An Empirical Investigation of the Lucas Hypothesis: the Yield Curve and Non Linearity in the Money-Output Relationship

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Abstract

Existing evidence about the effectiveness of money growth to stimulate economic activity has been criticized from different perspectives. In addition, high correlation between money and output is not helpful to detect the direction of causality. From a policy perspective, in fact, positive correlation may arise from two opposite policy conducts: either the monetary authority sets the supply of money to influence future output fluctuations, or the central bank controls money growth as a reaction to the recent evolution of macro variables. In this work the relationship between money and output is analysed within a non linear framework that ascribes a primary role to expectations. In particular, we find evidence that the Lucas (1973) hypothesis, that exists an inverse correlation between the variance of nominal shocks and the magnitude of output response to nominal shocks, is supported by data evidence when the yield curve is either flat or downward sloping. We also provide evidence suggesting that the Friedman (1977) hypothesis, that the variability of inflation exerts a negative effect on the natural level of output, holds when a positive risk premium is incorporated in an upward sloping term structure of interest rates.

JEL classification: C01, C22, E32, E44, G12.

Keywords: *Term Structure, Kalman Filtering, Expectations, Output Growth.*

1. Introduction

Investigating the relationship between money and output has always been a major concern of macroeconomists. The classical dichotomy about whether money influences the future level of output or, viceversa, whether output fluctuations influence money supply, is still an unresolved puzzle. Economists affiliated to the monetarist school believe that money growth will be merely reflected in the future price level, at least this is assumed to be true over long horizons. However, the monetarist view that money does not affect real output may be a weak argument in the short run. On the other hand, Keynesian economists think that short run policies may well influence the level of economic activity.

In this contribution we would like to investigate whether there is a direct relationship between money, or a measure of it, and output; alternatively, we examine whether the effect of money on the business cycle is contingent to other factors, such as, for instance, the shape of the yield curve.

We analyse the relationship between money and output within a non linear empirical framework that allows for rational expectations. In particular, we investigate whether the Lucas (1973) hypothesis, that there is a negative relationship between the variance of nominal shocks and the magnitude of output response to nominal shocks, is supported by data evidence. The contrarian argument, that the conditional variance of money forecast errors negatively affects the natural level of output, has been proposed by Friedman (1977).

Working with US post-war data, we find evidence that conditioning the examination of the money-output relation to the shape of the yield curve gives the opportunity of reconciling the aforementioned opposite views[1]. Evidence seems to suggest that the variability of inflation, captured by the conditional variance of money forecast errors, exerts a negative influence on output when the yield curve is upward sloping. There is also some evidence that the Friedman hypothesis holds when the linear model is estimated over the entire sample (from 1967 to 2007).

However, in the regime characterized by a flat or downward sloping yield curve the Lucas hypothesis seems to prevail. Interpreting the variance of nominal shocks as agents' perception of monetary policy uncertainty, there seems to be an inverse correlation between aggregate uncertainty and output response. Aggregate risk displays greater effect the more agents internalize uncertainty. In that agents' actions are driven by a large amount of precaution resulting in mild output changes.

We thus provide evidence supporting the Lucas hypothesis which is usually rejected by data in linear analysis (Kim and Nelson, 1989).

[1] Estrella and Mishkin (1997), as well as Estrella and Hardouvelis (1991) and Wright (2006), show, in fact, that a negative spread increases the likelihood of a recession in the near future.

The weaknesses of traditional models (King and Plosser, 1984; Ravn and Sola, 2004) coupled with the difficulty of detecting a unique direction for causality call for an approach that emphasizes the role of expectations. Hence, the contribution of this study is also methodological. We propose to examine expectations exploiting a two-level structure: *micro* and *macro*.

At a *micro* level (*bottom* level) agents' expectations focus on the central bank operational procedure regarding the supply of money. The *micro* mechanism of processing available information is based on the Kalman filter which implies a continuous refinement of expectations on the basis of past prediction errors, i.e. deviations between *ex ante* expected and *ex post* observed values of the money stock. In such a setting agents form expectations according to a Bayesian iterative sequence that combines the re-elaboration of past prediction errors with the analysis of new flows of information. Moreover, what is peculiar in the *micro* analysis of expectations is that Kalman filtering allows separating the expected from the unexpected component of money growth. The adoption of a time-varying approach for the policy rule is also consistent with recent evidence. Cogley and Sargent (2006), as well as Boivin (2006), document important time variation in the response of the monetary authority to the state of the economy. Also Sims and Zha (2006) point out that the changing view of the Fed about the economy has been gradual; they argue it could be attributed to the changes of shocks' variance. Finally, Primiceri (2008) provides evidence that the reaction of monetary policy to the changes in both inflation and unemployment has become more and more aggressive in the last decades. Last, but certainly not least, we focus on a policy rule expressed in terms of money supply since our sample is characterized by periods of inflation instability; Bernanke and Mishkin (1993) argue, in fact, that central banks tend to adopt targets in terms of money growth when the inflation rate threatens to be out of control.

At a *macro* level (*top* level) expectations focus on the future economic outlook, as reflected by the dynamics of the term structure of interest rates. The *macro* perspective captures the sentiment regarding the future evolution of key macro variables as well as institutional or socio-political factors, or, eventually, technological changes. There is in fact large evidence that the information content of term structure could be used to make inference about the future state of the economy (Estrella and Hardouvelis, 1991; Estrella and Mishkin, 1997). Several studies point out that she slope of the yield curve help forecasting output fluctuations; more recent evidence focuses on the predictive role of term premia implied by the term structure (Hamilton and Kim, 2002; Favero, Kaminska, and Soderstrom, 2005; Ang, Piazzesi, and Wei, 2006; Rudebusch, Sack, and Swanson, 2008; Modena, 2008 a). Finally, Modena (2008 b, 2011) draws attention to the fact that also curvature, and not only extreme points of the yield curve, may be related to output movements.

Furthermore, our methodology partially accounts for the criticism moved by Amato and Swanson (2000). Despite some evidence suggests monetary aggregates help predict future output (Stock and Watson, 1989; Becketti and Morris, 1992; Feldstein and Stock, 1994), Amato and Swanson point out that such evidence might be contingent upon the nature of the dataset. Using real time, rather than revised, data they document a substantial reduction of the marginal predictive power of money. The threshold approach adopted in this work implies non linearity in the dataset thus breaking time continuity; for this reason specifically, it allows reducing the impact of the historical track.

The rest of the paper is organized as follows. The next Section contains a brief survey of the literature and discusses motivations. In Section 3 we present some evidence about causality. In Section 4 we outline the structure of expectations at the *micro* level. Empirical evidence is discussed in Section 5. Finally, Section 6 concludes. Data are presented in *Appendix I*. In *Appendix II* we report auxiliary estimations.

2. Motivations and Literature Review

Whether and to what extent money growth is capable of contributing to the determination of real output is still an unresolved puzzle in economics. The monetarist view that money growth induces a proportional change in the price level leaving real output unaffected is acknowledged to work in the medium-long run. However, as Keynesian theory suggests, monetary disturbances are believed to have some real effects in the short run. Although the contribution of monetary shocks on permanent income growth is limited or absent, managing money supply is a useful instrument under control of the central bank for stabilizing or stimulating the economy across the business cycle.

The classical dichotomy of money neutrality, i.e. nominal variables are unable to affect real variables, has been initially investigated by means of the following equation, which is known as the Saint Louis equation since it has been introduced by economists of that Federal Reserve District:

$$\Delta IP_{t,t-3} = \alpha + \beta_0 \Delta M_{t,t-3} + \beta_1 \Delta M_{t-3,t-6} + \beta_2 \Delta M_{t-6,t-9} + \beta_3 \Delta M_{t-9,t-12} + \beta_4 trend + \varepsilon_t \quad (1)$$

The LHS variable is the quarterly change in the industrial production index; while, the quarterly changes of the money stock over the last year are explanatory variables. The above regression also includes a constant and a time trend (to account for eventual trend in money and output growth). Different monetary aggregates have been considered: M1, M2, and the U.S. Fed Board of Governors monetary base (MB). The analysis is performed on monthly data from 1967 to 2007. Empirical results are reported in Table 1.

Saint Louis Equation - IP grw (3)							
	α	β_0	β_1	β_2	β_3	β_4	R^2
M1	1.2986	-0.1155	0.1551	0.0822	0.0686	0.0013	0.027
t-stat	(0.1318)	(0.0963)	(0.0312)	(0.2534)	(0.3221)	(0.5490)	
NW	(0.5261)	(0.2251)	(0.1359)	(0.4039)	(0.4407)	(0.7734)	
W	(0.2081)	(0.1465)	(0.0504)	(0.2466)	(0.2651)	(0.5728)	
M2	-0.0020	-0.1238	0.1447	0.1191	0.0593	0.0005	0.094
t-stat	(0.0710)	(0.0016)	(0.0007)	(0.0051)	(0.1292)	(0.0262)	
NW	(0.4226)	(0.0069)	(0.0027)	(0.0361)	(0.2146)	(0.2781)	
W	(0.1166)	(0.0007)	(0.0001)	(0.0062)	(0.1138)	(0.0364)	
MB	-0.0093	0.4487	0.5870	0.3514	0.3628	0.0001	0.019
t-stat	(0.4713)	(0.2062)	(0.1004)	(0.3263)	(0.3101)	(0.3757)	
NW	(0.7033)	(0.3348)	(0.1775)	(0.3638)	(0.4721)	(0.6656)	
W	(0.4594)	(0.1458)	(0.0428)	(0.1478)	(0.2115)	(0.3926)	

p-values in parenthesis. **NW**: Newey-West correction. **W**: White correction.

IP grw (3): 3-month growth of the seasonally adjusted industrial production index (dependent variable). Regressors appear in the equation as quarterly rate of growth: **M1**: monetary aggregate M1; **M2**: monetary aggregate M2; **MB**: monetary base.

Table 1

The contemporaneous effect on output exerted by M2 turns out to be negative ($\beta_0 < 0$); however, more generally, results suggest that the rate of growth of M2 over the last three quarters have a positive influence on IP growth ($\beta_1, \beta_2 > 0$). The rate of growth of M1 has a marginal, though significant, effect on the current growth of the IP index.

In order to check whether there is a significant influence of money on output we have looked at the jointly significance of estimated coefficients in each equation. In all the equations, coefficients turn out to be jointly significant supporting the influence of money on output. In addition, we have also run a Wald test to check the following restrictions: $\sum_{i=0}^{3} \beta_i = 0$, i.e. to check whether money growth does explain output. The null hypothesis is rejected in two cases since the probability value associated to the test is *zero* for both M2 and MB equations, thus suggesting the influence of these monetary aggregates on real activity. M1 growth, instead, does not seem informative about output.

As a forecasting exercise we have estimated the above regression using the future determination of output in the LHS:

$$\Delta IP_{t+3,t} = \alpha + \beta_0 \Delta M_{t,t-3} + \beta_1 \Delta M_{t-3,t-6} + \beta_2 \Delta M_{t-6,t-9} + \beta_3 \Delta M_{t-9,t-12} + \beta_4 trend + \varepsilon_t \qquad (2)$$

Results reported in Table 2 suggest that M2, rather than M1 and the monetary base, is effective in influencing both the current and the future level of industrial production. In the equation for M2 both coefficients β_0 and β_1 are statistically significant indicating that future output is influenced by money growth up to six months before. The goodness of fit is definitely poor; however, the M2 equation returns a much better fit than the other equations. The dynamics of M1 over the most recent quarter has a marginal impact on the IP growth.

St. Louis Equation - IP grw (+3)

	α	β_0	β_1	β_2	β_3	β_4	R^2
M1	0.0228	0.0304	0.0190	0.0215	-0.0181	0.0005	0.025
t-stat	(0.2877)	(0.0796)	(0.2883)	(0.2321)	(0.2941)	(0.3033)	
NW	(0.6378)	(0.2752)	(0.4457)	(0.2973)	(0.4842)	(0.6076)	
W	(0.3499)	(0.0812)	(0.2828)	(0.1945)	(0.2622)	(0.3162)	
M2	-0.0074	0.2269	0.2916	0.0989	-0.0191	0.0001	0.075
t-stat	(0.0101)	(0.0223)	(0.0068)	(0.3563)	(0.8461)	(0.0081)	
NW	(0.2359)	(0.1031)	(0.0492)	(0.4050)	(0.8891)	(0.1906)	
W	(0.0197)	(0.0126)	(0.0101)	(0.3488)	(0.8458)	(0.0112)	
MB	-0.0011	0.1584	0.1056	0.1041	0.0086	0.0004	0.017
t-stat	(0.7192)	(0.0771)	(0.2396)	(0.2474)	(0.9238)	(0.4298)	
NW	(0.8391)	(0.1954)	(0.2347)	(0.3739)	(0.9369)	(0.6890)	
W	(0.6976)	(0.0422)	(0.0674)	(0.1437)	(0.9073)	(0.4349)	

p-values in parenthesis. **NW**: Newey-West correction. **W**: White correction.

IP grw (+3): quarterly growth of the seasonally adjusted industrial production index over the next 3-month period (dependent variable). Regressors appear in the equation as quarterly rate of growth: **M1**: monetary aggregate M1; **M2**: monetary aggregate M2; **MB**: monetary base.

Table 2

The macroeconomic debate has further focused on the asymmetric effect of monetary policy on output. Models with sticky prices or financial constraints suggest that interest rate changes generate greater effect on real activity during recessions. Similarly to Romer and Romer (1994), Garcia and Schaller (1999) find evidence in line with this conjecture arguing that monetary policy is more effective during recessions. Ravn and Sola (2004) find evidence corroborating the hybrid traditional Keynesian asymmetry, that is only small negative monetary policy shocks tend to influence real output. In order to account for this effect we estimate the above equations including dummy variables to distinguish the effect of positive rather than negative money growth rates.

$$\Delta IP_{t,t-3} = \alpha + \beta_0 D^{(+)} \Delta M_{t,t-3} + \beta_1 D^{(+)} \Delta M_{t-3,t-6} + \beta_2 D^{(+)} \Delta M_{t-6,t-9} + \beta_3 D^{(+)} \Delta M_{t-9,t-12} + \varepsilon_t \quad (3)$$

$$\Delta IP_{t,t-3} = \alpha + \beta_0 D^{(-)} \Delta M_{t,t-3} + \beta_1 D^{(-)} \Delta M_{t-3,t-6} + \beta_2 D^{(-)} \Delta M_{t-6,t-9} + \beta_3 D^{(-)} \Delta M_{t-9,t-12} + \varepsilon_t \quad (4)$$

Where $D^{(+)}$ indicates a positive quarterly growth rate of the monetary aggregate, while $D^{(-)}$ indicates a negative growth rate. Results strongly support the hypothesis advanced by Romer and Romer (1994). Coefficients are statistically significant only in equation (3), as reported in the top panel of

Table 3, suggesting that only stimulus to economic activity seem to be effective. Moreover, the goodness of fit (0.033) of equation (3) is much larger than that (0.006) of equation (4) -bottom panel of Table 3-. The *F*-test suggests coefficients are jointly significant only in equation (3). The Wald test confirms that the null hypothesis $\sum_{i=0}^{3}\beta_i = 0$ cannot be rejected for equation (3) solely.

Dummies D(+) - IP grw (3)						
	α	β_0	β_1	β_2	β_3	R^2
M1	1.5221	-0.0924	0.1143	0.1369	0.1591	0.033
t-stat	(0.0009)	(0.1307)	(0.0538)	(0.0212)	(0.0094)	
Dummies D(-) - IP grw (3)						
	α	β_0	β_1	β_2	β_3	R^2
M1	2.2632	0.1056	0.0615	0.0301	-0.0415	0.006
t-stat	(0.0000)	(0.1877)	(0.4278)	(0.6977)	(0.6020)	

p-values in parenthesis.

IP grw (3): 3-month growth of the seasonally adjusted industrial production index (dependent variable). Regressors appear in the equation as quarterly rate of growth: **M1**: monetary aggregate M1; **D(+)**: dummy variables that capture positive changes of the quarterly growth of the monetary aggregate; **D(-)**: dummy variables that capture negative changes of the quarterly growth of the monetary aggregate.

Table 3

Results are similar when replacing the dependent variable $\Delta IP_{t,t-3}$ with its future realization $\Delta IP_{t+3,t}$:

Dummies D(+) - IP grw (+3)						
	α	β_0	β_1	β_2	β_3	R^2
M1	0.3606	0.0352	0.0335	0.0323	-0.0157	0.032
t-stat	(0.0019)	(0.0220)	(0.0247)	(0.0301)	(0.3046)	
Dummies D(-) - IP grw (+3)						
	α	β_0	β_1	β_2	β_3	R^2
M1	0.5663	0.0910	0.0747	-0.0459	0.0242	0.005
t-stat	(0.0000)	(0.6507)	(0.7008)	(0.8131)	(0.2246)	

p-values in parenthesis.

IP grw (+3): quarterly growth of the seasonally adjusted industrial production index over the next 3-month period (dependent variable). Regressors appear in the equation as quarterly rate of growth: **M1**: monetary aggregate M1; **D(+)**: dummy variables that capture positive changes of the quarterly growth of the monetary aggregate; **D(-)**: dummy variables that capture negative changes of the quarterly growth of the monetary aggregate.

Table 4

Although not reported, and coherently with the estimations above, the joint estimation of an equation including all dummy variables, both $D^{(+)}$ and $D^{(-)}$, returns significant coefficients only for dummies $D^{(+)}$ denoting an increase of the monetary aggregate.

The main drawback of the above equations is that they are not sufficient to establish any causality relation running from money to output. King and Plosser (1984) observe that monetary aggregates such as M1 and M2 are determined by the interaction between the high-powered money, a liability of the central bank, the behaviour of both firms and households, and the efficiency of the financial system through the strategies of the banking sector. Therefore it is possible to observe changes in the money stock that anticipate output movements without causing them.

Endogeneity is the second problem associated with both equations (1) and (2) and equations (3) and (4). The high correlation eventually captured by the coefficients of the equations may well derive from the conduct of the monetary authority that sets the future supply of money in response to past output fluctuations. The chronological sequence of a tight monetary policy which follows growing GDP, like a reduction of the rate of money growth whose final goal is to curb economic activity, and of an accommodative policy to tackle falling GDP preserves high correlation between money and output but with important implications for reverse causation. In addition, from a policy point of view, it is impossible to ascribe to monetary policy the effect of money on output without simultaneously considering the effect on GDP exerted by fiscal policies. The poor good of fit obtained for the above regressions is, in fact, a sign of misspecification; in particular, some relevant variables may be omitted.

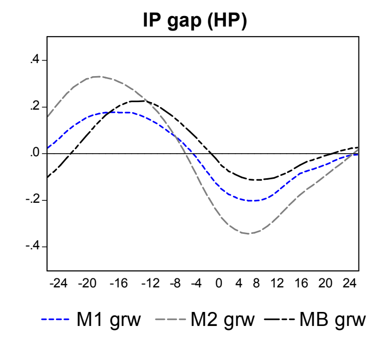
Finally, the time series analysis performed by estimating the above equations might be affected by shifts in money demand since financial innovations contributes to changing agents' preferences. In particular, as Ravn and Sola (2004) argue, the instability of M1 demand may underlie the poor fit of the M1 equation; furthermore, and specifically in this analysis, the monthly frequency of data may, in principle, accentuate the effect of M1 volatility.

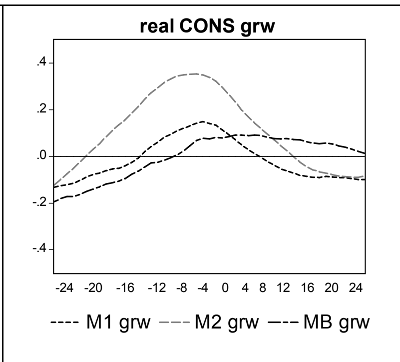
The aforementioned intrinsic difficulties of detecting the effect of money on output coupled with the devastating effect of the Lucas critique call for an empirical method based on dynamic expectations as that implied by Kalman filtering. Expectations are subject to continuous refinement as long as new information becomes available; in addition, agents revisit their expectations on the base of past prediction errors. So far, in fact, we have not discriminated between anticipated and unanticipated money growth which is a core distinction in economics. In this vein, prediction errors work like a proxy for unanticipated money supply. Our approach will be deeply motivated later.

Before presenting in details the methodology adopted in this chapter, next Section provides some more evidence regarding the money-output relation.

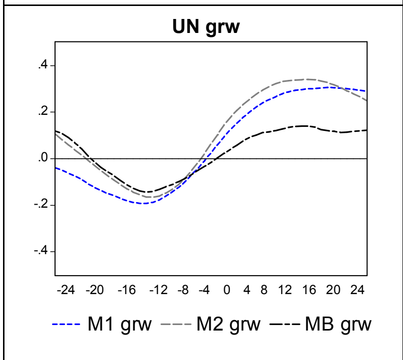
3. Preliminary Evidence on Causality

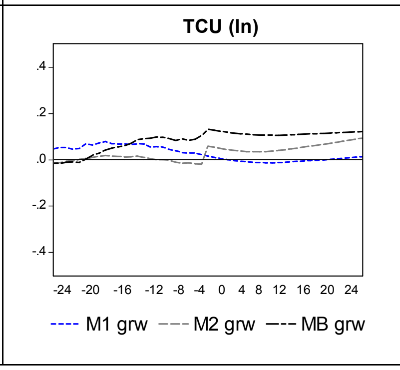
In this Section we focus on the causality issue characterizing the empirical relationship between real variables and monetary aggregates. Using monthly data from January 1967 to December 2007, we start by looking at dynamic short-run correlations. Each panel of Figure 1 shows correlations between a measure of real activity and different monetary aggregates (M1, M2, and MB, the monetary base). The top-left diagram indicates that all monetary aggregates are positively correlated with the Hodrick-Prescott detrended series of industrial production at lags, but negatively correlated at leads. Hence, booms (high IP relative to trend) tend to be preceded by high values of money growth; while positive values of IP relative to its HP trend tend to be followed by low values of money growth. This evidence is in line with the idea that money supply acts as a stimulus to real economic activity; while in response to fast-growing economy, and to the associated threat of mounting inflation, the monetary authority inverts the sign of the monetary policy conduct. The bottom-left diagram shows the correlations between monetary aggregates and the annual change in the unemployment rate. Consistently with the above story, all monetary aggregates are negatively correlated at lags with the growth in unemployment, i.e. a reduction in unemployment tends to be preceded by high values of money supply; on the other hand, monetary aggregates are positively correlated with the increase of unemployment at leads.





IP gap (HP): industrial production gap obtained with the Hodrick-Prescot filter. **real CONS grw**: annual growth of the seasonally adjusted real personal consumption expenditures. **M1 grw**: annual growth of monetary aggregate M1. **M2 grw**: annual growth of monetary aggregate M2. **MB grw**: annual growth of the monetary base.





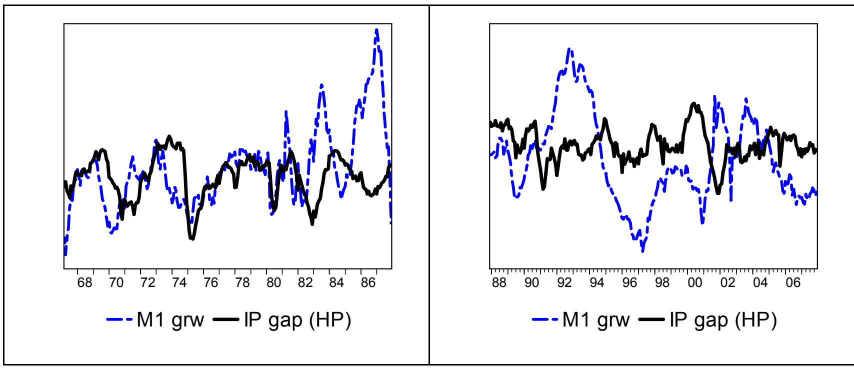
UN grw: annual growth of the unemployment rate. TCU (ln): *log* total capacity utiliation. **M1 grw**: annual growth of monetary aggregate M1. **M2 grw**: annual growth of monetary aggregate M2. **MB grw**: annual growth of the monetary base.

Figure 1

The top-right diagram of Figure 1 shows the pattern of short run correlations between the change in real consumption expenditures and monetary aggregates. Private consumption seems to be positively related to money growth in the short run at both leads and lags. The smoother pattern of the monetary base correlations is consistent with the theory put forward by King and Plosser (1984); they find that inside money, i.e. the internal monetary measures which represent the liabilities of the banking sector as a component of monetary aggregates, rather than outside money, i.e. the external monetary measures which represent the liabilities of the central bank, are positively correlated with real activity.

Finally, although lower with respect to other real indicators, the bottom-right panel shows that the correlation between monetary aggregates and the (*log*) total capacity utilization is positive at both lags and leads; thus both past money supply and expectations of future important money supply tend to positively affect the employment of the factors of production.

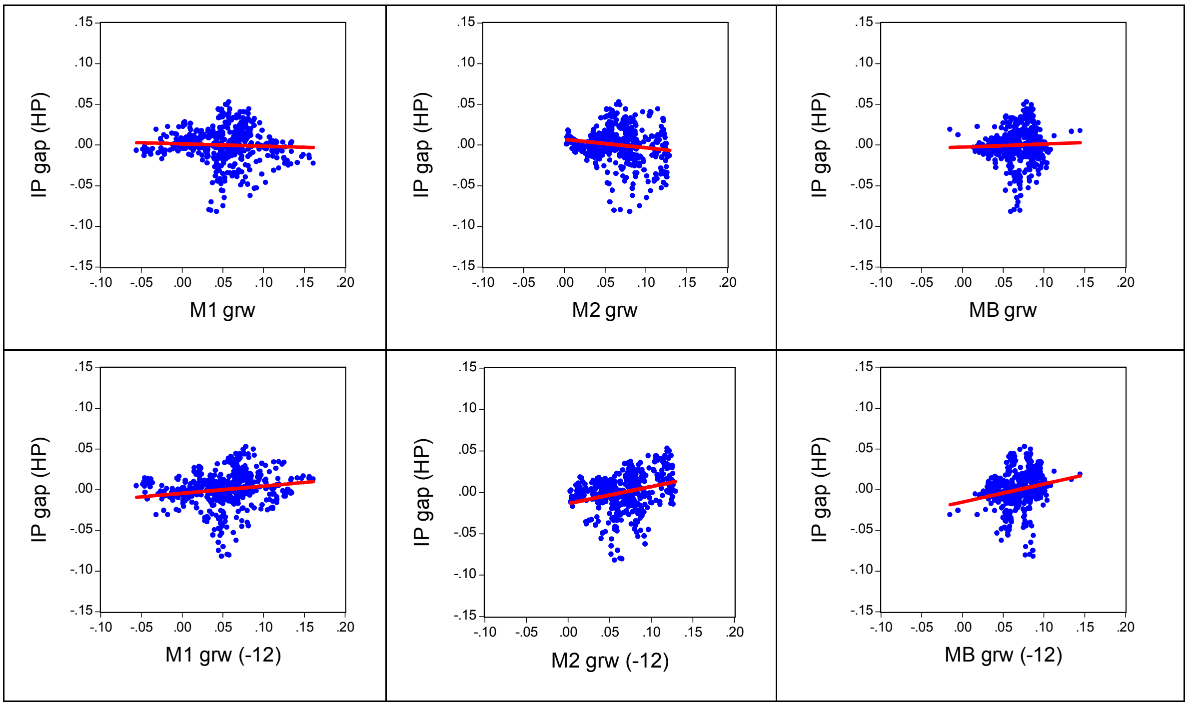
Previous evidence is provided by Friedman and Schwartz (1963) in a classical contribution about the monetary history of the United States; they find that money growth rate changes lead changes in real GDP. The left diagram of Figure 2 shows that the rate of growth of M1 systematically anticipates business cycle movements between 1967 and the mid 1980s. Falling money growth precedes slowdowns in economic activity; while increasing money stocks anticipate both recoveries and booms. However, more recent evidence presented in the right diagram is more controversial: starting from 1985, in fact, the relationship between money and output is not as close as before, both the length and the magnitude of cyclical fluctuations do not reproduce the preceding dynamics of the monetary aggregate. The different pattern of the relationship between M1 and the IP gap might be due to financial innovations which affected the demand for money. The greater variability of the rate of growth of M1 might also reflect greater difficulty of the money stock to influence output from 1986 and 1997. In addition, Choudhry (2002) finds that the stated monetary act of 1980 considerably affected the income and interest rate demand elasticities of both M1, M2 and their components in U.S. Moreover, he argues that the fall in the M1 interest rate elasticity may well indicate M1 as possibly a more effective monetary policy tool after 1980.



M1 grw: annual growth of monetary aggregate M1. **IP gap (HP)**: industrial production gap obtained with the Hodrick-Prescot filter.

Figure 2

In line with the above evidence, the following scatter diagrams suggest a stronger effect displayed on both current and future output (detrended IP) by M2 rather than by M1 or MB[2]. The top diagrams show the scatter between the IP gap and the contemporaneous growth rate of the monetary aggregates. In the bottom panels we report the scatter plots between the actual HP-detrended IP and the rate of money growth 12 months before.



Scatter plot diagrams. **IP gap (HP)**: industrial production gap obtained with the Hodrick-Prescot filter. **M1 grw**: annual growth of monetary aggregate M1. **M2 grw**: annual growth of monetary aggregate M2. **MB grw**: annual growth of the monetary base. **(-12)** means the twelfth lag (monthly frequency of data).

Figure 3

As Walsh (2003) points out "*while suggestive, evidence based on time patterns and simple correlations may not indicate the true casual role of money. Since the Federal Reserve and the banking sector respond to economic developments, movements in the monetary aggregates are not exogenous, and the correlation patterns need not reflect any casual effect of monetary policy on economic activity*". Notwithstanding the above mentioned stronger influence on real variables exerted by M2 than other aggregates, the Granger causality tests suggest M1 being the only source capable of affecting the future level of both industrial production and total capacity utilization. The null hypothesis that the rate of growth of M1 does not cause the IP gap cannot be rejected, as well as the null that M1 does not cause (*log*) TCU. The real personal consumption expenditure seems to be caused in the Granger sense by all monetary aggregates. Results are reported in Table 5.

[2] The regressing line associated to M2 is, in fact, always steeper. The only exception occurs in the mid-bottom panel, the regressing line associated to MB turns out to be marginally steeper than the one associated to M2. However, in the former case (MB) there is a greater vertical dispersion of observations around the regressing line; in the latter case (M2), instead, observations more closely concentrated around the regressing line along its entire length.

Granger Causality Test					
	lags	**IP gap (HP)**	**TCU (*log*)**	**unemp**	**r-cons**
M1	**3**	(0.0432)	(0.0296)	(0.8090)	(0.3827)
	6	(0.0577)	(0.0158)	(0.8598)	(0.0063)
	12	(0.7340)	(0.6834)	(0.1455)	(0.0066)
M2	**3**	(0.1971)	(0.1568)	(0.6825)	(0.0013)
	6	(0.2398)	(0.4621)	(0.2978)	(0.0005)
	12	(0.4150)	(0.2986)	(0.1740)	(0.0050)
MB	**3**	(0.5173)	(0.5698)	(0.7667)	(0.0321)
	6	(0.7928)	(0.9175)	(0.7430)	(0.1002)
	12	(0.2935)	(0.2558)	(0.3408)	(0.0436)

Null Hypothesis: the monetary aggregate does not Granger-cause the real variable. Tests *p-values* in parenthesis.

IP gap (HP): industrial production gap obtained with the Hodrick-Prescott filter; **TCU (*log*)**: *log*-series of the total capacity utilization; **unemp**: unemployment rate; **r-cons**: real personal consumption expenditures. **M1**: monetary aggregate M1; **M2**: monetary aggregate M2; **MB**: monetary base.

Table 5

The Granger tests are also employed to investigate whether lagged levels of the real variables help to predict the future path of monetary aggregates. Results are significantly supportive in this respect, as shown in Table 6. Such statistical evidence about causality is compatible with a Taylor-type monetary policy reaction function implying the monetary authority to raise the policy rate when the pace of economic growth is as fast as to create undesired inflationary pressures.

Granger Causality Test				
	lags	**M1**	**M2**	**MB**
IP gap (HP)	**3**	(0.0002)	(0.0003)	(0.0765)
	6	(0.0231)	(0.0003)	(0.0239)
	12	(0.0914)	(0.0004)	(0.1145)
unemp	**3**	(0.0100)	(0.0002)	(0.1510)
	6	(0.0408)	(0.0006)	(0.3795)
	12	(0.0606)	(0.0136)	(0.4834)
TCU	**3**	(0.0055)	(0.0170)	(0.4816)
	6	(0.0321)	(0.0053)	(0.0407)
	12	(0.1117)	(0.0136)	(0.1669)
r-cons	**3**	(0.0046)	(0.0005)	(0.0778)
	6	(0.1192)	(0.0008)	(0.1646)
	12	(0.3280)	(0.0340)	(0.6395)

Null Hypothesis: the real variable does not Granger-cause the monetary aggregate. Tests *p-values* in parenthesis.

M1: monetary aggregate M1; **M2**: monetary aggregate M2; **MB**: monetary base. **IP gap (HP)**: industrial production gap obtained with the Hodrick-Prescott filter; **TCU (*log*)**: *log*-series of the total capacity utilization; **unemp**: unemployment rate; **r-cons**: real personal consumption expenditures.

Table 6

We sum up the preliminary evidence discussed in this Section by saying that results regarding the effect of money on output turn out to be somehow ambiguous. Short run correlations tend to suggest a positive influence of lagged money on actual output; however, the Granger tests partially contradict this evidence by suggesting a causality relationship working in the opposite direction. In addition, despite the existing evidence suggesting that money helps to predict future output (Stock and Watson, 1989; Becketti and Morris, 1992; Feldstein and Stock, 1994), Amato and Swanson (2000) argue that results are someway misleading because they crucially depends on revised, rather than real time, monetary aggregates data.

The aforementioned ambiguity can be dealt with by introducing a new element in the analysis; we thus attribute a role of primary importance to agents' expectations and, in particular, to the associated expectations errors. To conclude, we recall that the choice of M1 as the benchmark reference aggregate for the monetary policy rule in the following analysis hinges on the results of the Granger tests reported in Table 5.

4. Empirical Methods for Expectations

In this Section we summarize the approach employed to derive agents' expectations about the future stance of monetary policy as captured by the rate of growth of M1.

King and Plosser (1984), in fact, suggest inside money, a component of M1 representing the liabilities of the banking sector, being highly correlated with business cycle movements. In addition, Bernanke and Mishkin (1993) argue that the monetary authority tends to define targets in terms of money growth if there is a concrete likelihood that inflation gets out of control. Our sample is characterized by periods of high and volatile inflation. Finally, we justify the time-varying approach by observing that there is substantial evidence highlighting that both the monetary policy conduct and the variance of nominal shocks have changed over time (Cogley and Sargent, 2006; Boivin, 2006; Sims and Zha, 2006; Primiceri 2008).

We thus compute expectations by applying the Kalman filter, since it provides with an effective formalisation of the mechanism through which agents form expectations rationally. Moreover, the Kalman approach gives the opportunity of deriving a measure of innovations which overcomes the criticism traditionally moved to the VAR approach. In what follows we briefly outline the main features of Kalman filtering.

The observation equation, or measurement equation, of the state-space system is:

$$\Delta M_t = a + x_{t-1}\beta_t + u_t \tag{5}$$

Actual quarterly money growth is a function of the changes of the T-bill rate, of the price level, and of the money stock over the previous quarter; u_t is a stochastic $i.i.d.(0,\sigma_u)$ noise. The specification of the money equation come from Mishkin (1982) and Weintraub (1980); it has been successively considered by Kim and Nelson (1989). The only difference is that we rule out the fiscal variable, because of the superior independence achieved by the monetary authority in recent times. The state, or transition, equation captures the evolution of coefficients over time:

$$\beta_t = \mu + F\,\beta_{t-1} + v_t \tag{6}$$

Where v_t is an idiosyncratic disturbance $i.i.d.(0,\sigma_v)$. Following standard practice, we impose matrix *F* to be the identity matrix since we assume that the regressing coefficients follow random walk processes (Kim and Nelson, 1998; Kim and Nelson, 2006; Boivin, 2006). The Kalman filter is an iterative algorithm based on updating the informative set with most recent available information and predicting future movements of the variable under examination. The coefficients covariance matrix conditional on information available up to time *t-1* is:

$$P_{t|t-1} = E\left[\left(\beta_t - \beta_{t|t-1}\right)\left(\beta_t - \beta_{t|t-1}\right)^I\right] \tag{7}$$

Equation (8) provides the prediction of money growth based on information available up to time *t* given that economists know the econometric relationship linking the core variable to the explanatory variables till time *t-1*.

$$\Delta M_{t|t-1} = x_t \beta_{t|t-1} \tag{8}$$

Once the actual contemporaneous value of the core variable is observed, agents can compute the prediction error according to the following

$$\eta_{t|t-1} = \Delta M_t - x_t \beta_{t|t-1} = \Delta M_t - \Delta M_{t!t-1} \tag{9}$$

Finally, equation (10) represents the conditional variance of money growth prediction errors:

$$h_{t|t-1} = E\left[\eta^2_{t|t-1}\right] = x_t P_{t|t-1} x_t^I + \sigma^2_\varepsilon \tag{10}$$

According to (10) Kalman filtering allows two sources of uncertainty generating the conditional variance of the forecast error ($h_{t|t-1}$). One source depends on the evolutionary behaviour of

estimated coefficients through the coefficients covariance matrix, thus capturing the gradual change of the policy regime over time; the other source is a random noise related to future disturbances, like unpredictable institutional or technological shocks. The assumption of a constant variance of nominal shocks to money growth seems too severe since aggregate M1 is regarded to respond on a great variety of shocks. First, M1 trivially depends upon the monetary policy conduct through the money supply (high-powered money). Second, M1 is affected by the interaction between money supply and money demand, so that a demand shock, rather than a supply shock, may influence aggregate M1. For instance financial innovations as well as deregulation may affect M1 in the medium-short run. Finally, M1 depends also on the strategic decisions of the banking system and on the credit market conditions. Therefore, a measure of variance which is conditional to the state of the economy provides with a more realistic picture of aggregate risk.
An alternative method to compute the time-varying conditional variance is to estimate an autoregressive model for money growth (either AR or VAR, depending on the nature of the analysis), and then to compute the squared of fitted residuals[3] (Piazzesi, 2003).

5. Empirical Results

In this Section we provide evidence that the Lucas hypothesis, i.e. that exists a negative relationship between the variance of nominal shocks and the magnitude of output response to nominal shocks, usually rejected in linear model, holds when the likelihood of a slowdown in economic activity is warned by financial indicators. Evidence also suggests that the alternative Friedman hypothesis, that the augmenting variability of inflation exerts a negative effect on output, tends to hold when financial indicators anticipate a thriving pace of economic growth.
A crucial issue involved in testing the Lucas hypothesis is the examination of the conditional variance of nominal shocks over time. Hence, the analysis starts with the estimation of time-varying monetary policy function expressed in terms of money rather than in terms of the rate of interest. The inverse relationship tying money supply and interest rates is trivially respected in any modern economy; moreover, the extensive sample period covered in this analysis calls for a generic version of the monetary policy rule.
After estimating a time-varying specification of equations (5) and (6), we obtain a series of prediction errors ($\eta_{t|t-1}$) and a measure of forecast errors' conditional volatility ($h_{t|t-1}$). The basic equation to test the Lucas *versus* the Friedman hypothesis is the following:

[3] In *Appendix II* we provide with a comparison between the conditional variances of money growth prediction errors obtained both by Kalman filtering and by autoregressive modeling.

$$gap_t = \alpha_0 + \alpha_1 \eta_{t|t-1} + \alpha_2 h_{t|t-1} + \alpha_3 gap_{t-1} + v_t \qquad (11)$$

Where the output gap is the HP de-trended series of *log* IP. In addition, coefficient α_1 is set equal to $\alpha_1 = \gamma_0 + \gamma_1 h_{t|t-1}$. The functional form of α_1 is motivated with the aim of reducing the effect of multi-collinearity in the OLS regression. Equation (11) thus becomes

$$gap_t = \alpha_0 + \gamma_0 \eta_{t|t-1} + \gamma_1 \left(\eta_{t|t-1} \cdot h_{t|t-1}\right) + \alpha_2 h_{t|t-1} + \alpha_3 gap_{t-1} + v_t \qquad (11')$$

The inclusion of the first lag of the dependent variable, which is highly correlated with its actual level, certifies the robustness of other coefficients estimates. In addition, different computational methods for the variance-covariance matrix have been employed in order to obtain consistent estimates (White, 1980; Hansen and Hodrick, 1980; Newey and West, 1987; and, finally, the simplified Hansen and Hodrick).

The theory advanced by Lucas is satisfied when both $\gamma_0 > 0$ and $\gamma_1 < 0$; while, testing the Friedman hypothesis is equivalent to detecting whether coefficient α_2 is negative ($\alpha_2 < 0$)[4]. The assumption here is that the conditional variance of money forecast errors acts as a proxy for the variability of inflation. The original idea put forward by Friedman (1977), in fact, is that the variability of inflation, rather than that of money, reduces the natural level of output since it disturbs the allocation efficiency of the price system.

The linear model (11'), estimated over the entire sample 1967-2007, does not reveal any particular information about the way nominal shocks affect business cycle fluctuations. Table 7 shows estimation results for different measures of the business conditions. The dependent variable in the top panel is the Hodrick-Prescott measure of the IP gap; (from the top to the bottom) in the second panel the dependent variable is the *log* total capacity utilization; in the third panel the dependent variable is the rate of unemployment; finally, in the bottom panel, the dependent variable is the rate of change of unemployment[5].

There is weak evidence supporting the Friedman hypothesis. Although coefficient α_2 is inversely related to the dynamics of real variables, it is either marginally or not significant with the only exception holding for unemployment. The Lucas hypothesis is definitely rejected. Coefficient γ_0 is

[4] Trivially it holds the opposite sign of coefficients when the dependent variable is either the unemployment or its rate of change.

[5] Coefficient α_3 in equation (11') multiplies the first lag of the respective dependent variable.

not statistically significant. In two cases coefficient γ_1 turns out to be significantly positive thus contradicting the Lucas hypothesis[6].

joint estimation - IP gap (HP)						
	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	-0.1198	0.0436	-0.0158	-0.0182	0.9444	0.906
t-stat	[-1.423]	[1.238]	[-1.143]	[-1.544]	[64.75]	
W	[-1.346]	[1.485]	[-1.325]	[-1.410]	[54.71]	obs 468
HH (12)	[-1.501]	[1.884]	[-1.552]	[-1.702]	[95.31]	
NW (12)	[-1.342]	[1.634]	[-1.449]	[-1.531]	[55.74]	
s-HH	[-0.656]	[1.474]	[-1.181]	[-0.812]	[22.92]	

joint estimation - TCU						
	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	-0.9839	0.0163	0.0283	-0.0003	0.9780	0.974
t-stat	[2.893]	[0.958]	[2.181]	[-0.846]	[127]	
W	[-2.816]	[0.971]	[2.107]	[-1.123]	[124]	obs 468
HH (12)	[-1.894]	[0.880]	[2.409]	[-1.835]	[83.27]	
NW (12)	[-1.839]	[0.873]	[2.291]	[-1.360]	[80.98]	
s-HH	[-0.918]	[1.267]	[1.174]	[-0.848]	[40.26]	

joint estimation - Unemployment						
	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	-0.2986	-0.1251	0.0012	**0.0077**	0.9870	0.985
t-stat	[-0.805]	[-0.301]	[0.522]	[2.458]	[173]	
W	[-0.771]	[-0.298]	[0.786]	[2.895]	[149]	obs 468
HH (12)	[-0.472]	[-0.462]	[0.973]	[2.991]	[83.63]	
NW (12)	[-0.501]	[-0.339]	[0.881]	[3.005]	[88.19]	
s-HH	[-0.257]	[-0.398]	[0.555]	[1.323]	[53.01]	

joint estimation - Unemployment grw						
	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	0.1502	-0.0852	0.0861	0.0102	0.9634	0.933
t-stat	[0.027]	[-0.867]	[1.710]	[0.134]	[76.10]	
W	[0.027]	[-0.812]	[2.189]	[0.137]	[59.75]	obs 468
HH (12)	[0.031]	[-1.700]	[5.122]	[0.208]	[89.14]	
NW (12)	[0.027]	[-0.968]	[2.819]	[0.158]	[49.62]	
s-HH	[0.013]	[-1.139]	[1.719]	[0.075]	[25.62]	

t-statistics in square brackets.

OLS: *ordinary least squares* estimation; **W**: White correction; **HH**: Hansen Hodrick correction; **NW**: Newey-West correction; **s-HH**: simplified Hansen-Hodrick correction (see Cochrane and Piazzesi, AER 2005).

Table 7

The linear estimation of equation (11') is not entirely reliable though, since the pattern of residuals series is affected by heteroscedaticity in all cases. In addition, a strong ARCH effect is found after

[6] In *Appendix II* we report the estimation of equation (11') using an alternative measure of the conditional variance of money growth forecast errors leading to similar results. In particular, after estimating an unrestricted VAR (9) model of money growth, inflation, and the change in the 3-month T-bill rate, we obtain the conditional variance of money forecast errors as the squares of residuals from the money growth equation.

performing the Engle (1982) test. The recursive residuals and the CUSUM square of residuals reveal the instability of coefficient estimates as reported in Figure 4[7]. Finally, also the Hansen tests (1992, 2000) highlight the presence of non-linearity.

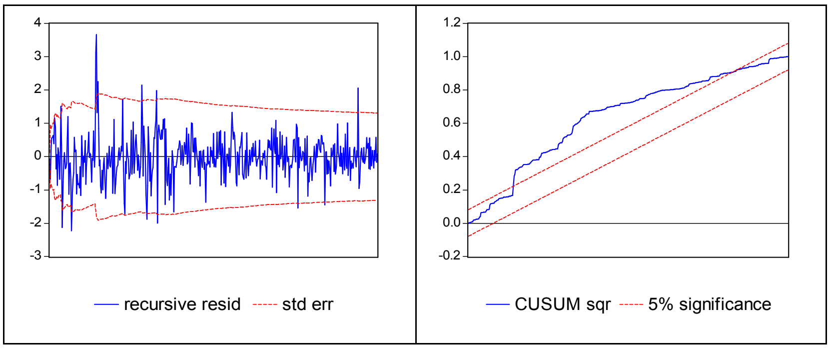


Figure 4

Therefore, we consider a non linear version of model (11') allowing for two different regimes determined by slope of the term structure, i.e. the difference between the 10-year and the 3-month yields[8]. The slope of the yield curve is believed to reflect agents' expectations about the future stance of monetary policy and, thus, it is thought to anticipate business cycle movements (Estrella and Hardouvelis, 1991; Estrella and Mishkin, 1997). Conditioning the test of the Lucas hypothesis to the slope of the term structure means to relating agent's expectations to a leading economic indicator (Stock and Watson, 1989). In particular, the yield curve represents a link between monetary policy, the financial sector and the real economy. We recall that the peculiar aspect of this methodology is to consider expectations on a double level. At the first level, the *micro* level, expectations are modelled through Kalman filtering the money supply function in order to isolate prediction errors (Section 4). At the second level, the *macro* level, expectations have a *forward-looking* nature in that they are intended to capture the future evolution of the economy as reflected by the dynamics of the term structure of interest rates. *Macro* expectations interpret a perspective sentiment present throughout the economy, a broad view regarding the economic conditions, including political as well as institutional, or technological, factors.

A crucial issues to be pointed out is that the *micro* level analysis of expectations is performed on a monetary policy function expressed in terms of money supply. While, the *macro* level expectations are inferred by the evolution of the yield curve, whose dynamics depends not only on the determination of the policy rate, or a measure closely related to it as it may be the effective federal funds rate, but also on the abovementioned factors. The choice of the monetary policy function in terms of the money stock is thus intended to separate two different levels of expectations' analysis.

[7] Figure 4 reports tests when the dependent variable is the HP filtered IP series; tests for other equations with different real variable offer very similar results.

[8] Similar results are obtained if the threshold variable is the spread between the 5-year and the 3-month yields.

We aim at distinguishing the expectations regarding the operational procedure of the monetary authority in setting the money supply from the overall movements displayed by the yield curve. Although there exists an unquestionable inverse relationship linking money supply and short rates, the evolution of the yield curve, as well as the determination of expectations at a *macro* level, depend on a greater variety of factors. So that we believe our approach is immune from the criticism that the *micro* and *macro* structures for expectations share a common root[9].

The threshold methodology implies that the same equation is estimated in two different regimes depending on the values assumed by a predetermined variable (τ), i.e. the yield spread which is a measure of the slope of the term structure.

$$\begin{cases} gap_t = \alpha_0 + \gamma_0 \eta_{t|t-1} + \gamma_1 \left(\eta_{t|t-1} \cdot h_{t|t-1}\right) + \alpha_2 h_{t|t-1} + \alpha_3 gap_{t-1} + v_t & if \quad \tau \leq \hat{\tau} \\ gap_t = \alpha_0 + \gamma_0 \eta_{t|t-1} + \gamma_1 \left(\eta_{t|t-1} \cdot h_{t|t-1}\right) + \alpha_2 h_{t|t-1} + \alpha_3 gap_{t-1} + v_t & if \quad \tau > \hat{\tau} \end{cases} \tag{12}$$

Regime 1 is determined by values of the yield spread below the estimated threshold ($\hat{\tau}$); hence, the first regime is characterized by a flat or downward sloping yield curve. The conventional view tends to associate such a regime to an imminent slowdown in economic activity. On the other hand, regime 2 is defined on high values of the spread (positive slope of the yield curve) reflecting an accommodative stance of monetary policy. Estrella and Hardouvelis (1991), as well as Estrella and Mishkin (1997), present important evidence that the yield spread is related to the future evolution of real activity. In particular, a downward sloping or a flat yield curve augments the odds of a recession in the near future. Along the same line Wright (2006) finds a link between the shape of the yield curve and the probability of future economic slowdowns.

Estimation results for regime 1 (below the estimated threshold) are show in Table 8. The Lucas hypothesis seems to be respected regardless the variable used to measure the business cycle. The conditional variance of money growth affects real variables through the coefficients of the prediction-error term (γ_1). The direct influence of the conditional variance implied by the Friedman hypothesis is not significant with the only exception for the IP gap equation, where surprisingly the effect of the conditional variance appears to work in the opposite direction. However, this is far from being paradoxical as long as when the economy is going toward a recession, a peak in the variability of inflation, captured by the conditional variance of money growth, might act as a stimulus to economic activity, or might be interpreted as a sign that the recession is neither severe nor long-lasting.

[9] If we had chosen to apply the Kalman filter to a monetary policy rule (Section 4) expressed in terms of the policy rate the criticism might have been appropriate.

REGIME 1 - IP gap (HP)						
	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	0.4327	**0.0291**	**-0.0118**	**0.0080**	0.9039	0.860
t-stat	[1.603]	[1.524]	[-1.809]	[2.246]	[19.60]	
W	[2.083]	[2.245]	[-1.893]	[2.358]	[21.60]	obs 74
HH (12)	[2.857]	[2.644]	[-1.838]	[5.067]	[19.49]	
NW (12)	[2.222]	[2.004]	[-1.693]	[2.533]	[18.95]	
s-HH	[0.912]	[1.975]	[-1.557]	[1.630]	[12.74]	

REGIME 1 – TCU						
	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	-0.6857	**0.0242**	**-0.0119**	0.0017	0.9844	0.968
t-stat	[-1.584]	[2.651]	[-2.632]	[0.971]	[100]	
W	[-1.531]	[2.373]	[-2.188]	[1.017]	[97.12]	obs 355
HH (12)	[-1.474]	[2.134]	[-1.993]	[1.448]	[93.61]	
NW (12)	[-1.463]	[2.256]	[-2.080]	[1.194]	[92.99]	
s-HH	[-0.809]	[2.928]	[-2.967]	[0.777]	[51.40]	

REGIME 1 – unemployment						
	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	0.2949	**-0.0472**	**0.0232**	-0.0063	0.9967	0.985
t-stat	[0.071]	[-2.188]	[2.177]	[-1.547]	[146]	
W	[0.073]	[-1.914]	[1.790]	[-1.611]	[145]	obs 362
HH (12)	[0.081]	[-1.761]	[1.708]	[-1.842]	[129]	
NW (12)	[0.084]	[-1.875]	[1.812]	[-1.692]	[143]	
s-HH	[0.033]	[-2.399]	[2.409]	[-1.356]	[70.28]	

REGIME 1 - unemployment grw						
	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	-0.1966	**-0.0126**	**0.0622**	0.0174	1.0198	0.913
t-stat	[-2.558]	[-2.175]	[2.148]	[1.553]	[47.98]	
W	[-2.615]	[-2.551]	[2.484]	[1.566]	[43.52]	obs 248
HH (12)	[-6.869]	[-2.325]	[2.147]	[4.158]	[37.65]	
NW (12)	[-3.198]	[-2.257]	[2.189]	[1.843]	[43.36]	
s-HH	[-1.532]	[-2.305]	[2.357]	[1.217]	[39.38]	

t-statistics in square brackets.

OLS: *ordinary least squares* estimation; **W**: White correction; **HH**: Hansen Hodrick correction; **NW**: Newey-West correction; **s-HH**: simplified Hansen-Hodrick correction (see Cochrane and Piazzesi, AER 2005).

Table 8

Generated variables in the above regression might invalidate inference procedures. To handle with it not only we propose alternative measures of the standard errors (White, 1980; Hansen and Hodrick, 1980; Newey and West, 1987; the simplified Hansen and Hodrick), but also we perform a Monte Carlo simulation. Results show a clear convergence of both Lucas parameters towards the true values just after few thousands replications (γ_0 in the left panel; γ_1 in the right panel). The top diagrams show the simulation with 15000 replications, while in the bottom diagrams we run 50000 replications.

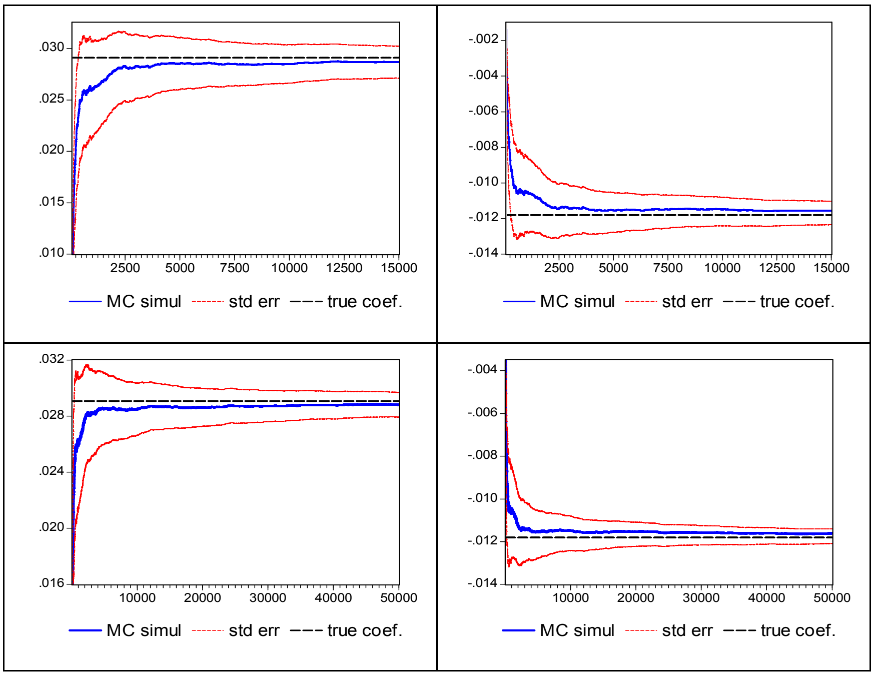


Figure 5

The estimation of regime 1 has also been performed after ruling out the conditional variance of money prediction errors. Results are reported in Table 9. There is a clear confirmation that the Lucas hypothesis is not rejected by the data when the yield curve is either flat or downward sloping.

REGIME 1 - IP gap (HP)						
	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	0.0414	**0.0245**	**-0.0123**		0.9554	0.843
t-stat	[1.152]	[2.963]	[-3.007]		[58.41]	
W	[1.143]	[2.599]	[-2.449]		[49.35]	obs 74
NW (12)	[1.143]	[2.438]	[-2.279]		[50.81]	
REGIME 1 – TCU						
	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	-0.5250	**0.0254**	**-0.0125**		0.9878	0.968
t-stat	[-1.261]	[2.852]	[-2.861]		[104]	
W	[-1.183]	[2.529]	[-2.342]		[98.19]	obs 355
NW (12)	[-1.126]	[2.391]	[-2.207]		[93.54]	

t-statistics in square brackets.

OLS: *ordinary least squares* estimation; **W**: White correction; **NW**: Newey-West correction.

Table 9 (a)

REGIME 1 – Unemployment						
	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	-0.1997	**-0.0532**	**0.0265**		0.9998	0.985
t-stat	[-0.513]	[-2.503]	[2.525]		[153]	
W	[-0.535]	[-2.243]	[2.124]		[147]	obs 362
NW (12)	[-0.557]	[-2.151]	[2.124]		[152]	

REGIME 1 - Unemployment grw						
	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	-0.0843	**-0.0179**	**0.0522**		1.0085	0.911
t-stat	[-3.209]	[-1.891]	[1.841]		[50.35]	
W	[-3.237]	[-1.941]	[1.833]		[45.45]	obs 248
NW (12)	[-3.298]	[-1.872]	[1.778]		[45.00]	

t-statistics in square brackets.

OLS: *ordinary least squares* estimation; **W**: White correction; **NW**: Newey-West correction.

Table 9 (b)

Before showing results for the second regime we would like to offer an insight about the above results. When the variance of nominal shocks tends to be relatively high, it is also true that the actions of the monetary authority are governed by a considerable amount of uncertainty, since predictions errors are quite volatile. Thus, if the central bank cannot easily recognize which is the most suitable strategy to be implemented given the current economic scenario, economic agents internalize such insecurity and assume more balanced and measured behaviors, for instance consumers stop spending money on superfluous goods, and firms refrain from investing in capital goods. Aggregate actions result in output responses that are mild, i.e. of lower magnitude. Hence, there is an inverse relationship between variance of nominal shocks and the magnitude of business cycle fluctuations. This relationship is also true when the variance of nominal shocks tends to be relative small. Suppose the yield curve is flat due to severe monetary policy correctly expected by agents (again this case perfectly fits regime 1), in such an atmosphere output would react heavily falling below the steady state level thus preserving the inverse correlation highlighted by Lucas (1973).

Regime 2 is characterized by high, and positive, values of the yield spread. Regime 2 estimates are reported in Table 10.

REGIME 2 - IP gap (HP)

	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	-0.2375	0.0059	0.0009	**-0.0371**	0.9436	0.904
t-stat	[-2.683]	[0.170]	[0.064]	[-2.949]	[57.55]	
W	[-2.902]	[0.220]	[0.096]	[-3.213]	[48.44]	obs 394
HH (12)	[-2.731]	[0.214]	[0.090]	[-3.033]	[62.92]	
NW (12)	[-2.685]	[0.194]	[0.082]	[-3.045]	[52.57]	
s-HH	[-1.558]	[0.136]	[0.054]	[-2.332]	[39.26]	

REGIME 2 - TCU

	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	0.4740	0.0351	-0.0076	**-0.0041**	**1.012**	0.982
t-stat	[0.787]	[0.815]	[-0.480]	[-2.502]	[73.79]	
W	[0.690]	[1.023]	[-0.694]	[-3.385]	[64.84]	obs 113
HH (12)	[1.948]	[1.039]	[-0.814]	[-4.501]	[181]	
NW (12)	[0.911]	[1.044]	[-0.711]	[-3.227]	[85.78]	
s-HH	[0.930]	[0.603]	[-0.363]	[-2.308]	[86.72]	

REGIME 2 – Unemployment

	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	0.3876	-0.0263	0.0131	**0.0096**	0.9874	0.938
t-stat	[0.394]	[-0.223]	[0.301]	[2.114]	[77.32]	
W	[0.246]	[-0.249]	[0.412]	[3.314]	[46.66]	obs 106
HH (12)	[0.476]	[-0.237]	[0.439]	[5.850]	[63.61]	
NW (12)	[0.325]	[-0.214]	[0.381]	[4.067]	[50.27]	
s-HH	[0.452]	[-0.157]	[0.217]	[1.904]	[95.88]	

REGIME 2 - Unemployment grw

	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	0.1910	-0.0211	0.0983	**0.0169**	0.9416	0.953
t-stat	[2.664]	[-0.842]	[1.026]	[1.745]	[63.27]	
W	[2.906]	[-0.780]	[1.119]	[2.050]	[51.02]	obs 220
HH (12)	[6.212]	[-0.752]	[1.164]	[4.223]	[39.94]	
NW (12)	[3.392]	[-0.767]	[1.165]	[2.377]	[45.15]	
s-HH	[1.618]	[-0.627]	[0.803]	[1.468]	[58.52]	

t-statistics in square brackets.

OLS: *ordinary least squares* estimation; **W**: White correction; **HH**: Hansen Hodrick correction; **NW**: Newey-West correction; **s-HH**: simplified Hansen-Hodrick correction (see Cochrane and Piazzesi, AER 2005).

Table 10

The monetary accommodation reflected in the upward sloping term structure is usually expected to stimulate economic activity thus pushing the economy on an expansionary path. On the other hand, a positive slope of the yield curve implies a positive risk premium required by investors to move to longer horizons. In particular, Ang, Bekaert and Wei (2008) find evidence that the slope of the nominal term structure is due to a positive inflation risk premium. In case of perfect foresight about future spot rates, in fact, the arbitrage mechanism would equalize holding period returns along the entire spectrum of maturities implying a flat yield curve. In a context characterized by imperfect information, uncertainty causes the term structure to deviate from its risk-neutral implied shape. The joint effect of uncertainty and economic growth is reflected in a greater variability of expected inflation, and, thus, in the dynamics of the conditional variance of money growth (Barro, 1976; Friedman, 1977). Therefore, the threshold estimation of regime 2 should return a significant coefficient α_2 stressing out the inverse relationship between the money conditional variance and the economic cycle. Coefficient α_2 is, in fact, negative in the equations expressed in terms of the IP gap and the total capacity utilization; while, coefficient α_2 turns out to be positive when the dependent variable is unemployment or its rate of change. As Rudebusch, Sack and Swanson (2007) have found, this result can be interpreted in the sense that there exists an inverse correlation between term premia and output growth.

6. Concluding Remarks

In this work we propose and implement an innovative method for expectations in order to investigate the relationship between money and output. Previous evidence tends to support the view that money is effective in stimulating real economic activity; however, some contrarian evidence suggests the issue is still controversial. In particular, the high correlation between money and output does not reveal an unambiguous direction of causality. From a policy perspective, in fact, the aforementioned positive correlation can derive from two opposite phenomena. On the one hand, the monetary authority can govern the supply of money to influence the future economic conjuncture; on the other hand, the central bank can manage the dynamics of monetary aggregates in response of past macroeconomic conditions. In addition, the weaknesses associated to traditional approaches call for an effective role of expectations.

After estimating a time-varying monetary policy rule where expectations are analysed at a *micro* level, we condition the examination of the money-output relationship to the shape of the yield curve. In particular, in this paper we test the Lucas (1973) hypothesis against Friedman's (1977). Within a non linear approach, we find evidence that the conditional variance of money growth affects real output through the coefficients on the forecast error term in the Lucas-type output equation only when the flat shape of the term structure reflects expectations of a slowdown in economic activity. Moreover, the conditional variance of money growth, which is used as a proxy for inflation variability, appears to affect directly output, thus corroborating Friedman's theory, when the term structure is upward sloping, i.e. investors require a positive term premium.

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Appendix I - Data

All data have monthly frequency; the sample starts in January 1966. The core econometric analysis, after Kalman filtering, is thus performed from January 1967 since, prudently, we rule out the first 12 observations. The U.S. series of seasonally adjusted industrial production is from the FRED database (Federal Reserve Economic Data). The seasonally adjusted unemployment rate series (civilian unemployment), as well as the total capacity utilization index, are from the FRED database; the source is the U.S. Department of Labour (Bureau of Labour Statistics) indeed. The series are covariance stationary as suggested by both the augmented Dickey-Fuller test and the Kwiatkowski-Phillips-Schmidt-Shin test.

The ADF test rejects the null hypothesis of unit root; while the null hypothesis of stationarity cannot be rejected by the KPSS test. To match the monthly frequency of data, the rule of thumb selected number of lags in the auxiliary regression is either 11 or 12. The automatic lag selection based on different criteria (Akaike, Schwarz, Hannan-Quinn) is consistent with our choice. Unit root test results obtained with the automatic lag selections are similar. The critical values of the KPSS test are 0.739 (1%), 0.463 (5%), and 0.347 (10%) when the intercept is included in the auxiliary model.

Stationarity

	adf (aic) [lag]	adf (sic) [lag]	adf (hq) [lag]	pp (b)	pp (q)	kpss (b)	kpss (q)
ffr	(0.0253)* [16]	(0.0942)* [2]	(0.0883)* [13]	(0.1576)*	(0.1397)*	0.2652*	0.2161*
y3m grw3	(0.0001)* [17]	(0.0000)* [7]	(0.0002)* [16]	(0.0000)*	(0.0000)*	0.0521*	0.0454*
M1 grw3	(0.0548)* [17]	(0.0452)* [12]	(0.0452)* [12]	(0.0000)*	(0.0000)*	0.5193*	0.4683*
M1 grw	(0.0043)* [12]	(0.0210)* [13]	(0.0043)* [12]	(0.0000)*	(0.0000)*	0.7103*	0.9431*
M2 grw	(0.0036)* [17]	(0.1160)* [13]	(0.1160)* [13]	(0.0465)*	(0.0440)*	0.9783*	1.3169*
MB grw	(0.0879)* [13]	(0.0879)* [13]	(0.0879)* [13]	(0.0088)*	(0.0163)*	0.4564*	0.5296
infl(3) (cpi grw3)	(0.0751)* [16]	(0.0729)* [13]	(0.0729)* [13]	(0.0000)*	(0.0001)*	0.5873*	0.5244*
IP gap (HP)	(0.0000)* [9]	(0.0000)* [3]	(0.0000)* [3]	(0.0001)*	(0.0000)*	0.3804*	0.4213*
unemp	(0.0494)* [4]	(0.0494)* [4]	(0.0494)* [4]	(0.1347)*	(0.1429)*	0.1849*	0.2589*
unemp grw	(0.0001)* [16]	(0.0058)* [12]	(0.0001)* [16]	(0.0011)*	(0.0007)*	0.2707*	0.3901*
***log*(tcu)**	(0.0006)* [9]	(0.0066)* [3]	(0.0066)* [3]	(0.0176)*	(0.0189)*	0.4238*	0.5549*

Sample: jan 1967 - dec 2007; * Exogenous: intercept.

adf:augmented Dickey-Fuller test; **pp**: Phillips-Perron test; kpss: Kwiatkowski-Phillips-Schmidt-Shin test ; **aic**: Akaike; **sic**: Schwarz; **hq**: Hannan Quinn; **b**: Barlett; **q**: quadratic special kernel.

ffr: effective federal fund rate;**y3m grw3**: 3-month rate of growth (**grw**) of the 3-month yield; **M1 grw3**: 3-month rate of growth of monetary aggregate M1; **M1 grw**: annual rate of growth of monetary aggregate M1; **M2 grw**: annual rate of growth of monetary aggregate M2; **MB grw**: annual rate of growth of monetary base; **infl(3) (cpi grw3)**: 3-month rate of growth of the seasonally adjusted consumer price index; **IP gap (HP)**: industrial production gap obtained with the Hodrick-Prescott filter; **unemp**: unemployment rate; **unemp grw**: annual change in the rate of unemployment;***log*(tcu)**: *log*-series of the total capacity utilization.

Table 11

In Table 12 the stationarity tests are carried out on both the forecast error and the conditional variance series obtained from the monetary policy function expressed in terms of money growth. As mentioned above in the text both the time-varying parameter model estimated by Kalman filtering and the vector autoregressive model of order 9 have been employed in this paper.

Stationarity							
	adf (aic) [lag]	**adf (sic) [lag]**	**adf (hq) [lag]**	**pp (b)**	**pp (q)**	**kpss (b)**	**kpss (q)**
forecast error VAR (9)	(0.0000)* [11]	(0.0000)* ^{m} [3]	(0.0000)* [11]	(0.0000)*	(0.0000)*	0.2636*	0.2681*
conditional variance VAR(9)	(0.0000)* [4]	(0.0000)* [1]	(0.0000)* [4]	(0.0000)*	(0.0000)*	0.3061*	0.3221*
forecast error (kf)	(0.0000)* [6]	(0.0000)* [4]	(0.0000)* [4]	(0.0000)*	(0.0000)*	0.0733*	0.1093*
conditional variance (kf)	(0.0000)* [3]	(0.0000)* [3]	(0.0000)* [3]	(0.0000)*	(0.0000)*	0.1686*	0.1728*

Sample: jan 1967 - dec 2007; * Exogenous: intercept. Supersctipt *m indicates the* modified version of the test.

adf:augmented Dickey-Fuller test; **pp**: Phillips-Perron test; kpss: Kwiatkowski-Phillips-Schmidt-Shin test ; **aic**: Akaike; **sic**: Schwarz; **hq**: Hannan Quinn; **b**: Barlett; **q**: quadratic special kernel.

VAR(lags): Vector Auto-Regression; **kf**: kalman filtering.

Table 12

Appendix II

In Table 13 we report the estimations of equation (11') after deriving the prediction error and the respective conditional variance series from a vector autoregressive model of order 9. The conditional variance series is obtained computing the square of the residuals in the money equation of the VAR(9) system (Piazzesi, 2003). Money growth, the inflation rate, and the quarterly change of the T-Bill rate are the endogenous variables; the constant is the only exogenous variable. The number of lags has been selected on the basis of the Akaike and Schwarz criteria. Results tend to support the Friedman hypothesis rather than the Lucas' one. The dynamics of the real variables seems to be lowered by the conditional variance of money growth, which is regarded to be a proxy for price volatility.

joint estimation - IP gap (HP) - VAR(9)

	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	0.0473	-0.0491	0.0190	**-0.0194**	0.9497	0.905
t-stat	[0.150]	[-0.283]	[0.139]	[-1.074]	[66.56]	
W	[0.153]	[-0.264]	[0.259]	[-2.332]	[55.23]	obs 468
HH (12)	[0.110]	[-0.234]	[0.211]	[-1.791]	[188.0]	
NW (12)	[0.109]	[-0.232]	[0.224]	[-2.239]	[62.41]	
s-HH	[0.045]	[-0.271]	[0.229]	[-0.722]	[23.57]	

joint estimation - TCU - VAR(9)

	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	0.0819	0.0190	-0.0002	**-0.0346**	0.9813	0.974
t-stat	[2.472]	[0.101]	[-0.106]	[-1.748]	[130]	
W	[2.329]	[0.096]	[-0.202]	[-3.611]	[122]	obs 468
HH (12)	[1.600]	[0.096]	[-0.178]	[-2.412]	[84.32]	
NW (12)	[1.549]	[0.089]	[-0.184]	[-2.935]	[81.73]	
s-HH	[0.778]	[0.096]	[-0.174]	[-1.225]	[40.95]	

joint estimation - Unemployment - VAR(9)

	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	0.0580	-0.0344	-0.0008	**0.0776**	0.9903	0.985
t-stat	[1.656]	[-0.747]	[-0.236]	[1.625]	[176]	
W	[1.499]	[-0.669]	[-0.403]	[3.233]	[147]	obs 468
HH (12)	[0.854]	[-0.784]	[-0.520]	[4.039]	[81.21]	
NW (12)	[0.909]	[-0.681]	[-0.433]	[3.691]	[86.19]	
s-HH	[0.504]	[-0.733]	[-0.398]	[1.088]	[53.49]	

joint estimation - Unemployment grw - VAR(9)

	α_0	γ_0	γ_1	α_2	α_3	R^2
OLS	-0.0397	-0.0411	0.0038	**0.0253**	0.9626	0.934
t-stat	[-0.202]	[-0.381]	[0.458]	[2.245]	[79.95]	
W	[-0.207]	[-0.377]	[0.808]	[4.536]	[64.59]	obs 468
HH (12)	[-0.150]	[-0.571]	[1.449]	[4.195]	[102]	
NW (12)	[-0.146]	[-0.513]	[1.086]	[3.948]	[54.29]	
s-HH	[-0.061]	[-0.365]	[0.755]	[1.538]	[26.23]	

t-statistics in square brackets

Table 13