Monetary Asset Substitution in the Euro Area

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Abstract

We study the relation between holdings of monetary assets and government securities in the Euro area. We estimate time-varying elasticities of substitution between monetary assets using the semi-nonparametric method of Gallant (1981). The empirical elasticities are then tested for structural breaks using the framework of Bai and Perron (1998). Since our sample starts from January 2000, we discuss explicitly the implications of the recent financial turmoil of 2007.

The estimated elasticities are consistent with the assumption of imperfect substitution between monetary assets. Our results suggest that three main types of episodes affected the stability in the dynamics of monetary assets, including the 2001-2003 episode of heightened risk aversion, the outbreak of the recent financial crisis, and the Lehman bankruptcy.

Keywords: money demand, nonparametric methods, elasticity estimation.
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> "*In a world involving no transaction friction and no uncertainty, there would be no reason for a spread between the yield on any two assets, and hence there would be no difference in the yield on money and on securities. (...) In such a world securities themselves would circulate as money and be acceptable in transactions; demand bank deposits would bear interest, just as they did in this country in the period of the twenties.*"

Samuelson (1947), p. 123

1 - Introduction

The monetary policy strategy of the European Central Bank (ECB) contemplates a monetary pillar. Various measures of money growth are monitored or 'money gap' are monitored to extract information on inflation for the medium and long term (see ECB, 2008 and 2009). Quite a lot of empirical evidence is available on the interpretation of the structural determinants of money growth in the Euro area. For instance, Stracca (2004) studies the properties of quarterly divisia monetary aggregates. His results suggest that this monetary indicator has predictive power for output and inflation.

The available literature has however presented few empirical facts on how the relative holdings of alternative monetary assets contribute to changes in money demand over time (e.g., see Stracca 2004). This is somewhat surprising because knowledge of the degree of substitution between monetary assets could shed light on the policy options of the ECB during the recent period of financial market turmoil started in 2007. As Walsh (2004) notes, the available literature on monetary policy at low nominal interest rates suggests that

> *(t)he possibility that altering the relative supply of short-term and long-term securities will have real effects is based on the idea that different assets are imperfect substitutes. (…) (T)he empirical evidence for imperfect asset substitutability is limited. What evidence does exist suggests a relatively high degree of substitutability."*

The experience of the European Economic and Monetary Union (EMU) has pointed to relevant episodes of substitution between assets even before 2007. In particular, between 2001 and 2003, a large increase in M3 took place following portfolio shifts of approximately 180 billion Euros from equity to money balances. This happened in the wake of financial market instability due to the persistent weakness of stock markets worldwide (see Issing et al., 2005, p. 70). As a result, Euro area investors increased the share of safe and liquid assets. ECB (2003, 2004) shows that these portfolio shifts corresponded with an increase in risk aversion.

In this paper, we provide a formal investigation of instabilities in the dynamics of monetary assets in the Euro area. First of all, we provide a measure of substitution elasticity between the assets that form the M1 and M2 aggregates, and the changes to holdings of short- and long-term public debt. As a second point, by using a sample that covers the entire period of existence of the Euro, we study the impact of the recent financial turmoil of 2007 on changes in monetary assets.

In our empirical investigation, we follow an approach that is very different from that of previous studies based on parametric models for money demand, such as univariate regressions and cointegrated vector-autoregressions. We estimate time-varying elasticities of substitution between assets using the semi-nonparametric framework proposed by Gallant (1981).

Since the advent of divisia aggregates proposed by Barnett (1978, 1980), the search for an appropriate functional form for monetary demand system has been controversial. Mainly due to the simplicity of their intepretation, household utility functions based on Cobb-Douglas or constant elasticity of substitution have been used widely (see Barnett, Offenbacher and Spindt 1981). Uzawa (1962) has suggested that these functions lack the desirable analytical properties for measuring asset substitution. For this reason, flexible functional forms, such as the so-called translog function (e.g., see Ewis and Fischer, 1984), have been introduced to estimate elasticities at approximation points (e.g., see Barnett et al., 1992). While gaining in terms of accuracy, these functions tend to violate global regularity conditions for optimization in large regions. An innovation in this respect is the semi-nonparametric flexible form that possesses global flexibility. Furthermore, Serletis (2007) suggests that asymptotic inferences are potentially free from specification errors for the semi-nonparametric function.

We measure the asset substitution through the so-called Morishima elasticities. This has already been employed in other studies of money demand (e.g., see Davis and Gauger, 1996). The intuition for this measure is that it takes into account the cross-substitution effects arising from changes in the prices of a given asset, while holding all the other prices at a constant level.

We provide two main sets of results. Our estimates of elasticities are consistent with the assumption of imperfect substitution between monetary assets. In particular, they values are equal to one, on average. The temporal variation of the estimates suggests that structural breaks provide an important empirical role. Thus, we test for the number and timing of structural changes using the framework of Bai and Perron (1998, 2003a and 2003b). There are three episodes that affect the stability of substitution between monetary assets. First, our results support the discussion of ECB (2003, 2004) about the impact of changes in market risk aversion in the period 2001-2003. The beginning of the recent financial crisis in August 2007 has also generated a shift in the substituion between monetary aggregates. Finally, a similar effect has taken place after the Lehman bankruptcy of September 2008.

The paper is organized as follows. Section 2 outlines the empirical model and the concept of elasticity of substitution used in the paper. Section 3 discusses both the construction of the dataset. Section 4 presents the estimation results. In Section 5, we propose some concluding remarks.

2 - The demand-system approach

The standard approach for the estimation of substitution elasticities relies on the specification of a conditional demand system of assets from the two-stage utility maximization problem of a representative consumer. The first-stage problem consists in the choice of the expenditure level for each asset. The second problem, instead, is based on utility maximization for a given aggregate expenditure level. The solution of the second step yields a conditional Marshallian demand function and a conditional indirect utility. The indirect utility function measures consumer’s utility at a given price and wealth level. This is then used to calculate the elasticities of substitution (see Barnett, Fisher and Serletis, 1992).

The demand system for assets is obtained from the household’s indirect utility function $g(x, \theta)$, where x is a vector of normalized asset prices or user costs, and θ is a parameter vector. Roy’s identity can be used to compute the expenditure share of each asset.[†] Instead of specifying a functional form for indirect utility, Gallant (1981) introduces the semi-nonparametric Fourier approximation

$$g^{*}(x,\theta) = u_0 + b'x + \frac{1}{2}x'Cx + \sum_{\alpha=1}^{A}\left[u_{0,\alpha} + \sum_{j=1}^{J}\left(u_{j,\alpha}\cos(jk'_{\alpha}x) - v_{j,\alpha}sin(jk'_{\alpha}x)\right)\right] \quad (1)$$

where $C = -\sum_{\alpha=1}^{A} u_{0,\alpha}k_{\alpha}\, k'_{\alpha}$, the parameter vector is $\theta \coloneqq \{b, u_{0,\alpha}, u_{j,\alpha}, v_{j,\alpha}\}$ for $\alpha = 1,\ldots A$ and $j = 1,\ldots J$, and k_{α} denotes the partial differentiation of utility fuction. After applying Roy's identity to equation 1, a Fourier system of shares

$$s_i(x,\theta) = \frac{x_i b_i - \sum_{\alpha=1}^{A}\left[u_{0,\alpha}x'^{k_{\alpha}} + 2\sum_{j=1}^{J}(u_{j,\alpha}\cos(jk'_{\alpha}x) - v_{j,\alpha}\sin(jk'_{\alpha}x))\right] k}{b'x - \sum_{\alpha=1}^{A}\left[u_{0,\alpha}x'^{k_{\alpha}} + 2\sum_{j=1}^{J}(u_{j,\alpha}\cos(jk'_{\alpha}x) - v_{j,\alpha}\sin(jk'_{\alpha}x))\right]} \quad (2)$$

is obtained. By stacking a set of equations with the form (2) for each asset, we obtain a system of expenditure shares that can be taken to the data.

In the empirical application, I estimate a system of four share equations

$$s_t = f(x_t,\theta) + e_t \quad (3)$$

with additive errors $e_t = \rho e_{t-1} + \epsilon_t$ and ϵ_t is a white noise with constant covariance matrix. Since the shares sum up to unity, the covariance matrix of the error term is singular. Hence, maximum-likelihood estimates can be obtained by dropping any equation.

[†] Roy's identity determines a Marshallian demand function from an indirect utility function.

2.1 Measuring the elasticity of substitution between assets

There is a large literature on measures for the degree of substitution in asset demand systems for the U.S. economy. Davis and Gauger (1996) show that, when more than two assets are included in the household's budget constraint, the standard Hicksian elasticity of substitution should not be used. The reason is that this form of elasticity can be used to study the impact of a price change when a consumer's utility is held constant. This provides a net measure of elasticity that is relevant when we are interested in examining the effect on a single asset. As a result, it neglects the role of cross-assets substitution effects.

In this paper, we use the Morishima elasticity between assets *i* and *j*. This is equal to

$$\mathrm{ME}_{i,j} = s_i\left(\sigma_{j,i} - \sigma_{i,i}\right) \tag{4}$$

The term $\sigma_{j,i}$ denotes the Allen-Uzawa elasticity, which is equal to the ratio between the Hicksian elasticity and the asset expenses share. The MES measures the percentage change in relative quantities, or ratios, with respect to a percentage change in one price. Like the Hicksian and Allen measures, the Morishima elasticity is a measure of net substitution that examines preferences revealed by curvature of the indifference curve (see Blackorby and Russell, 1989).

Computing the elasticities of substitution requires differentiating the Fourier flexible form. This means that the elasticities are a function of both the parameters and the expenditure shares. Hence, the flexible Fourier form produces point estimates for the elasticities over the available sample, and can be used to compute a time series for elasticities. The detailed derivations for the expressions of the elasticities are presented by Gallant (1981).

3 - Dataset

The dataset consists of four assets for the Euro area at a monthly frequency, measured as end-of-period stocks: M1, consisting of currency, demand deposits, and interest-bearing checkable deposits in M1 (denoted as 1); the non-term assets in M2, consisting of savings deposits, money market accounts and money market mutual funds (denoted as 2); the outstanding amounts of central government's short and long-term debt securities (denoted as 3 and 4 respectively). The sample spans from January 1995 to June 2007.

We deflate the asset quantities by an index for the price level. For each point in time, the normalized user costs are defined as

$$x_i = (R - r_i)/(1 + R) \tag{5}$$

where R is the interest rate on the benchmark asset, and r_i is the asset's own rate (see Barnett, 1978). We assume that the benchmark rate is the yield on a 3-month government bond, and that the own price of M1 is the average rate on demand deposits for the Euro area. The user cost of M2 is the return on money market funds. Finally, I use the yields on 1 and 10-year bonds as proxies for the prices of short and long-term bonds. The data on quantities are available from the ECB's Statistical Data Warehouse. We have obtained the user costs on non-bond assets from Datastream.[‡]

Estimating elasticities of substitution between monetary assets implies studying the empirical properties of money demand. This can also be accomplished by formulating parametric models, such as vector error-correction models, that account explicitly for the persistence of the series. Since we pursue a different methodological avenue in this paper, we follow Davis and Gauger (1996) and disregard the issue of persistence of our time series.

4 - Empirical results

Figure 1 reports the estimated series of Morishima elasticity. Several regularities emerge. The first one is that there are common patterns between alternative elasticities. The elasticity between M2 and long-term bonds

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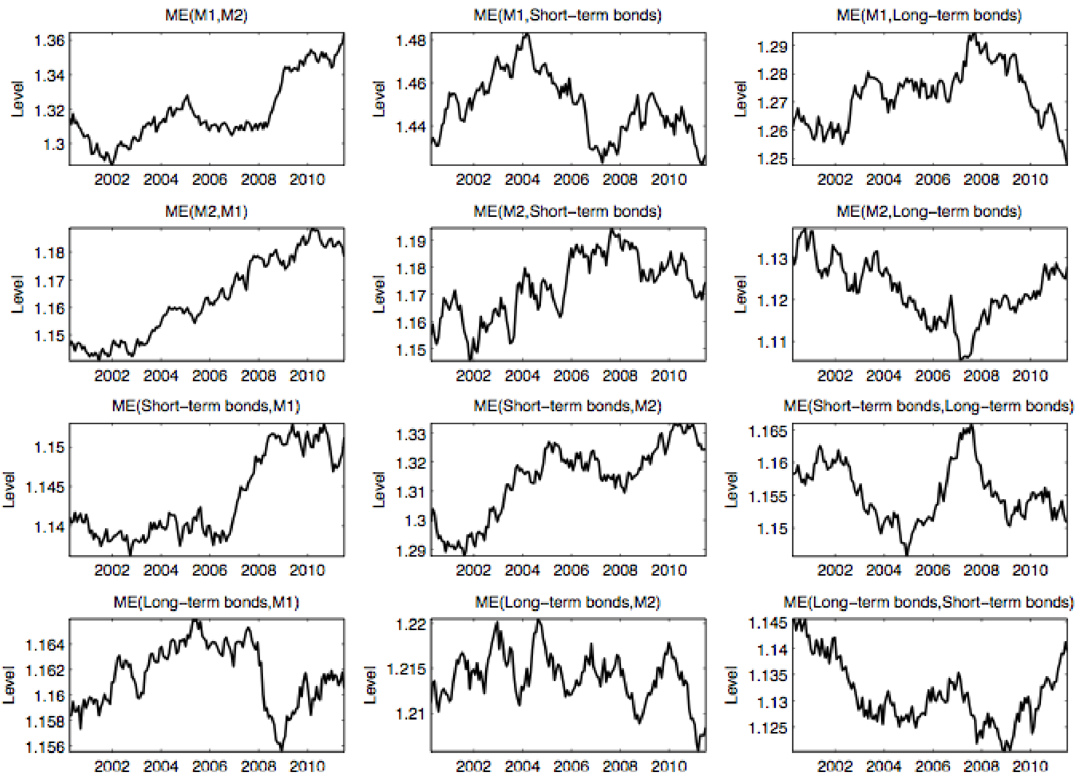
follows a trend similar to that of the elasticity between long- and short-term bonds. This is also the case for the elasticities between M1 and M2, and between short-term bonds and M1. The estimated elasticities are characterized by fluctuations not very wide, as they change within limited bounds since the creation of the Euro. There is an evident role for structural breaks and instabilities in some of the elastiticities, for instance between M2 and short-term bonds. There are also estimated elasticities for which the presence of a single time trend is evident, like in the case of the elasticity between M2 and M1.

Figure 1: Estimated Morishima elasticities



All the elasticities of substitution between M1 and M2 are estimated close to one, These figures broadly in line to those of Jones et al. (2008) for the U.S. Furthermore, the degree of substitution between bonds, as well as between bonds and the liquid assets is not as high in absolute value. The overall picture indicates that the assumption of imperfect substitution between bonds and money is supported by the data.

How is the demand for money affected by shocks to short- and long-term bonds? Short-term securities react to changes in M1 in a way opposite to the elasticitiy of long-term bonds. The elasticity of short-term security holdings to changes of M1 jump in 2007, at the onset of the financial crisis. The relation to M2 moves to an upward trend in 2008, before the bankruptcy of Lehman Brothers. On the other hand, the elasticity between long-term bonds and M1 falls until the middle of 2008, and then increases after the Lehman episode.

What is the role of instabilities in the determination of money demand? As suggested earlier, the ECB (2004) provides evidence for a surge in monetary assets due to a decrease in equity holdings between 2001 and 2003. Figure 1 suggests that M1 and M2 react in different ways to changes in government bonds. Moreover, there are varied patterns of fluctuations over 2001-2003. This is consistent with the presence of structural break in the relationship between government bond holdings and liquid assets.

Table 1 reports some descriptive statistics on the estimated elasticities. Our results are characterized by volatility to a small extent. The estimates are only modestly skewed, suggesting that they are characterized by a symmetric empirical distribution. This means that the shocks driving money demand substitution are not characterized by a large one-off event. Rather, the evidence suggests there money demand is driven by a series of shocks that generate fluctuations around the mean.

Table 1: Statistics of estimated Morishima elasticities

Elasticity	Mean	Std. Dev.	Minimum	Maximum	Skewness	Kurtosis
ME(M1, M2)	1.3183	0.3187	1.2874	1.3645	0.7269	2.4249
ME(M1, Short-term bonds)	1.4495	0.3146	1.4219	1.4834	0.1655	2.3034
ME(M1, Long-term bonds)	1.2726	0.1102	1.2475	1.2946	-0.0096	2.3630
ME(M2, M1)	1.1632	0.2149	1.1407	1.1887	0.0931	1.5833
ME(M2, Short-term bonds)	1.1730	0.2117	1.1455	1.1946	-0.2777	2.1739
ME(M2, Long-term bonds)	1.1220	0.5071	1.1054	1.1370	-0.1587	2.6529
ME(Short-term bonds, M1)	1.1434	0.5053	1.1362	1.1530	0.5523	1.6421
ME(Short-term bonds, M2)	1.3139	0.7126	1.2876	1.3332	-0.5574	2.2251
ME(Short-term bonds,	1.1843	0.0945	1.1744	1.1951	0.3353	2.4390
ME(Long-term bonds, M1)	1.1615	0.1025	1.1744	1.1660	-0.2545	2.1167
ME(Long-term bonds, M2)	1.2140	0.3920	1.2057	1.2205	-0.2899	3.2728
ME(Long-term bonds,	1.1201	0.4160	1.1092	1.1342	0.5739	2.5722

Table 2: test statistics for UD max and sup $\hat{F}_T(n)$

Elasticity	sup $\hat{F}_T(1)$	sup $\hat{F}_T(2)$	sup $\hat{F}_T(3)$	sup $\hat{F}_T(4)$	sup $\hat{F}_T(5)$	UD max
ME(M1, M2)	52.22*	66.70*	51.30*	46.47*	38.83*	52.22*
ME(M1, Short-term bonds)	90.30*	81.42*	79.71*	51.95*	44.00*	70.59*
ME(M1, Long-term bonds)	69.40*	87.75*	73.29*	60.08*	38.83*	90.19*
ME(M2, M1)	64.07*	71.75*	71.30*	36.11*	41.05*	61.05*
ME(M2, Short-term bonds)	72.20*	59.66*	51.30*	44.04*	30.80*	72.20*
ME(M2, Long-term bonds)	75.80*	67.49*	62.57*	46.47*	38.83*	90.48*
ME(Short-term bonds, M1)	70.79*	70.13*	64.90*	41.47*	30.04*	88.10*
ME(Short-term bonds, M2)	74.11*	72.80*	67.00*	42.06*	30.59*	90.37*
ME(Short-term bonds, Long-term bonds)	70.30*	71.15*	68.09*	31.83*	39.20*	71.53*
ME(Long-term bonds, M1)	73.27*	71.80*	67.90*	31.51*	30.18*	63.98*
ME(Long-term bonds, M2)	69.10*	72.03*	72.90*	44.60*	39.90*	70.58*
ME(Long-term bonds, Short-term bonds)	73.01*	72.01*	69.90*	44.50*	39.85*	67.92*

Legend: * test statistic is significant at the 5% level based on the asymptotic critical values of Bai and Perron (2003a).

How stable are the estimated elastiticities? And, in case we identify structural breaks from a statistical point of view, when do they occur? In the remainder of this section, we apply the methodology developed by Bai and Perron (1998, 2003a and 2003b) for finding multiple structural breaks in time series and testing for their statistical significance. These tests are generalizations of Andrews (1993) test for the single structural change case, and are shown to be robust to serial correlation and heterogeneity of the residuals under the null. We start by introducing the model

$$y_t = z_t' \delta_j + u_t \tag{6}$$

for $= T_{j-1} + 1, \dots T_j$, and j=1,… m and δ_j is a vector of coefficients. The term m denotes the number of of breaks. The break points $T_1, \dots T_m$ are treated as unknown, with $\lambda_i = T_i/T$ and $0 < \lambda_1 < \cdots < \lambda_m < 1$. The model is estimated using a least-squared method. For each m-partition $(T_1, \dots T_m)$ denoted as $\{T_j\}$, the estimate of δ_i is obtained by minizing the sum of squared residuals with the constraint $\delta_i \neq \delta_{i+1}$. The break-point stimators are global minimizers of the objective function.

We consider the test for structural stability against a fixed number of breaks proposed by Bai and Perron (1998). The test statistics can be written as

$$F_T(\lambda_1, \dots \lambda_n; q) = \frac{1}{T}\left(\frac{T-(n+1)q}{nq}\right)\hat{\delta}' R' \left(R\hat{V}(\hat{\delta})R'\right)^{-1} R\hat{\delta} \tag{7}$$

The term R denotes a matrix such that $(R\delta)' = (\delta_1' - \delta_2', \dots \delta_n' - \delta_{n+1}')$, and $\hat{V}(\hat{\delta})$ is an estimate of the variance-covariance matrix of $\hat{\delta}$. This expression is used to compute the ‘sup’ test statistics

$$\sup F_T(n; q) = F_T(\hat{\lambda}_1, \dots \hat{\lambda}_n; q) \tag{8}$$

where the break fraction estimates $(\hat{\lambda}_1, \dots \hat{\lambda}_n)$ minimize the global sum of squared residuals of equation (6). Pai and Perron (1998) also introduce a test of no structral change against an unknown number of breaks, given an uppper bound M for m. The so-called maximum test UDmax is defined for a set of weights $\{a_1, \dots a_M\}$ as

$$\max F_T(M, q, a_1, \dots a_M) = \max_{1 \le m \le M} a_M \, F_T\left(\hat{\lambda}_{1,\dots} \hat{\lambda}_1; q\right) \tag{9}$$

In this paper, we impose equal weights across possible breaks, i.e. $a_m = 1$. Finally, we apply a sequential test for l structural breaks against $l+1$ breaks, with the test statistics

$$\sup F_T(l+1|l) = \left\{S_T\left(\hat{T}_1, \dots \hat{T}_l\right) - \min_{1 \le i \le l+1} \inf_{\tau \in \Lambda_i} S_T\left(\hat{T}_1, \dots \hat{T}_l\right)\right\} / \hat{\sigma}^2 \tag{10}$$

The term Λ_i defines a neighbourhood of a given break date. The term $S_T\left(\hat{T}_1, \dots \hat{T}_l\right)$ denotes the sum of squared residuals resulting from the least-squares estimation from each *m*-partition, and $\hat{\sigma}^2$ is a consistent estimator under the null hypothesis.

In our empirical application, we apply the tests in the following way. First, we compute the UDmax test to check for evidence of at least one structural break. We then run a series of sequential tests, with a number of breaks consistent with the UDmax test. In the comments of the results, we use the asymptotic critical values reported by Bai and Perron (1998) for a 95% significance levels.

Table 2 reports the results for the UD max and sup $\hat{F}_T(n)$ for up to five breaks. All the estimated test statistics are significant, indicating that the series presents at least one break. The number of structural changes is then identified through the sequential test sup $\hat{F}_T(l+1|l)$. First of all, the largest number of breaks used in our application is supported by the test results. There is evidence for a different number of breaks across the various elasticity estimates. For instance, the elasticities between short-term bonds and the two money aggregates M1 and M2 are characterized by the largest number of breaks, which is equal to four.

The estimated break dates with 95% confidence intervals are reported in Table 3. The statistical significance is judged against the critical values computed by Bai and Perron (2003a). All the dates are estimated precisely with intervals covering a few months before and after. There are three types of major events that affect the elasticities of substitution between most of the assets. The beginning of the recent financial turmoil in 2007 falls within the statisticaly-significant bands of break dates. The bankruptcy of Lehman Brothers of September 2008 represents an additional event of relevance. Finally, our estimates of shifts in money aggregates capture the instabilities that affected the Euro area between 2001 and 2003, as documented by ECB (2003), among others.

Table 3: Bai and Perron's recursive test statistics

Elasticity	sup $\hat{F}_T(2\|1)$	sup $\hat{F}_T(3\|2)$	sup $\hat{F}_T(4\|3)$	sup $\hat{F}_T(5\|4)$
ME(M1, M2)	32.30*	46.17*	2.13	7.16
ME(M1, Short-term bonds)	31.70*	5.30	8.57	2.15
ME(M1, Long-term bonds)	50.30*	6.11	1.05	1.05
ME(M2, M1)	32.30*	31.70*	1.29	8.50
ME(M2, Short-term bonds)	30.57*	46.17*	2.13	2.18
ME(M2, Long-term bonds)	32.30*	46.17*	5.07	7.16
ME(Short-term bonds, M1)	38.50*	45.66*	32.30*	4.04
ME(Short-term bonds, M2)	33.29*	47.15*	2.57*	6.17
ME(Short-term bonds, Long-term bonds)	40.05*	7.80	5.82	5.53
ME(Long-term bonds, M1)	41.50*	31.32*	1.18	3.58
ME(Long-term bonds, M2)	38.50*	42.60*	5.50	5.58
ME(Long-term bonds, Short-term bonds)	35.50*	9.50	5.85	7.24

Legend: * test statistic is significant at the 5% level based on the asymptotic critical values of Bai and Perron (2003a).

Table 4: Estimated break dates with 95% confidence intervals (in brackets)

Elasticity	$\hat{T}_1$	$\hat{T}_2$	$\hat{T}_3$	$\hat{T}_4$
ME(M1, M2)	2002:3 (2001:6-2002:7)	2004:12 (2004:2-2005:6)	2008:10 (2008:2-2009:4)	-
ME(M1, Short-term bonds)	2004:1 (2003:5-2004:7)	2007:3 (2006:6-2007:9)	-	-
ME(M1, Long-term bonds)	2002:10 (2001:7-2003:4)	2007:9 (2007:5-2008:6)	-	-
ME(M2, M1)	2003:1 (2002:4-2003:7)	2005:6 (2004:11-2005:9)	2007:2 (2006:8-2007:9)	-
ME(M2, Short-term bonds)	2001:10 (2001:7-2002:6)	2003:10 (2003:2-2004:5)	2005:11 (2005:4-2006:5)	-
ME(M2, Long-term bonds)	2002:3 (2001:11-2002:8)	2006:5 (2006:1-2007:2)	2007:8 (2007:1-2008:3)	-
ME(Short-term bonds, M1)	2002:10 (2002:2-2003:4)	2005:10 (2005:1-2006:2)	2006:11 (2006:5-2007:3)	2008:7 (2008:1-2009:1)
ME(Short-term bonds, M2)	2002:4 (2001:11-2002:8)	2004:2 (2003:9-2004:9)	2006:1 (2005:5-2006:8)	2007:7 (2007:1-2008:1)
ME(Short-term bonds, Long-term bonds)	2005:1 (2004:4-2005:9)	2007:7 (2007:1-2008:3)	-	-
ME(Long-term bonds, M1)	2002:3 (2001:8-2002:11)	2003:2 (2002:11-2003:7)	2007:6 (2007:2-2007:11)	-
ME(Long-term bonds, M2)	2004:5 (2004:1-2005:2)	2008:3 (2007:7-2008:9)	2009:10 (2009:4-2010:3)	-
ME(Long-term bonds, Short-term bonds)	2007:3 (2006:11-2008:1)	2009:2 (2008:7-2009:7)	-	-

5 - Conclusions

We characterize breaks in money demand in the Euro area through the lenses of substitution elasticity between monetary assets. We use the semi-nonparametric Fourier approximation of the indirect utility function proposed by Gallant (1981) to estimate elastiticities of substitution between four type of monetary aggregates. For this purpose, we compute Morishima elasticities which, differently from other specifications, account for the cross-substitution effects between assets. Our results support the empirical assumption of imperfect substitution between monetary assets. Since structural breaks appear as an important empirical feature of the elasticities, we test for structural breaks. We provide statistical evidence suggesting that the breaks take place around 2001-2003, August 2007 and September 2008.

The analysis presented in this paper represents only the first step towards a better understanding of the implications of asset substitution in the Euro area. In fact, our findings can be extended along several relevant directions. First of all, we can provide a formal investigation on the role of shocks to monetary substitution across different assets. We set up a structural-vector autoregression with the elasticities, and compute impulse responses to different shocks. It would be relevant to estimate elasticities of substitution between monetary aggregates and alternative non-monetary assets, such as equity. Furthermore, we should consider the issue of national diversification in monetary asset substitution in the countries of the Euro area. This is an important topic that can shed additional light on the role of geographical diversition of monetary developments.

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