gamma_0 + \boldsymbol{\alpha} \left(\sum_{i=0}^m \boldsymbol{\xi}_i P_{i,T}(t) \right)' \mathbf{Z}_{t-1} + \sum_{j=1}^{p-1} \boldsymbol{\Gamma}_j \Delta \mathbf{Z}_{t-j} + \boldsymbol{\varepsilon}_t, \quad t = 1, \dots, T,$$

which can be rewritten as

$$\Delta \mathbf{Z}_t = \gamma_0 + \boldsymbol{\alpha} \boldsymbol{\xi}' \mathbf{Z}_{t-1}^{(m)} + \sum_{j=1}^{p-1} \boldsymbol{\Gamma}_j \Delta \mathbf{Z}_{t-j} + \boldsymbol{\varepsilon}_t, \quad t = 1, \dots, T, \quad (8)$$

where

$$\boldsymbol{\xi} = (\boldsymbol{\xi}'_0 \boldsymbol{\xi}'_1 \dots \boldsymbol{\xi}'_m)' \text{ and } \mathbf{Z}_{t-1}^{(h)} = \mathbf{Z}_{t-1} P_{h,T}(t), \text{ for } h = 0, \dots, m.$$

The null-hypothesis of TI cointegration corresponds to $\boldsymbol{\xi} = (\boldsymbol{\xi}'_0, \mathbf{0}_{r \times km})'$ so that $\boldsymbol{\xi}' \mathbf{Z}_{t-1}^{(m)} = \boldsymbol{\xi}'_0 \mathbf{Z}_{t-1}$. We can estimate both (8) and its time invariant counterpart (5), which is equivalent to (8) with $m = 0$, by Maximum Likelihood. Thus a likelihood ratio test for time-invariant cointegration emerges naturally as

$$LR_T = -2 \left[\hat{\ell}_T(r, m=0) - \hat{\ell}_T(r, m) \right], \quad (9)$$

where $\hat{\ell}_T(r, \cdot)$ are the log-likelihoods computed in the estimated values of the parameters. When $m = 0$ we are in a TI case, when $m > 0$ we are in a TV case. In both cases r is the cointegration rank which we assume to be given. Bierens and Martins (2010b) prove in Theorem 1 that, as $T \rightarrow \infty$,

$$LR_T \xrightarrow{d} \chi_{(rmk)}^2, \quad (10)$$

i.e. LR_T is asymptotically distributed as a Chi-squared distribution with rmk degrees of freedom. Bierens and Martins (2010a) compute empirical critical values via Monte Carlo simulations and show that already for $T = 100$ these values are very close to the asymptotic critical values. We thus base our test on the asymptotic distribution (10) and its quantiles.

5 Results

All the results presented in this section are obtained for one cointegration relation $r = 1$ and one Chebyshev polynomial $m = 1$. The first choice is motivated by the fact that in the Bierens and Martins (2010b) approach, a proper cointegration rank test is not available. When focusing on the case $r = 1$, the set of identifying restrictions reduces to a simple normalization choice. We leave to further research the identification problem of the number of cointegrating vectors. Still, likelihood-ratio tests for $r > 1$ already show that allowing for additional cointegration relationships strengthens instability. The second choice is motivated by Figure 3, being $m = 1$ the lightest admissible source of time variation. Not surprisingly, results from the likelihood-ratio test show that money demand instability is all but strengthened when allowing for more Chebyshev polynomials. Moreover, since we want our results to be the most “unaffected” as possible by turmoils in money and stock markets of 2008-2009, we choose to conduct the empirical analysis on quarterly data from 1980:Q1 to 2007:Q4. Again, including 2008 data could only strengthen the outcome in terms of instability. Finally, we show here only results for $p = 2$. Additional results for the likelihood-ratio test for different values of r , m and p are available upon request.

5.1 Time-varying cointegration test

We present here results obtained by applying the Bierens and Martins (2010b) test on the eight specifications presented in Section 3. Table 1 is dedicated to the three models taken as the starting point of the analysis, respectively by Calza *et al.* (2001), Coenen and Vega (2001) and Vlaar (2004). The second column recalls the vector \mathbf{X}_t for each of the specifications, the third column reports the sample considered, and the last two columns display, respectively, the computed LR_T statistics and the corresponding p -value under the asymptotic distribution (10). We first check models stability on the sample 1980:Q1-2001:Q4. Indeed, if the test rejects the null hypothesis even on the horizon in which money demand is recognized to be stable, one can cast doubts on its power, arguing that we are able to reject money demand stability just *by construction* and not because of the real nature of the process. Consistently with expectations, each specification is stable on the sample 1980:Q1-2001:Q4, with large p -values of 0.66, 0.79 and 0.49 respectively. On the contrary, when extending the estimation sample to 1980:Q1-2007:Q2, we reject the TI-VECM null-hypothesis.

Table 2 presents the same structure, but for papers claiming to find a stable relation once extending the explanatory set \mathbf{X}_t to new motives. The first row shows that the Dreger and Wolters (2009) hypothesis of achieving stability simply relaxing the short run homogeneity constraint is rejected by data, with a p -value of 0.00 both on data from 1983:Q1 to 2004:Q4 - the original sample in their contribution - and from 1980:Q1 to 2007:Q2. The second row presents an analogous result for the stock market downswing money demand in Carstensen (2006), exhibiting a p -value of 0.00 for the original sample as well as for the extended one. Also, results clearly do not support the hypothesis of equity and labour markets suggested by De Bondt (2010). Indeed, both the unrestricted specification, i.e. the one comprehensive of wealth and precautionary motives and the one based only on unemployment are rejected with a p -value equal to 0.00. When testing for the restricted specification which imposes a zero unemployment coefficient (i.e the “pure financial” specification), we find a p -value of 0.03 on the sample 1983:Q1-2007:Q2 and of 0.05 on 1980:Q1-2007:Q2. Finally, in the Beyer (2009) specification relying on financial wealth, and, in particular on the share of wealth allocated to

housing, stability is not supported by data. The LR_T statistics has again a p -value of 0.00.

The last and more interesting case is the one concerning the international portfolio allocation in De Santis *et al.* (2008). This is the only specification which supports the null hypothesis of a TI-VECM. Table 3 presents results of the test for this case. The first row reports results for the money demand equation as identified in De Santis *et al.* (2008), which is specified in terms of EA and US long run interest rates and price-earnings ratios. As the short-term interest rate is found to be not significant by De Santis *et al.* (2008) it is omitted and this result suggests a marginal role for monetary policy in controlling for $M3$ demand. The LR_T statistics is equal to 5.44, for a corresponding p -value of 0.49. From the second to the last row we proceed by sequentially excluding variables from the original specification, in order to identify the key determinants, if any. When removing ℓ_t^{US} the p -value is 0.77. When excluding both ℓ_t^{US} and ℓ_t^{EA} the p -value is 0.67. Hence, long term interest rates seem to be not fundamental. On the contrary, when ruling out either $(q_t - d_t)^{US}$ or $(q_t - d_t)^{EA}$, p -values drop to 0.01. This finding seem to be in favour of a decisive role by domestic and foreign stock markets in getting a stable EA money demand specification. We discuss in detail these results and their implications in the next Sections.

5.2 Long-run coefficients

As said above, when $m = 1$ the cointegrating vector is identified up to a normalization constant. Therefore, we can plot the time-varying long run elasticities β_t^y , respectively for the classical specifications in Calza *et al.* (2001), Coenen and Vega (2001) and Vlaar (2004) and for the “new motives” specifications in Dreger and Wolters (2009), Carstensen (2006), De Bondt (2010) and Beyer (2009). Some remarks are due.

First of all, we do not explicitly compute confidence intervals for the coefficient. Actually, as already highlighted in (7), our time-varying coefficients are given by the sum of two components: $\beta_t^y = \xi_0 + \xi_1 P_{1,T}(t)$, where the first term on the right-hand-side coincides with the time invariant estimate obtained in the standard time-invariant approach, while the second stems from Chebyshev polynomial of order one. Since significance of ξ_0 is assured by the analysis presented in the existing literature, we are only left to consider the significance of the second term. But then, as the LR_T test rejects the null hypothesis of time-invariant cointegration, we interpret this as evidence of $\xi_1 P_{1,T}(t)$ being statistically significant, otherwise we would accept the null hypothesis.

The long run income elasticity β_t^y is monotonically increasing in time along the whole the sample considered. Instability is immediately perceived in the classical specifications, i.e. Calza *et al.* (2001), Coenen and Vega (2001) and Vlaar (2004) (see Figure 4). Furthermore, Figure 5 shows β_t in Dreger and Wolters (2009), which is just based on relaxing the short run homogeneity constraint. It can be thus interpreted as an *intermediate* framework standing between classical and financial specifications. Indeed, we observe a long-run income elasticity varying from 2 to 10, while, in the same time the inflation coefficient varies from -0.03 to -0.5 .⁵ This time-varying behaviour of the inflation coefficient is totally consistent with our findings pointing to a key role for financial variables in stabilizing the long-run relation. In-

⁵Dreger and Wolters (2009) obtain a point estimate $\beta^y = 1.24$ and $\beta^\pi = -5.16$. Not allowing for a time-varying coefficients they overestimate β_t^π .

deed, β_t^π captures part of the instability stemming from the omission of stock market variables. However, from our results this is not sufficient (see Table 2). Finally, back to Figure 5 we observe almost constant long-run interest rates (semi)-elasticities. Analogous paths are found every time interest rates enter a specification and for all other variables.

From these results, we might think of a structural change in long-run income elasticity, which in turn could be linked to micro-parameters describing household preferences (e.g. risk aversion with respect to holding real balances, see Galí, 2008a). However, we believe this interpretation to be misleading and simply due to an omitted variable bias.

Actually, this bias is reduced, even if not eliminated, when considering extended sets of explanatory variables. When looking at Figure 6, β_t^y has a reduced range of variation, approximatively going from 1 to 2.6 in Carstensen (2006) specification and from 1 to 4 in De Bondt (2010) and Beyer (2009) formulations. Hence, we can say that using financial variables to explain money holdings helps in reducing instability, but it does not solve the problem as the time-invariant framework is rejected by the test. In Figure 8 we show the rate of increase of β_t^y for the considered specifications. The period of maximum increase is around 2001. This is precisely the period when the classical specifications start to reject the hypothesis of a stable money demand equation. A slowdown in the increase of income elasticity is observed in more recent years from 2002 to 2007.

6 Money demand in an open economy

As highlighted in the previous Section, we find no evidence of time-varying behavior when properly specifying money demand motives by linking money to domestic and foreign financial variables, namely long-term interest rates and price-earnings ratios in EA and US economy (see Figure 7 and Table 3). De Santis *et al.* (2008) consider jointly the utility maximization problem of domestic (EA) and foreign (US) households.⁶ They find evidence of four effects driving domestic (EA) money demand:

1. a monetary policy effect linked to interest rates;
2. a wealth effect linked to income;
3. a size effect which depends on relative wealth;
4. an international portfolio effect which depends on relative Sharpe ratios, proxied by price-earnings ratios.

The first two effects are present in all specifications considered in this paper. The third effect is due to a more attractiveness of domestic (EA) non-monetary assets which increases foreign (US) demand, the fourth effect is caused by an expected rise in domestic (EA) Sharpe ratios which increases the demand for money in the economy as a whole.

In a closed economy we cannot call for the two latter effects in order to stabilize EA money demand. Indeed, if the domestic (EA) Sharpe ratio increases any attempt to buy non-monetary assets would imply just a transfer of money from purchasers to sellers, leaving the

⁶Technical details of the model are in Section 2 of De Santis *et al.* (2008).

aggregate stock of M3 unchanged. Therefore, every time we observe a time-varying long-run income elasticity we are facing a phenomenon which is not driven by changes in structural parameters, but it is instead caused by the omission of domestic and foreign financial markets variables from the analysis. In particular, since we observe an increase in M3 growth rate, money is moving in from US to EA. Indeed, although M2 in US increases as well, its rate of growth is lower than in EA (see Figure 9.a).⁷ According to the model by De Santis *et al.* (2008), this flow is the effect of an increase of the EA price-earnings ratio relative to the US one, which makes Euro stocks more attractive for US investors, thus increasing the stock of M3 (see Figure 9.b).

Once the time-varying cointegration test has provided us with no evidence of time variation, we are allowed to limit ourselves to the standard time-invariant framework. As De Santis *et al.* (2008) we estimate the TI-VECM (5). We have

$$\mathbf{X}_t = [\ell_t^{EA} \quad \ell_t^{US} \quad (q_t - d_t)^{EA} \quad (q_t - d_t)^{US}]'$$

Once imposing the De Santis *et al.* (2008) restrictions, i.e. the perfect comovement between EA and US long-term interest rates and price-earnings ratios, the first identified cointegration relation is the international portfolio allocation money demand, (*t*-statistics in parenthesis)

$$\begin{aligned} (m_t - p_t) &= 1.87 y_t &+& 0.42 (q_t - d_t)^{EA} &-& 0.42 (q_t - d_t)^{US} &+ \\ &(35.23) &&(11.85) &&(-11.85) & \\ &+ 1.06 \ell_t^{EA} &-& 1.06 \ell_t^{US}. &&& \\ &(2.44) &&(-2.44) &&& \end{aligned}$$

A second and a third cointegrating relations, not shown here, are also estimated and denote EA and US financial markets equilibrium, respectively. The imposed overidentifying restrictions are accepted by the likelihood-ratio test statistics, distributed as $\chi_{(6)}^2$, with a value of 5.28 and *p*-value of 0.51.

Next, we test for exclusion of ℓ_t^{EA} and ℓ_t^{US} from the first cointegrating vector and we obtain (*t*-statistics in parenthesis)

$$\begin{aligned} (m_t - p_t) &= 1.78 y_t &+& 0.37 (q_t - d_t)^{EA} &-& 0.37 (q_t - d_t)^{US}. \\ &(36.79) &&(11.43) &&(-11.43) \end{aligned}$$

Also in this restricted case the overidentifying restrictions test statistics (including all three cointegrating relations) can not reject the null hypothesis of a restricted specification, the test is distributed as $\chi_{(7)}^2$ with a value of 7.99 and *p*-value equal to 0.33. This result is in line with our finding obtained by means of sequentially testing for time-varying stability of money demand (see Table 3). EA money demand stability is mainly driven by international portfolio allocations, and only indirectly by monetary policy (present in the second and third cointegrating relations). The disequilibrium paths of real balances in the De Santis *et al.* (2008) model and in the restricted model are reported in Figures 10 and 11 respectively and both show the requested mean-reverting behaviour.

⁷Since 2006 M3 in US is no more recorded, we thus follow the literature by considering M2 as the broadest measure of liquidity (see e.g. Favara and Giordani, 2009).

Finally, since our time-varying specification is considered only for just one cointegration relation, we repeat the previous analysis after conditioning on a cointegrating rank $r = 1$. The estimated cointegrating vector for the restricted case is (t -statistics in parenthesis)

$$(m_t - p_t) = 1.83 y_t + 0.25 (q_t - d_t)^{EA} - 0.34 (q_t - d_t)^{US}.$$

(38.16) (7.23) (-10.76)

The test for overidentifying restrictions is distributed as $\chi^2_{(3)}$ with a value of 1.78 and p -value equal to 0.62. Thus, once more we can not reject the hypothesis that monetary policy has not a direct effect on money demand, which is instead strongly affected by international portfolio choices. Therefore, we conclude in favour of a strong linkage between money and asset prices in a multicountry environment, which sounds promising to properly analyze financial stability issues.

7 Policy implications for financial stability

A natural question arising from previous Sections concerns policy implications of our money demand specification, if any. Indeed, the debate on money demand is a policy debate. How can we use the insights of this paper for supporting the need of a macro-prudential analysis? Recently, De Grauwe and Gros (2009) argue on the possibility that the ECB faces a trade-off between price stability and financial stability. They show that, if this is the case, a policy aimed at targeting price level or inflation may contribute to amplify asset prices bubbles, which eventually may lead to crashes. In particular, when technological developments trigger booms in asset markets or when “animal spirits” create a cycle of booms and busts a trade-off between price stability and financial stability arises. However, while price stability boils down to controlling Consumer Price Index, providing a clear definition of financial stability is not that straightforward. Here we follow Borio and Lowe (2002) in stating that the odds of financial instability are increased by sustained credit growth, measured by loans supplied by banks to non-financial corporations, joint to big booms in asset prices.

We study the link between M3 and loans: although monetary and credit aggregates refer to two different sides of banks balance sheet, they tend to comove. Figure 12 shows the growth of rates of nominal M3 and of nominal loans: the observed correlation is 0.71 on the sample 1980:Q1-2007:Q2. Interestingly, during the recent financial crisis, i.e. adding data until 2010:Q1, the correlation increases up to 0.77.

Estimating a short-run relation between real loans growth rate $\Delta(l_t - p_t)$ and real M3 growth rate $\Delta(m_t - p_t)$ would provide us with some evidence on the possible policy design. Hence, we model loans as a function of real balances, representing bank deposits (see Bernanke and Blinder, 1988; Driscoll, 2004), and interest rates (t -statistics in parenthesis):

$$\Delta(l_t - p_t) = -0.00 \Delta s_t + 0.74 \Delta(m_t - p_t), \quad \bar{R}^2 = 0.34.$$

(-0.59) (2.17)

This estimate shows a direct influence of money demand on loans supplied by banks but no appreciable effect of interest rates. Given results of previous Section, we can then run a regression of loans on the determinants of real balances, i.e. lagged income and price-earnings

ratios (t -statistics in parenthesis):

$$\Delta(l_t - p_t) = \underset{(4.72)}{1.58} \Delta y_{t-1} + \underset{(2.34)}{0.04} \Delta(q_{t-1} - d_{t-1})^{EA} - \underset{(-1.95)}{0.06} \Delta(q_{t-1} - d_{t-1})^{US}, \quad \bar{R}^2 = 0.53.$$

This close relation between loans and money demand explanatory variables suggests that, had ECB taken into account international financial markets that drive M3 growth diverging path, it would have detected the same warning signs of an upcoming financial bubble as those coming from credit aggregates. Similarly to our results, Giannone *et al.* (2009) point towards a role of M3 as an early predictor of financial crisis, by means of a policy counterfactual experiment. Finally, Cappiello *et al.* (2010) find that money demand affects loans supply.

8 Concluding remarks

Classical long-run money demand equation fails to explain the observed diverging path of M3 after 2001. Recent studies complement this literature by adding financial and labour market variables, claiming to find a stable relation. In this paper we systematically check for stability of the long-run money demand equation by using the Bierens and Martins (2010b) time-varying cointegration likelihood-ratio test.

Using this new methodology, we find evidence of time-varying coefficients for all (but one) the specifications considered. When taking into account the role of international portfolio allocation (De Santis *et al.*, 2008), we cannot reject a time-invariant relation. Moreover, our results suggest that the key determinant of the observed increase in M3 growth rate are EA and US price-earnings ratios and only indirectly long-term interest rates. This result can be explained in terms of international portfolio flows. Not considering these latter in the empirical analysis produces a time-varying long-run income elasticity due to flows of money in or out the EA.

Our findings have a double policy implication. First, while the interest rate is effective in controlling inflation in the goods market, it has lower influence on money and credit markets. Second, information coming from monetary aggregates and loans, can be employed in a macroprudential control perspective in order to achieve financial stability and consequently economic stability. In particular, our analysis on money demand shows that to avoid potential dangers to financial stability, the ECB should not be looking only at EA variables but also at international financial developments. In this sense, we provide empirical support for the New Two-Pillars Strategy proposed by De Grauwe and Gros (2009). Further theoretical and empirical research is desirable to deepen this issue.

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Table 1: Likelihood-ratio test. Classical specifications.

Ref.	\mathbf{X}_t	Sample	LR_T	p-value
Calza et al. (2001)	$(s_t - o_t), (\ell_t - o_t)$	1980:Q1-2001:Q4	2.39	0.66
		1980:Q1-2007:Q2	11.87	0.02
Coenen and Vega (2001)	$(\ell_t - s_t)$	1980:Q1-2001:Q4	1.05	0.79
		1980:Q1-2007:Q2	11.27	0.01
Vlaar (2004)	$(\ell_t - s_t), \pi_t$	1980:Q1-2001:Q4	3.44	0.49
		1980:Q1-2007:Q2	25.34	0.00

The TV long run money demand equation considered is

$$(m_t - p_t) = \beta_0 + \beta_t^y y_t + \beta_t^X \mathbf{X}_t, \text{ where}$$

$(m_t - p_t)$ = log-real balances;

y_t = log-income;

s_t = short-term interest rate;

ℓ_t = long-term interest rate;

o_t = own rate;

π_t = q-o-q inflation rate;

p -values under the null-hypothesis of LR_T distributed as $\chi_{(rmk)}^2$, with $r = 1$, $m = 1$ and k is the size of \mathbf{X}_t plus 2.

Table 2: Likelihood-ratio test. Financial and labor market specifications.

Ref.	\mathbf{X}_t	Sample	LR _T	p-value
Dreger and Wolters (2009)	s_t, ℓ_t, π_t	1983:Q1-2004:Q4	27.47	0.00
		1980:Q1-2007:Q2	25.13	0.00
Carstensen (2006)	$(s_t - o_t), (e_t - o_t), v_t$	1980:Q1-2003:Q2	16.63	0.00
		1980:Q1-2007:Q2	19.18	0.00
De Bondt (2010)	w_t, o_t, e_t, u_t	1983:Q1-2007:Q2	41.85	0.00
		1980:Q1-2007:Q2	33.65	0.00
	o_t, u_t	1983:Q1-2007:Q2	24.69	0.00
		1980:Q1-2007:Q2	18.39	0.00
	w_t, o_t, e_t	1983:Q1-2007:Q2	12.63	0.03
		1980:Q1-2007:Q2	10.96	0.05
Beyer (2009)	s_t, h_t, o_t, π_t	1980:Q2-2007:Q2	43.36	0.00

The TV long run money demand equation considered is

$$(m_t - p_t) = \beta_0 + \beta_t^y y_t + \beta_t^X \mathbf{X}_t, \text{ where}$$

- $(m_t - p_t)$ = log-real balances;
- y_t = log-income;
- s_t = short-term interest rate;
- ℓ_t = long-term interest rate;
- o_t = own rate;
- π_t = q-o-q inflation rate (in Beyer (2009) y-o-y inflation rate);
- e_t = equity return rate;
- v_t = volatility of returns;
- w_t = log-wealth;
- u_t = y-o-y unemployment growth rate;
- h_t = y-o-y housing wealth growth rate;

p -values under the null-hypothesis of LR_T distributed as $\chi_{(rmk)}^2$, with $r = 1$, $m = 1$ and k is the size of \mathbf{X}_t plus 2.

Table 3: Likelihood-ratio test. International portfolio allocation specification.

Ref.	\mathbf{X}_t	Sample	LR _T	p-value
De Santis et al. (2008)	$(q_t - d_t)^{EA}, (q_t - d_t)^{US}, \ell_t^{EA}, \ell_t^{US}$	1980:Q1-2007:Q2	5.44	0.49
	$(q_t - d_t)^{EA}, (q_t - d_t)^{US}, \ell_t^{EA}$	1980:Q1-2007:Q2	2.57	0.77
	$(q_t - d_t)^{EA}, (q_t - d_t)^{US}$	1980:Q1-2007:Q2	2.38	0.67
	$(q_t - d_t)^{EA}$	1980:Q1-2007:Q2	11.22	0.01
	$(q_t - d_t)^{US}$	1980:Q1-2007:Q2	12.72	0.01

The TV long run money demand equation considered is

$$(m_t - p_t) = \beta_0 + \beta_t^y y_t + \beta_t^X \mathbf{X}_t, \text{ where}$$

$(m_t - p_t)$ = log-real balances;

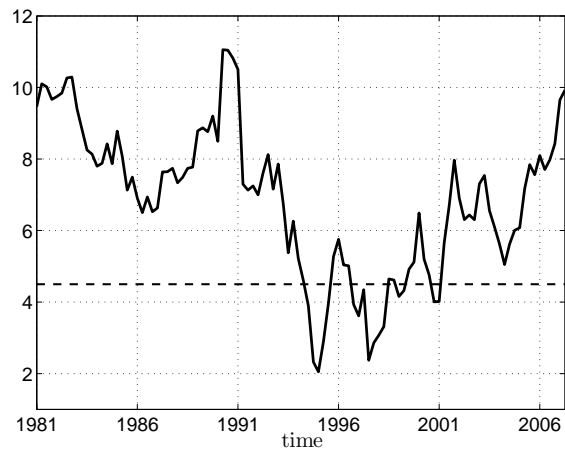
y_t = log-income;

$(q_t - d_t)$ = log-price to earnings ratio (EA or US);

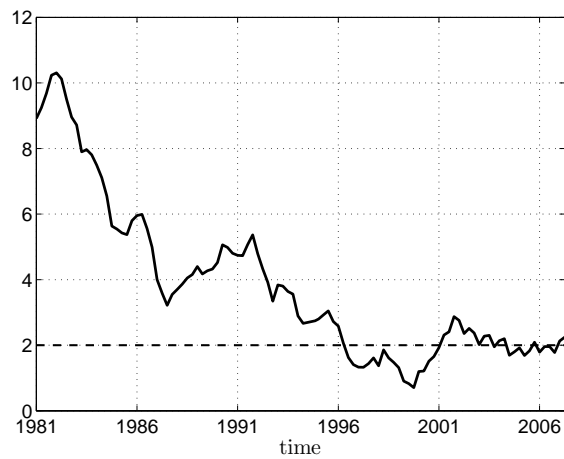
ℓ_t = long-term interest rate (EA or US);

p -values under the null-hypothesis of LR_T distributed as $\chi_{(rmk)}^2$, with $r = 1$, $m = 1$ and k is the size of \mathbf{X}_t plus 2.

Figure 1: Money growth and inflation in the Euro Area from 1980:Q1 to 2007:Q2.



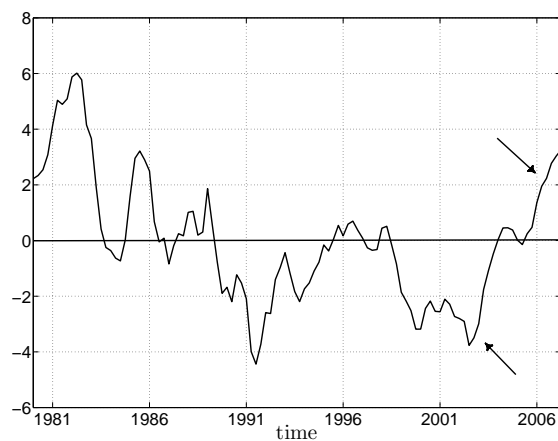
(a) Nominal M3 growth rate



(b) Inflation (GDP deflator growth rate)

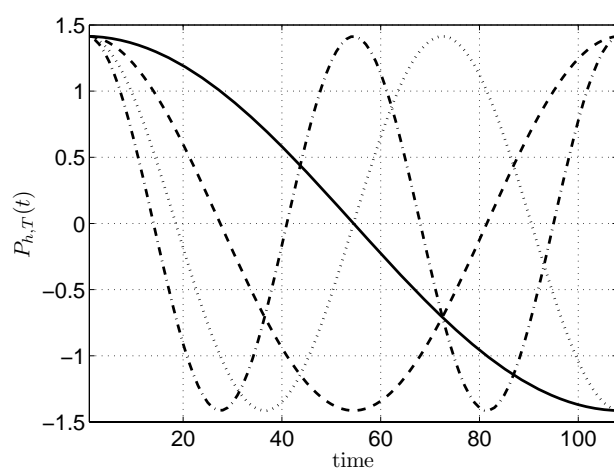
Solid lines: nominal M3 or inflation annual growth rate; dashed lines: reference value for M3, 4.5%, target value for infalction 2%. Annual percentage changes on the vertical axis.

Figure 2: Cointegration residual in Calza et al. (2001).



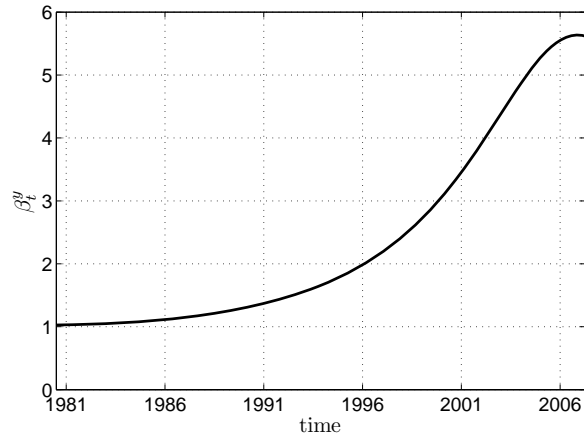
Note: the disequilibrium is computed as $(m_t - p_t) - 1.34y_t + 0.86(s_t - o_t)$. Values are re-scaled to average zero over the sample period, percent values on the vertical axis.

Figure 3: Chebyshev polynomials $P_{h,T}(t)$.

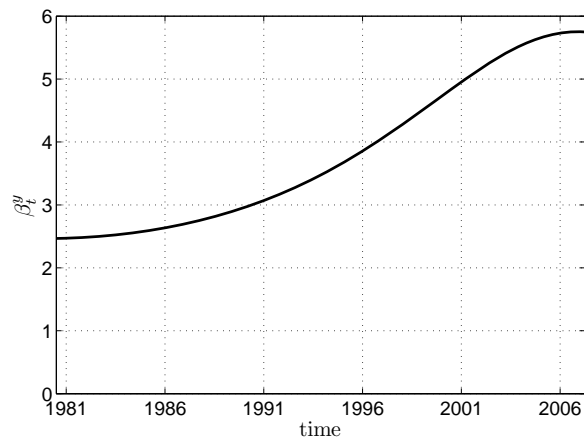


Solid line: $h = 1$. Dashed line: $h = 2$. Dotted line: $h = 3$. Dashed-dotted line: $h = 4$.

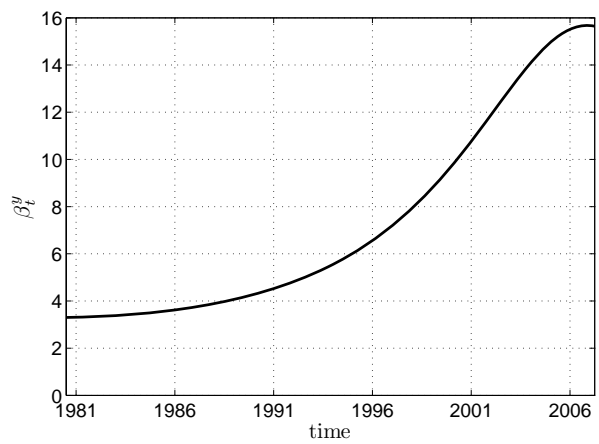
Figure 4: Estimated β_t^y . Classical specifications.



(a) Calza et al. (2001)

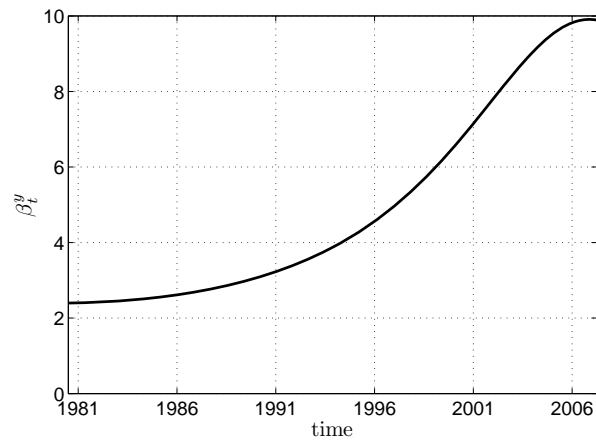


(b) Coenen Vega (2001)

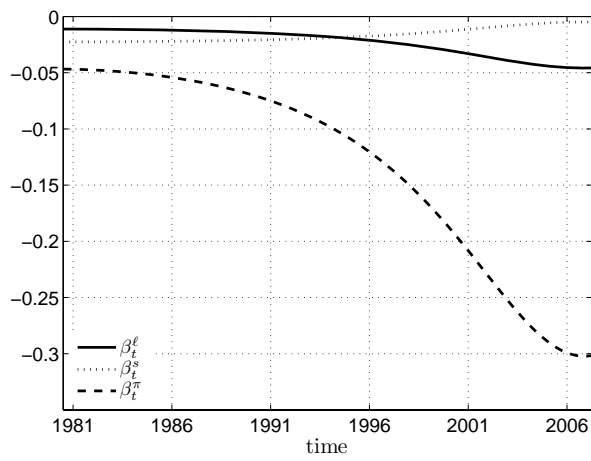


(c) Vlaar (2004)

Figure 5: Estimated β_t . Dreger and Wolters (2009) specification.

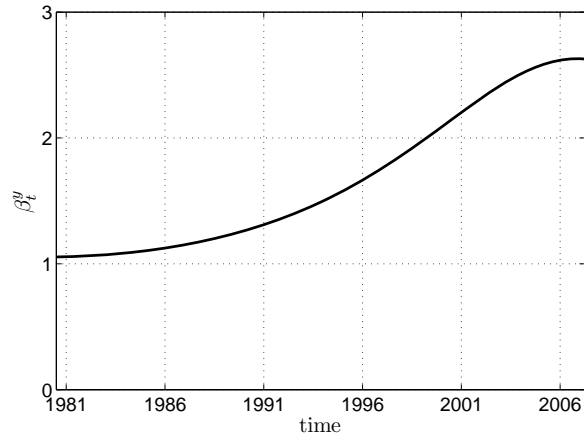


(a) β_t^y

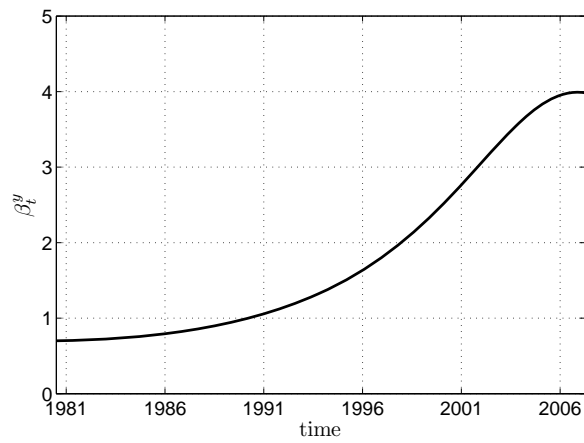


(b) β_t^x

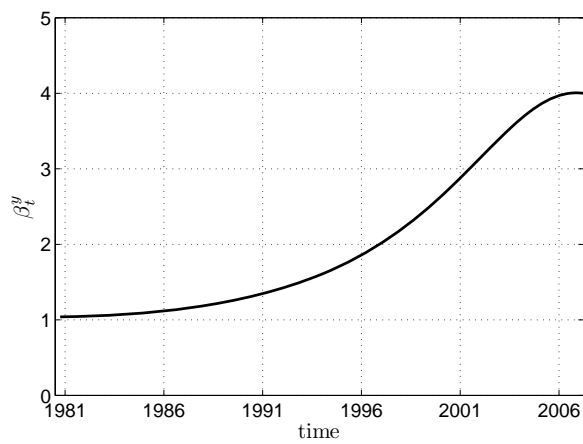
Figure 6: Estimated β_t^y . Financial and labor market specifications.



(a) Carstensen (2006)



(b) De Bondt (2010)



(c) Beyer (2009)

Figure 7: Estimated β_t^y . De Santis et al. (2008) specification.

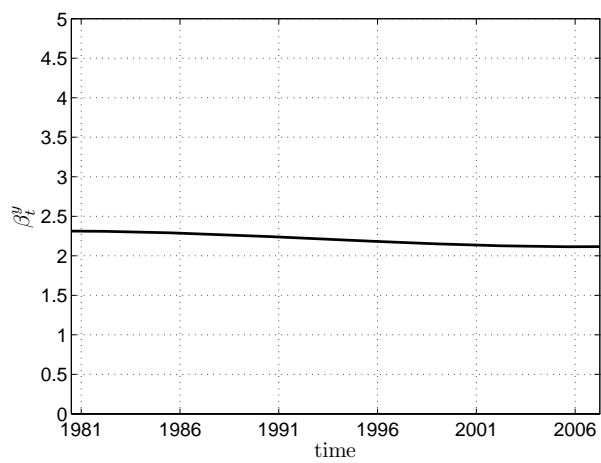
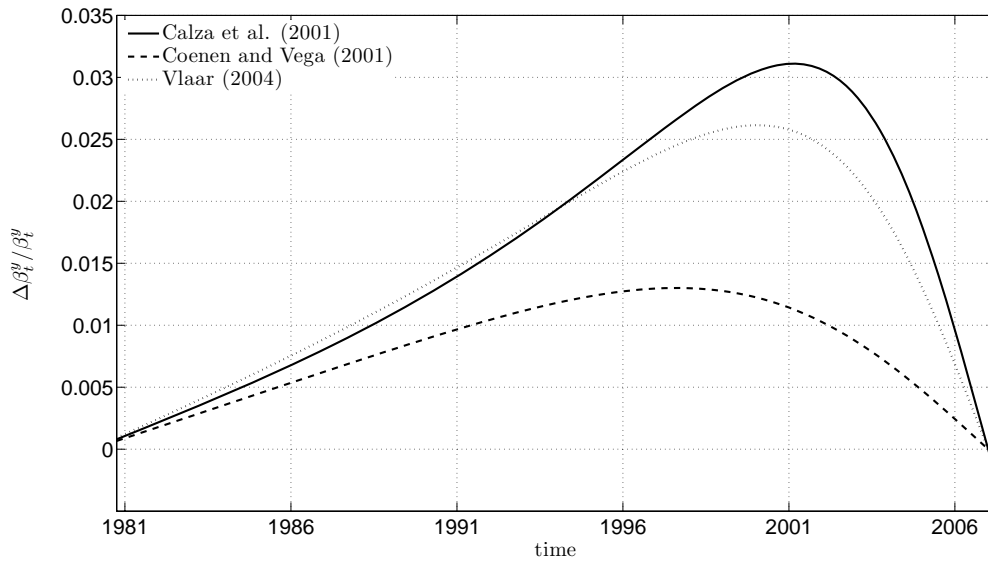
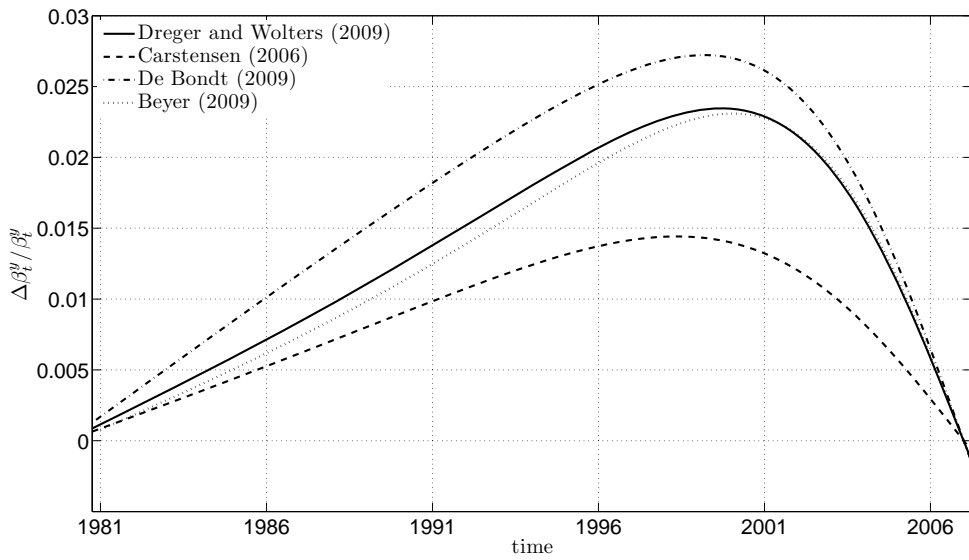


Figure 8: Estimated rate of increase: $\Delta\beta_t^y/\beta_t^y$.

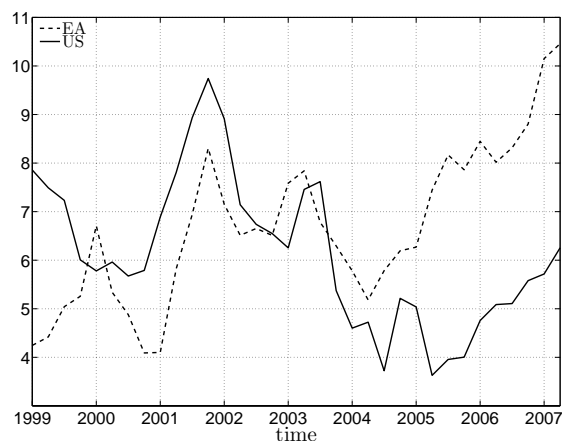


(a) Classical specifications.



(b) Financial and labor market specifications.

Figure 9: Money growth and price-earnings ratios from 1999:Q1 to 2007:Q2.



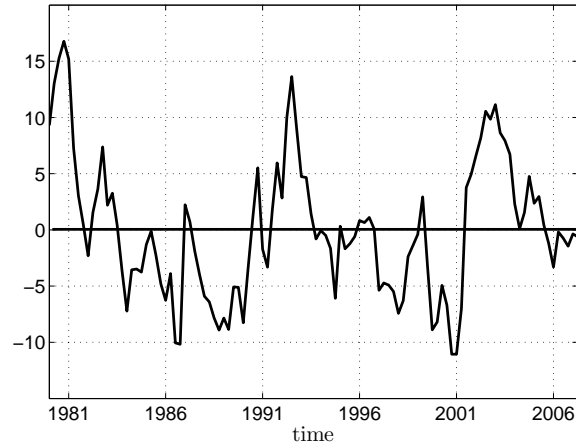
(a) Nominal money growth rates



(b) Relative price-earnings ratio

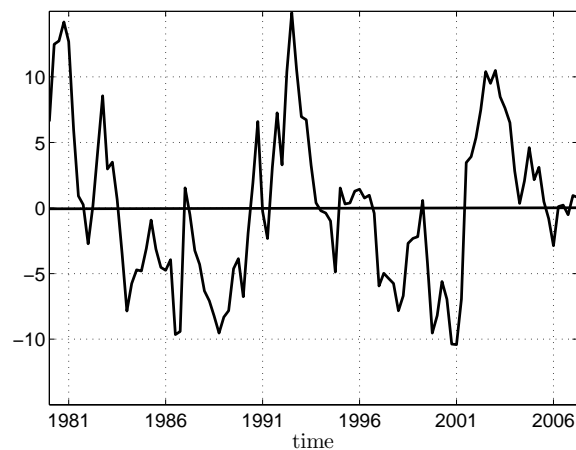
Top: solid line, US nominal M2 annual growth rate; dashed line, EA nominal M3 annual growth rate. Annual percentage changes on the vertical axis. Bottom: relative price-earnings ratio $(Q_t/D_t)^{EA}/(Q_t/D_t)^{US}$.

Figure 10: Cointegration residual in De Santis et al. (2008).



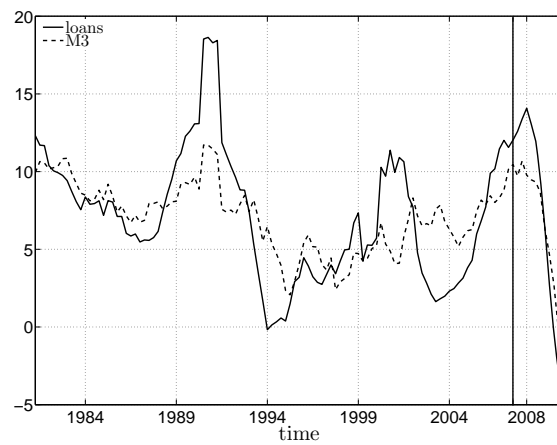
Note: the disequilibrium is computed as $(m_t - p_t) - 1.87y_t - 0.42(q_t - d_t)^{EA} + 0.42(q_t - d_t)^{US} - 1.06\ell_t^{EA} + 1.06\ell_t^{US}$. Values are re-scaled to average zero over the sample period, percent values on the vertical axis.

Figure 11: Cointegration residual in a pure stock market driven model.



Note: the disequilibrium is computed as $(m_t - p_t) - 1.78y_t - 0.37(q_t - d_t)^{EA} + 0.37(q_t - d_t)^{US}$. Values are re-scaled to average zero over the sample period, percent values on the vertical axis.

Figure 12: Loans and money demand growth rates from 1980:Q1 to 2010:Q1.

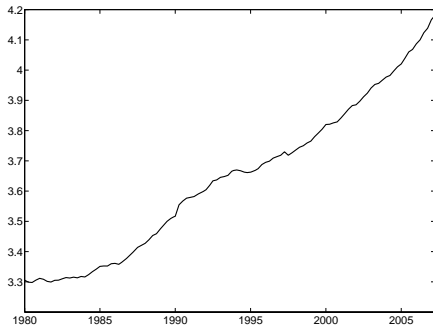


Solid line: nominal loans annual growth rate. Dashed line: nominal M3 annual growth rate. Annual percentage changes on the vertical axis. The vertical line indicates the end of our estimation sample 2007:Q2.

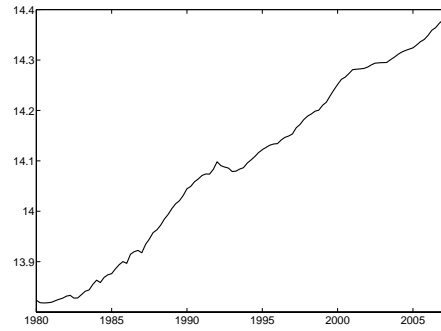
Appendix - Data description and transformation

- $(m_t - p_t) = \log(M_t/P_t)$ real balances, where
 - $m_t = \log(M_t)$ is M3 from ECB. Outstanding amounts at the end of the period. Series ID: BSI.M.U2.Y.V.M30.X.1.U2.2300.Z01.E;
 - $p_t = \log(P_t)$ is GDP deflator (YED) from Area Wide Model Database, see Fagan et al. (2001);
- $y_t = \log(Y_t)$ is real GDP (YER) from Area Wide Model Database, see Fagan et al. (2001);
- s_t is short-term interest (STR) from Area Wide Model Database, see Fagan et al. (2001);
- ℓ_t is long-term interest (LTR) from Area Wide Model Database, see Fagan et al. (2001);
- o_t is own rate based on ECB calculations, see e.g. CGR (2001) for details;
- $\pi_t = \Delta p_t$ or in Beyer (2009) $\pi_t = \Delta_4 p_t$ is inflation, where p_t is defined above;
- e_t is equity return based on German DAX30 for 1980-1986 and the Dow Jones Euro Stoxx50 for 1987-2007, from Datastream, see Carstensen (2006) for details;
- v_t is volatility of equity returns based on German DAX30 for 1980-1986 and the Dow Jones Euro Stoxx50 for 1987-2007, from Datastream, see Carstensen (2006) for details;
- $(q_t - d_t)^{EA} = \log(Q_t^{EA}/D_t^{EA})$ is US price to earnings ratio based on Datastream constituents for the EA, see De Santis et al. (2008) for details;
- $(q_t - d_t)^{US} = \log(Q_t^{US}/D_t^{US})$ is US price to earnings ratio based on Datastream constituents for the US, see De Santis et al. (2008) for details;
- s_t^{US} is 3-months money market rate on T-bills end of the month, from Federal Reserve;
- ℓ_t^{US} is 10-year US Treasury notes and bonds yields end of month, from Federal Reserve;
- $u_t = \Delta_4 U_t$, where U_t is unemployment rate (URX) from Area Wide Model Database, see Fagan et al. (2001);
- $w_t = \log W_t$, where W_t is wealth (WLN) from Area Wide Model Database, see Fagan et al. (2001);
- $h_t = \Delta_4(\log H_t)$, where H_t net household housing wealth at current replacement costs, ECB estimates, see Beyer (2009) for details.
- $(l_t - p_t) = \log(L_t/P_t)$ real loans, where
 - $m_t = \log(M_t)$ is loans from ECB. Outstanding amounts of loans to non-financial corporations at the end of the period. Series ID: BSI.Q.U2.N.A.A20.A.1.U2.2240.Z01.E;
 - $p_t = \log(P_t)$ is GDP deflator (YED) from Area Wide Model Database, see Fagan et al. (2001);

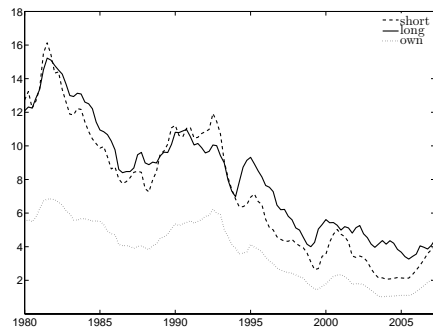
Figure 13: Data used in the analysis, for details see the Appendix and references therein.



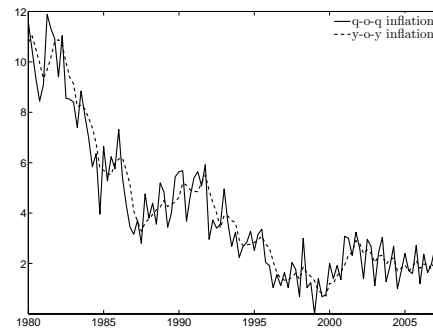
(a) real balances $(m_t - p_t)$



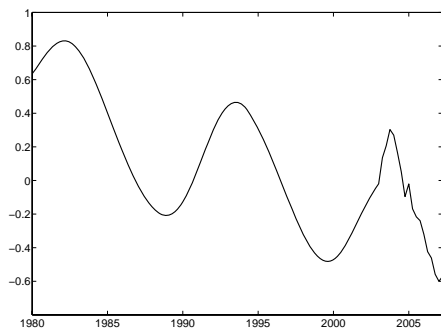
(b) GDP y_t



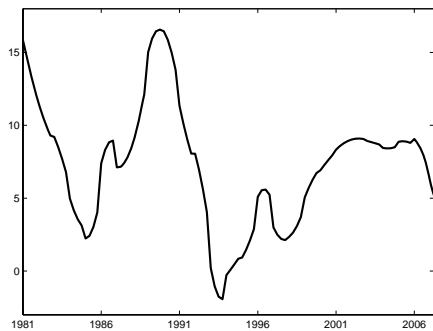
(c) EA interest rates s_t and l_t



(d) Inflation π_t

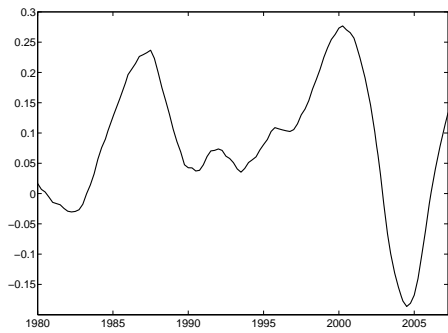


(e) Unemployment growth rate u_t

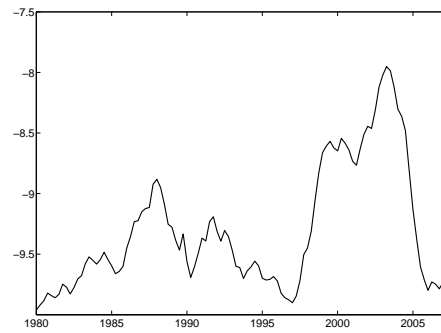


(f) Housing wealth growth rate h_t

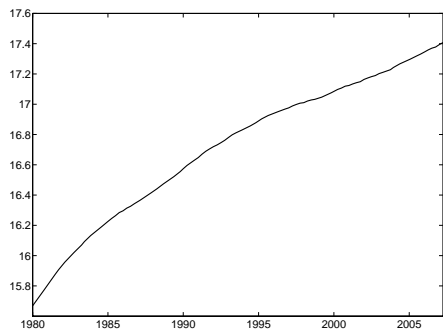
Figure 14: Data used in the analysis, for details see the Appendix and references therein.



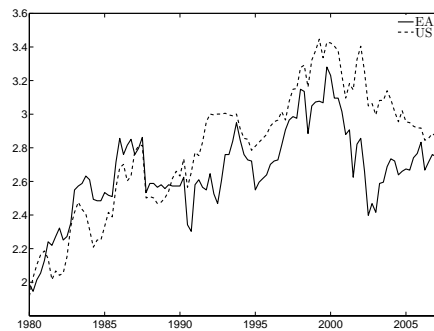
(a) Equity returns e_t



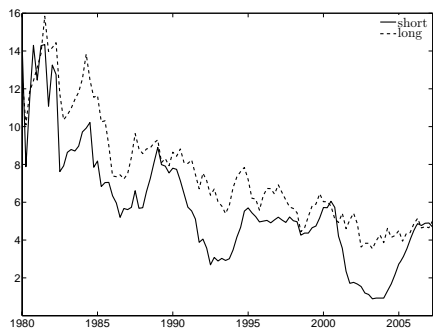
(b) Volatility of equity returns v_t



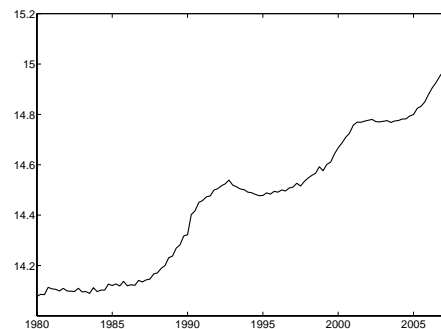
(c) Wealth w_t



(d) Price to earnings ratios $(q_t - d_t)$



(e) US interest rates s_t^{US} and l_t^{US}



(f) loans $(l_t - p_t)$