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Stock market fluctuations and money demand in Italy, 1913-2003

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STOCK MARKET FLUCTUATIONS AND MONEY DEMAND IN ITALY, 1913-2003

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Abstract

This paper examines the impact of stock market fluctuations on money demand in Italy from a long-run perspective. The money demand function estimated by Muscatelli and Spinelli (2000) for a long time span is utilised as a benchmark, adding to the specification information on share prices from the Milan Stock Exchange Reform of 1913 to recent years. For a shorter time period (1938-2003), annual observations on stock market capitalisation and turnover velocity are also considered. The empirical findings suggest that stock market fluctuations help to explain temporary movements in liquidity preference, rather than its secular patterns. Overall, a positive association emerges between an index of stock market prices that includes dividends and real money balances; however, the estimated long-run relationship is unstable. In a dynamic, short-term specification of money demand the estimated coefficient of deflated stock prices is positive, and therefore compatible with a wealth effect, in the years 1913-1980, while in the last two decades a substitution effect has prevailed and the correlation between money and share prices has been negative. This is likely to reflect a change in financial structure and the increasing role of opportunity costs defined over a wider range of assets. These results are confirmed by data on stock market capitalisation. Moreover, in the recent period stock market turnover and money growth are positively correlated.

JEL classification: E41, E44, N14, N24.

Keywords: long-run money demand function, asset prices volatility.

Contents

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1. Introduction[1]

Is the demand for money independent from fluctuations in stock prices? Estimated money demand functions usually consider income or, less often, wealth as a scale variable and control for the opportunity cost of holding money balances, but ignore the stock market. However, since money and shares are main components of aggregate portfolios, a zero restriction should be tested rather than imposed. Earlier studies of the effects of stock prices on money demand include Milton Friedman's work (1988) on the income velocity of money in the US and Choudhry's contribution (1996) on long-run money demand in the US and Canada. Caruso (2001) reports empirical results based on annual observations for a panel of 25 countries and time-series evidence on quarterly data (1960-1998) for 6 industrial countries (Japan, UK, France, Germany, Switzerland and Italy). Bruggeman, Donati and Warne (2003) study the influence of the recent movements in stock prices on money growth in the Euro area. Overall, these findings suggest that periods of asset inflation and deflation have systematic influences on the pattern of monetary aggregates.

This paper evaluates the impact of fluctuations in stock market prices on money demand in Italy in the long-run (1913-2003). Stock market capitalisation has grown rapidly in recent years; it represented 18.2 per cent of nominal income in 1995 and reached 69.7 per cent in the year 2000. At the end of 2003, total shares (comprising equities that are not quoted) accounted for 22.4 per cent of financial assets of households and 46.1 per cent of financial assets of firms; their domestic component is preponderant (foreign shares owned directly by families and firms amount to 2.5 and 11.7 per cent of their financial portfolios, respectively). In recent years, the demand for Italian quoted shares of families has risen quickly, although they still represent only 5.1 per cent of their financial assets (in December 2003); investment fund units (all kinds) of households amount to 12.3 per cent. In the past the Italian stock market

[1] I am grateful to the anonymous referees and Giuseppe Grande for their helpful comments on a earlier draft. I also wish to thank Fabio Panetta for having kindly provided a share price index including dividends for this sample period, Antonio Di Cesare for useful information on stock market statistics, and Federico Barbiellini Amidei for his suggestions about the historical sources.

was undoubtedly less important and investment in shares much less widespread. However, several arguments suggest that a long-run perspective is appropriate.

Firstly, the available empirical evidence indicates that long-run money demand specifications are more informative than short-run estimated demand functions. Stock and Watson (1993) find on annual data a stable long-run money demand in the US in the period 1900-1989. Recent work by Muscatelli and Spinelli (2000), based on co-integration analysis, evaluates the long-run properties of the demand for money in the Italian economy and thus gives an appropriate benchmark for appreciating or discarding an additional role for stock market fluctuations in the behaviour of monetary aggregates. A second reason concerns the role of the stock market as an efficient mechanism for embodying currently available information and anticipating events. Following the present value model, higher share prices are reflected in lower dividend yields that require an expected decline in real interest rates and/or an increase in expected future economic growth rates in order to match long-run equilibrium. This aspect is potentially important in a long-run money demand function, since stock prices may represent a bridge between observed income and the theoretically relevant scale variable, permanent income.

Moreover, it is interesting to consider explicitly an empirical implication acknowledged about forty years ago by Brunner and Meltzer (1963, p. 324) with regard to money demand in the United States: "Data for 1941-50 have been excluded… During these ten years, bond prices were pegged by the Federal Reserve… a … stable demand-for-money equation can be obtained by using the yield on bonds as a measure of the weighted average of the yields on a variety of financial and real assets. When bond prices are controlled and rates of return on real assets are free to fluctuate, this assumption is patently false. Precisely these conditions prevailed from 1946 through March, 1951. Bond prices were controlled, but commodity prices were free to fluctuate". In the past, the stabilisation of interest rates was common practice among central banks, whereas stock market fluctuations were not dampened.

The inclusion of stock prices in a long-run money demand specification may contribute to improved functional relations.

Milton Friedman (1988, pp. 222-3 and footnote 3) suggested several explanations for the observed correlation between stock prices and money in the US economy: (a) a wealth effect: "A rise in stock prices means an increase in nominal wealth… The higher wealth-to-income ratio can be expected to be reflected in a higher money-to-income ratio"; (b) a substitution effect: "The higher the real stock price, the more attractive are equities as a component of the portfolio"; and (c) risk considerations: "A rise in stock prices reflects an increase in the expected return from risky assets relative to safe assets… The resulting increase in risk could be offset by increasing the weight of relatively safe assets in an aggregate portfolio". A fourth channel of influence may also concern (d) the role of stock market fluctuations as (noisy) expectations of future income flows. Friedman indicates that in the US economy the relationship between money and the stock market is basically time-varying; both wealth and substitution effects show up (Friedman 1988, p. 221): "Annual data for a century suggest that the apparent dominance of the wealth effect is the exception, not the rule". Panetta (2002) on Italian data (1979-1994) finds that the relation between the stock market and macroeconomic forces is unstable, and that the sensitivity of equity returns to the macroeconomic state variables often changes sign. Caruso (2001) in a multi-country study notices that wealth effects on money demand are widespread across countries; an inverse relationship between deflated stock prices and money per unit of nominal income, reflecting a substitution effect, also prevails in some countries and time periods, including Italy.

A long-term perspective can shed more light on these issues. To this end, the careful work of data reconstruction and econometric analysis by Muscatelli and Spinelli (2000) on the Italian long-run money demand function has been merged with statistical information on share prices available after the stock exchange reform of March 1913 (introduced by Law 272) that has regulated and modernised the Italian stock market.

2. The long-run behaviour of deflated stock prices and real money balances

This paper considers the pattern of per capita real money balances, deflated stock prices and the opportunity cost of money in the period 1913-2003; variables are plotted in Figures 1 and 2. Annual data on money (M2), real per capita income, prices (CPI) and interest rates until 1996 are culled from the Appendix of the Muscatelli and Spinelli (2000) paper; observations have been updated with comparable data obtained from the statistical appendix of recent Bank of Italy Annual Reports and from the IMF International Financial Statistics (IFS) tape.

Broad money (M2) is defined as total monetary base and bank deposits (the reader is referred to the Muscatelli and Spinelli contribution for detailed data description); it has been updated with the growth rates of money culled from IFS data (M2, national definition, line 39m) and, in the last three years, with the annual variation in the Italian component of the M2 aggregate in the Euro area (from the Bank of Italy Annual Report, Statistical Appendix, Table AD4). Real GDP, population and the price variable, a cost-of-living index, have been updated with observations taken from the IFS (lines 99bvr, 99z, 64) and Istat data. The government bond yield and the average rate on bank deposits (a measure of the own rate for M2) have also been complemented with comparable recent data (IFS lines 61b and 60l).

Real money balances per capita (Figure 1) show three main patterns in the period 1913-2003; moderate growth rates can be observed from 1913 to 1942 (with a sudden drop due to inflation in 1943-45), very rapid growth from the end of World War II until the first oil shock in 1973 and slower increases in the last 30 years. Money per unit of nominal income (the reciprocal of money velocity) fluctuates widely between the World Wars; after a minimum in 1947 it reaches a peak in the 1970s, goes down in the 1980s and tends to stabilise in the 1990s (Figure 2).

Annual observations of stock prices are merged from three sources; they are deflated by the CPI and refer to the month of December (in order to match with end-of-period money data). I have culled data from 1913 to 1938 from Rosania (1954,

Table 1); they are based on a Bank of Italy sample of about 40 shares. Observations from 1939 to 1957 are taken from Mondani (1978, Table 1) and refer to a Mediobanca sample starting from 88 quoted stock prices. Data from 1958 to 2003 are culled from the IFS tape (line 62) and are based on the MIB index of the Milan Stock Exchange (1990=100). These three time series have been merged backwards (using the overlapping years of the samples, 1958 and 1938); they do not include dividends.

The behaviour of deflated stock prices (Figure 1) in a time span of 90 years shows (approximately) four prolonged phases of growth in real terms (1932-1943, 1949-1961, 1977-1986, 1993-2000), two periods of high variability (1922-1931, 1946-1948), an abrupt drop (1944-1945) and four periods of declining prices (1913-1921, 1962-1976, 1987-1992, and 2001-2002). The recent asset price fluctuations in real terms (on a logarithmic scale) are not particularly large by historical standards.

The volatile behaviour of stock prices contrasts with the smoother pattern of interest rates, especially between the two World Wars, when the bond yield ranged from 4.1 to 5.8 per cent and the deposit rate oscillated from 2.3 to 3 per cent, with the opportunity cost showing an upward tendency. The yield on money balances started to rise in the 1960s, lowering the opportunity cost; after the first oil shock it follows the bond yield fairly closely. The opportunity cost widens again in the period 1987-1995 and very recently.

I have also collected data on the real capitalisation and turnover velocity of the Italian stock market for a shorter time period (1938-2003). Wealth or substitution effects may be more directly linked to developments in market capitalisation, rather than prices. The inclusion of the stock market turnover in a money demand function makes it possible to test the hypothesis, suggested by Friedman (1988 p. 235), that money responds to changes in the volume of financial transactions (following the "transaction" motive of money demand). Moreover, to the extent that stock market volatility and turnover are positively correlated, a higher volume of transactions may

Figure 1 - Deflated stock prices and per capita real money balances, 1913-2003

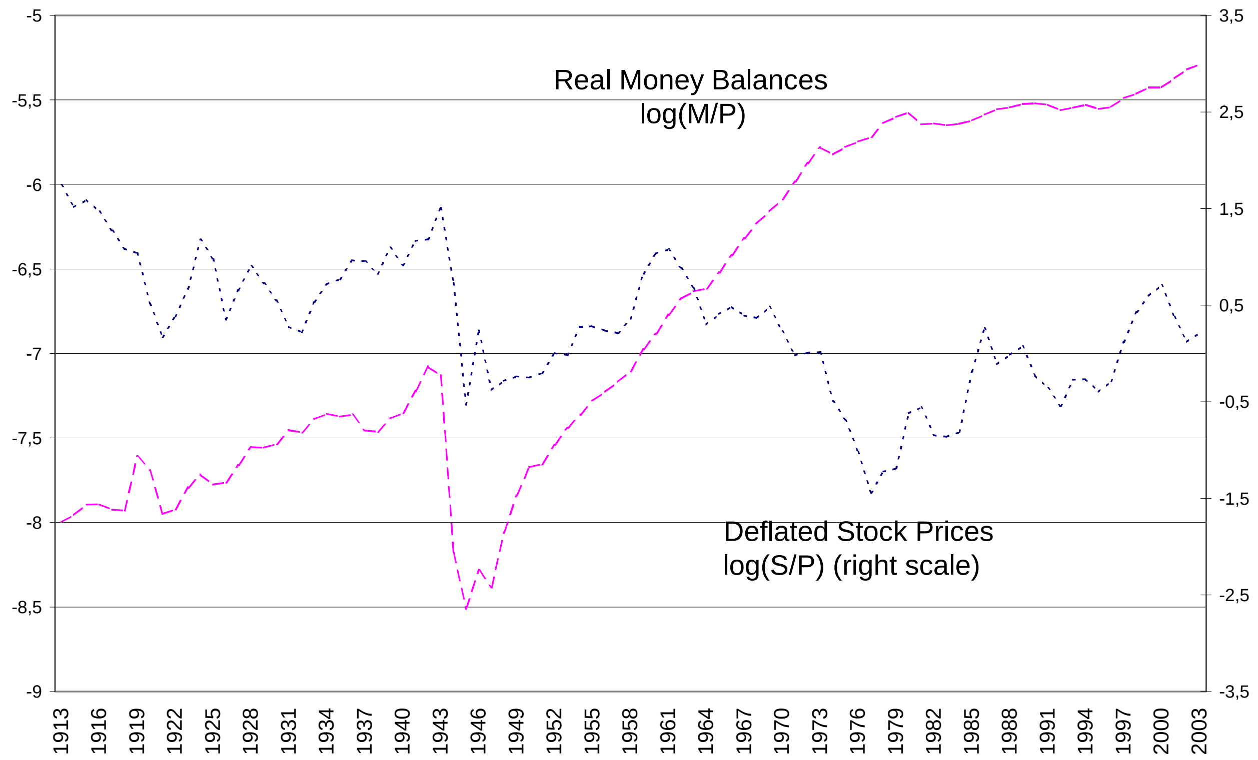


Figure 2 - Money per unit of nominal income and the opportunity cost, 1913-2003

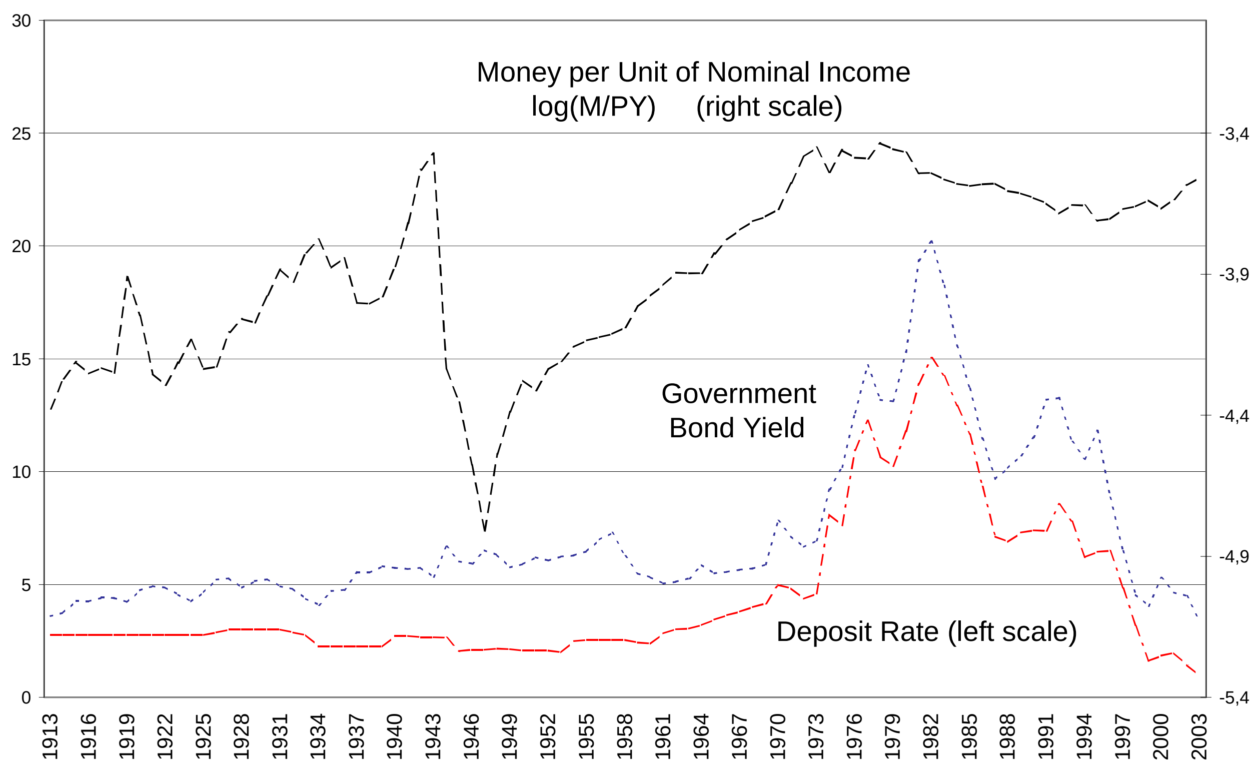
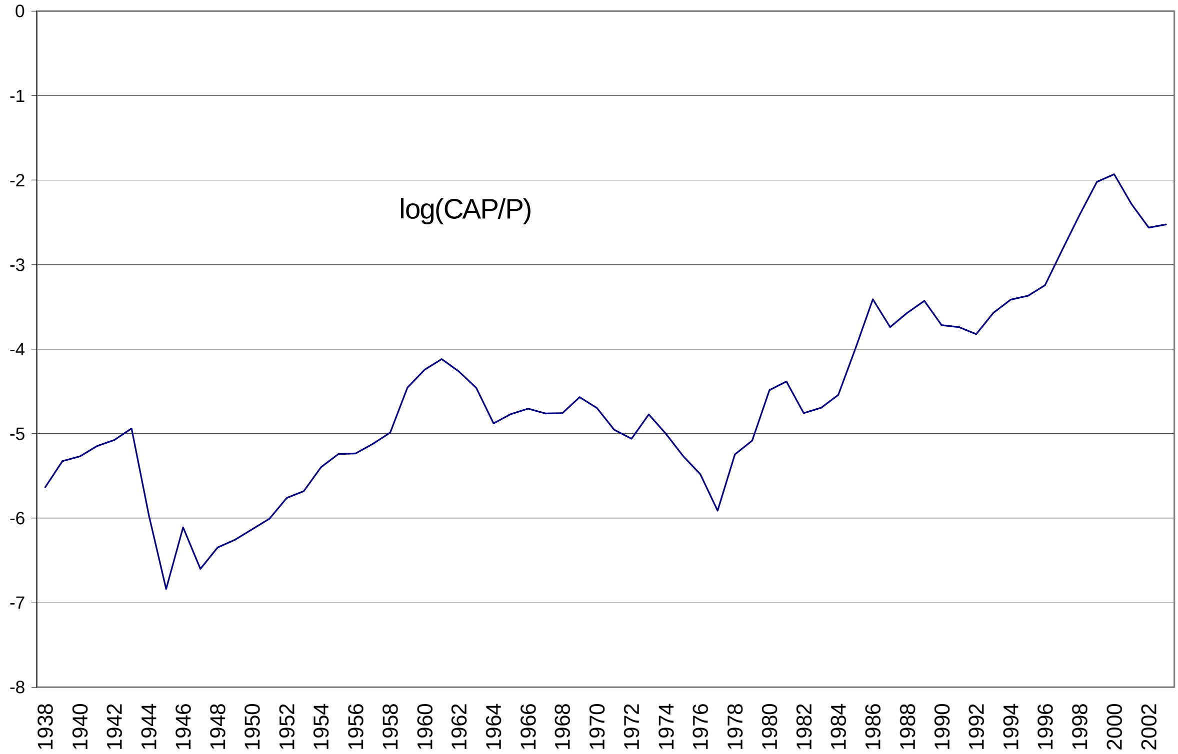


Figure 3 - Stock market capitalisation and turnover velocity, 1938-2003



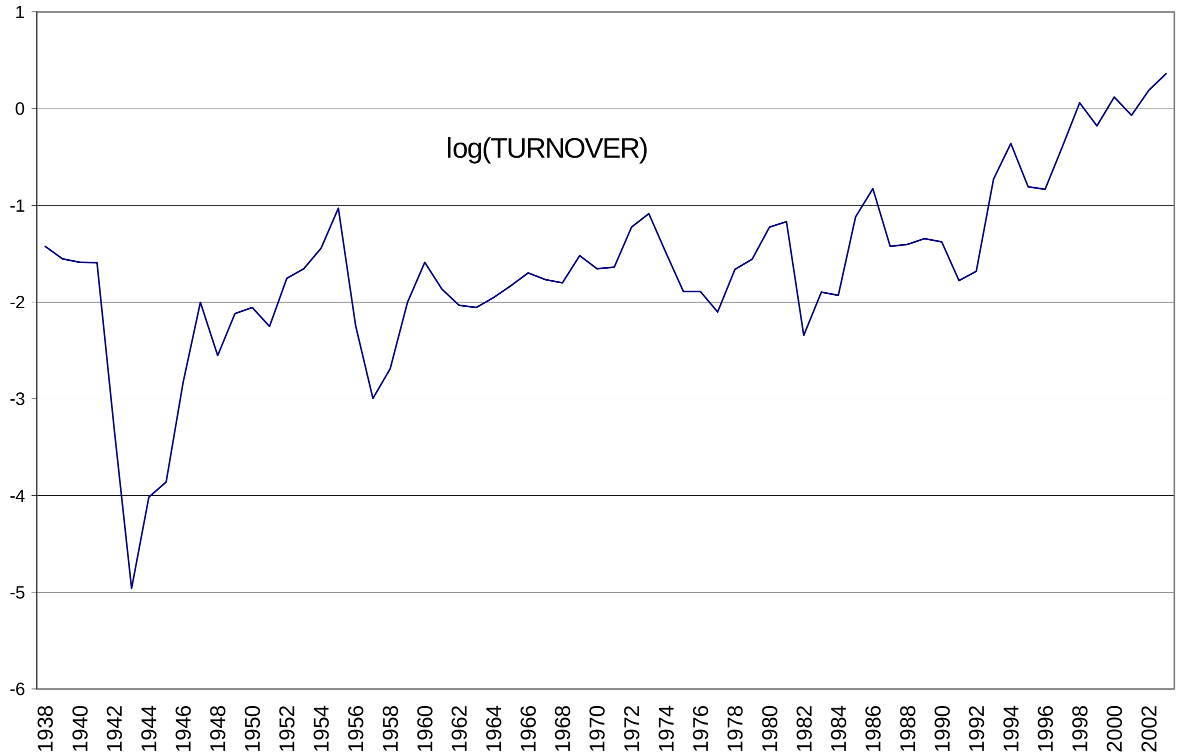
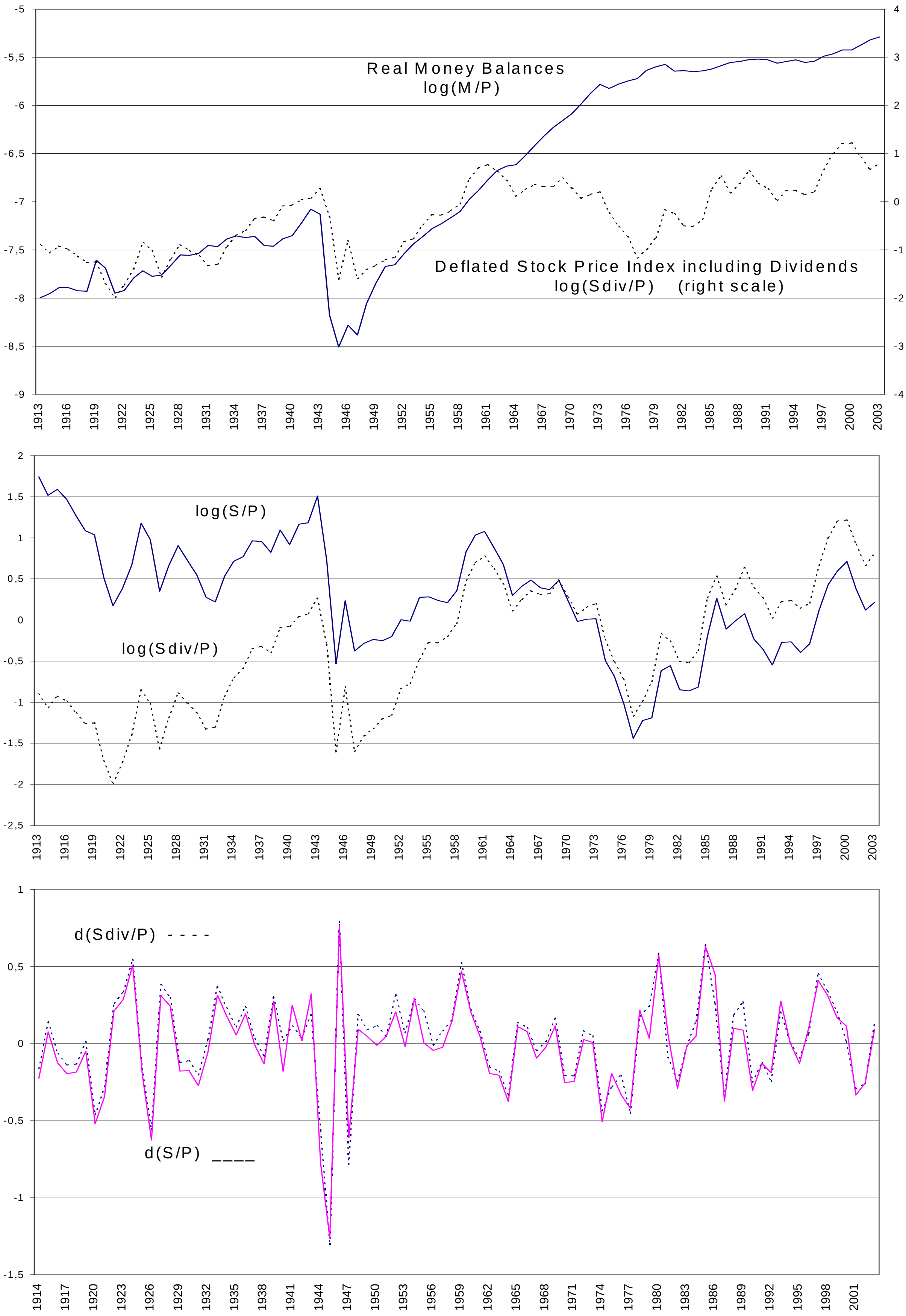


Figure 4 - A comparison with a stock market index that includes dividends



also be related to "precautionary" money holdings.[2] In any case, it must be considered that observations on capitalisation and turnover are culled from heterogeneous sources and are more liable than stock prices to measurement errors; for these reasons, they are included in the analysis mainly as a check on the overall results.[3]

The impact of stock prices on money could be better evaluated on the basis of a gross stock price index than on a capital index. Panetta and Violi (1999) introduce and discuss a price index that includes dividends; its pattern is plotted in Figure 4. For the recent years (1996-2003) this series has been updated utilising the growth rates of a comparable gross index (Datastream source). The capital and gross indexes fluctuate in a very similar manner, but the index that includes dividends (differently from deflated prices) shows a long-run upward trend due to the contribution of the dividend flow. In practice, the difference between the annual growth rates of these indexes is small because the large variations in stock prices dominate the relatively smooth pattern of dividends.[4]

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[2] Fong (2003), among others, studies the links between trading volume and stock price volatility. Daily US data (1980-1999) indicate that past volatility shocks have a greater positive effect on current volume than past volume shocks have on current volatility. This suggests that stock market volatility may cause a higher turnover; of course, at annual frequencies a contemporary positive relationship between volume and volatility is more likely.

[3] Data on stock market capitalisation (2003-1975) have been culled from Table 1, "Borsa Italiana – Fatti e cifre sul 2003". They have been merged backwards with: Bank of Italy data (1974-1971, "Relazione del Governatore", various years); annual observations (1970-1961) from Table 8, line c, Pivato and Scognamiglio (1972); annual data (1960-1948) from Table 5, Barbiellini Amidei and Impenna (1999); and Istat data (1947-1938) culled from "Annuario Statistico Italiano, 1949-50" (an issue with statistical information on the Italian stock exchange from 1938). Data on turnover velocity (the ratio of exchanged shares to total market capitalisation) have been obtained from the same sources (they refer to Table 10, instead of Table 5, in Barbiellini Amidei and Impenna, 1999). The heterogeneity of sources requires a note of caution; these series are useful proxies - rather than exact measures - of the underlying stock market patterns.

[4] In the last 90 years the correlation between the growth rate of the deflated net index of Italian stock prices d(S/P) discussed in this paper and an index including dividends d(Sdiv/P) introduced by Panetta and Violi (1999) is 0.973. However, log(Sdiv/P) has an upward trend, while log(S/P) has not, due to the flows of rewards to the asset owners (Figure 4). The difference in the growth rates of the two indexes is 3.6 per cent per annum, with a standard error of 0.7 per cent (1914-2003). On average, dividends have been higher in past years. It can be shown that splitting the sample size in two, the measure d(Sdiv/P)-d(S/P) is equal to about 5 per cent per annum (standard error 1.1 per cent) in the period 1914-1959; it goes down to 2.1 per cent (standard error 1.0 per cent) in the years 1960-2003. Based on annual data, in the period 1939-2003 the correlation coefficient between the variations in the stock market capitalisation in real terms d(CAP/P) and deflated stock prices d(S/P) is equal to 0.935.

Some univariate properties of the data are described in Table 1. Per capita real money balances, real income and interest rates are integrated of first order, on the basis of a Dickey Fuller (DF) or Augmented Dickey Fuller (ADF) test (in this sample of annual data, the lag order has been selected by the Schwarz Bayesian criterion); they are non-stationary in levels (both excluding and including a linear trend) and stationary in first differences. Inflation (the first difference of the price level) and the opportunity cost are stationary. These results are analogous to the findings of Muscatelli and Spinelli (2000, pp.723-4) for a longer sample period (1861-1996).

Regarding the stock market, it can be noted that capitalisation in real terms (1938-2003) is integrated of first order; the presence of a unit root cannot be excluded in levels, while market capitalisation is stationary in first differences. Turnover velocity is stationary around a linear deterministic trend. Instead, results for deflated stock prices (1913-2003) are less clear-cut. The ADF test rejects non-stationarity around the sample mean, but at the 6 per cent level only. This result suggests the presence of an important mean reverting component in deflated stock prices (S/P); however, it cannot be excluded at usual confidence levels that they are integrated of first order, I(1).[5] Non-stationarity cannot be ruled out in the case of the deflated gross stock price index (Sdiv/P), which comprises the contribution of the dividend flow; this variable is distributed I(1).

Spectral density estimates at zero frequency (Table 1, column 4) indicate that inflation, the deposit rate and, to a lesser extent, the bond yield have more persistent patterns than the average. The estimated size of the unit root is higher for these nominal variables, suggesting near-integrating behaviour in the long run (inflation and interest rate changes are formally distributed as I(0), but are closer to I(1) than the other variables in this sample). The persistence of nominal interest rates can be

[5] Analogously to the ADF test, a Phillips-Perron test for unit root applied to the level of deflated stock prices (1913-2003) rejects non-stationarity at about the 8 per cent level: Z(t)=-2.673 (McKinnon approximate p-value=0.0789). This empirical finding matches results by Fama and French (1988) for the US market (1926-1985); they show that stock prices have a mean reverting (and partially predictable) component, more evident at horizons of 3 to 5 years.

Table 1 - Unit root tests and persistence measures

	Levels		First differences		
	(1)	(2)	(3)	(4)	Test
Deterministic	DF or ADF test	DF or ADF test	DF or ADF test	Persistence	outcome
component:	Constant	C+ linear trend	Constant	Constant	
Variables:					
Per capita real	-.474 ADF(1)	-1.899 ADF(1)	-7.148 ADF(1)	.876	I(1)
money balances	(.898)	(.655)	(.000)	(.474)	
	(p-value)	(p-value)	(p-value)	(s. e.)	
Real income	.121 DF	-2.001 DF	-7.922 DF	1.157	I(1)
per capita	(.967)	(.597)	(.000)	(.626)	
Price level	-1.380 ADF(1)	-.939 ADF(1)	-4.767 DF	1.918	I(1)
	(.591)	(.952)	(.000)	(1.038)	
Nominal income	-1.126 ADF(1)	-.0700 DF	-4.894 DF	2.632	I(1)
per capita	(.705)	(.974)	(.000)	(1.359)	
Govt bond yield	-1.363 DF	-.754 DF	-5.976 ADF(1)	1.545	I(1)
	(.599)	(.970)	(.000)	(.836)	
Deposit rate	-1.035 DF	-.492 DF	-6.682 ADF(1)	2.047	I(1)
	(.741)	(.985)	(.000)	(1.108)	
Opportunity cost	-3.524 ADF(1)	-3.515 ADF(1)	-9.542 ADF(1)	.296	I(0)
(GBY-DR)	(.007)	(.038)	(.000)	(.160)	Mean-
					reverting
Real interest rate	-4.697 ADF(1)	-4.837 ADF(1)	-10.639 ADF(1)	.130	I(0)
(GBY-dP)	(.001)	(.001)	(.000)	(.069)	Mean-
					reverting
Deflated share	-2.777 ADF(1)	-2.879 ADF(1)	-8.950 ADF(1)	.469	5%: I(1)
prices (S/P)	(.062)	(.169)	(.000)	(.242)	10%: Mean-
					reverting
Share prices in real	-1.720 ADF(1)	-2.950 ADF(1)	-9.044 ADF(1)	.370	I(1)
terms including	(.421)	(.147)	(.000)	(.200)	
dividends (Sdiv/P)					
Market capitalisation	-.610 ADF(1)	-2.158 ADF(1)	-7.105 ADF(1)	.628	I(1)
in real terms	(.870)	(.515)	(.000)	(.360)	
Stock market	-1.816 ADF(1)	-3.689 ADF(1)	-6.970 ADF(1)	.337	I(0)
turnover velocity	(.372)	(.023)	(.000)	(.193)	Trend-
					stationary

Note: Dickey-Fuller or Augmented D-F (ADF) test. The lag order ADF(n) has been selected by the Schwarz Bayesian criterion. MacKinnon interpolated p-values in parentheses. Persistence measures are Bartlett estimates of spectral density at zero frequency and window width of 15 years (asymptotic standard errors in parentheses). Variables in levels are in logs, with the exception of the interest rates (per centage points). Sample period 1913-2003 (T=91); for two variables (stock market capitalisation and turnover) the sample period is 1938-2003 (T=66).

ascribed to sticky inflation and also to some extent to direct or indirect stabilisation of nominal yields before the first oil shock, both by the central bank (the bond rate) and the commercial banks (deposit rates).

3. Empirical findings based on a co-integration analysis

In this Section, the hypothesis of a stationary long-run relationship between real money balances, the traditional determinants of money demand (income, prices and interest rates) and stock market variables will be tested. To this end, the co-integration analysis proposed by Muscatelli and Spinelli (2000) is introduced as a benchmark. However, the use of different econometric techniques may be of help in ascertaining the robustness of the results, and some alternative empirical approaches, based on multiple co-integration and VAR in levels or assuming the presence of broken deterministic trends, will also be followed.

The co-integration approach to long-run money demand (results are presented in Table 2) follows Muscatelli and Spinelli (2000) and is based on the dynamic ordinary least squares (DOLS) methodology developed by Stock and Watson (1993). These authors show that the estimation of long-run elasticities is asymptotically efficient when leads and lags of the first differences of the right-side variables are introduced in a co-integrating regression in levels. Allowing for two leads and lags of growth rates, the estimation period is 1916-2001 (86 years).

Muscatelli and Spinelli (2000), as well as Stock and Watson (1993), treat inflation as a non-stationary variable, distributed I(1) and introduce the change of the price level π=dP in the long-run co-integrating regression, while the first difference of inflation dπ (accelerating or decelerating prices) is utilised in the second-stage, as an explanatory variable in the dynamic specification. It is convenient to have a common benchmark, and I will do the same here. Several papers find that the price level is particularly sticky, a variable integrated of second order, I(2), and it is difficult to judge if the stationarity result for the inflation rate reported in Table 1 is due to a

Table 2 - Long-run money demand and stock prices in Italy: DOLS regressions

DYNAMIC OLS (DOLS): STOCK-WATSON METHODOLOGY										
β_P	β_Y	β_{BR}	β_{DR}	β_{dP}	β_{WAR1}	β_{WAR2}	β_{RSP}	β_{RCAP}	$_cR^2$ SER	Cointegration Test
(1) Dependent variable: log(M), 1916-2001 (allowing for two leads and lags)										
.924	1.557	-.033	.054	-	-.253	.287			.9994 .0897	-4.321*
.017	.066	.017	.018		.062	.074				DF
(.000)	(.000)	(.053)	(.005)		(.000)	(.000)				
(2) Dependent variable: log(M/P), 1916-2001										
-	1.282	-.076	.103	-.178	-.249	.188			.9889 .1019	-4.171*
s.e.	.026	.015	.016	.166	.078	.102				ADF(2)
	(.000)	(.000)	(.000)	(.290)	(.002)	(.070)				
Including deflated stock prices (S/P):										
(3) Dependent variable: log(M), 1916-2001										
.910	1.590	-.035	.050	-	-.233	.316	-.059		.9996 .0910	-4.366*
.021	.076	.018	.020		.069	.080	.048			DF
(.000)	(.000)	(.060)	(.018)		(.001)	(.000)	(.226)			
(4) Dependent variable: log(M/P), 1916-2001										
-	1.286	-.075	.106	-.147	-.262	.181	.037		.9886 .1032	-4.059
s.e.	.029	.017	.017	.196	.084	.109	.048			ADF(2)
	(.000)	(.000)	(.000)	(.456)	(.003)	(.103)	(.447)			
Including deflated stock prices, gross index (Sdiv/P):										
(5) Dependent variable: log(M), 1916-2001										
.933	1.456	-.038	.066	-	-.252	.210	.061		.9994 .0901	-4.598*
.018	.086	.018	.020		.065	.085	.036			DF
(.000)	(.000)	(.040)	(.002)		(.000)	(.017)	(.095)			
(6) Dependent variable: log(M/P), 1916-2001										
-	1.191	-.079	.114	-.095	-.247	.091	.084		.9894 .0996	-4.168
s.e.	.050	.016	.017	.184	.077	.110	.039			ADF(3)
	(.000)	(.000)	(.000)	(.606)	(.002)	(.412)	(.036)			
Including stock market capitalisation in real terms:										
(7) Dependent variable: log(M), 1938-2001 (allowing for two leads and lags)										
.876	1.754	-.025	.040	-	-	.356		-.048	.9996 .0467	-5.064**
.024	.081	.013	.015			.111		.026		DF
(.000)	(.000)	(.056)	(.010)			(.003)		(.069)		
(8) Dependent variable: log(M/P), 1938-2001										
-	1.482	-.063	.084	-.566	-	.719		-.064	.9801 .1323	-4.680*
s.e.	.069	.011	.013	.193		.165		.031		ADF(1)
	(.000)	(.000)	(.000)	(.006)		(.000)		(.047)		

Note: Estimated long-run coefficients and standard errors (p-values in parentheses). The residual-based tests for co-integration are Dickey-Fuller or Augmented Dickey Fuller tests applied to the demeaned equilibrium errors (calculated from the long-run coefficients); lags are selected according to the Schwarz Bayesian criterion (0 to 3 lags, chosen lag order in parentheses). Phillips and Ouliaris (1990, Table 2b page 190) report asymptotic critical values for demeaned variables. With four non-deterministic variables (regressions 1-2) critical values are -3.959 (15 per cent level), -4.157 (10), -4.454 (5), -5.073 (1 per cent). With five non-deterministic variables (regressions 3-8) critical values are –4.236 (15 per cent level), -4.431 (10), -4.710 (5), -5.281 (1 per cent). Equilibrium errors where the null of a unit root (no co-integration) is rejected at the 1, 5 and 10 per cent level or better are marked by *******, ****** and *****, respectively.

break in the sample (a discontinuity in the price level following the Second World War); the inflation rate has the highest persistence measure across the sample variables and this may justify the introduction of the price level changes in the long-run regression.

The long-run demand for money is modelled as in Muscatelli and Spinelli (2000, Table 2); the levels of deflated stock prices are added to the specification (and their growth rates contribute to the leads and lags):

(1) $(M)_t = c + \beta_{wars}(dummywar)_t + \beta_P(P)_t + \beta_Y(Y)_t + \beta_{BR}(RL)_t + \beta_{DR}(RD)_t + \beta_{dP}(INF)_t + \beta_{SP}(S/P)_t$ + two leads and lags of growth rates of the explanatory variables + ε_t

(2) $(M/P)_t = c + \beta_{wars}(dummywar)_t + \beta_Y(Y)_t + \beta_{BR}(RL)_t + \beta_{DR}(RD)_t + \beta_{dP}(INF)_t + \beta_{SP}(S/P)_t$ + two leads and lags of growth rates of the explanatory variables + ε_t

The estimated long-run regression (1) in Table 2 considers nominal money log(M) as the dependent variable and indicates co-integration between money, income, the price level and interest rates (at the 10 per cent level). The homogeneity assumption applied to prices (log(P) equal to one) is rejected in this sample; however, the estimated coefficient is not too different from unity, suggesting that a reference to money balances in real terms is an acceptable working hypothesis (Table 2, regression 1); the estimated coefficient β_P is .92 with a standard error of about .02 (Muscatelli and Spinelli, 2000 Table 2 estimate $\beta_P = 1.07$ with a standard error of .05 in 1864-1994). Income and interest rate elasticities have the correct signs; dummies for the two world wars are also included.[6]

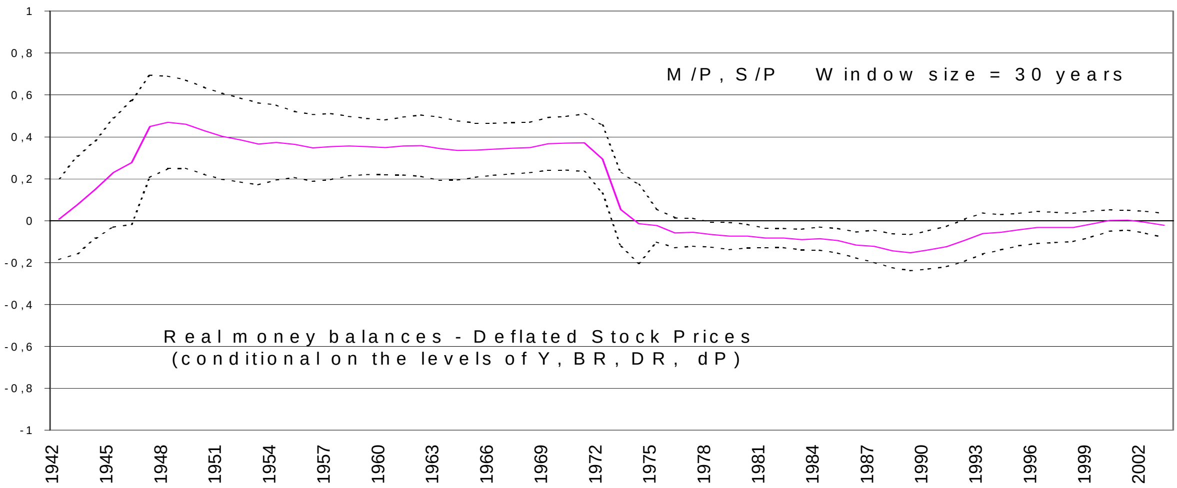
[6] Dropping the dummies, which signal a monetary overhang during the Second World War and lower than average money balances in First World War, the β_P coefficient goes up to .96 with a standard error of .02, a result that gives support to the homogeneity assumption. Results presented in Table 2 follow the specification proposed by Muscatelli and Spinelli (2000) and are qualitatively robust to the introduction of yearly dummies for the periods 1915-18 and 1940-45 (controlling for changes in the relationship during the war years). Regarding the contribution of the deflated gross stock price index (Sdiv/P) to the equation for real money balances (Table 2, regression 6), the estimated coefficient is positive (.071 with standard error .038) and significant at the 6.7 per cent level when annual dummies are introduced.

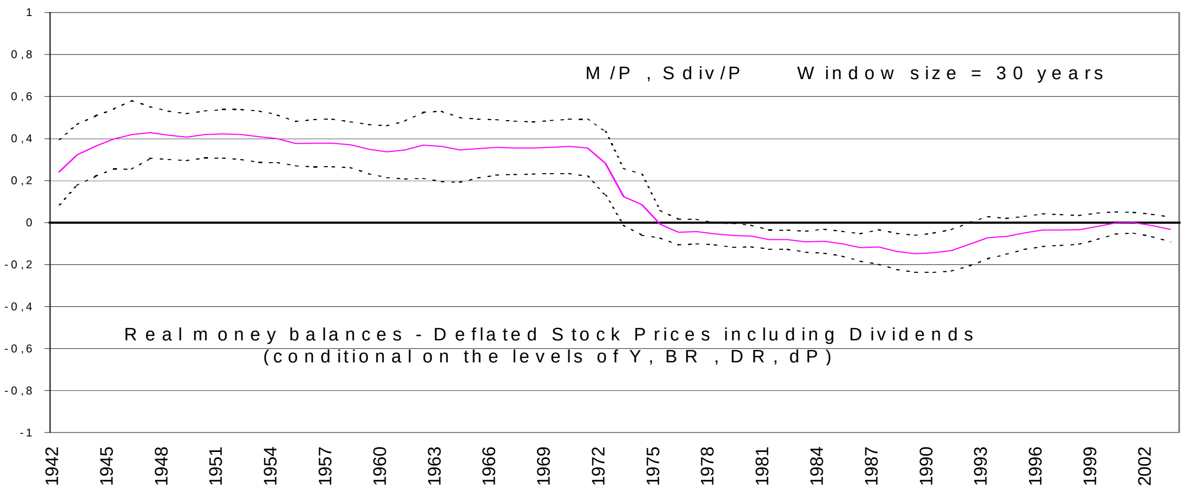
Regression (2) in Table 2 refers to real money balances and adds the growth rate of prices in the co-integrating relationship, which has the correct negative sign (agents try to escape the inflation tax by lowering their money holdings) but is imprecisely estimated. The coefficient of real income is higher than one (1.28 with a standard error of about .03) but lower than the income elasticity found by Muscatelli and Spinelli (2000, Table 2 pag. 727) for the years 1864-1994 (1.94 with a standard error of .21), implying a slower overall downward trend of money velocity in the more recent period.

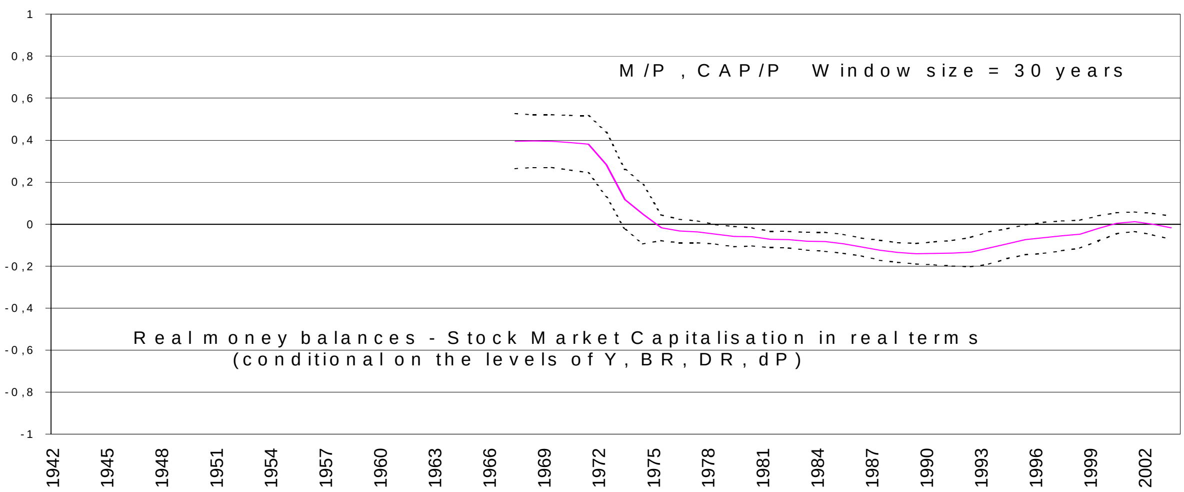
Regressions (3-4) in Table 2 add the level of deflated stock prices to the estimated long-run demand for money. Overall, the empirical results are not favourable to the introduction of information from equity prices in the long-run money demand equilibrium relationship. First, it is unclear whether deflated stock prices are I(1) in levels and are thus legitimate candidates for joining the co-integrating regression (see Table 1). Furthermore, assuming also that they are non-stationary, the estimated coefficient β_{SP} has large standard errors and uncertain sign; in one case (the regression for long-run real money balances) the residuals are not stationary, ruling out co-integration.

The role of the gross index of deflated stock prices (Sdiv/P) is examined in the long-run regressions (5-6). An index that includes dividends is a better proxy for financial wealth and, as is to be expected, a wealth effect shows up more clearly in these estimates. In the full sample, deflated stock prices and the flow of dividends seem to have a positive influence on long-run money demand. However, the introduction of the levels of the variable (Sdiv/P) allows some slight improvement in the standard error of the long-run regression (6), but the fitted regression for real money balances does not reach conventional significance levels in a co-integration test. A lower common order of integration is confirmed for nominal money in regression (5), but in this case the positive coefficient of the gross stock index is less

Figure 5 - Rolling regressions of estimated single-equation long-run relationships







precisely estimated. The reason for these results is that the link between share prices and money is time-varying over the sample period (see Figure 5).

Stock market capitalisation in real terms enters the equilibrium money demand function with a negative sign in the years 1938-2003; the estimated coefficients are significant at about the 7 and 5 per cent level in the DOLS regression with nominal and real money balances as dependent variables, respectively. This finding suggests that, overall, in this period money and shares mainly competed with each other in investors' portfolios. A substitution effect in a money market–stock market equilibrium, controlling for other determinants of money demand, is also not in contrast with Friedman's empirical results (1988) for the US economy based on a century of data, indicating that "the wealth effect is the exception, rather than the rule".

Figure 5 reports rolling estimates, computed with a window size of 30 years, for the coefficients of the levels of the stock market variables in the fitted long-run regressions (dependent variables are real money balances and results take into account the contribution of real income, interest rates and inflation). Regardless of the specific stock market variable, it is apparent that a wealth effect shows up in the first two-thirds of the sample (until the 1970s) while substitution effects are prevalent in recent years. The positive influence of the gross index (Sdiv/P) is more precisely estimated until Second World War (a likely explanation is that the contribution of the flow of dividends is larger in the first half of the sample) but later the contribution of the gross and net indexes looks very similar.

Until the 1970s, the stock market variables capture a wealth effect that reflects the store of value function of the monetary aggregate (the level of stock prices proxies the pattern of financial wealth). Later, the gradual introduction of financial innovations (including the spread of investment funds) favoured the prevalence of substitution effects between monetary and stock market assets; in recent years, share prices have entered money demand mainly as an opportunity cost. The transition from a monetary aggregate held primarily as a store of value to real money balances that reflect more directly a transaction motive and the presence of opportunity costs defined over a larger range of assets help to explain the discontinuity.[7]

The Stock-Watson methodology is based on a single-equation approach, while the potential multiplicity of co-integrating relationships among the endogenous variables cannot be excluded a priori. A system analysis following Johansen's methodology may be more informative (Johansen, 1995). However, it requires specification of each variable entering the model, not just the long-run money demand function; a complete modelling of the behaviour of real money balances, inflation, real income, interest rates and share prices over the last 90 years would be an ambitious task.[8] For these reasons, I report here some preliminary findings for a smaller system, including four endogenous variables only (real income, money, the long-term interest rates and deflated share prices). The specification that includes the gross share price index, inflation and the deposit rate (or a stationary opportunity cost) does not seem robust to changes in the lag structure and its results are more difficult to interpret; controlling for larger systems, for the contribution of the dividend flow and for additional exogenous variables is left to further research.

The empirical results obtained following Johansen's approach to co-integration are shown in Table 3. The estimated (4x4 variables) co-integrated VAR system is:

(3) CVAR (M/P, Y, S/P, RL)

Two stationary (co-integrating) vectors can be detected in the system (M/P, Y, S/P, RL) at usual confidence levels, according to both the maximal eigenvalue and the

[7] Angelini, Hendry and Rinaldi (1994) observe that money demand in Italy is better modelled (on quarterly data) utilising net financial wealth as a scale variable in the years 1975-79, while in 1983-91 the appropriate scale factor is found to be domestic demand; this is interpreted as reflecting the growing role of money as a transaction medium, instead that as a store of value.

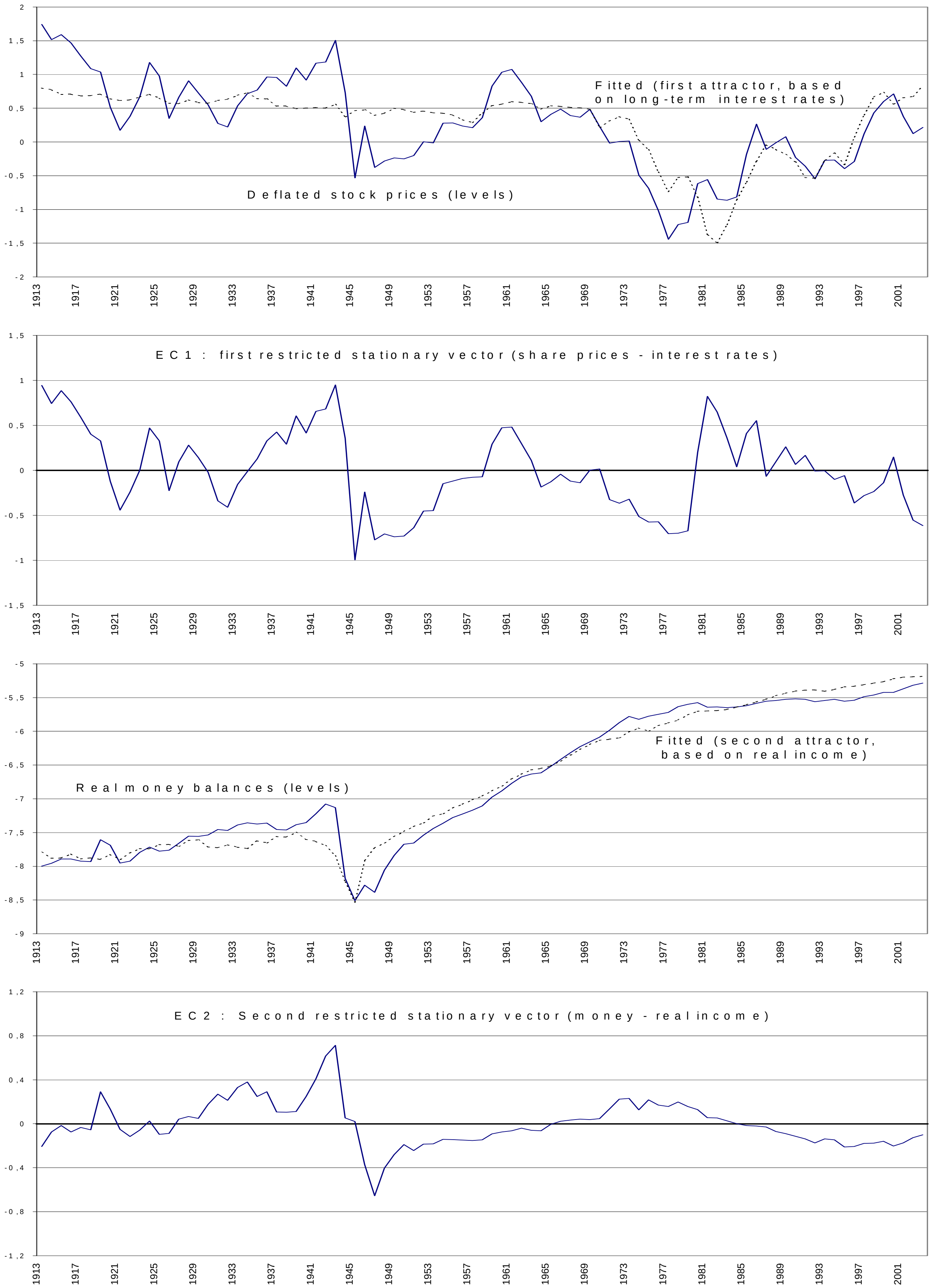
[8] Muscatelli and Spinelli (2000 pp. 725-6) also note that Johansen's approach to co-integration is sensitive to the choice of the lag order.

Table 3 – Multiple co-integrating relationships: a four-variables system

System Statistics:	1 lag	2 lags	3 lags	4 lags	5 lags	
AIC Akaike information criterium	30.57	43.77	68.94	63.71	59.97	
LR lag exclusion test (p-value)	(.000)	(.000)	(.056)	(.094)	(.301)	
Co-integration tests:	Maximal Eigenvalue test , 3 lags			Trace test, 3 lags		
(Johansen's methodology)	Null r=0	Null r<=1	Test	Null r=0	Null r<=1	Test
(1913-2003, T=91)	Alt r=1	Alt r=2	result	Alt r>=1	Alt r>=2	result
Bivariate tests:						
M/P, S/P	11.67	-	0 vector	12.54		0 vector
M/P, RL	4.07	-	0 vector	4.29	-	0 vector
M/P, Y	11.92	-	0 vector	12.27	-	0 vector
S/P,RL	14.81*	-	1 vector	19.69**	-	1 vector
S/P,Y	9.83	-	0 vector	9.86	-	0 vector
RL,Y	3.52	-	0 vector	3.57	-	0 vector
Multivariate tests:						
M/P,RL,Y, S/P	26.14*	21.92**	2 vectors	57.16**	31.02*	2 vectors
Unrestricted stationary vector 1	(.3729 RM -.5548 Y -.0331 RL -.2902 S/P) (α_1M/P -α_2Y -α_3RL -α_4S/P)					
Unrestricted stationary vector 2	(.6284 RM -.8325 Y +.0135 RL +.1421S/P) (β_1M/P -β_2Y +β_3RL +β_4S/P)					
Proposed normalisation:	Vector 1: Asset pricing equation (α_4=1)					
	Vector 2: Money demand equation (β_1 =1)					
Imposed joint restrictions:	α_4=1 α_1=0 α_2=0 β_1=1 β_3=0 β_4=0					
LR test of restrictions:	CHSQ(2)	1.4594	(.482)	Test outcome: restrictions not rejected		
Estimated restricted vectors:	Vector 1: S/P = -.138 RL + EC1			Vector 2: M/P = 1.379 Y + EC2		
(asymptotic standard error)		(.023)			(.044)	
	E-C models (1914-2003, T=90) Dependent variables:					
	(1)	(2)	(3)	(4)	(5)	(6)
Explanatory variables:	d(S/P)$_t$	d(S/P)$_t$	d(S/P)$_t$	d(M/P)$_t$	d(M/P)$_t$	d(M/P)$_t$
EC1$_{t-1}$	-.219	-.128	-.153	-.017	-.013	-.019
(standard error)	(.098)**	(.069)*	(.067)**	(.043)	(.030)	(.028)
EC2$_{t-1}$	-.032	.380	.259	-.220	-.242	-.303
(standard error)	(.220)	(.161)**	(.160)	(.097)**	(.066)***	(.062)***
DUMMYW2	(condi-	-.010	.187	(condi-	.067	.111
(standard error)	tional on	(.136)	(.148)	tional on	(.065)	(.063)*
d(S/P)$_t$	lagged	-	-	lagged	.106	.100
(standard error)	informa-			informa-	(.047)**	(.041)**
d(M/P)$_t$	tion set)	.778	.555	tion set)	-	-
(standard error)		(.241)***	(.245)**			
d(Y)$_t$	-	.898	1.094		.753	.807
(standard error)		(.514)*	(.499)**		(.208)***	(.196)***
d(RL)$_t$	-	-.090	-.089		-.004	-
(standard error)		(.025)***	(.024)***		(.011)	
d(P)$_{t-1}$	-	-	-.453		-.152	-.135
(standard error)			(.161)***		(.072)**	(.068)*
(RL-DR)$_t$						-.032
(standard error)						(.011)***
$_cR^2$	.294	.360	.409	.376	.487	.538
SER	.265	.249	.239	.117	.104	.099
Serial correlation (p-value)	(.528)	(.146)	(.475)	(.838)	(.286)	(.365)
Functional form (p-value)	(.078)	(.079)	(.107)	(.000)	(.000)	(.000)
Normality (p-value)	(.307)	(.416)	(.636)	(.000)	(.000)	(.000)
Heteroscedasticity (p-value)	(.295)	(.128)	(.265)	(.000)	(.000)	(.000)

Note. Regressions include a constant, not reported. P-values in parentheses. ***, ** and * = 1, 5 and 10 per cent.

Figure 6 - Two co-integrating vectors in a four-variable system, 1913-2003



trace test statistics.[9] The appropriate identifying restrictions, among several plausible alternatives, can be chosen according to empirical testing and some basic economic priors. A check based on bivariate co-integration tests among all the endogenous variables indicates the existence of a stationary long-run link between share prices and interest rates. Another stationary vector can be based on a money demand relationship, linking real money balances and real income (in this case, however, the bivariate test presented in Table 3 does not reach conventional significance levels).

These two unrestricted co-integrating vectors may represent: a) a mean reverting asset price equation (between interest rates and stock prices) and: b) a stationary velocity function (between money and income). The imposed joint restrictions are not rejected (Table 3) and the estimated restricted co-integrating vectors are: a) (S/P) = -.138 (RL) + EC1 and b) (M/P) =1.379 (Y) + EC2.

The long-run features of this four-variable system are plotted in Figure 6. The negative correlation between interest rates and stock prices is expected a priori. Assuming stable expected future cash flows in a standard present value model of stock prices and constant risk patterns, when the yield on bonds goes up the yield on shares must rise to maintain equilibrium, and the stock price goes down. Rising long-term interest rates may also be regarded as "bad news" for the stock market, because they may imply worsening prospects for inflation, that are reflected in declining stock prices.[10]

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[9] Three lags are chosen, according to the Akaike information criterion, a plausible lag structure with annual data.

[10] Bagliano and Beltratti (1997) offer a cointegration analysis of the determinants of stock returns based on Italian quarterly data (1963-1995) and find evidence of a long-run equilibrium negative relation between the inflation rate and a real stock price index. They note that the money stock can be considered weakly exogenous in their estimated system.

The second co-integrating vector indicates that in the long-run per capita real money balances rise more than per capita real income, a steady state money velocity that points to a role of broad money as a store of wealth in equilibrium. Stationary fluctuations around these two "long-run attractors" are shown in Figure 6 and are named EC1 and EC2, respectively.

Results shown at the end of Table 3 ascertain how these estimated error correction mechanisms are at work in this system. Regression (1) posits deflated stock prices changes d(S/P) as a function of lagged EC1 and EC2 and three lags of the changes in all the variables in the system.[11] Previous deviations from the long-run share prices – interest rate relation $EC1_{t-1}$ enter this dynamic equation at conventional significance levels. In regression (2) the set of lagged variables is substituted by two contemporary, pro-cyclical variables ($d(M/P)_t$ and $d(Y)_t$) and by annual interest rate changes $d(RL)_t$.

Regression (3) also adds a predetermined variable, the inflation rate $d(P)_{t-1}$. Interestingly, this equation explains about 40 per cent of the variance of yearly variation in share prices, with satisfactory diagnostics (according to standard tests for serial correlation, functional form, normality and heteroskedasticity of the residuals).

Share price movements are positively correlated with changes in real income and real money balances; they are negatively linked to variations in long-term interest rates and to inflation and an inverse relation with the level of interest rates can also be noted in an error-correction form ($EC1_{t-1}$). There are signs that $EC2_{t-1}$ (a stationary variable) may enter positively the regression for $d(S/P)_t$. In periods when real money balances are relatively abundant with respect to real income, this "wealth effect" may bring rising share prices, other things being equal. However, the statistical evidence is rather weak (EC2 is significant in regression 2 but not in regression 1, and shows up at the 11 per cent level only in regression 3).

[11] A dummy for the Second Word War years has also been introduced in the dynamic equations (not in the long-run analysis); its contribution is scarcely significant, with the exception of regression (6).

Regarding the monetary side of this small system (regressions 4-6), it can be noted that the dynamics of real money balances are influenced by $EC2_{t-1}$, the equilibrium error between money and income, but not by the error-correction mechanism $EC1_{t-1}$, estimated from the asset market, that links long-term interest rates and deflated share prices. Because the focus of this empirical work is on secular money growth only, this result supports the view that, at least to a first approximation, a single equation approach to co-integration centred on the long-run money demand is appropriate.

Regression 6 in Table 3 posits money growth $d(M/P)_t$ as a function of the error-correction term with the level of real income $EC2_{t-1}$, the first differences of output $d(Y)_t$ and deflated share prices $d(S/P)_t$, and it adds the opportunity cost $(RL\text{-}DR)_t$, a stationary variable (see Table 1). These variables are significant and have the sign expected a priori; share prices enter with a positive coefficient, pointing to a wealth effect on money dynamics. However, the diagnostic is not satisfactory; residuals are not serially correlated but they do not pass the other tests. For this reason, in Section 5 a deeper analysis of short-term money dynamics will be offered, along the lines of Muscatelli and Spinelli (2000), and some instrumental variable estimates will be also presented and discussed.

Summing up, co-integration analysis does not seem to capture a role of the stock market as a determinant of long-run money demand; the relationship is time-varying. The lack of a co-integrating relationship between real money balances and deflated stock prices is confirmed when a system approach is followed, while the presence of an equilibrium pattern between interest rates and share prices cannot be ruled out.

4. An application of tests for breaks at an unknown date and some VAR results

In this Section, a different assumption on the long-run behaviour of stock prices and the other series is investigated, considering trend stationarity in the presence of a (one-time) structural break. Under the unit root hypothesis, implicit in the co-integration approach of the previous Section, shocks have permanent effects on the level of the variables; stationarity of the single series can be induced by first differencing or a lower (common) order of integration can be achieved by means of co-integration with other non-stationary variables. Instead, under the breaking trend hypothesis series fluctuate in a stationary manner around points of change, in the mean, trend or both, that are not modelled explicitly. The only shock with permanent effects is the infrequent, exogenous change summarised by a deterministic discontinuity; other endogenous, stochastic shocks have temporary effects. Evidence of an important mean-reverting component in deflated share prices in the last 90 years (Table 1) suggests that inference conditional on points of structural change is worth pursuing.

An empirical example in the seminal paper by Perron (1989) on the role of break points in tests of the unit root hypothesis is concerned with the appropriate univariate representation of US stock prices in the years 1870-1970. He shows (pp. 1382-85 and Table 7) that assuming a drop in the level of the series in concomitance with the great crash of 1929 and a different trend thereafter, the unit root hypothesis for stock prices can be rejected. Later research has developed on this subject, considering tests that do not require prior knowledge about the time of the change.

Zivot and Andrews (1992), Vogelsang and Perron (1998), among others, generalise Perron's first tests to the case in which the breakpoint is estimated rather than fixed.[12]

The analysis of the trend-break hypothesis in this paper follows the test proposed by Perron (1989) for stock prices and the sequential procedure introduced by Zivot and Andrews (1992). It allows for a simultaneous change in the mean and in the following trend (Model C in Perron, 1989). A Dickey-Fuller tests for unit roots is applied to the residuals of the following regression:

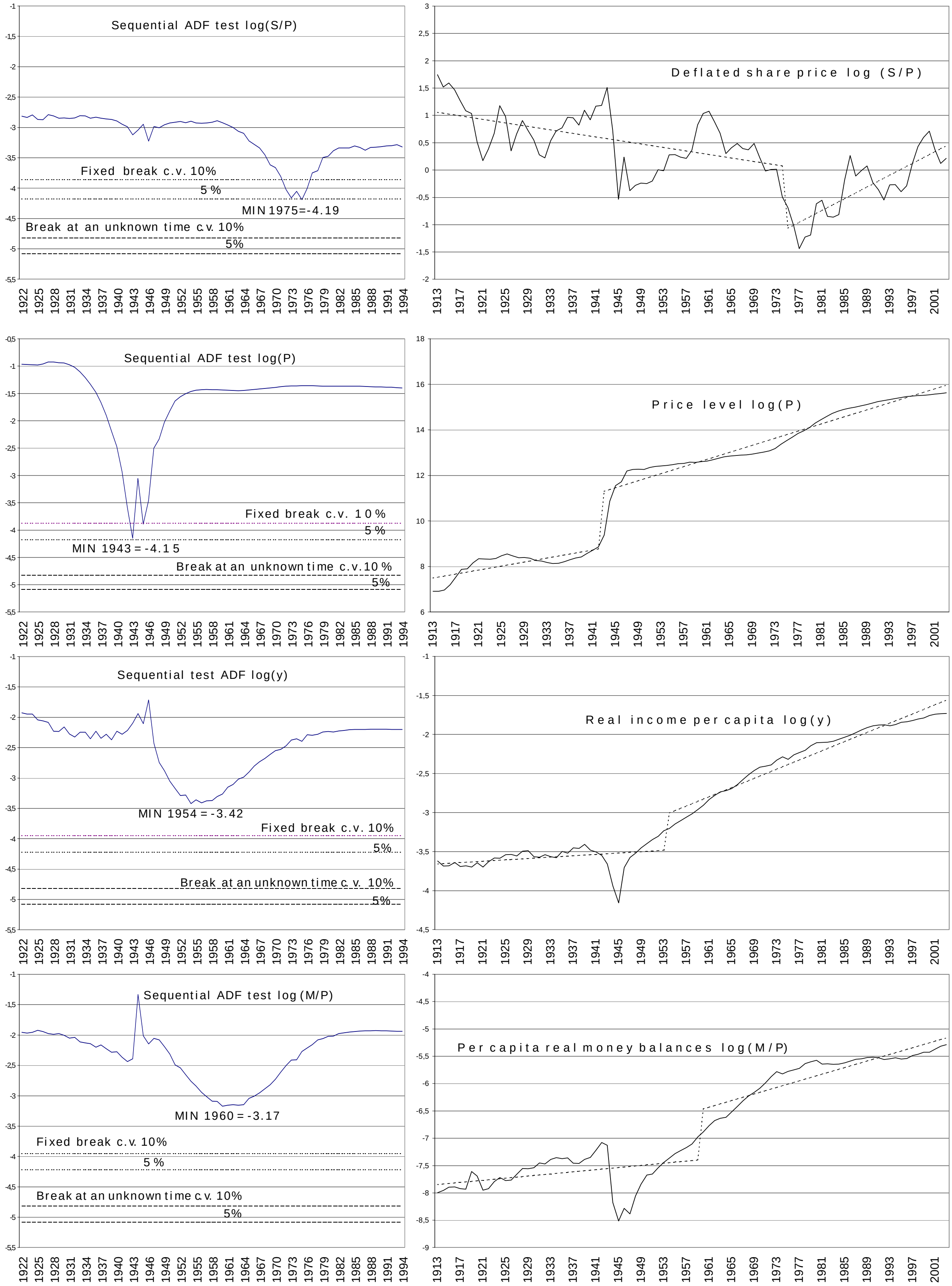
$$(4)\ Y_t = \mu_1 C_t + (\mu_2 - \mu_1) C_{breakpoint} + \mu_3 TIME_t + (\mu_4 - \mu_3) TIME_{breakpoint} + \varepsilon_t$$

Perron (1989) proposes a unit root test conditional on discontinuities chosen a priori, with the great crash (1929) or the first oil price shock (1973) as fixed breakpoints. Asymptotic critical values differ somewhat according to the time of break relative to total sample size; for Model C (Perron 1989, Table 6.B), at the 5 per cent level they are -4.24 for a break positioned at the middle of the sample and -4.18 for a change point positioned at 0.7 of sample size. The null of a random walk is rejected for higher values (in absolute value).

Zivot and Andrews (1992) allow for a shift at an unknown time and apply a sequential procedure. Model (4) is estimated for each possible breakpoint (chosen inside an interval that excludes the first and last portion of the sample for gaining degrees of freedom) and the minimum Augmented Dickey-Fuller statistics is selected. This estimated breakpoint gives the highest possible weight to the trend-stationary

[12] Zivot and Andrews (1992, Table 6, pp. 259-61) apply their test for a break at an unknown date to a century of US stock price data and reject at the 1 per cent level the unit root hypothesis in favour of the trend-break alternative, with an estimated discontinuity in 1936. Chaudhuri and Wu (2003) apply the Zivot-Andrews sequential test for analysing the long-run properties of stock prices in a sample of 17 emerging markets and find that for a majority of countries the random walk can be rejected against a changing trend alternative. An extension of these tests to two or more breakpoints raises difficult inferential problems and is left for further research. Caruso (2004) offers a study of the role of trend breaks in shaping the relationship between cyclical fluctuations and secular output growth in a large sample of countries.

Figure 7 – Test for a break at an unknown date: estimated segmented trends



alternative across all the estimated regressions. As in Perron (1989) the null is a unit root (with no breaks), but because the choice of the breakpoint is the outcome of an estimation procedure, asymptotic critical values differ. Zivot and Andrews (1992 Table 4) indicate that the largest (in absolute value) t-statistics of the Dickey-Fuller test must be -5.08 and -4.82 for rejecting, respectively at the 5 and 10 per cent level, a random walk taking account of the data-dependent previous choice of the break year. Small sample critical values for this sequential test have been computed by Vogelsang and Perron (1998, Table 2, panel A, lag choice based on the Akaike statistics); assuming 100 observations these values are -5.20 and -4.89, at the 5 and 10 per cent level.

Results of the sequential test procedure are reported in Figure 7 and in Table 4, panel (A).[13] The level of deflated stock prices shows a discontinuity after the first oil shock; a slowly declining trend is estimated in the first two-thirds of the sample, followed by a drop in the mid-1970s and a more vigorous upward trend thereafter. In this segmented-trend representation, the recent tendency of rising stock prices is due to exogenous structural changes (for instance, different preferences of investors after the oil price shock, in a climate of gradual financial liberalisation, growing openness of the economy and larger participation in the stock market). The minimum ADF test is reached in 1975 and is equal to -4.19; taking into account that the break point is not chosen independently of the data, the null of a unit root cannot be rejected. However, according to Perron's (1989) fixed-break critical values, allowing for the impact of the first oil shock, the random walk is rejected at the 5 per cent level. An economist with a strong prior (not due to pre-testing) that a large, exogenous inflationary impulse has had a lasting impact on financial markets and that permanent shocks occur much less

[13] Break points are searched in the 1922-94 time interval, excluding the first and last one-tenth of the sample. Augmented Dickey-Fuller tests of the residuals of equation (4) have been computed sequentially, up to three lags. In this sample of annual data, one lag is sufficient for whitening the residuals and the reported results refer to an ADF(1) test; overall, the lag choice has a limited influence on these empirical findings.

Table 4 – Estimated trend breaks and VAR systems, 1913-2003.

PANEL (A): ESTIMATED SEGMENTED TRENDS				
Dependent variable:	Regressors (T=91):			
Deflated stock prices	1.151 C	-.016 T	-1.194 C75	.070 T75
(standard error)	(.122)***	(.003)***	(.193)***	(.010)***
Price level	7.229 C	.044 T	2.459 C43	.034 T43
(standard error)	(.173)***	(.008)***	(.162)***	(.008)***
Real income	-3.685 C	.004 T	.450 C54	.025 T54
(standard error)	(.047)***	(.002)***	(.052)***	(.002)***
Real money balances	-7.906 C	.010 T	.906 C60	.021 T60
(standard error)	(.093)***	(.003)***	(.114)***	(.004)***
Minimum ADF test	Stock prices	Price level	Real income	Money
(break year)	-4.19 (1975)	-4.15 (1943)	-3.42 (1954)	-3.17 (1960)
***, ** and * denote significance at the 1, 5 and 10 per cent level, respectively.				

PANEL (B): VAR(3) - LEVELS OF THE ENDOGENOUS VARIABLES				
	System statistics (T=88, allowing for 3 lags):			
		Log-likelihood	AIC	SBIC
Dummies: DW1 D19 D21 D42		879.87	-17.27	-13.89
D43 D44 D45 D99	Lag exclusion	LAG(1)	LAG(2)	LAG(3)
		970.88 (.000)	229.20 (.000)	35.76 (.075)
	System diagnostic:			
Observations: 88	Serial correlation	AR(1)	AR(2)	AR(3)
(allowing for 3 lags)		33.61 (.117)	23.80 (.531)	26.72 (.370)
	$_{c}R^{2}$	SER	NORMALITY	CAUSALITY
Real money balances	.997	.052	4.90 (.086)	82.22 (.000)
Inflation	.959	.039	3.93 (.140)	36.09 (.000)
Real income	.998	.039	1.99 (.370)	69.91 (.000)
Real interest rate	.959	.039	5.23 (.073)	45.31 (.000)
Deflated stock prices	.888	.216	2.25 (.324)	30.52 (.002)
P-values in parentheses.				

PANEL (C): VAR(3) – DEVIATIONS FROM SEGMENTED TRENDS				
	System statistics (T=88, allowing for 3 lags):			
		Log-likelihood	AIC	SBIC
Dummies: DW1 D19 D21 D42		866.19	-16.62	-12.82
D43 D44 D45 D99	Lag exclusion	LAG(1)	LAG(2)	LAG(3)
D54 D60 D75		927.13 (.000)	286.32 (.000)	69.93 (.000)
	System diagnostic:			
Observations: 88	Serial correlation	AR(1)	AR(2)	AR(3)
(allowing for 3 lags)		32.51 (.144)	26.93 (.359)	19.24 (.785)
	$_{c}R^{2}$	SER	NORMALITY	CAUSALITY
Real money balances	.956	.057	.51 (.773)	28.19 (.005)
Price level	.983	.046	4.06 (.131)	62.07 (.000)
Real income	.935	.031	1.31 (.520)	33.91 (.001)
Real int. rate (not detrended)	.943	.046	3.88 (.144)	81.55 (.000)
Deflated stock prices	.679	.231	.36 (.846)	40.80 (.000)
P-values in parentheses.				

frequently than the random walk implies would find in-sample confirmation for this assumption. Of course, a simple extrapolation of these tendencies could be misleading, because in this representation stock prices fluctuate around secular trend patterns that change with the next large, infrequent and unforeseen permanent shock. There is almost observational equivalence with the co-integration result between interest rates and deflated stock prices shown in Section 3; the stationary pattern of share prices around segmented trend lines interrupted by the effects of the oil shock (Figure 7) looks similar to the stationary vector between long-term interest rates and stock prices (Figure 6). Discriminating further between these hypotheses is beyond the scope of this paper.

A breakpoint during Second World War (in 1943) is apparent for the price level, with a minimum ADF statistics of -4.15. The upward price trend following the War is steeper than the previous one; prices fluctuate around these segmented trends with fairly regular cycles (Figure 7). A unit root cannot be rejected by the sequential procedure. Similarly to stock prices, however, a difference-stationary model (a unit root in the price level) implicit in the commonly used inflation rate representation of observed price movements would instead be rejected, conditional on this break, with Perron (1989) critical values by a researcher who holds a genuine “WWII prior” as a main structural break of the price level without previously examining the data-generating mechanism.

Results for real income and money balances, with minimum estimated ADF statistics of -3.42 and -3.17, respectively, are less favourable to the breaking-trend alternative. Estimated break years are 1954 for output and 1960 for the growth path of broad money.

The estimated trend slopes have gone up for both variables after the breakpoints (Figure 7); the last decade shows low output growth, with a persistent departure from the underlying trend. Real money balances, after slowing in the late 1980s and early 1990s, are returning to the segmented trend line in recent years. However, for these series a unit root cannot be rejected and, by inspecting the data, a higher number of breaks seems more likely than a single change-point. Overall, these results are not implausible; in particular, fluctuations in real money balances around the "most stationary" alternative across all break dates (Figure 7) resemble fairly closely the co-integrating relationship between money and income reported in the previous Section (Figure 6). Both representations may have useful information on the pattern of long-run money velocity over time. The co-integration finding indicates that permanent shocks to money and income (two integrated variables) cancel out, allowing only temporary deviations from a single stochastic trend (money velocity). The breaking trend hypothesis suggests that, assuming that a rare event in the middle-1950s (for instance, a productivity shock) permanently raised the trend growth of output, eventually money growth moved in the same direction in the early 1960s for exogenous reasons (say, changes in preferences or reflecting the infrequent structural output adjustment), stabilising money velocity. Since a unit root in money or income is not rejected conditional on these breaks, the co-integration result is strengthened while the segmented trend alternative is not. For this reason, the findings reported in this Section are a complement, not a substitute, for the previous co-integration analysis.

Most of the empirical VAR literature on this topic concentrates on the dynamic response of stock prices to unanticipated movements in interest rates and monetary policy shocks, rather than on the opposite relation, the sensitivity of monetary aggregates to impulses arising from the stock market. The simple VAR model presented here is able to account for the significant, negative response of stock prices to the interest rate shocks noted by several authors. The full system effects will be evaluated in order to ascertain the magnitude and timing of the response of broad money to unanticipated movements in stock prices.

For purposes of comparison, two VAR systems have been estimated on the same set (5x5) of endogenous variables (real money balances, real income, inflation, deflated stock prices, the long-term real interest rate, computed as the yearly government bond yield minus the inflation rate):[14]

(5) VAR (M/P, Y, dP, S/P, RL)

A first VAR (panel B, Table 4) is estimated on the levels of the series, introducing the appropriate lag structure for whitening the residuals, a standard practice in applied econometrics. A second VAR (panel C, Table 4) has been estimated conditional on the reported breakpoints and segmented trends, fitting the de-trended series. A VAR system has the form of a seemingly unrelated regression model with predetermined variables and equal regressors in each equation, and consistent and asymptotically efficient estimates can be obtained by OLS. Provided that the residuals are not auto-correlated, the computed innovations are labelled "news" or "surprises" in the data, conditional on the information set of endogenous variables considered. The estimation technique traces the dynamic response of each variable in the system to these impulses or "shocks", measuring the impact of the unanticipated fluctuations in the series. The estimated VAR model in levels, with three lags (a reasonable lag structure with annual data) has satisfactory statistical properties (Table 4, panel B); the system residuals are not auto-correlated and pass tests for normality; exceptions are some departures from the Gaussian (at 7-9 per cent significance) detected for the equations estimated for real interest rates and money balances. Granger causality tests reject exogeneity for all the variables in the system at usual

[14] A VAR system that includes the nominal instead of the real interest rate or an opportunity cost (the government bond yield minus the deposit rate) shows a less satisfactory diagnostic and overall fit. Data on inflationary expectations are not available for this sample and the real rate is thus measured ex-post. Some yearly dummies (significant at conventional levels in the VAR) have also been introduced. They take into account the worsening of economic conditions in concomitance with the two world wars and, following Muscatelli and Spinelli (2000, pp. 731-2), the removal of price controls in 1919 and the sharp monetary contraction in 1921. A dummy 1999, the start year of the recent convergence process towards EMU, has been successfully tested. The VAR system conditional on segmented trends (panel C, Table 4) also includes dummies for the estimated break dates (1954, 1960, 1975).

confidence levels. Orthogonalised impulse response functions (obtained by Choleski's decomposition) and 10 per cent confidence bands are plotted in Figure 8.[15]

An unanticipated upward movement in the real interest rate has a persistent, negative impact on real money balances (Figure 8); this effect peaks after two years.[16] A positive interest rate "surprise" also has a depressing effect on the stock market; the largest negative impact of higher interest rates on share prices occurs after one year. Unanticipated increases in broad money balances seem to have an overall positive impact on stock prices, but this effect does not reach conventional significance levels. The dynamic response of M2 to an unexpected, positive shock in deflated stock prices shows both a short-term wealth effect, significant at a one-year horizon, and the presence of a persistent, negative substitution effect in the long-run, significant at forecasting horizons between four and ten years. This suggests that a positive correlation shapes the relationship between stock prices and the monetary aggregate at

[15] These results may depend on the ordering of the variables; for this reason, generalised impulse responses (Pesaran and Shin, 1998) have also been computed (Figures 8-9). Generalised impulses are invariant to the ordering of the data; confidence bands are not available for them, but their estimated impact at different forecasting horizons is similar to the effects of the orthogonalized impulses, corroborating the robustness of the findings. It can also be observed that, in the empirical VAR literature, monetary policy shocks can be identified by a Choleski's decomposition. This method involves selecting a first group of variables that has useful information for policy-makers but does not respond contemporaneously to policy shocks, a policy variable (usually a short-term interest rate that the monetary authorities influence closely) and a second group of variables that respond at horizon zero to policy but to which the policy variable does not respond contemporaneously. This recursive approach is not attempted, because at yearly frequencies both macroeconomic and policy variables can be supposed to be interdependent. Instead, the focus is on the generalized dynamics of the system and on the long-run responses to unanticipated impulses arising from the stock market.

[16] At forecasting horizon zero (contemporary impact) real money balances show a positive response to interest rate shocks; however, this counter-intuitive effect is temporary and the dynamic response of the monetary aggregate from one to eight years is negative and persistent. The contemporary impact may signal the influence of common cyclical patterns or the role of some external factors. For instance, higher interest rates may determine an immediate exchange rate appreciation; a higher relative price of the domestic currency may bring a capital inflow that tends to sustain, for a while, the growth of the domestic monetary aggregate. After one year, M2 gradually reaches a permanently lower steady-state value following an unanticipated interest rate shock; transaction costs, lags in learning, habits or the presence of long-run relationships between banks and their clients may also tend to delay a negative response and an immediate portfolio adjustment, which takes some years to complete.

Figure 8 - Var system in levels: impulse response functions

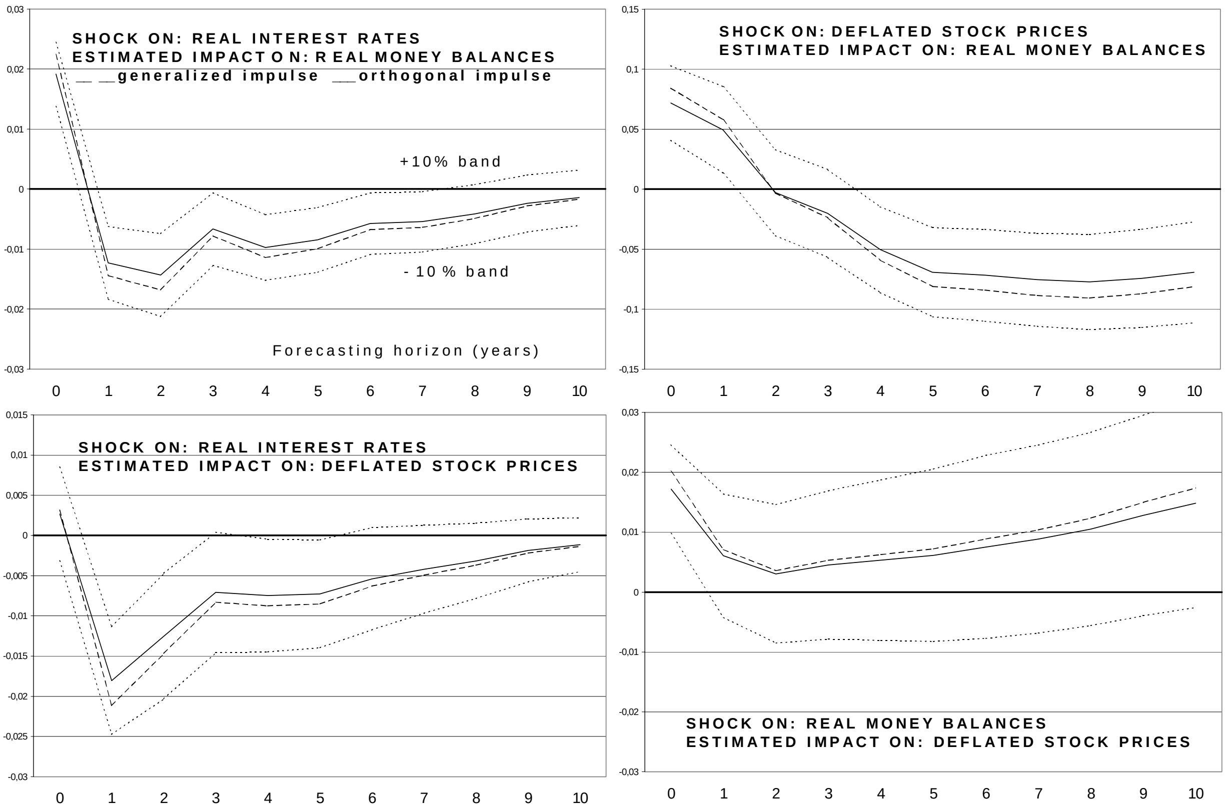
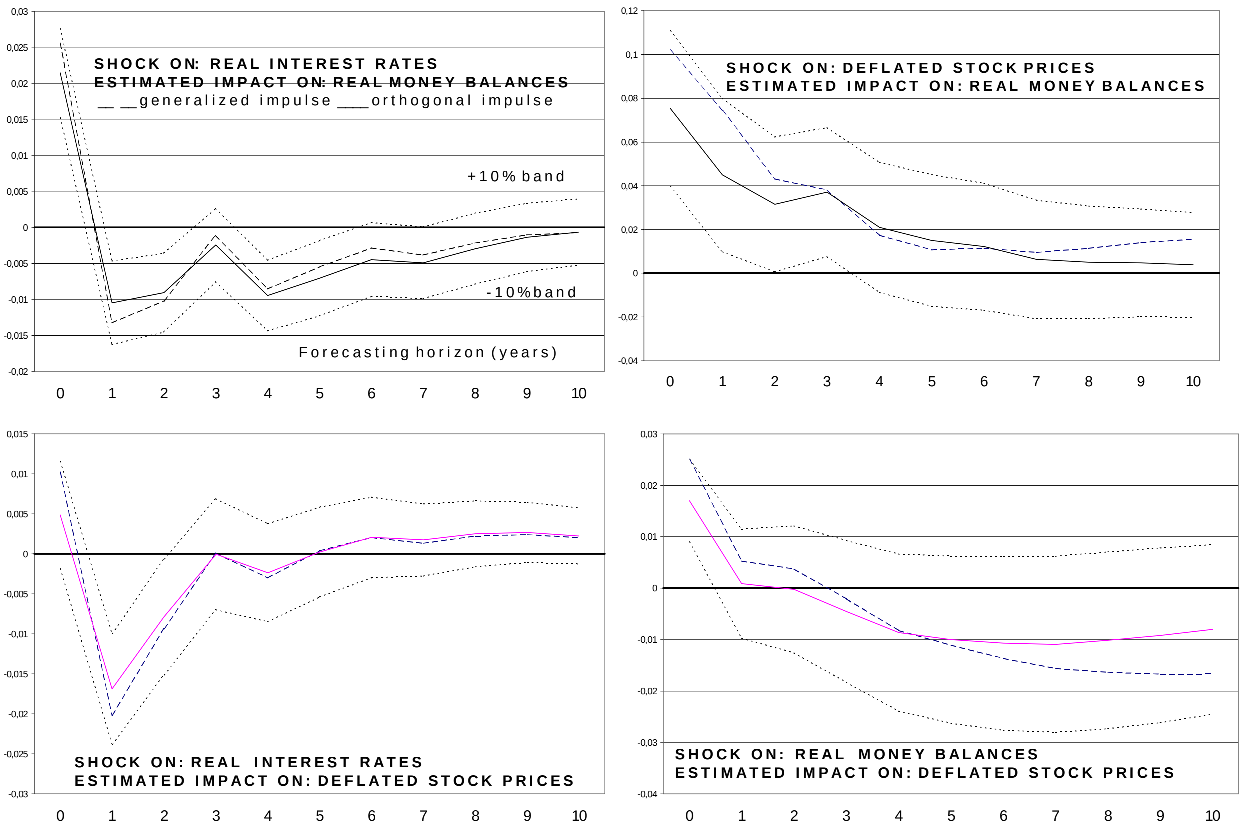


Figure 9 - Var conditional on segmented trends: impulse response functions



short or medium-term horizons, while at lower frequencies money and shares are found empirically to be competitors in investors' portfolios.[17]

VAR results for the de-trended variables (deviations from the estimated segmented trends, with the exception of the real interest rate, not de-trended) are shown in Table 4, Panel (C). The system's diagnostic is favourable; residuals are not auto-correlated and are normally distributed. In this VAR, results are conditional on exogenous structural changes that occur infrequently in the sample period and these events are separated from the sequence of stochastic innovations that drives the dynamics of the system. By comparing the results with the VAR in levels, it is possible to ascertain whether these rare shocks have effects that do not vanish but persist over long horizons.

The estimated impulse response functions and their confidence bands are reported in Figure 9. Real money balances and deflated stock prices respond negatively to interest rate shocks. The monetary response is persistent and the stock market response temporary, with a peak after one year.[18] The impact of unanticipated movements in de-trended share prices on the dynamics of real money balances indicates that the short-run, positive wealth effect is remarkably robust to the removal of infrequent shocks, while the negative, long-run substitution effect tends to disappear. It cannot be excluded that the shifts between money and shares in investors' portfolios at long horizons depend on the impact of exogenous structural changes (rare inflationary episodes, opening of financial markets, changes in investors' preferences) that may drive the secular pattern of the variables.

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[17] The long-run response of stock prices to impulses to their own growth rate is not significantly different from zero.

[18] Neri (2004) finds on monthly data (1985:1-2000:12) that a contractionary monetary policy shock (an unanticipated increase in short-term interest rates) has a significant negative and transitory effect on stock market indexes in the G7 countries and Spain, with varying magnitudes across countries.

5. A dynamic money demand function including information from the stock market

The lack of a stable, long-run stationary equilibrium between stock market fluctuations and money does not rule out that in the period 1913-2003 stock prices may have affected the short-run dynamics of per capita real money balances.

An empirical strategy in this case is to estimate a specification in first differences (an error-correction framework) employing the long-run equilibrium error between money, income, prices and interest rates, which does not consider information from the stock market, in order to introduce the changes in deflated stock prices and ascertaining their significance at different points in the sample. Another possibility, which is left for further research, would require modelling explicitly the possible presence of one or more structural breaks in a co-integrating system including stock prices and money, and the underlying learning process followed by agents.

It can be shown that deflated stock prices enter the short-run money demand function in a time-varying manner, rather than being extraneous to the relationship. The following dynamic regression (an error-correction model) describes the short-term adjustment of money demand towards the long-run equilibrium:

$$(6) \quad d(M/P)_t = c + \gamma_{wars}(\text{dummies})_t + \gamma_{ec}(EC)_{t-1} + \gamma_x(dX)_{t,t-n} + \gamma_{sp}\, d(S/P)_t + \varepsilon_t$$

Annual changes in real money balances per capita are explained by a few step dummies controlling for unusual monetary conditions following the war years,[19] a stationary error correction mechanism that reverses temporary deviations from money market equilibrium, $(EC)_{t-1}$ (derived from a long-run money demand function that does not include shares, (2) in Table 2), a vector of macroeconomic control variables in first differences $(dX)_{t,t-n}$, and stock price changes in real terms $d(S/P)_t$.

[19] Muscatelli and Spinelli (2000, pp. 731-2) introduce dummies for the Second World War, the inflation episode of 1944 and the removal of price controls in 1919 and observe that 1921 represents an outlier due to a sharp monetary contraction. Accordingly, I have introduced dummies for the years 1919, 1921, as well as for 1940-43, 1944 and 1945, in order to take into account the worsening of economic conditions in concomitance with the Second World War.

Results for the short-run money demand function are reported in Table 5. These regressions deal with stationary variables only and show satisfactory explanatory power and diagnostic, on the basis of tests for serial correlation, normality, heteroskedasticity, Chow's stability test and, for the instrumental variables estimates, Sargan's test for the adequacy of instruments.[20]

Regression 1 in Table 5 posits d(M/P) function of the error correction term from the long-run co-integrating relationship (Table 2, equation 2), which shows as expected a negative sign, an opportunity cost (interest rate changes) and adds information from the stock market. The estimated coefficient of share price changes in real terms is positive and significant at the 1 per cent level, pointing to a wealth effect.

In order to make sure that this result cannot be ascribed to outliers or to a single dominant sub-period, inference has been based on rolling regressions; results are presented in Figure 10. Specification (1) in Table 5 is based on five regressors only (the constant, the error correction mechanism, interest and share price changes) and a short window size can be selected (20 years); it cannot be expected that coefficients are precisely estimated with this method, but the procedure is useful for ascertaining empirically how results may be sensitive to different sub-samples.

Results shown in Figure 10 are consistent with the existence of a wealth effect on money growth arising from the stock market over most of the sample period. Occasional drops (the Second World War years) or periods of larger-than-average correlation can be noticed (during the 1950s and 1960s), but overall a positive elasticity prevails, with values that are broadly comparable across sub-samples, over a time span of about 65-70 years, until the late 1970s and early 1980s. Instead, in the last two decades a negative correlation between real money and share prices shows up, signalling a substitution effect across assets.

[20] There are signs of non-normality in the residuals of regression 4, but they do not extend to the other specifications. Regression 1 shows auto-correlation (at the 9 per cent level), and regressions 5 and 9 have heteroskedastic disturbances. Substantial stability is found across the relationships, with the exception of regressions 4 and 10.

Regressions (2) and (3) perform a sample split in 1980-81; encouraging diagnostic and an improvement in specification (1) is obtained.[21] Controlling for the adjustment process to a long-run standard money demand equation, the stock market variable influences money growth; a wealth effect is estimated at the 1 per cent level in the years 1914-1980, while a negative impact of stock prices on money is found, at the 10 per cent level of confidence, in the recent period (1981-2003).

The role of stock market capitalisation and turnover velocity in the money demand function is also evaluated (regressions (4) and (5) in Table 5 and rolling coefficients in Figure 10). Turnover is introduced as deviations from its upward trend because this variable is trend-stationary (see the univariate analysis in Table 1). The contribution of the annual changes in capitalisation in real terms is similar to that of deflated stock prices, and the substitution effect is precisely estimated in the recent period.

In the last twenty years, the turnover of the stock market and money growth are positively correlated. A higher trading volume may require larger amounts of money for transactions; moreover, to the extent that turnover velocity proxies for stock market volatility and uncertainty (because investors may trade more in order to rebalance their portfolio risks and/or for speculating on their available information under uncertain circumstances), broad money goes up mainly for precautionary purposes.

Estimated regressions (4) and (5) in Table 5 also evaluate whether a modelled variance of stock price changes enters the money demand function. Applications in finance use widely the ARCH and GARCH specifications for modelling the conditional variances of asset prices. While the variance of stock prices can be predicted to some extent, it changes over time and thus in the stock market some

[21] A Chow test applied to the year 1981 rejects stability of regression (1) at the 1 per cent level, while an analogous test at the middle point of the sample (1958) finds mild signs of instability (p-value=.109) but does not formally reject; this gives support to the existence of a break point in the early 1980s, as signalled by the analysis based on rolling regressions, which focuses more directly on the coefficient of interest.

periods are more volatile (and riskier) than others. The ARCH model summarises the persistence of the conditional variance (its auto-correlation) and traces the volatility clustering of stock price movements. Clusters of higher-than-average volatility may be associated with larger money holdings for precautionary motives, which are likely to diminish in less turbulent times.

Briefly, it can be shown that the residuals in the regression for stock price changes presented in Table 3 (equation 3) follow a first-order ARCH model. The parameters that explain the conditional variance of the error term are (0.046 Constant + 0.312 e^2_{t-1}), with asymptotic t-statistics of 4.574 and 1.851, respectively. The estimated conditional variance of the annual growth of deflated stock prices is plotted in Figure 10. Particularly volatile periods can be observed after the Second World War, from the mid-1970s until the 1980s, and in the most recent years. However, in the money demand specification (Table 5) the modelled variance does not reach conventional significance levels (contemporary or lagged) and the contribution of stock prices, capitalisation and turnover is robust to its introduction.[22]

Specifications (6)-(7) in Table 5 are chosen following a general-to-specific approach; they are similar to the preferred dynamic model estimated by Muscatelli and Spinelli (2000, p. 731) on a longer time period. These regressions introduce an information set of macroeconomic variables in first differences and may represent a more demanding benchmark for evaluating the contribution of stock prices. Regressions (6)-(7) reproduce in these shorter sub-samples the Muscatelli and Spinelli (2000, equation 6) preferred model in growth rates, which includes changes in real income, long-term interest rates and a lagged dependent variable. In these regressions the non-significant variables are dropped and the first difference in the own rate on money (proxied by the deposit rate) is introduced. They explain a considerable

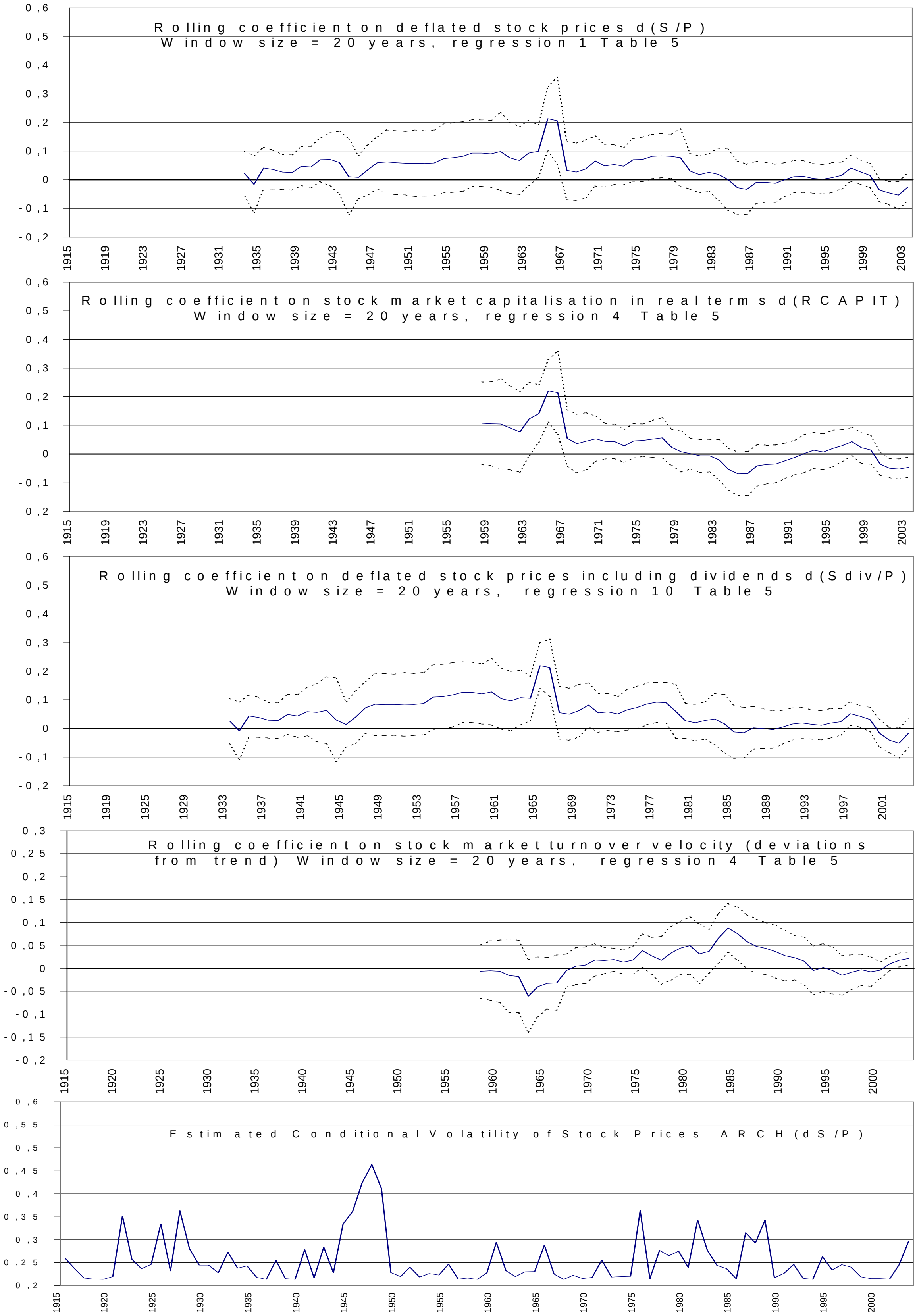
[22] Similar results are obtained by estimating the parameters of the conditional heteroskedastic model on a simple first-order auto-regression of stock price changes; in this case, the conditional variance of the error term is: (0.059 Constant + 0.305 e^2_{t-1}), with asymptotic t-statistics of 5.217 and 2.214, respectively. Also this volatility proxy does not have an influence on money demand. A more complete analysis of temporal dependencies in share price movements is left for further research; it would require other information sets, asymmetries in volatility and higher-frequency data to be considered.

Table 5 – Explaining the dynamics of per capita real money balances

Estimation period	1914-2003	1914-1980	1981-2003	1939-1980	1981-2003	1914-1980	1981-2003	1916-1980	1981-2002	1914-2003	1916-1980	1981-2002
Dependent variable: $d(M/P)_t$												
Estimation method	OLS	OLS	OLS	OLS	OLS	OLS	OLS	IV	IV	OLS	IV	IV
Explanatory v.:	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
EC_{t-1} (Tav. 2 eq. 2)	-.220	-.224	-.177	-.244	-.191	-.071	-.139	-.092	-.185	-.219	-.090	-.190
standard error	(.047)	(.046)	(.122)	(.097)	(.094)	(.030)	(.058)	(.042)	(.082)	(.045)	(.043)	(.093)
p-value	(.000)	(.000)	(.164)	(.017)	(.060)	(.022)	(.026)	(.034)	(.040)	(.000)	(.040)	(.059)
Money												
$d(M/P)_{t-1}$						.393		.373			.371	
standard error						(.059)		(.066)			(.067)	
p-value						(.000)		(.000)			(.000)	
Real Income												
$d(Y)_{t+1}$								.031	-.217		.022	-.287
standard error								(.130)	(.496)		(.131)	(.537)
p-value								(.814)	(.668)		(.865)	(.600)
$d(Y)_t$						-.043						
standard error						(.071)						
p-value						(.550)						
Inflation												
$d\pi_t$						-.603		-.577			-.582	
standard error						(.058)		(.073)			(.075)	
p-value						(.000)		(.000)			(.000)	
Interest rate												
dBR_t	-.014	-.038	-.008	-.026	-.009	-.019	-.009	-.020	-.009	-.011	-.019	-.009
standard error	(.009)	(.016)	(.005)	(.019)	(.004)	(.010)	(.003)	(.014)	(.003)	(.009)	(.015)	(.003)
p-value	(.133)	(.024)	(.097)	(.191)	(.022)	(.053)	(.004)	(.157)	(.011)	(.199)	(.214)	(.011)
Deposit rate												
dDR_t	.026	.035	.0003	.018	.002	.013		.021		.024	.019	
standard error	(.012)	(.016)	(.012)	(.0200	(.009)	(.009)		(.012)		(.011)	(.012)	
p-value	(.027)	(.032)	(.981)	(.368)	(.801)	(.157)		(.090)		(.033)	(.135)	
Stock prices												
$d(S/P)_t$	.087	.110	-.041			.054	-.050	.051	-.054			
standard error	(.025)	(.028)	(.023)			(.018)	(.019)	(.024)	(.021)			
p-value	(.001)	(.000)	(.094)			(.004)	(.015)	(.038)	(.022)			
Stock p. with div.												
$d(Sdiv/P)_t$										.103	.042	-.046
standard error										(.024)	(.024)	(.024)
p-value										(.000)	(.091)	(.071)
Capitalisation												
$dRCAPIT_t$				.107	-.058							
standard error				(.039)	(.018)							
p-value				(.010)	(.005)							
Turnover velocity												
$(TURN\text{-}T)_t$				.001	.022		.019		.020			.019
standard error				(.017)	(.008)		(.007)		(.011)			(.011)
p-value				(.960)	(.012)		(.021)		(.076)			(.108)
Conditional volat.												
$ARCH(dS/P)_{t-1}$				.038	.117			-.033	.144		-.025	.206
standard error				(.195)	(.110)			(.062)	(.136)		(.063)	(.146)
p-value				(.848)	(.300)			(.597)	(.306)		(.699)	(.180)
cR^2	.844	.891	.468	.911	.666	.964	.608	.946	.607	.854	.962	.527
SER	.058	.055	.021	.059	.017	.032	.018	.033	.019	.056	.033	.021
Serial correlation	(.090)	(.784)	(.113)	(.253)	(.271)	(.990)	(.642)	(.466)	(.413)	(.030)	(.475)	(.655)
Normality	(.503)	(.719)	(.538)	(.001)	(.508)	(.106)	(.518)	(.245)	(.463)	(.438)	(.215)	(.583)
Heterosched.	(.528)	(.509)	(.844)	(.638)	(.001)	(.436)	(.451)	(.424)	(.016)	(.512)	(.403)	(.065)
Chow stability t.	(.109)	(.344)	(.409)	(.027)	(.175)	(.301)	(.140)	-	-	(.086)	-	-
Sargan's IV test	-	-	-	-	-	-	-	(.181)	(.349)	-	(.159)	(.458)
T	90	67	23	42	23	67	23	65	22	90	65	22

Note: Regressions include the constant and dummies for the years 1919, 1921, 1940-43, 1944, 1945 (significant at usual confidence levels). Chow's tests are computed splitting each sub-sample in two parts of equal size. In order to save on degrees of freedom, for regressions 5, 7, 9, 12 only the significant variables are considered in the sample split.

Fig. 10 – Estimated rolling coefficients of short-run money demand and the conditional volatility of stock prices



fraction of the annual variation in per capita real money balances, show satisfactory diagnostic and control for the dynamics of the opportunity cost (dBR_t, dDR_t) and the variation in the inflation rate $d\pi$ (accelerating or decelerating prices, introduced as in Muscatelli and Spinelli, 2000 p. 731 or Caruso, 2001), which enters with a significant negative coefficient in the years 1914-80, as can be expected (agents try to escape the inflation tax by lowering their money holdings), and is not significant thereafter.

Deflated stock price changes have influenced the dynamics of real money balances in this sample. Regardless of the exact specification, a wealth effect in the earlier years and a substitution effect in the recent period show up quite clearly in these estimates. Regressions (8) and (9) in Table 5 also employ an instrumental variable (IV) technique; it is likely that causality may run from money to the stock market, especially on annual data, and that stock price movements are not weakly exogenous but must be considered an endogenous variable.[23] Share price changes may enter the money demand function as an expected variable (an opportunity cost) and this justifies the use of an instrumental variable technique (Chow, 1983, Chapter 11). The set of instrumental variables employed in Table 5, regressions (8) and (9), consists of a constant, time, time squared, the yearly dummies (1919, 1921, 1940-43, 1944, 1945), population and some predetermined (with a one year lag) macroeconomic variables, which include the lagged first difference and levels of money, real income, prices, interest rates, stock prices and the ratios of consumption and exports over national income. The overall satisfactory performance of the IV

[23] Most papers consider the reverse causation (from the money market to the stock market). Caruso (1996) offers empirical results on the effects of money velocity and interest rate shocks on share prices in Italy and six main stock exchanges. Lastrapes (1998) evaluates the response of real equity prices to money supply innovations in eight industrial countries, comprising Italy. Neri (2004) reports the effects of exogenous monetary policy shocks on stock prices in the G-7 countries and Spain.

estimates is supported by the diagnostic and by a Sargan's test, which does not reject the validity of the instruments.[24]

The IV regressions confirms at usual confidence levels a role for the own rate of M2, bond yield and share price movements in explaining the dynamics of per capita real money balances. Moreover, in the recent years, the turnover velocity of the stock market and M2 growth are positively correlated. Analogous results are obtained introducing in the specification the annual variations in the gross stock price index d(Sdiv/P); they are shown in Table 5 (regressions 10-12) and Figure 10.

6. Conclusions

Money demand is inherently forward-looking and its changes reflect portfolio shifts between assets with different degrees of liquidity; both reasons suggest that the accumulation of monetary balances should respond to developments in stock markets. This paper notes that in the last 90 years fluctuations in stock prices have influenced money growth in the Italian economy. However, the levels of deflated share prices seem to be extraneous to the long-run, equilibrium relationship linking together real money balances and the traditional explanatory variables of the money demand function (real income, interest rates, inflation). The proposed explanation is empirical; the failure to find a common steady state including money and stock prices (although there are signs that an overall positive association prevails between real money balances and a stock price index that includes dividends) is due to the fact that in the sample period the correlation with the broad monetary aggregate is basically time-varying. Over the years, the inflation process has represented an implicit cost on

[24] The instrumental variable C/Y is defined as nominal consumption over nominal income; it is culled from Rey (1991, pp. 215-16) until 1950 and updated with IFS data (lines 96f.c/99a). Missing data for the War (1942-45) have been supplemented by the average C/Y for the years 1941 and 1946 (79.9 per cent). The variable E/Y (nominal exports of goods over national income) is reported by Rey (1991 pp. 215-16) and is computed on IFS data from 1951 (lines 70/99a). In the years 1942-47 the average E/Y ratio for 1941 and 1948 (4.2 per cent) has been used instead (instrumental variables cannot contain missing data). Instruments are predetermined, with one lag, C/Y_{t-1}, E/Y_{t-1}, $d(C/Y)_{t-1}$, $d(E/Y)_{t-1}$. In order to save on degrees of freedom, regressions 9 and 12 are estimated with a sub-set of instruments only. Regression (9) in Table 5 estimated by instrumental variables also introduces the lead of real income growth dY_{t+1} in order to ascertain whether share price changes proxy for expected output growth; the overall results are unaffected.

money holdings, and the gradual spread of financial innovations (together with the growing openness of the economy) has allowed agents to choose between a wider range of assets. Both effects have favoured a portfolio reallocation and a change in the relationship between money and stock prices. A positive association in earlier years reflects the store of value function of money and the level of stock prices mainly represents a broad proxy of financial wealth; later, the negative sign of the stock market variable signals the prevalence of an opportunity cost. Referring to monthly and quarterly estimates of money demand in Italy, Angelini, Hendry and Rinaldi (1994) note that in the years 1975-79 the appropriate scale variable is a measure of financial wealth (following the portfolio motive of holding money), while in the period 1983-91 domestic demand, a variable that proxies the volume of transaction, yields better results. This is likely to reflect a more intense transaction role of desired money balances, following the gradual evolution of the financial environment.

In a dynamic specification of money demand the contribution of stock prices is not nil but changes over time. A wealth effect emerges in the period 1914-1980, but it reverses to a substitution effect in the last twenty years. This finding matches previous empirical results based on different samples. Referring to annual data on money velocity and the level of deflated stock prices, Friedman (1988, Table 4) shows that in the US economy the relationship is also time-varying and that a substitution effect marked the years 1886-1939 and 1951-1973, while a wealth effect is evident in the time period 1974-1985. Caruso (2001) finds on quarterly data (1960-98) evidence of a substitution effect between stock price changes and money in Italy, and to a lesser extent also in France and Germany, while wealth effects seem to be prevalent across countries. These results are also not in contrast with the literature which gives ample evidence of instability in the connection between stock returns and macroeconomic factors (Panetta, 2002), and with a recent econometric analysis carried out at the European Central Bank (Bruggeman, Donati and Warne, 2003), which notes how over the years 1980-2001 Euro area stock prices did not matter for the long-run M3 demand but are useful when studying the short-run dynamics of money balances.

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