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QUANTITY THEORY IS ALIVE

THE ROLE OF INTERNATIONAL PORTFOLIO SHIFTS

by Roberto A. De Santis

In 2012 all ECB publications feature a motif taken from the €50 banknote.

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Abstract

We challenge the view that the relationship between money and prices is too loose in countries with low inflation rates and argue that cross-border portfolio shifts are the root cause of the volatility in real money balances. The novelty of this paper is that we model jointly in the euro area and the United States (i) the equilibrium in the money market that takes into account the cross-border portfolio shifts, and (ii) the equilibrium in the domestic asset markets, by finding a no-arbitrage relation between domestic long-horizon expected stock and bond returns. We estimate a stable money demand in the long-run and find that the short-run correlation between annual inflation and model-based excess money growth is not statistically different from unity in both the euro area and the United States. We also find that the resulting long-run equity risk premium comoves counter-cyclically with quarterly real GDP growth in both economies.
Keywords: Money demand, asset prices.
JEL classification: E31, E41, E51, E52, G58, F40

Executive Summary

Money can help forming policymakers' opinions, if one can detect the causes of monetary growth, if one can determine the stock of money holdings which is consistent with price stability in the long term and possibly if one can find a link between inflation and excess money growth in the short run, because if excess money growth is highly variable, it provides little useful information about the near-term outlook for spending or inflation.

The economic profession has shown that the quantity theory of money, which links money to goods prices, broke down after the 1980s in the United States due to changes in the monetary policy regime, in these times reacting sufficiently aggressively to incipient inflationary pressures. This is because central banks have increasingly stabilized inflation at low rates, while shocks to transaction technologies have made money growth more volatile.

We challenge the view that the relationship between money and prices is too loose in countries with low inflation rates and argue that cross-border portfolio shifts are the root cause of the volatility in real money balances.

We provide a new insight explaining why broad money demand in the euro area and the United States was considered to be unstable. The key message is that, with financial globalization and increased share of household savings in the asset markets, money velocity has become much more sensitive to domestic and foreign risky asset prices and, therefore, money balances have been fluctuating above what standard scale variables and domestic interest rates could capture.

The data show that there is a clear upward trend between 2001 and 2008 in the (inverse) euro area M3 velocity growth, which then turned negative after Lehman's bankruptcy, and that this trend is positively correlated with net capital flows. The correlation is also visible in the case of the United States from 1995 onwards.

Using this evidence, we model jointly in both the euro area and the United States the equilibrium in the money market, by finding a relation between real money balances, output and the difference between the yield curves of the United States and the euro area, and the equilibrium in the domestic asset markets, by finding a no-arbitrage relation between long-horizon expected returns in stock and bonds. We show that a link between money growth and inflation in both the euro area and the United States can be re-established in the long run as well as at relatively short-run frequencies once taking explicitly into account domestic and cross-border portfolio shifts.

1 Introduction

Sargent and Surico (2011) show that the quantity theory, established since Friedman (1956, 1959) and later renowned by Lucas (1980), broke down after the 1980s in the United States due to changes in the monetary policy regime, in these times reacting sufficiently aggressively to incipient inflationary pressures. Similarly, Teles and Uhlig (2010) find that while the theory is useful in high inflation OECD countries, the relationship between inflation and the growth rate of money in low-inflation OECD countries is tenuous at best in more recent times. Sargent and Surico (2011) and Teles and Uhlig (2010) assert that central banks have increasingly stabilized inflation at low rates, while shocks to transaction technologies have made money growth more volatile.[1]

In this paper, we show that a link between money growth and inflation can be re-established in the long run as well as at relatively short-run frequencies once taking explicitly into account domestic and cross-border portfolio shifts. With financial globalization and increased share of household savings in the asset markets, money velocity has become much more sensitive to global developments. Therefore, standard money demand equations fail because they have not properly treated the international portfolio channel.

Quantity theory builds upon the following identity:

$$m_t + v_t = p_t + y_t, \tag{1}$$

which states that (the log of) money (m_t) plus (the log of) velocity (v_t) is equal to (the log of) prices (p_t) plus (the log of) real output (y_t). The de-link between m_t and p_t is due to the difficulty of modelling v_t.

It is typically assumed that velocity is function of output (y_t) to capture the precautionary motive and technology developments in the payments process, and the interest rate (i_t) to capture the speculative motive:

$$v_t = (1 - \gamma^y)\, y_t + \gamma^i i_t - \gamma + \varepsilon_t. \tag{2}$$

where γ is a constant level-shift and ε_t denotes shocks to money velocity.

Given that monetary authorities use the interest rate as an instrument to achieve their objectives, money supply is demand-determined. Therefore, substituting (2)

[1]The inability to identify a stable M2 demand function in the United States after the mid-90s is widely recognized (see Carlson, et al., 2000; Duca, 2000; Choi and Oh, 2003; Duca and VanHoose, 2004; Choi and Cook, 2007). On the contrary, there are several studies showing a cointegrating relatioship between US M1, GDP and interest rates (among the most recent Choi and Oh, 2003; Ireland, 2009; Calza e Zaghini, 2010). We look at broad money, which include non-cash components that can usually be converted into cash very easily. We use M2 for the United States and employ M3 for the euro area, given its prominent role within the ECB as an indicator to assess risks to price stability.

into (1) yields the standard money demand specification estimated in the literature as well as by Teles and Uhlig (2010)[2]:

$$m_t - p_t = \gamma + \gamma^y y_t - \gamma^i i_t - \varepsilon_t. \tag{3}$$

It is useful to point out that γ^y is not restricted to be unity. Only if the precautionary motive and technology developments have a constant impact on velocity captured by the intercept, money velocity would be independent of output. If instead $\gamma^y > 1$, households demand more real balances that they really need to purchase consumption goods to account for unexpected circumstances; namely, the precautionary motive would prevail.[3] If instead $0 < \gamma^y < 1$, velocity would rise reflecting mainly technology developments. Clearly, γ^y cannot be negative, otherwise given the growth rate of money, a positive supply shock would be associated with higher inflation. γ^i is expected to be positive given that i_t is the opportunity cost of holding money.

Critics of the theory argue that money velocity is not stable, because ε_t is not stationary, partly owing to the fact that financial intermediation has grown in complexity and sophistication in the last thirty years.

The link between money growth and inflation in traditional money demand models relies on the hypothesis of a stationary velocity growth. Figure 1 shows that there is a clear upward trend between 2001 and 2008 in the (inverse) euro area M3 velocity growth, which then turned negative after Lehman's bankruptcy. This trend is positively correlated with net capital flows. The correlation is also visible in the case of the United States. However, we have to stress that the correlations are weakened by various factors. Most importantly, one should use the net capital flows in non-monetary financial institutions, as cross-border transactions between monetary institutions (e.g. banks) are not relevant for money, but these data are not available for the United States and are available for the euro area only. Nevertheless, this stylized fact suggests that international portfolio allocation could be a key explanation for the instability of traditional money demand specification (De Santis, et al., 2008).

Traditionally, the literature estimates money demand equations with a vector of

[2] Teles and Uhlig (2010) employ panel econometric techniques with random fixed effects to capture a permanent country-specific level-shift due to transaction technology developments. They investigate international cross-sections of countries, as carried out by McCandless and Weber (1995), and then concentrate on a pool of 14 countries, whose inflation rate was below 12 percent. To visually show the results, Teles and Uhlig (2010) plot the cross-section of the (1970-2005) average inflation rate against the cross-section of the (1970-2005) average excess money growth. They find that the relationship between inflation and the growth rate of money in low-inflation OECD countries is tenuous at best in more recent times.

[3] Hromcova (1998), for example, shows that after a technology shock individuals demand more real balances that they really need to purchase consumption goods due to the generated uncertainty. Their cash-in-advance constraint may become non-binding and the discrepancy between the growth rate of output and real balances is such that the income velocity is higher than unity.

domestic and foreign asset prices aiming at capturing the opportunity cost of holding money, include bond yields and the level of domestic stock prices (Friedman, 1988; Choudhry, 1996) or 3-year average of domestic quarterly stock returns (Carstensen, 2006).

Cross-border portfolio flows are ruled out in this literature because with complete information and complete markets asset prices immediately adjust. However, if one assumes asymmetric information (Brennan and Cao, 1997; Froot and Ramadorai, 2008) or international differences in opinions (Dumas, et al., 2011), capital flows become function of asset prices, ultimately affecting money holding.

The novelty of this paper is that we model jointly in the euro area and the United States (*i*) the equilibrium in the money market that takes into account the cross-border portfolio shifts, and (*ii*) the equilibrium in the domestic asset markets, by finding a no-arbitrage relation between domestic long-horizon expected stock and bond returns. In this respect, the suggested model is a step forward with a new model specification.

The key results of the paper over the sample period 1980Q1-2010Q4 are summarized in Figure 2 and Table 1, which look at the relationship between goods prices and excess monetary liquidity in the long and short run, respectively. Figure 2 plots the cointegrating residuals of the money demand equations as estimated in this paper (DS, henceforth). The plots show that in the long run excess monetary liquidity mean-reverts in both the euro area and the United States, as expected by quantity theory. Table 1 shows that in the short run the estimated slope between annual goods price inflation and annual excess money growth is strikingly not statistically different from unity in both the United States and the euro area. These results suggest that quantity theory is alive also in countries characterized by low inflation rates.

The remaining session of the paper are structured as follows: Section 2 describes the DS model and the traditional alternative specifications and looks at the main results; Section 3 analyses the statistical features of the DS model and presents additional results of the suggested two-country model; Section 4 concludes.

2 The empirical models

2.1 The DS model

Money demand functions typically include bond yields, stock prices or stock returns to control for the opportunity cost of holding money (i.e. Friedman, 1988; Choudhry, 1996; Carstensen, 2006). However, asset prices and cross-border portfolio shifts are not explicitly modelled.

The novelty of this paper is that we model jointly the money market equilibrium together with the asset price equilibrium between bonds and stocks in both the euro area and the United States.

2.1.1 The money market equilibrium

In order to identify the long run relatioships, vector error correction models ought to be parsimonious.

On the supply side, given that monetary authorities use the interest rate as an instrument to achieve their objectives, money supply is assumed to be perfectly elastic and, therefore, money is fully demand-determined.

On the demand side, to capture the cross-border portfolio shifts, we assume that assets are imperfect substitutes and investors make long-run decisions based on economic fundamentals. Insofar that the yield curve helps predicting future economic growth, as found in many studies (Estrella and Hardouvelis, 1991; Estrella, 2005; Bordo and Haubrich, 2008),[4] the yield curve in the United States and the euro area are good candidates to explain cross-border portfolio shifts and, as a result, money demand in these two economies. It should also be stressed that two-thirds of the global portfolios are invested in debt securities (De Santis, 2010) and historically the yield curve is one of the most important drivers of return in fixed-income portfolios. Therefore, we rely on the difference between 10-year and 3-month government bond yields to be a good proxy for the yield curve.

On this basis, the money demand specifications include the yield curve in the United States and the euro area as well as the price-earnings ratios:

$$\begin{aligned} m_t^j - p_t^j = \gamma^j + \gamma^{j,y} y_t^j + \gamma^{US,R}\left(R_{n,t}^{US} - i_t^{US}\right) + \gamma^{EA,R}\left(R_{n,t}^{EA} - i_t^{EA}\right) \\ + \gamma^{US,q}\left(q_t^{US} - e_t^{US}\right) + \gamma^{EA,q}\left(q_t^{EA} - e_t^{EA}\right) + \eta_t^{m,j}, \end{aligned} \tag{4}$$

where j denotes the country (euro area and United States), i_t^j is the short-term interest rate, $R_{n,t}^j$ is the n-period coupon bond yield and $q_t^j - e_t^j$ the price-earnings ratio.[5]

[4] Some authors have suggested that the yield curve is not longer predicting economic growth in the more recent period (Haubrich and Dombrosky, 1996; Stock and Watson, 2003; Giacomini and Rossi, 2006). For example, under a credible monetary regime with low persistence of inflation, a nominal shock will increase short interest rate, while only marginally affecting long-term interest rates. The yield curve is twisted, but this does not imply a negative impact on economic growth (Bordo and Haubrich, 2004). Nevertheless, it is fact that since 1990, the US yield curve has twisted 5 times and was always followed by lower economic growth in the United States. This evidence however could be due to developments in short-term yields, which are reduced in recessions in an effort to stimulate economic activity. Ang, et al. (2005) find that the short-term interest rate has more predictive power than any term spread.

[5] If cross-border portfolio assets matter for the price level, an alternative approach would be to include in the model the quantity of these assets. However, this would imply a general equilibrium approach and high quality data on portfolio assets, which is outside of the scope of this paper.

2.1.2 The no-arbitrage asset price equilibrium in the domestic market

To model asset prices, given the long run focus of the analysis, we make use of present-value models which link asset prices with long-horizon asset returns. In the specific, we consider the dynamic dividend growth model of Campbell and Shiller (1988) and the n-period coupon bond yield model of Shiller (1979).

Using a log-linear approximation to the returns on the stock market, Campbell and Shiller (1988) express the log of the stock price-dividend ratio in market j at the beginning of period t, $q_t^j - d_t^j$, as a linear function of (i) the future discounted dividend growth, Δd_{t+1}^j and (ii) the future expected equity returns, $E_t s_{t+1}^j$:

$$q_t^j - d_t^j = \frac{\kappa^j}{1-\rho^j} + E_t \left[\sum_{\tau=0}^{\infty} \rho^{j^\tau} \left(\Delta d_{t+1+\tau}^j - s_{t+1+\tau}^j \right) \right], \tag{5}$$

where $\kappa^j = -\log\left(\rho^j\right) - \left(1-\rho^j\right) \log\left(1/\rho^j - 1\right)$, $\rho^j = 1/\left(1 + \exp\left\{\overline{d^j - q^j}\right\}\right)$ and $\overline{d^j - q^j}$ is the steady-state level of the dividend-price ratio. Clearly, if the dividend-price ratio is high, and dividends are expected to growth only slowly, then stock returns are expected to increase.

Following Shiller (1979), the n-period coupon bond yield $(R_{n,t}^j)$ satisfies

$$R_{n,t}^j = \frac{1-\lambda^j}{1-\lambda^{j^n}} E_t \sum_{\tau=0}^{n-1} \lambda^{j^\tau} b_{n-\tau,t+1+\tau}^j + \Phi_{n,t}^j, \tag{6}$$

where $b_{n,t+1}^j$ is the one-period bond return on an n-period coupon bond, $\Phi_{n,t}^j$ is the term premium, $\lambda^j = 1/\left(1 + \overline{R^j}\right)$ and $\overline{R^j}$ is the steady-state level of the redemption yield. Short-term government bonds are a component of broad money, while long-term government bonds are perceived as substitutes for cash and short-term debt instruments, which give rise to a term premium.[6] Given the term premium, this relation says that the current yield to maturity should predict future changes in bond returns. Since the value of bonds depends on the money market conditions and given the impact of monetary policy on the term premium (Cox, et al., 1985; Campbell et al., 1997; Canzoneri et al., 2007), we assume that $\Phi_{n,t}^j = \delta^j i_t^j$. If the increase in the interest rate reduces expected inflation, the term premium declines, $\delta^j < 0$. However, when the short-term interest rate rises and investors prefer to allocate their funds into short-term debt instruments, there is an incentive to hold long-term bonds, if a higher term premium is offered. In this case, $\delta^j > 0$.[7]

[6] A long-term government bond must pay a higher risk premium, because both the inflation rate and the interest rate become more difficult to predict farther into the future. Such risk materialises only if the bondholder sells before maturity. Nevertheless, there is an opportunity cost, since the long-term bondholder forfeits the higher unexpected interest.

[7] Campbell et al. (1997, pg. 437) show that bond risk premia are a linear function of interest rates, whose sign can be either positive or negative, depending upon the covariance between consumption

Assuming a stationary dividend growth, a constant relation between the dividend- and earnings-yield (inverse of the price-earnings ratio),[8] and n relatively large, the use of (5) and (6) yields the long-horizon equity risk premium, namely the difference between a weighted average of expected future one-period stock returns and a weighted average of expected short-term interest rates:

$$E_t \sum_{\tau=0}^{\infty} \rho^{j^{\tau}} s^j_{t+1+\tau} - E_t \sum_{\tau=0}^{n-1} \lambda^{j^{\tau}} b^j_{n-\tau,t+1+\tau} = \mu^j - \left(q^j_t - e^j_t\right) - \frac{1}{1-\lambda^j} R^j_{n,t} + \frac{\delta^j}{1-\lambda^j} i^j_t, \quad (7)$$

where $\mu^j = \frac{1}{1-\rho^j}\left[\kappa^j + \Delta d^j + \left(1-\rho^j\right)\psi^j\right]$, Δd^j and ψ^j are the average growth rate in dividends and the payout ratio of stock j, and $q^j_t - e^j_t$ is (the log of) the price-earnings ratio of stock j.

Expression (7) resembles the "FED model" used by practitioners to measure expectations in the stock market (see Lander et al., 1997; Koivu et al., 2005; Bekaert and Engstrom, 2010).[9] If (the log of) the current price-earnings ratio is above the level requested by the current long-term interest rate, given the short-term interest rate, stock prices and/or long-term bond yields are expected to decline. This generates a portfolio reallocation between stocks and bonds until arbitrage opportunities across assets are cancelled. Therefore, an equilibrium condition between equity and bond markets is such that the long-horizon expected equity returns is equal to the long-horizon expected bond returns in country j:

$$q^j_t - e^j_t = \mu^j - \frac{1}{1-\lambda^j} R^j_{n,t} + \frac{\delta^j}{1-\lambda^j} i^j_t + \eta^{s,j}_t, \quad (8)$$

namely investors reallocate assets such that stock and bond prices move in the direction that mean-reverts $\eta^{sj}_t = E_t \sum_{\tau=0}^{n-1} \gamma^{j^{\tau}} b^j_{n-\tau,t+1+\tau} - \sum_{\tau=0}^{\infty} \rho^{j^{\tau}} E_t s^j_{t+1+\tau}$. η^{sj}_t are assumed to have mean zero and are normally distributed. The no-arbitrage condition in the domestic asset market is consistent with the idea that economic agents are rational and do not make systematic errors when they make portfolio decisions at time t based on information available at time *t-1*. The additive inverse of η^{sj}_t is the long run equity risk premium, which we can estimate.

innovations and revisions in expected future consumption growth. If such covariance is positive (negative), then a positive consumption shock drives up expected future consumption growth and increases (decreases) interest rates; the resulting fall (increase) in bond prices makes bonds covary negatively (positively) with consumption and gives them negative (positive) risk premia.

[8] The cointegration test between earnings yields and dividend yields in both the euro area and the US supports the hypothesis that the pay out ratio is stationary. The unit root tests indicate that dividend yield growth is I(0). The results are available from the authors upon request.

[9] The FED model states that if the price-earnings ratio is above the bond yield, equity prices are expected to decline until the long-run equilibrium between the two variables is re-established. This regularity was used as an input by Alan Greenspan in a famous speech on market's irrational exuberance in December 1996 (http://www.federalreserve.gov/boarddocs/speeches/1996/19961205.htm).

2.1.3 The VAR

Having established the long-run equilibrium in the money market, given by expression (4), and the long-run equilibrium in the asset markets, given by expression (8), we propose an empirical model capable of analyzing simultaneously the long-run equilibria in these markets in both the euro area and the United States based on the specification of the following VAR:

$$\mathbf{X}_t = \mathbf{A}(L)\mathbf{X}_{t-1} + v_t$$
$$\mathbf{X}_t' = \begin{bmatrix} m_t^{EA} - p_t^{EA}, y_t^{EA}, i_t^{EA}, q_t^{EA} - e_t^{EA}, R_t^{EA}, \\ m_t^{US} - p_t^{US}, y_t^{US}, i_t^{US}, q_t^{US} - e_t^{US}, R_t^{US} \end{bmatrix}. \tag{9}$$

where $m_t - p_t$ is the (log of the) real money balances, y_t is the (log of) real GDP, i_t is the short-term interest rate on bills, R_t is the yield to maturity of long-term bonds, $q_t - e_t$ is the (log of the) price-earnings ratio and v_t are independent normally distributed with mean zero and variance Ω. The Appendix describes briefly the dataset.

The aim is to identify for the euro area and one for the United States a money demand that would resemble (4), and an equilibrium condition between the log of the price-earnings ratio and short- and long-term bond yields that would resemble (8).

2.2 The alternative models

To benchmark the results of the proposed empirical model, we run standards alternative specifications also considered by Teles and Uhlig (2010). To save space and given that they do not add value to the analysis, we do not report the results based on the Baumol-Tobin and Miller-Orr money demand equations, which assume that money velocity is function of half and one third of the interest rate, respectively (Baumol, 1952; Tobin, 1956; Miller and Orr, 1966). As alternative specifications, following Lucas (2000) and Ireland (2009), we consider what we label log-log and semilog money demand equations.

The log-log money demand equation considers money velocity function of the log of the interest rate. Therefore, the long-run specification of the log-log error correction model takes the following form:

$$m_t^j - p_t^j = \gamma^j + y_t^j - \beta^j \log i_t^j + \eta_t^{l,j}.$$

The semilog money demand equation considers money velocity function of the interest rate. Therefore, the long-run specification of the semilog error correction model takes the following form:

$$m_t^j - p_t^j = \gamma^j + y_t^j - \beta^j i_t^j + \eta_t^{s,j}.$$

Lucas (2000) and Ireland (2009) use M1 given their focus on the welfare cost of inflation and, therefore, they employ the short-term interest rate as a measure of the opportunity cost of holding money. The use of broad money in the paper calls for employing the long-term interest rate. The results suggest that the interest rate is statistically significant and with the correct sign for the euro area, but insignificant for the United States (standard errors in brackets):

		Euro area	United States
Log-log	$\widehat{\beta}$	32.18	17.92
		(3.86)	(13.35)
Semilog	$\widehat{\beta}$	4.40	-1.31
		(0.66)	(1.45)

The results are very similar when using the short-term interest rates confirming the general difficulty in modelling US M2, also because standard test statistics reject the null that the income elasticity is one. Using M1, Lucas (2000) finds that the interest rate elasticity is equal to 0.5 and the semi-elasticity is equal to 7, while Ireland (2009) focusing on the post-1980 period (the sample period we also consider) finds that the interest rate elasticity is equal to 0.9 and that the semi-elasticity of the interest rate is equal to 1.8. Given these results, the long-run parameters estimated by Ireland (2009) are imposed on the alternative log-log and semilog specifications.

2.3 Comparing the results among alternative specifications

To make the models comparable, the lag length is set equal to 2 for all specifications chosen according to the Schwarz information criterion.

The main results are summarized in Figure 3, which shows GDP deflator against estimated excess liquidity that is defined as nominal money minus estimated real money demand.

The first panel plots good prices and nominal broad money corrected for real GDP. In quantity theory's terms, the series would be correlated if velocity is constant, as suggested by the Cambridge equation. The second and third panels plot good prices and excess monetary liquidity defined as nominal broad money minus real money demand as estimated using respectively the log-log and the semilog money demand specifications. They all indicate that the link between good prices and money is weak particularly since the beginning of the 1990s. As for the United States, although the log-log and semilog excess liquidity measures mean revert, the mean reversion is too slow. Therefore, the difference between the GDP deflator and excess liquidity is characterized by a unit root process, which makes money demand unstable when estimated using traditional approaches. The fourth panel plots good prices and excess monetary liquidity defined as nominal broad money minus real money demand as estimated in this paper, where velocity is function of key determinants of portfolio allocation, namely the difference between the euro area and the United States yield curve. The plots show that excess monetary liquidity estimated using the DS model fluctuates rapidly around good prices making the system stationary (see next section for statistical properties of the model).

It is very important to indicate that the change in trend in excess liquidity after Lehman's bankruptcy in 2008Q3 is due to the fall in money demand driven by the fall in real output in both the euro area and the United States (see also Section 3.1). This result is opaque using the alternative specifications.

Notice also that the rise in money velocity from 1990 through about 1995 in the United States was temporary given that excess liquidity mean-reverts in 1993. On the contrary, Carlson, et al. (2000) suggested that fall in money was permanent due to financial deregulation and innovation in the first half of 1990s.

Therefore, the analysis summarized in Figure 3 shows that the instability of money demand in the euro area and the United States is mainly due to the domestic and cross-border portfolio channels. When these channels are taken into consideration as in the DS specification, the system is stationary. Unexpected changes affecting the precautionary motive and technology shocks do have an impact on the relations between money, output and asset prices, but their effects are transitory as in the first half of 1990s and the end of 2000s.

Overall, the DS model is superior to the more standard specifications of the money demand equation, because the residuals of the DS model are "more stationary". It could be argued that this is not surprising given that the DS specification has simply more regressors, and that in the other specifications the coefficients are fixed, and not estimated. The estimated coefficients for the other specifications produce residuals that are not stationary. Moreover, the larger number of regressors in the DS model implies a larger number of restrictions needed to identify the system of equations, as it is explained in the next section. Therefore, there is no advantage in increasing the number of variables in the system.

The quality of the fit improves substantially also when looking at the short dynamics in real money growth. Table 2 reports the variance of the residuals as a fraction of the demeaned real money growth at various subsamples. The residuals of the DS specification have the lowest variance in both the euro area and the United States. During the 1990s, when it is believed that financial innovation in the United States spurred growth in velocity, the residuals are the smallest. This suggests that it is financial globalization with its implication on cross-border portfolio shifts and only partly financial innovation that spurred trend growth in velocity in the 1990s.

The next sections focus on the statistical properties and the results obtained using the DS model.

3 The cointegrated VAR model of money and asset prices

3.1 The long run equilibrium of the DS model

The standard unit root tests, such as Phillips-Perron (1998) tests, applied to each of the ten variables taken into consideration indicate that the null hypothesis of a unit root cannot be rejected. These results (available upon request) pave the way for tests of cointegration. The baseline empirical specification is the VAR (9). Suppose that (*i*) $|\mathbf{A}(\mathrm{L})|= 0$ has a unit root or roots greater than one,[10] (*ii*) $\mathbf{A}_1 = -\alpha\beta'$ has rank equal to r, which is lower than the number of variables, and (*iii*) I(2) processes are absent, then the VAR system can be cast in an isomorphic vector error correction form, where $\Pi = \alpha\beta'$ is the standard long-run matrix to be restricted in order to find r long-run economic relations $\beta'\mathbf{X}_t$ that can be made stationary. The idea is to eliminate common trends among variables by suitable linear combination, such that the long run economic relationships are stationary.

The Johansen's (1988) trace and max-eigenvalue tests on the number of cointegrating vectors allowing for a deterministic trend in the long-run equilibria both indicate five cointegrating vectors at the 5% significance level over the sample 1980Q1-2010Q4 (see Table 3). This implies that we have to find five linearly independent cointegrating relations in the cointegration space. Having determined the rank, two criteria are adopted to assess whether the model supports the economic hypothesis of interest: first, the restrictions on the cointegrating relations have to satisfy the Likelihood Ratio (LR) test, which is distributed as a $\chi^2(v)$ with degree of freedom given by $v = k - r(r-1)$, where k is the total number of linear restrictions in the cointegrating space; second, the residuals of the cointegrating relations ought to be individually stationary.

Therefore, we can write the model (9) in the vector error correction form as

$$\Delta\mathbf{X}_t = \varphi + \sum_{h=1}^{p-1}\Gamma_h\Delta\mathbf{X}_{t-h} + \Pi\mathbf{X}_{t-1} + u_t, \tag{10}$$

where φ is a vector of constant, Γ is a full coefficient matrix, $\mathbf{\Pi} = \alpha\beta'$ is the standard long-run matrix to be restricted in order to find the long-run economic relations $\beta'\mathbf{X}_t$, and p is the lag lenght of the VAR. The elements of α are known as the adjustment parameters in the vector error correction model and each column of β is a cointegrating vector. As already mentioned, according to the Schwarz information criterion, the lag length is set equal to 2.

Given that $r = 5$, the system can be identified if there are maximum five free parameters for each long run equation and $k \geq 20$. We first try to identify a symmetric cointegrating structure for the specifications of money demand and asset prices in the

[10] The assumption that the roots of $|\mathbf{A}(\mathrm{L})|$ are outside the unit root disk implies that the reciprocal values, which are the eigenvalues of Π, are less than 1 in absolute value, hence inside the unit circle.

euro area and the United States to capture the cross-border portfolio shifts as follows:

$$\begin{aligned}
m_t^{US} - p_t^{US} &= \beta_{1,0} + \beta_{1,7} y_t^{US} + \beta_{1,10} \left(R_t^{US} - i_t^{US}\right) + \beta_{1,5} \left(R_t^{EA} - i_t^{EA}\right) \\
&\quad + \beta_{1,9} \left(q_t^{US} - e_t^{US}\right) + \beta_{1,4} \left(q_t^{EA} - e_t^{EA}\right) \\
m_t^{EA} - p_t^{EA} &= \beta_{2,0} + \beta_{2,2} y_t^{EA} + \beta_{2,10} \left(R_t^{US} - i_t^{US}\right) + \beta_{2,5} \left(R_t^{EA} - i_t^{EA}\right) \\
&\quad + \beta_{2,9} \left(q_t^{US} - e_t^{US}\right) + \beta_{2,4} \left(q_t^{EA} - e_t^{EA}\right) \\
q_t^{US} - e_t^{US} &= \beta_{3,0} + \beta_{3,8} i_t^{US} + \beta_{3,10} R_t^{US} \\
q_t^{EA} - e_t^{EA} &= \beta_{4,0} + \beta_{4,3} i_t^{EA} + \beta_{4,5} R_t^{EA} \\
\left(y_t^{EA} - y_t^{US}\right) &= \beta_{5,0} + \beta_{5,3} \left(i_t^{EA} - i_t^{US}\right) + \beta_{5,5} \left(R_t^{EA} - R_t^{US}\right).
\end{aligned}$$

Specifically, the first and the second cointegrating vectors are consistent with the long-run money demand (4), where the yield curves and the price-earnings ratios in the United States and the euro area play a key role in domestic and cross-border portfolio allocation, ultimately affecting money. The third and fourth cointegrating vectors correspond to the domestic asset price equilibrium as derived in (8). The fifth vector pins down the equilibrium condition between the two economies real economic activities and asset prices to pin down the cross-border portfolio flows. This structure is rejected by the LR test for overidentifying restrictions and some of the variables are not statistically significant.

However, the following model cannot be rejected (standard errors are in parenthesis):

$$\begin{aligned}
m_t^{US} - p_t^{US} &= \beta_{1,0} + \underset{(0.07)}{1.14} y_t^{US} + \underset{(1.12)}{10.88} \left(R_t^{US} - i_t^{US}\right) - \underset{(1.54)}{9.00} \left(R_t^{EA} - i_t^{EA}\right) - \underset{(0.05)}{0.23} \left(q_t^{EA} - e_t^{EA}\right) \\
m_t^{EA} - p_t^{EA} &= \beta_{2,0} + \underset{(0.06)}{2.43} y_t^{EA} + \underset{(0.64)}{4.83} \left(R_t^{US} - i_t^{US}\right) - \underset{(1.09)}{7.89} \left(R_t^{EA} - i_t^{EA}\right) - \underset{(0.02)}{0.35} \left(q_t^{US} - e_t^{US}\right) \\
q_t^{US} - e_t^{US} &= \beta_{3,0} - \underset{(2.52)}{10.48} R_t^{US} \qquad (11) \\
q_t^{EA} - e_t^{EA} &= \beta_{4,0} + \underset{(1.98)}{14.11} i_t^{EA} - \underset{(2.35)}{21.58} R_t^{EA} \\
y_t^{EA} - y_t^{US} &= \beta_{5,0} + \underset{(0.14)}{0.61} \left(i_t^{EA} - i_t^{US}\right) + \underset{(0.15)}{2.28} R_t^{EA}.
\end{aligned}$$

The LR ratio test for the validity of the over-identifying restrictions distributed as a $\chi^2(12)$ takes a value of 20.17 with an associated tail probability of 0.064. This is an important result given the large numbers of parameters to be estimated and the large number of restrictions.

In addition, given that the yield curve's semi-elasticity on the money demand equations have similar size and opposite sign, the LR ratio test for the validity of the over-identifying restrictions is associated with a tail probability of 0.088, if the restriction is imposed only on the US money demand specification and of 0.043 if the restriction is imposed only on the euro area money demand specification.

Therefore, we can also use the following more restricted model, as the results of the analysis remain broadly invariant, except for the magnitude of the residuals of the US money demand as discussed at the end of the section:

$$\begin{aligned}
m_t^{US} - p_t^{US} &= \beta_{1,0} + \underset{(0.11)}{1.68} y_t^{US} + \underset{(1.76)}{16.38} \left[\left(R_t^{US} - i_t^{US} \right) - \left(R_t^{EA} - i_t^{EA} \right) \right] - \underset{(0.07)}{0.84} \left(q_t^{EA} - e_t^{EA} \right) \\
m_t^{EA} - p_t^{EA} &= \beta_{2,0} + \underset{(0.07)}{2.46} y_t^{EA} + \underset{(0.72)}{6.41} \left[\left(R_t^{US} - i_t^{US} \right) - \left(R_t^{EA} - i_t^{EA} \right) \right] - \underset{(0.02)}{0.39} \left(q_t^{US} - e_t^{US} \right) \\
q_t^{US} - e_t^{US} &= \beta_{3,0} - \underset{(2.65)}{11.56} R_t^{US} \qquad (12) \\
q_t^{EA} - e_t^{EA} &= \beta_{4,0} + \underset{(1.30)}{11.41} i_t^{EA} - \underset{(1.72)}{18.76} R_t^{EA} \\
y_t^{EA} - y_t^{US} &= \beta_{5,0} + \underset{(0.14)}{0.68} \left(i_t^{EA} - i_t^{US} \right) + \underset{(0.15)}{2.21} R_t^{EA}.
\end{aligned}$$

Figure 4 shows the time series of deviations from the equilibria implied by the five identified cointegrating relations. They all mean revert and are stationary.

The first two equations identify the long-run money demand in the United States and the euro area, respectively. The income elasticity is 2.4 and 2.5 for the euro area and 1.1 and 1.7 for the United States, estimated using the system of equation (11) and (12), respectively. We interpret the large income elasticity, as the result of households holding more real money balances than needed for transaction and speculative purposes. The economic crisis that unexpectedly hit the world at the end of 2008 can be used to make the point. After Lehman's bankruptcy on September 15 2008, the amount of real money balances above the transaction and the speculative motives mean-reverted sharply in the subsequent two-three quarters in the euro area and in the United States (see upper panel of Figure 4). This suggests that money velocity is strongly affected by precautionary motives generated by uncertainty. This result, however, is less robust for the United States given that the unity income elasticity cannot be rejected when using the system (11).

The long-run evidence speaks in favour of the importance of portfolio shifts for the determination of euro area and US money demand specifications. In particular, the estimated coefficients of the long-run relations show an important role for the difference between the US and the euro area yield curve in determining euro area and US money demands, but also for the foreign price-earnings ratio. If the global economy is expected to growth (i.e. steeper US yield curve), international capital flows into both the United States and the euro area, and this increases money holdings. If the price-earnings ratio of the other country rises, capital seems to move out of the country, thus reducing money holdings. Although the volatility of the money demand cointegrating residuals is far larger in the United States, the two series comove highly with a correlation coefficient amounting to 91% over the entire sample period. This suggests that they are driven by a common factor. Specifically, given the role of the United States in the global economy, in particular in the financial sector, the US yield curve relative to the euro area is often considered to be an important global benchmark that influences international capital flows and, therefore, money holdings.

The third and the fourth equations correspond to the long-run asset market equilibrium under no-arbitrage in the United States and the euro area, respectively, as derived in (8), and further investigated in Section 3.2. We can back out the structural parameters from the asset price equations. From equation (8), the mean value of $\frac{1}{1-\lambda^j}$ is equal to 18.76 for the euro area and 11.56 for the United States. Similarly, the mean value of $\frac{\delta^j}{1-\lambda^j}$ is equal to 11.41 for the euro area and nil for the United States. This implies that

Parameters	Euro area	United States
	System 11	
λ^j	0.95 (0.94, 0.96)	0.90 (0.82, 0.94)
δ^j	0.65 (0.39, 1.07)	0
	System 12	
λ^j	0.95 (0.93, 0.95)	0.91 (0.84, 0.94)
δ^j	0.61 (0.40, 0.91)	0

with the 95% confidence interval reported in brackets. As desired, the discount factors $\widehat{\lambda}^j$ are close to unity, with a lower value for the United States, suggesting that US consumers are relatively more impatient than euro area consumers. The response of the term premium to the short-term interest rate $\widehat{\delta}^j$ is positive for the euro area and nil for the United States. This implies that when the short-term interest rate rises in the euro area, there is an incentive to hold euro area long-term bonds, only if a higher term premium is offered.

The fifth equation suggests that if US real GDP rises above the euro area real GDP, the resulting US current account deficit can be financed if the US short-term interest rate rises relative to the euro area interest rates.

To investigate the issue of structural stability of our estimates, we use the Nyblom (1989) test, which evaluates the time-invariance of the entire parameter vector in the cointegrating space. It suggests that the system is stable at all possible sample splits, with the p-value of SupQ(t/T) equal to 0.54 and the p-value of meanQ(t/T) equal to 0.60. Thus, these results provide evidence for the stability of the parameters determining the long-run solution and for the validity of the identifying restrictions.

All variables respond significantly to some of the disequilibrium (see Table 4), which imply that they play an important role for the correct specification of the model. To address the role of each individual variable we can impose coefficient restrictions on the short run equation (α), additional coefficient restrictions on the long run vectors (β), and we can look at the adjusted R^2.

The test of a zero row in α is a condition needed to test whether a variable $\mathbf{X}_t$ can be considered exogenous for the long run stochastic path of the other variables in the system: $\alpha_{x,1} = \alpha_{x,2} = \alpha_{x,3} = \alpha_{x,4} = \alpha_{x,5} = 0$. The results suggest that none of the variables is exogenous at 5% level of significance:

	$m_t^{EA} - p_t^{EA}$	y_t^{EA}	i_t^{EA}	$q_t^{EA} - e_t^{EA}$	R_t^{EA}
$\chi^2(20)$	60.28	59.65	53.68	45.33	51.38
$p - value$	0.000	0.000	0.016	0.002	0.000
	$m_t^{US} - p_t^{US}$	y_t^{US}	i_t^{US}	$q_t^{US} - e_t^{US}$	R_t^{US}
$\chi^2(20)$	44.46	67.57	47.46	35.58	42.89
$p - value$	0.007	0.000	0.001	0.017	0.005

The test of a zero row in β is the equivalent of testing whether a variable $\mathbf{X}_t$ can be omitted altogether from the long-run relations: $\beta_{x,1} = \beta_{x,2} = \beta_{x,3} = \beta_{x,4} = \beta_{x,5} = 0$. The results suggest that none of the variables can be excluded from the cointegrated space at 5% significance level.

As for the explained variance of the variables, the adjusted R^2 ranges between 10% and 12% for the dynamics of US asset prices and between 15% and 36% for the dynamics of euro area asset prices; the adjusted R^2 is equal to 47% for euro area real GDP growth, 49% for US real GDP growth, 41% for euro area real M3 growth and 32% for US real M2 growth. The outcome that asset prices have a lower adjusted R^2 is consistent with the fact that they are more volatile than macroeconomic variables.

In summary, both domestic and foreign asset prices do play an important role to identify a stable money demand equation, because cross border portfolio flows and relative asset returns influence each other (Cohen, et al., 2001; Froot and Ramadorai, 2005 and 2008) and all these variables are fundamental for the dynamics of the system.[11]

Our empirical results also support the view put forward by Adrian, et al. (2010), who argue that banks and other intermediaries influence macroeconomic fluctuations through the determination of asset prices, as they found strong evidence that balance sheet aggregates of some financial intermediaries are informative to the evolution of asset prices. They conclude pointing out that the evolution of macroeconomic aggregates and risk premia are closely tied together via the functioning of financial intermediaries. Similarly, we show that balance sheet quantities (i.e. real money balances)

[11] As for the identification approach, I have also tested whether the USD/EUR exchange rate should enter the long-run specification. The hypothesis, however, is strongly rejected. I have also conducted a counterfactual experiment by orthogonalizing the exchange rate shocks with respect to all other innovations and by generating counterfactual time series via dynamic simulation of all variables in the system when exchange rate shocks are set to zero. The results show that the idiosyncratic exchange rate shocks have virtually no impact on all other variables in the system. Therefore, the exchange rate could be even validly excluded from the empirical model in that the two main sources of its variability are either idiosyncratic shocks, which do not have a significant impact on all other variables in the system, or an endogenous shock absorber response to disequilibria in the system that would not obviously cause any omitted variable problem.

matter in the determination of the evolution of risk premia through the short-run dynamics as indicated by the adjustment coefficients in Table 4 (see coefficients on ECT 1 and ECT 2).

Are the systems (11) and (12) the most likely long-run model, or there are other long-run relations that are less consistent with the theoretical model? This is an important issue, which we have addressed looking at a buttery of statistical tests (Likelihood ratio tests, significance of coefficients, recursive estimations of the parameters) and the stationarity of the residuals. In principle, combining the first and forth co-integrating relations, US long run money demand could be written only in terms of interest rates. However, the estimation of the smaller system with 3 cointegrating vectors, which excludes the price-earnings ratios, fails to identify the imposed restrictions. This is because the gap between equity prices and bond yields generates an equity risk premium which is mean reverting and is influencing real money balances and real GDP (see coefficients on ECT 3 and ECT 4 of Table 4). The equity risk premia therefore need to be part of the system.

The only other alternative empirical model we could find is the following:

$$\begin{aligned}
m_t^{US} - p_t^{US} &= \beta_{1,0} + \underset{(0.05)}{1.08} y_t^{US} + \underset{(0.91)}{8.83} \left[\left(R_t^{US} - i_t^{US} \right) - \left(R_t^{EA} - i_t^{EA} \right) \right] - \underset{(0.05)}{0.31} \left(q_t^{EA} - e_t^{EA} \right) \\
m_t^{EA} - p_t^{EA} &= \beta_{2,0} + \underset{(0.07)}{2.43} y_t^{EA} + \underset{(0.73)}{7.39} i_t^{EA} - \underset{(0.50)}{7.55} \left(R_t^{EA} - R_t^{US} \right) \\
q_t^{US} - e_t^{US} &= \beta_{3,0} - \underset{(0.97)}{9.78} i_t^{US} - \underset{(0.97)}{9.78} R_t^{US} \\
q_t^{EA} - e_t^{EA} &= \beta_{4,0} + \underset{(1.51)}{12.70} i_t^{EA} - \underset{(1.89)}{21.26} R_t^{EA} \\
y_t^{EA} - y_t^{US} &= \beta_{5,0} + \underset{(0.15)}{0.94} \left(i_t^{EA} - i_t^{US} \right) + \underset{(0.12)}{1.86} R_t^{EA}.
\end{aligned} \tag{13}$$

The LR ratio test for the validity of the over-identifying restrictions distributed as a $\chi^2(15)$ takes a value of 21.62 with an associated tail probability of 0.12. The key difference in terms of parameters is that the income elasticity for the United States is smaller than in (12) and only marginally smaller than in (11), but the symmetry between the money demand equations in the euro area and the United States does not longer holds. Also the asset price equilibrium for the United States differs but only to the extent that a weighted average between short and long term bond yields plays a role. As for the cointegrating residuals, the residuals of the money demand specifications are smaller, but have the same developments, due to the income elasticity (see Figure 5). Conversely, the residuals of the asset price equilibrium are generally larger. The results of the remainder of the paper are broadly similar to the three identified specifications. Therefore, to save space, we report only the results based on (12).

3.2 The equity risk premium and real economic activity

Practitioners argue that the difference between the earnings yield and the long-term bond yield is an indicator for how much relative profit an investment in stocks yields for each unit invested compared with an investment in long-term bonds. The higher such gap, the higher the equity premium one would likely demand. We have formalized this principle in expression (8) and estimated it within the system of equations (12).

We have argued that the estimated cointegrating residuals can be used to compute the long-horizon equity risk premium. Given that the bond yield benchmark is a 10 year bond, the annualized long-horizon equity risk premium can be approximated dividing the cointegrating residuals by 10.

Panel A of Figure 6 indicates that long-horizon equity risk premia in the euro area and the United States comove strongly, as the correlation coefficient is equal to 71% over the entire sample period and rising to 82% after 1994. This implies that euro area and US long-horizon equity risk premia might be driven by a common factor, which could be related to financial globalization and the resulting massive cross-border portfolio flows.

High volatility and low comovement in euro area and US equity risk premia are evident in the early 1980s during the FED Charman Paul Volker's stabilization period and in 1992-1993 as a result of the European exchange rate mechanism's crisis. High volatility and strong comovement of the equity risk premium are instead evident during the Dotcom bubble at the turn of the century and since 2007 as a result of the global financial crisis, which impaired the functioning of the money market in early August 2007.

The results also suggest that the long-horizon equity risk premia have been rising since the burst of the equity bubble in 2000 reaching the highest level in 2008Q4 with the bankruptcy of Lehman Brothers and that were not low after 2003.[12]

Equity premia in the long run are expected to be countercyclical because investors do not like to take on risk in bad times. There is evidence that expected stock returns are countercyclical (Campbell and Diebold, 2009). If stocks provided insurance in the long run against current negative output shocks, the correlation between long-horizon equity risk premium and consumption growth would be negative.

Panel B of Figure 6 show that the estimated quarterly long-horizon equity risk premium (additive inverse) comove tightly with the quarterly real GDP growth in both the euro area and the United States, with a correlation coefficient amounting to -28.7% for the euro area and -43.4% for the United States over the whole sample

[12] It is often argued that risk premia were low over this period. We would agree only as regards the bond premia. The savings glut hypothesis put forward by the FED Chairman Bernanke ("The global saving glut and the U.S. current account deficit", remarks at the Sandridge Lecture, 19 March 2005) postulates that the global economy experienced a positive savings shock causing a reduction in bond premia.

period and rising (in absolute value) respectively to -60.3% and -59.2% after 1994. It is worth emphasizing that the long-horizon equity risk premium and real economic growth decoupled during the years before the 2008-2009 recession. Over the period 2002-2007, the premium justified lower real GDP growth, as then it turned out to be.

Panel C of Figure 6 shows that the estimated quarterly long-horizon equity risk premia (additive inverse) comove tightly also with the quarterly consumer confidence over the next 12 months (-53.9% for the euro area and 70.4% for the United States both calculated over the period 1985-2010 given the lack of data for the euro area in the first half of 1980s). This relationship is particularly tight for the United States with correlation coefficients of 83.2% after 1994. The different developments - vis-à-vis GDP growth - in the forward looking information provided by consumer confidence, which was very low in the United States in 2009 and 2010, might help explaining the recent decoupling in the long-horizon equity risk premia in the two economies.

It is important to mention that the estimated equity premia and quarterly real GDP growth share a similar stochastic process given that their standard deviations are of similar magnitude:

	Euro area	United States
1980Q1-2010Q4		
Equity Risk Premia	2.70	2.36
GDP growth	2.38	3.06
1980Q1-1994Q4		
Equity Risk Premia	2.19	1.20
GDP growth	2.36	3.47
1995Q1-2010Q4		
Equity Risk Premia	3.06	3.04
GDP growth	2.41	2.65

As pointed out by Campbell and Cochrane (1999, p. 207), a slowly time-varying countercyclical risk premium is important for matching asset pricing data.

3.3 Short-run dynamics: Inflation and money growth

The final exercise carried out to assess the quality of the results consists of investigating the relationship between inflation and money growth. The first panel of Figure 7 shows the scatter plot of inflation against nominal money growth corrected for real GDP growth. The relationship between the variables is very loose. If instead nominal money growth is corrected by real money demand growth estimated using the DS model, then one get a nice fit to the 45 degree line. The relationship is particularly tight when smoothing the quarterly series by taking the 8-quarter sum and, therefore, a medium-term perspective, as reported in the panel below.

It is useful to point out that two-points that are distant from the 45 degree line in the US scatter plot are related to the shocks in September 2001 (terrorist attacks) and September 2008 (Lehman bankruptcy). This evidence supports the hypothesis that uncertainty brings about an increase in money holding. Instead, the rapid growth in US M2 in 1983Q1 is due to the introduction of a new deposit account designed for investment. In response to the mandates of the Garn-St Germain Act, all depository institutions were authorized to offer money market deposit accounts (MMDA's) on December 14, 1982. Growth of MMDA's came at the expenses of large time deposits and money market mutual funds held by institutions, which are non included in M2 (Miller, et al., 1983). This is a good example of a transaction technology shock.

By simply excluding these three outliers, the scatter plots between US inflation and excess liquidity become very close to the 45 degree line (see Figure 8).

To test the link between inflation and excess money growth, we use annual observations and estimate the slope between inflation and excess money growth. The results are reported in Table 1 in the introductory section. First, we estimate the coefficients on nominal money growth and real money demand growth separately. The results show that they have the correct sign and the size of the coefficients – based on a Wald test restriction – is of the same magnitude. Then, we restimate the models on the excess money growth, namely the difference between nominal money growth and real money demand growth. The slopes are equal to 0.97 for the euro area and 0.84 for the United States and they are not statistically different from unity. The final specification controls for the three dummy variables to capture the uncertainty shocks in September 2001 (terrorist attacks) and September 2008 (Lehman bankruptcy) and the transaction technology shock in 1983Q1, when a new deposit account was introduced in the United States in response to the mandates of the Garn-St Germain Act. The results improve further, as the slopes are equal to 0.98 for the euro area and strikingly equal to 1 for the United States, suggesting that quantity theory is alive.

4 Conclusions

This paper shows that quantity theory is still alive also in countries with low inflation rates, challenging thereby the current view in the profession. We argue that the relationship between money and prices can be established in both the medium and long run, if domestic and cross-border portfolio shifts are considered.

The novelty of this paper is that we model jointly in both the euro area and the United States the equilibrium in the money market, by finding a relation between real money balances, output and the difference between the yield curves of the United States and the euro area, and the equilibrium in the domestic asset markets, by finding a no-arbitrage relation between long-horizon expected returns in stock and bonds. The two-country system is stationary in the post 1980-period.

We have provided a new insight explaining why broad money demand in the euro area and the United States was considered to be unstable. The key message is that, with financial globalization and increased share of household savings in the asset markets, money velocity has become much more sensitive to domestic and foreign risky asset prices and, therefore, money balances have been fluctuating above what standard scale variables and domestic interest rates could capture.

Three additional considerations might be useful: (i) the volatility of the excess monetary liquidity is far larger in the United States than in the euro area, explaining why money plays a more important role as an economic indicator in the euro area; (ii) the sharp change in trend in excess liquidity after Lehman's bankruptcy in September 2008 is due to the uncertainty generated by the sudden fall in real output; (iii) the fall in excess money at the beginning of the 1990s in the United States was not permanent contrary to the findings by Carlson, et al. (2000), who argued in favour of financial deregulation and innovation having a permanent effect on US M2 in the first half of 1990s.

Appendix: The dataset

The variables that enter the system are: money, output, price-earnings ratio, long-term bond yields and short-term interest rates for the euro area and the United States. We make use of historical series of quarterly data for the euro area and the United States over the period 1980Q1 to 2010Q4 for which high quality data for the euro area are available. Except for real GDP and GDP deflator, all other variables are measured as end-of-period. Real GDP, GDP deflator and money balances are seasonally adjusted. Except for the interest rates, variables are expressed in logarithms.

Given that cash and money market instruments are very close substitutes, we use broad money concepts, which include retail deposit sweep programs as well as short term debt instruments. As regards the euro area, the real M3 holdings are calculated as the nominal broad monetary aggregate M3 deflated by the euro area GDP deflator. With regard to the financial variables, the short-term interest rate for the euro area is a weighted average of the national three-month interbank interest rates up to end of 1998, and then Euribor afterwards. Similarly, the long-term interest rate is constructed as a weighted average of the yields on the national ten-year government bonds or their closest substitutes. The ECB is the source of these data.

For the United States, the real M2 holdings are calculated as the nominal broad monetary aggregate M2 deflated by the US GDP deflator, as provided by the US FED. The short- and long-term interest rates correspond respectively to the yields on the three-month US Treasury bills and to the yields on the ten-year US Treasury notes and bonds. The US yields are provided by the BIS. The price-earnings ratios for the euro area and the United States are obtained from Datastream.

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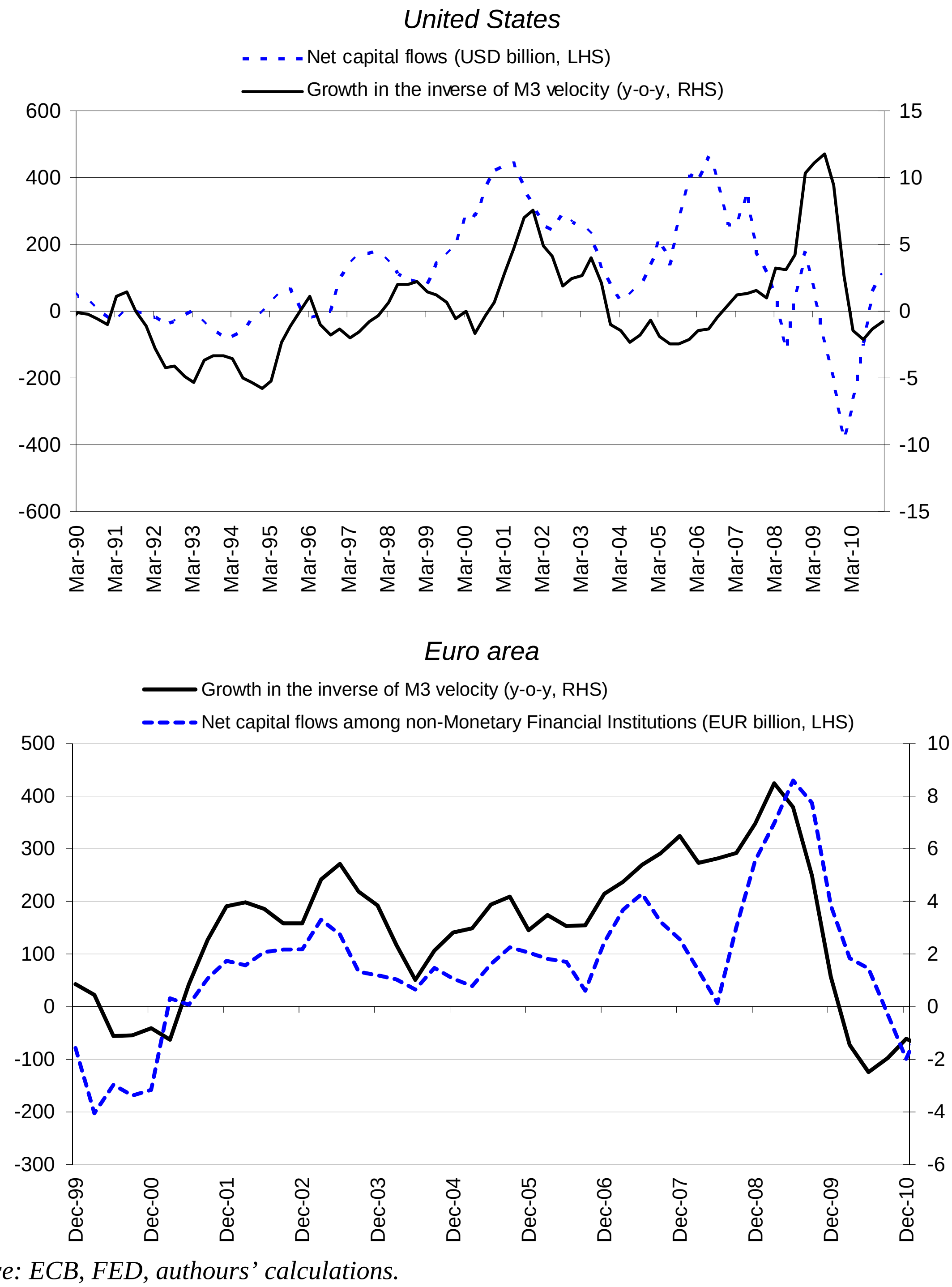
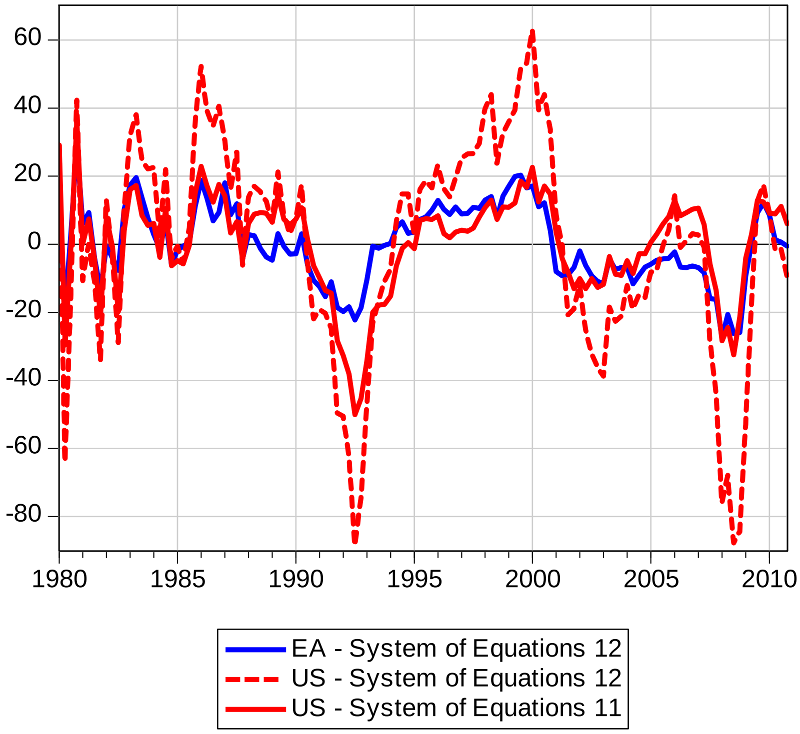
Figure 1. Broad money velocity growth and net capital flows in the euro area and the United States *(annual percentage changes; annual flows in USD or EUR billions, sample period: 1980Q1 – 2010Q4)*



Source: ECB, FED, authours' calculations.
Note: The capital flows for the euro area are only available from 1999. Broad money is defined as M3 for the euro area and M2 for the United States.

Figure 2. Cointegrating Residuals of Euro Area and the US Money Demand Equations
(indices, sample period: 1980Q1 – 2010Q4)

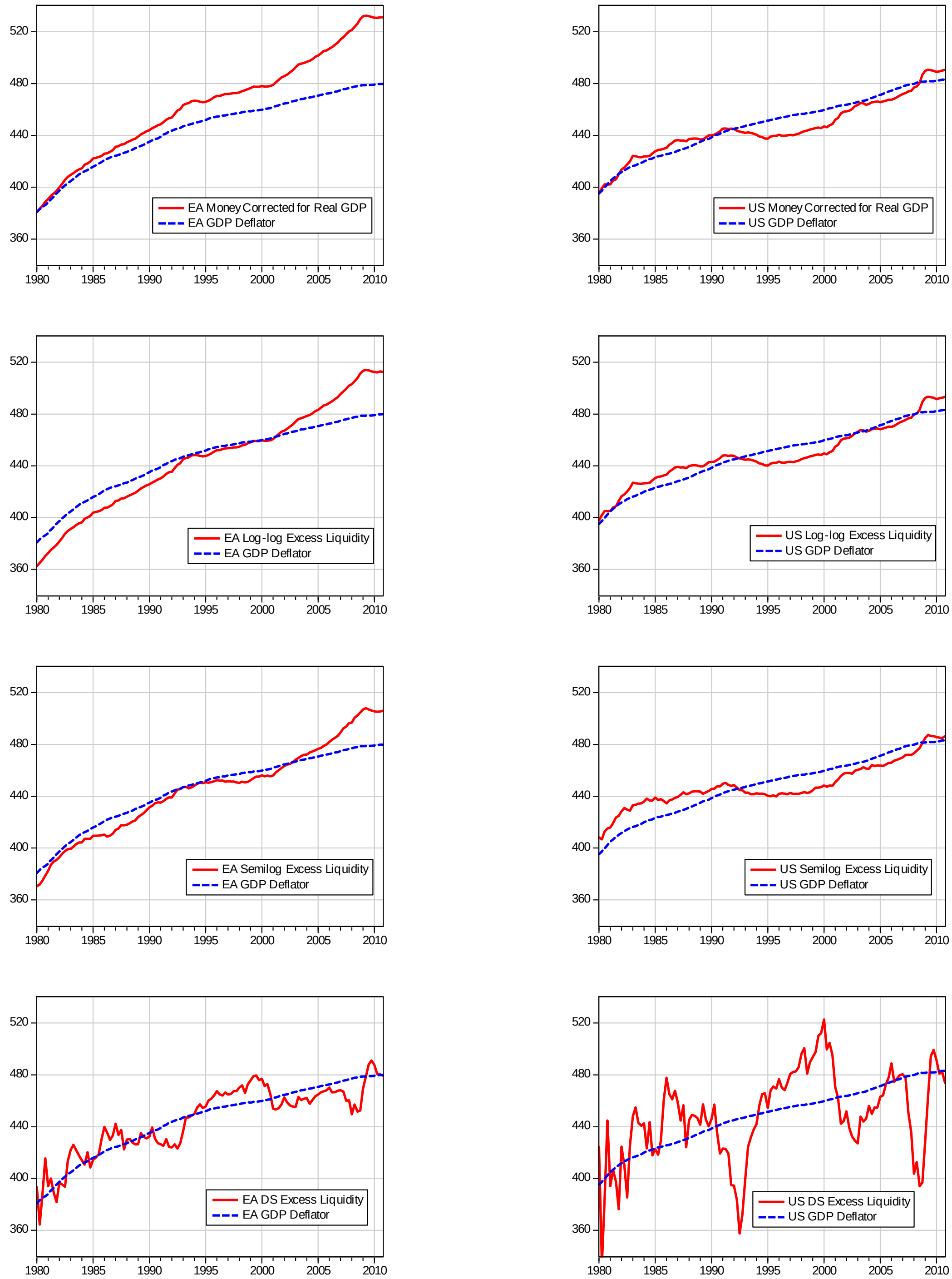


Note: The panel plots the cointegrating residuals of the money demand equations estimated using the DS model based on the system of equations 11 or 12. The cointegrating residuals of the euro area money demand equation estimated using the system of equations 11 is very similar to that of the system 12 and therefore here is not reported.

Figure 3. Excess Liquidity and Prices in the Euro Area and the United States under Alternative Specifications of Money Velocity

(indices, sample period: 1980Q1 – 2010Q4)

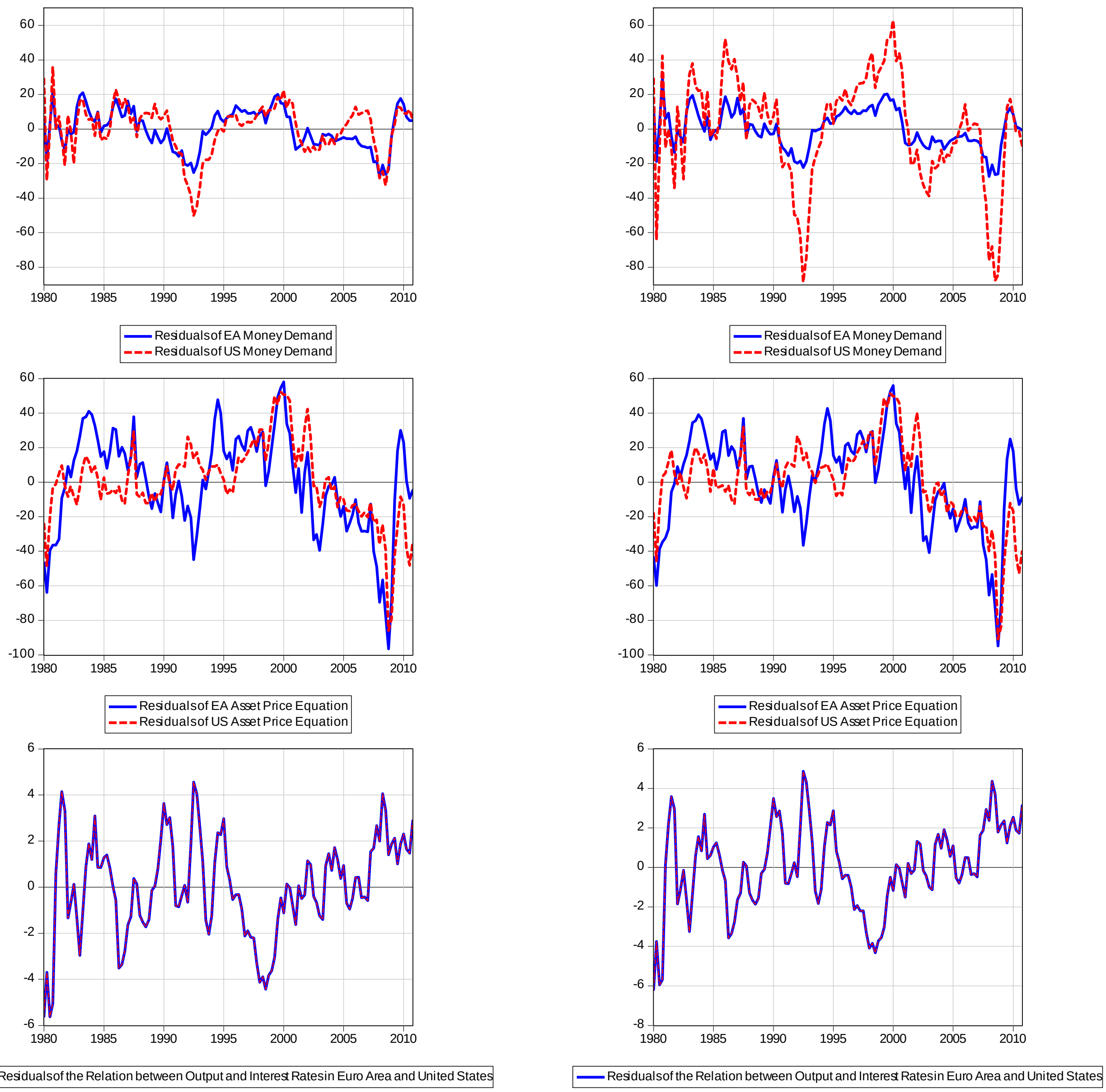
Euro Area | *United States*



Note: The first panel plots broad money minus real GDP vis-à-vis the GDP deflator. The second panel plots broad money minus real money demand estimated using the log-log money demand function with an interest rate elasticity equal to 0.08 vis-à-vis the GDP deflator. The third panel plots broad money minus real money demand estimated using the semilog money demand function with an interest rate semi-elasticity equal to 1.8 vis-à-vis the GDP deflator. The fourth panel plots broad money minus real money demand estimated using DS model vis-à-vis the GDP deflator based on the system of equations 12.

Figure 4. Cointegrating Residuals in the DS Model
(percentage points, sample period: 1980Q1 – 2010Q4)

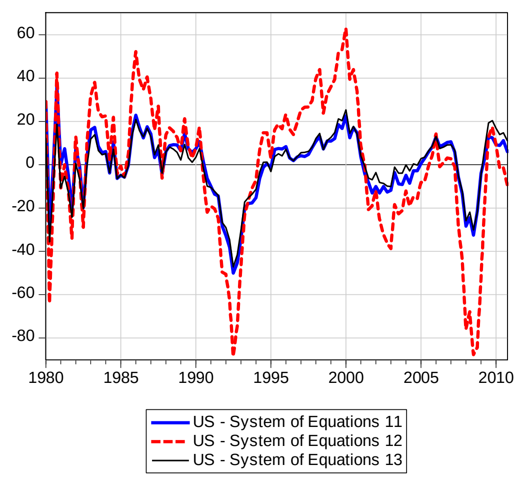
System of equations 11 *System of equations 12*



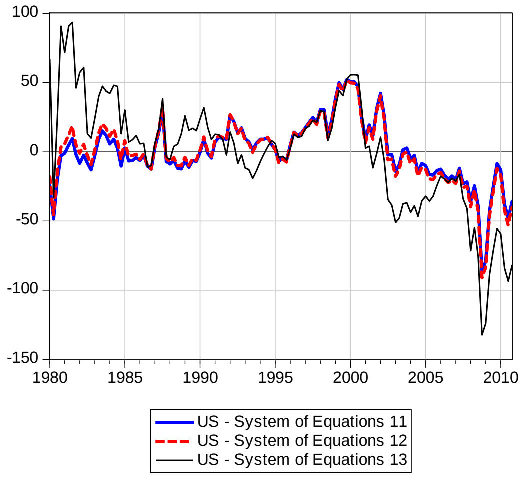
Note: The first panel plots the cointegrating residuals of the identified euro area and US money demand functions. The second panel plots the cointegrating residuals of the identified euro area and US asset price equations. The third panel plots the cointegrating residuals of the relation between euro area and US real and financial variables.

Figure 5. Cointegrating Residuals in the DS Model under Three Alternative Specifications *(percentage points, sample period: 1980Q1 – 2010Q4)*

Residuals of US money demand



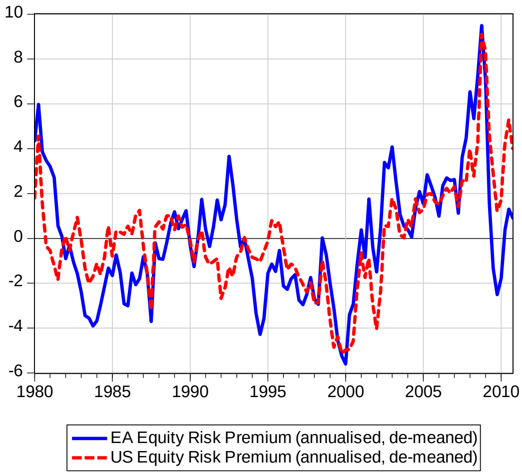
Residuals of US asset price equation



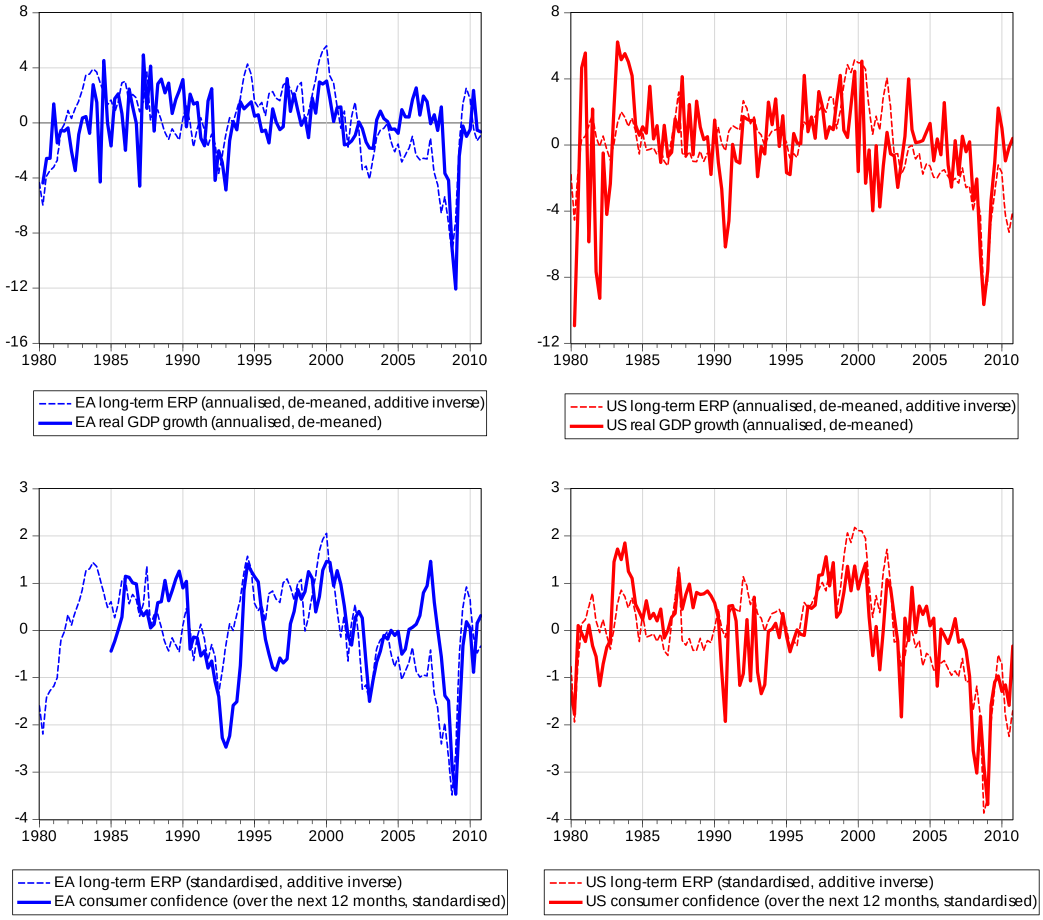
Note: The charts plot the cointegrating residuals of the identified US money demand functions and US asset price equilibria as estimated using the system of equations 11 and 12.

Figure 6. Countercyclicality of Euro area and US Equity Risk Premia in the DS Model
(percentage points, annualised, quarterly, sample period: 1980Q1 – 2010Q4))

Long-horizon Equity Risk Premium in the Euro Area and the United States



Long-horizon Equity Risk Premium and Real Economic Activity
Euro Area *United States*



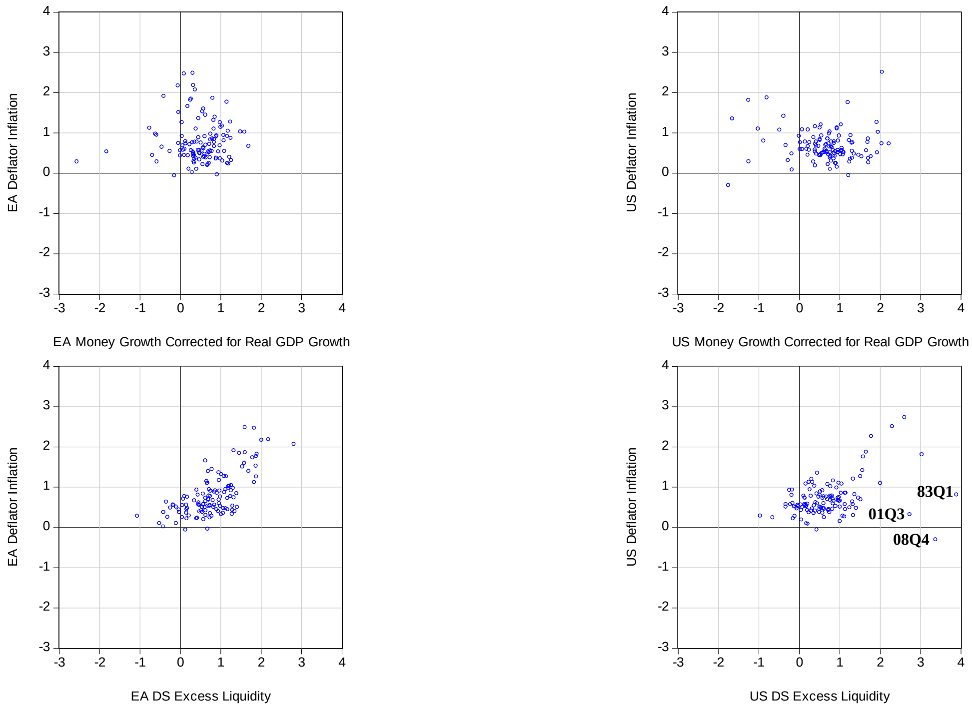
Note: The first panel plots the long-horizon equity risk premia estimated using the DS model (i.e. the respective cointegrating residuals are divided by -10). The second and third panels plot the long-horizon equity risk premia estimated using the DS model vis-à-vis respectively annualised quarterly real GDP growth and consumer confidence. The long-term equity risk premia are obtained using the system of equations 12.

Figure 7. Excess Money Growth and Inflation in the Euro Area and the United States

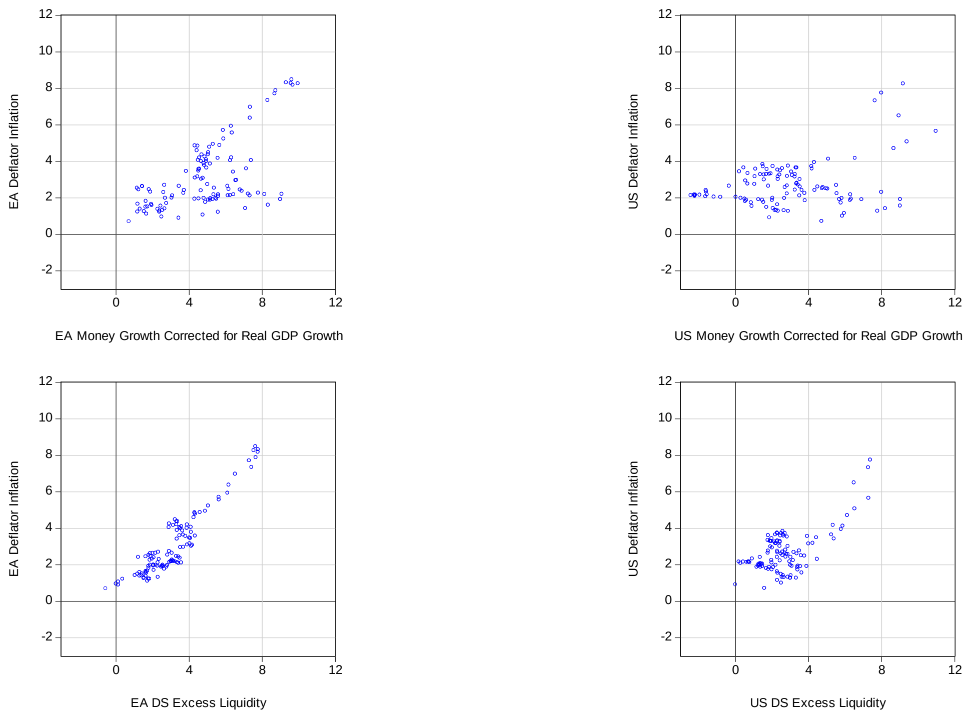
(percentage change, sample period: 1980Q1 – 2010Q4)

Euro area *United States*

Quarterly growth rates



Eight-Quarter Sum (annualized)

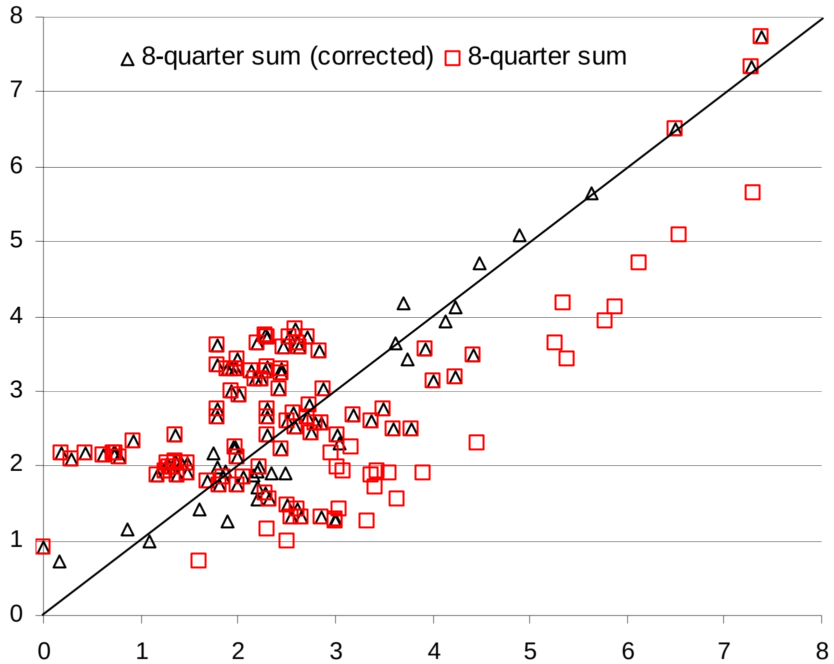


Note: The first panel plots broad money growth minus real GDP growth vis-à-vis GDP deflator inflation. The second panel plots broad money growth minus real money demand growth estimated using the DS model vis-à-vis GDP deflator inflation. The third and fourth panels report the same statistics based on 8-quarter sum. The DS results are based on the system of equations 12.

Figure 8. Corrected Excess Money Growth and Inflation in the United States

(percentage change, sample period: 1980Q1 – 2010Q4)

Eight-Quarter Sum (annualized)



Note: The panel plots broad money growth minus real money demand growth estimated using the DS model vis-à-vis GDP deflator inflation. The corrected time series is derived by smoothing out the three outliers in 1983Q1, 2001Q3 and 2008Q4. The results are based on the system of equations 12.

Table 1. Inflation versus Excess Money Growth in the Euro Area and the United States
(sample period: 1980 –2010, non-overlapping annual observations, Newey-West HAC Standard Errors)

				Euro area				
	coeff.	s.e.		coeff.	s.e.		coeff.	s.e.
Nominal money growth	0.96	0.09	Excess money growth	0.97	0.05	Excess money growth	0.98	0.07
Real money demand growth	-0.94	0.17						
						Dummy 2001	0.18	0.04
						Dummy 2008	-0.11	0.04
Adjusted R-squared	0.75			0.76			0.74	
Number of observations	31			31			31	
	F-statistics	P-value		F-statistics	P-value		F-statistics	P-value
Wald test (C1+C2 = 0)	0.06	0.81	Wald test (C1 = 1)	0.22	0.65	Wald test (C1 = 1)	0.07	0.79
Breusch-Godfrey (2)	1.31	0.29	Breusch-Godfrey (2)	1.26	0.30	Breusch-Godfrey (2)	1.10	0.35
				United States				
	coeff.	s.e.		coeff.	s.e.		coeff.	s.e.
Nominal money growth	0.80	0.17	Excess money growth	0.84	0.10	Excess money growth	1.00	0.06
Real money demand growth	-0.74	0.31				Dummy 1983	-0.74	0.10
						Dummy 2001	-0.69	0.07
						Dummy 2008	-1.20	0.11
Adjusted R-squared	0.26			0.28			0.52	
Number of observations	31			31			31	
	F-statistics	P-value		F-statistics	P-value		F-statistics	P-value
Wald test (C1+C2 = 0)	0.19	0.66	Wald test (C1 = 1)	2.64	0.11	Wald test (C1 = 1)	0.00	0.97
Breusch-Godfrey (2)	1.41	0.26	Breusch-Godfrey (2)	1.32	0.28	Breusch-Godfrey (2)	1.72	0.20

Note: GDP deflator inflation is regressed against excess money growth, defined as broad money growth minus real money demand growth estimated using the DS model. The null hypothesis of the Wald test is that the slope is equal to unity. The null hypothesis of the Breusch-Godfrey serial correlation LM tests of order 2 is that the residuals are not correlated. The results are based on the system of equations 12.

Table 2. The Unexplained Variance of Euro Area and US Real Money Growth

	Log-log	Semilog	DS
		Euro area	
1980-2010	**0.73**	**0.74**	**0.55**
1980-1990	0.65	0.69	0.40
1991-2000	0.73	0.74	0.66
2001-2010	0.75	0.73	0.57
		United States	
1980-2010	**0.72**	**0.70**	**0.59**
1980-1990	0.80	0.59	0.64
1991-2000	0.78	0.78	0.41
2001-2010	0.59	0.77	0.65

Note: Variance of residuals as a fraction of de-meaned real money growth. The coefficients used are estimated over the entire sample period 1980Q1-2010Q4. Interest rate elasticity (log-log specification) and semielasticity (semilog specification) are set equal to 0.9 and 1.8, respectively (Ireland, 2009). The DS results are based on the system of equations 12.

Table 3. Testing for Reduced Rank in the DS Model
(sample period: 1980Q1 – 2010Q4)

	Eigenvalue	Statistic	Critical Value	Prob.**
		Trace		
None *	0.551	410.525	251.265	0.000
At most 1 *	0.437	312.961	208.437	0.000
At most 2 *	0.415	242.981	169.599	0.000
At most 3 *	0.381	177.497	134.678	0.000
At most 4 *	0.345	118.949	103.847	0.004
At most 5	0.202	67.282	76.973	0.218
At most 6	0.135	39.804	54.079	0.480
At most 7	0.100	22.048	35.193	0.592
At most 8	0.055	9.152	20.262	0.722
At most 9	0.019	2.309	9.165	0.716
		Max-Eigenvalue		
None *	0.551	97.564	65.300	0.000
At most 1 *	0.437	69.980	59.240	0.003
At most 2 *	0.415	65.484	53.188	0.002
At most 3 *	0.381	58.548	47.079	0.002
At most 4 *	0.345	51.667	40.957	0.002
At most 5	0.202	27.477	34.806	0.287
At most 6	0.135	17.756	28.588	0.597
At most 7	0.100	12.896	22.300	0.567
At most 8	0.055	6.844	15.892	0.689
At most 9	0.019	2.309	9.165	0.716

*Note: Both Trace and Max-eigenvalue tests indicate 5 cointegrating equations at the 0.05 level, allowing for a deterministic trend in the long-run equilibria. *denotes rejection of the hypothesis at the 0.05 level. **MacKinnon-Haug-Michelis (1999) p-values.*

Table 4: The adjustment coefficients
(sample period: 1980Q1 – 2010Q4)

		Euro area				
		Real M3	*Real GDP*	*Short-term Yield*	*Price / Earnings*	*Long-term Yield*
ECT 1	coef.	0.024	0.013	0.008	-0.148	0.000
	s.e.	(0.005)	(0.005)	(0.005)	(0.109)	(0.004)
	t.stat.	[4.707]	[2.831]	[1.629]	[-1.362]	[-0.061]
ECT 2	coef.	-0.062	-0.007	-0.019	0.622	0.005
	s.e.	(0.012)	(0.01)	(0.012)	(0.245)	(0.009)
	t.stat.	[-5.288]	[-0.643]	[-1.63]	[2.537]	[0.509]
ECT 3	coef.	0.004	0.000	-0.006	0.036	-0.004
	s.e.	(0.002)	(0.002)	(0.002)	(0.043)	(0.002)
	t.stat.	[2.13]	[0.128]	[-2.703]	[0.829]	[-2.275]
ECT 4	coef.	-0.012	0.001	-0.002	-0.240	-0.005
	s.e.	(0.005)	(0.004)	(0.005)	(0.101)	(0.004)
	t.stat.	[-2.524]	[0.232]	[-0.437]	[-2.378]	[-1.31]
ECT 5	coef.	-0.028	0.025	-0.084	-0.427	-0.072
	s.e.	(0.031)	(0.028)	(0.031)	(0.648)	(0.024)
	t.stat.	[-0.893]	[0.905]	[-2.697]	[-0.658]	[-3.054]

		United States				
		Real M2	*Real GDP*	*Short-term Yield*	*Price / Earnings*	*Long-term Yield*
ECT 1	coef.	0.005	-0.021	-0.027	0.081	-0.010
	s.e.	(0.008)	(0.006)	(0.01)	(0.092)	(0.007)
	t.stat.	[0.655]	[-3.826]	[-2.684]	[0.88]	[-1.539]
ECT 2	coef.	0.021	0.023	0.035	-0.351	0.041
	s.e.	(0.019)	(0.013)	(0.022)	(0.208)	(0.015)
	t.stat.	[1.115]	[1.843]	[1.552]	[-1.689]	[2.684]
ECT 3	coef.	0.003	-0.010	-0.012	-0.069	-0.003
	s.e.	(0.003)	(0.002)	(0.004)	(0.037)	(0.003)
	t.stat.	[0.956]	[-4.549]	[-3.143]	[-1.871]	[-1.195]
ECT 4	coef.	-0.001	0.023	0.013	-0.004	-0.007
	s.e.	(0.008)	(0.005)	(0.009)	(0.085)	(0.006)
	t.stat.	[-0.093]	[4.376]	[1.378]	[-0.047]	[-1.137]
ECT 5	coef.	0.074	-0.178	-0.233	-1.454	-0.082
	s.e.	(0.049)	(0.033)	(0.059)	(0.549)	(0.041)
	t.stat.	[1.51]	[-5.317]	[-3.944]	[-2.65]	[-2.01]

Note:
ECT 1 is the cointegrating residual of the US money demand specification.
ECT 2 is the cointegrating residual of the EA money demand specification.
ECT 3 is the cointegrating residual of the US asset price specification.
ECT 4 is the cointegrating residual of the EA asset price specification.
ECT 5 is the cointegrating residual of the relation between output and interest rates in the euro area and the United States.
The DS results are based on the system of equations 12.