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"THE PREDICTIVE POWER OF REAL INTEREST RATES FOR FUTURE INFLATION"

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THE PREDICTIVE POWER OF REAL INTEREST RATES FOR FUTURE INFLATION

1 INTRODUCTION

Does a rise in short-term interest rates reflect a rise in real interest rates and, hence, a tightening of monetary policy or, alternatively, a rise in expected inflation? A standard view, commonly referred to as the Fisher effect, is that movements in short-term interest rates primarily reflect fluctuations in expected inflation and, as such, they have predictive ability for future inflation.

Although the Fisher effect is widely accepted for the period after the Fed-Treasury Accord in 1951 until October 1979 in the United States, this relationship between short-term interest rates and future inflation is not at all robust. Despite the general acceptance of the Fisher effect in theory, a stable one-for-one relationship between nominal interest rates and inflation has proven extremely difficult to establish empirically. Typically, estimated slope coefficients in regressions of nominal interest rates on various measures of expected inflation are substantially less than the hypothesized value of one, or, alternatively, real rates are negatively correlated with expected inflation.

The level of short-term interest rates has no ability to predict future inflation in the United States prior to World War II or after October 1979. In addition the Fisher effect is not found to be strong for many other countries.

The prediction of the path of future inflation rate is one of the most discussed macroeconomic topics. The variables that can be used as indicators of inflation fluctuations have been the theme of many empirical studies. Many researchers have tried to the forecast future inflation by using a great number of different methodologies and various econometric approaches.

Fama, Mishkin and a number of other researchers tried to find which are the factors that determine the movement of future inflation rates. *Fama* tested the predictive power of short-term interest rate on inflation, using as the only explanatory variable the nominal interest rate. *Mishkin (1990a, 1990b)* examined empirically the information that the term structure of nominal interest rates contains about the inflation rate. In these approaches the real interest rate is thought to be constant over time.

Real interest rates are among the most important economic variables and have been studied extensively. They figure prominently in discussions of the transmission mechanisms of monetary policy and also play a prominent role in explanations of business cycles and particular business cycle episodes. Real interest rates are a central element in savings-consumption decision and in debates about how to encourage savings. Real interest rates are also a critical explanatory variable for investment decisions since they represent the real cost of borrowing.

However the hypothesis that the real interest rate is constant over time is not a very attractive basis for trying to forecast the future path of inflation rate and is strongly rejected for most studied sample periods. The real interest rate is observed to vary over time. *Mishkin* (1981,1984,1991) himself has helped to document this variation freely.

Mishkin's thorough study (1981) strongly rejects the hypothesis of a constant real interest rate. Moreover *Frankel and Lown (1991)* apply a theoretical framework that allows the variation of the real short-term interest rate over time. Only in the unusual sample period from 1953 to 1971 studied by Fama, in which there was very little variation in the data, is the constancy of the real rate not rejected. Even in this sample period, the failure to reject constancy of the real rate should not be taken as strong evidence that the real interest rate was constant in that period, rather it appears to be a reflection of the lack of statistical power of tests on the beta coefficients.

Generally the international bibliography concludes that there is not a constant short-term real interest rate, but so far there is to our knowledge no empirical study that tries to connect the variation of the short term real interest rate with the future fluctuations of inflation.

Our paper will mainly focus on whether real interest rates contain useful information for the expected inflation rate. We analytically explain the theoretical framework that allows the real rate to vary over time and we derive a model that will estimate the relationship between real interest rates and future inflation. We also include the term structure in our regressions, in order to investigate whether the real rate contains useful information for inflation, over and above the information contained in the term structure. A series of papers have examined the information contained in the term structure about future inflation in the United States-see Fama and Gibbons (1982,1984), Fama (1990), Mishkin (1988,1990a, 1990b, 1991). The general finding is that the longer the time horizon the more information the term structure possesses about the future inflation. At the shortest end, i.e. for maturities up to six months, the term structure is found to have little predictive ability for inflation.

The structure of this paper is as follows: First we analyze the Fisher Hypothesis. Then we describe previous empirical studies (methodologies, data and results). In section 3 we explain the reasons that motivated or study and describe analytically the derivation of the model with which we believe that inflation can be forecasted. In section 4 we describe our data set. Section 5 contains the methodology used to test for unit root in our series and it's results. In section 6 we analyze our empirical results and section 7 contains the economic interpretation of our empirical results. This paper closes with the conclusions of our empirical analysis (section 8).

1.1 FISHER HYPOTHESIS

Ever since Irving Fisher's *"The Theory of Interest,"* the conjecture that nominal interest rates vary, *ceteris paribus*, one-to-one with expected inflation has become one of the most intensively studied topics in economics. Fisher argues that deterministic changes in inflation expectations cause equal changes in nominal interest rates, leaving the real interest rate unaffected. The Fisher effect is a cornerstone of many theoretical models that generate monetary neutrality and is important for understanding movements in nominal interest rates.

The Fisher equation is one of the oldest and simplest equilibrium asset pricing models. It relates the one period nominal interest rate to the sum of the ex-ante real interest rate and expected inflation:

$$i(t)=r(t+1) + \pi^{e}(t+1)$$

where

$i(t)$: one period nominal interest rate at time t,

$r(t+1)$: the corresponding ex ante real interest rate at time t+1, given information up to time t,

$\pi^{e}(t+1)$: the expected inflation rate between t and t+1 given information up to time t.

One of the most popular interpretations of the Fisher equation is that if the ex -ante real interest rate is determined by «technology and tastes», then the nominal interest rate is related one-to-one to expected inflation. In other words, we can say that the real rate of interest is independent of inflation expectations.

The Fisher equation draws attention to an important finding about money growth, inflation and interest rates. In the long run the real interest rate converges to it's full long-term average, r^*, and actual and expected inflation are equal. Using these two facts ($r(t)=r^*$, $\pi^e = \pi$), the long run relationship is expressed as $i = r^* + \pi$. With r^* given, this equation implies that when all adjustment has occurred, an increase in inflation is reflected fully in nominal interest rates. Nominal interest rates increase one-to-one with increase in inflation. The reason for such a strong inflation-nominal interest rate link is that in the long run the real interest rate is unaffected by monetary disturbances, which, however, do affect the inflation rate. Of course, the constancy of the real interest rate holds only in the long run equilibrium. During the adjustment process, the real interest rate does change and short run changes in the nominal interest rate reflect both changes in real rates and changes in inflationary expectations.

Because changes in the value of money redistribute purchasing power between debtors and creditors, a unity response of nominal interest rates to changes in expected inflation is required to avoid such re-distributions and insulate the real rate of interest. This is called "full" or "strict" (point-for-point) Fisher effect.

The Fisher Hypothesis is conventionally tested using the following regression[1]:

$$i(t) = \alpha + \beta \pi^{e}(t+1)$$

where $i(t)$ is the one period nominal interest rate, α is the constant real interest rate and π^{e} is the expected rate of inflation.

If the Fisher hypothesis holds, then β should be equal to unity. This implies that the nominal interest rate moves one-to-one with the rate of inflation. The weak form allows β to be greater than one due to Mundell-Tobin[2] effect (the decline in marginal product of capital due to the real balance response to inflation), under which the substitutability of bonds and real money balances results in a negative effect of anticipated inflation on real bond yields. When taxes on the nominal interest rate are included, β has been found to be less than one.

[1] See Fama (1975)
[2] See Levi and Makin(1978), Melvin(1982), Peek and Wilcox(1983)

1.2 HAS THE REAL INTEREST RATE PREDICTIVE POWER FOR INFLATION OVER AND ABOVE THE TERM STRUCTURE?

We have already seen that the Fisher equation is expressed as :

$i_{t,t+1} = r_{t,t+1} + \pi_{t,t+1}$. (a)

According to Fisher equation, the n-period nominal interest rate is equal to expected inflation over n periods plus the n-period real interest rate:

$i_{t,t+n} = r_{t,t+n} + \pi_{t,t+n}$. (b)

By subtracting (a) from (b) we obtain:

$i_{t,t+n} - i_{t,t+1} = (r_{t,t+n} - r_{t,t+1}) + (\pi_{t,t+n} - \pi_{t,t+1})$,

where

$i_{t,t+n} - i_{t,t+1}$: the term structure of nominal interest rates,

$r_{t,t+n} - r_{t,t+1}$:the real term structure,

$\pi_{t,t+n} - \pi_{t,t+1}$: the expected change of inflation from period 1 to period n.

If $r=r^*$ (r^*: the long run real interest rate) then $r_{t,t+n}= r_{t,t+1}$, which is the case considered in Mishkin's papers.

If r is time varying, then the term structure of the real interest rate must have predictive power for the future changes of inflation in addition to the term structure of nominal interest rates, provided that economic shocks have a different impact on the real rate than on the nominal rate of interest (real shocks may affect the real rates far more than the nominal rates).

2 PREVIOUS EMPIRICAL STUDIES

2.1 FAMA'S APPROACH

Eugene Fama (1975) tests the joint hypothesis that (i) expected real returns on U.S Treasury Bills are constant through time and (ii) The Treasury Bill market is efficient in the limited sense that interest rates on bills are based on assessments of expected future inflation rates that properly use the information in past inflation rates.

2.1.1 METHODOLOGY

Fama's main tests come from estimates of

$$\pi_t = \alpha + \beta R_t + e_t$$

Where

R_t: nominal interest rates on a Treasury Bill

π_t: actual percentage change in purchasing power over the term to maturity of the bill,

$\alpha = E(r)$, constant real interest rate

e_t: an error term

This equation is a test of Fisher hypothesis. If the expected real rate is constant and if correctly anticipated inflation is on average incorporated into the market rate of interest, then the coefficient β should be approximately one. Any trends in the expected inflation should show up in the interest rate. If the two variables do not follow common trends, the model must be rejected.

2.1.2 DATA & EMPIRICAL RESULTS FOR ONE MONTH TREASURY BILLS

The one-month nominal rate of interest R_t used in Fama's tests is the return from the end of month t-1 to the end of month t on the Treasury Bill that matures closest to the end of month t. The CPI is used to estimate π_t, the rate of change in the purchasing power of money from the end of month t-1 to the end of month t. The tests cover the period from January 1953 through July 1971. The data are from the quote sheets of Salomon Brothers. Fama extends the results to Treasury Bills with longer maturities.

To test the market efficient hypothesis Fama examines if the one-month nominal interest rate set in the market at the end of month t-1 is based on correct utilization of all the information about the expected value of π_t, which is in the time series of past values of π_t. The hypothesis would be meaningful only if past rates of change in purchasing power do indeed contain information about the expected future rate of change.

Unlike Fisher and most of the rest of literature, Fama's results suggest that, at least during the 1953-1971 period, there are definite relationships between nominal interest rates and rates of inflation subsequently observed. During the period 1953-1971, the bond market seems to be efficient in the sense that in getting one-to-six month nominal rates of interest, the market correctly uses all the information about future inflation rates that is in the time series of past inflation rates.

The hypothesis that equilibrium expected real returns on one to six-month bills are constant during the period cannot be rejected by Fama. When combined with the conclusion that the market is efficient, this means than one cannot also reject the hypothesis that all variation through time in one to six-month nominal rates of interest mirrors variation in correctly assessed one to six-month expected rates of change in purchasing power.

Finally, another interesting result is that substantial variation in nominal bill rates during the 1953-1971 period seems to be due entirely to variation in expected inflation rates; in other words, expected real returns on Treasury Bills seem to be constant during the period.

2.2 MISHKIN'S APPROACH

Mishkin in a series of papers examines empirically what the term structure of interest rates tells us about future inflation and if this information can be used to predict the path of future inflation rate.

2.2.1 METHODOLOGY

Mishkin uses the "inflation – change equation" which is the regression of change in the future m-period inflation rate from the n-period inflation rate ($\pi^m_t - \pi^n_t$) on the slope of the term structure ($i^m_t - i^n_t$).

$$\pi^m_t - \pi^n_t = \alpha_{m,n} + \beta_{m,n} (i^m_t - i^n_t) + \eta^{m,n}_t$$

where

π^m_t : inflation rate from time t to t+m.

π^n_t : inflation rate from time t to t+n.

i^m_t : m-period nominal interest rate at time t.

i^n_t : n-period nominal interest rate at time t.

$\alpha_{m,n}$: $rr^n - rr^m$, constant real term structure.

$\eta^{m,n}_t$: an error term.

2.2.2 DERIVATION OF THE INFLATION CHANGE EQUATION

According to Fisher equation, expected inflation over m periods is equal to the m-period nominal interest rate minus the m-period real interest rate:

$E_t\pi^m_T = i_t^m - r_t^m$

where,

E_t : expectations at time t.
π^m_t : inflation rate from time t to t+m.
i^m_t : m-period nominal interest rate at time t.
r_t^m: m-period ex-ante real interest rate.

The realized inflation rate over the next m periods can be written as the expected inflation rate plus the forecast error of inflation:

$\pi^m_t = E_t \pi^m_t + \varepsilon_t^m$ where $\varepsilon_t^m = \pi^m_t - E_t \pi^m_t$ is the forecast error of inflation.

So the equation $E_t\pi^m_t = i_t^m - r_t^m$ is written as

$$\pi^m_t = i_t^m - r_t^m + \varepsilon_t^m$$

In order to examine the information in the term structure about future changes in the inflation rate, Mishkin subtracts the above equation for the n-period inflation rate from the same equation for the m-year inflation rate to obtain:

$$\pi^m_t - \pi^n_t = i_t^m - i_t^n - r_t^m + r_t^n + \varepsilon_t^m - \varepsilon_t^n$$

This equation can be rewritten in the form of the already mentioned inflation-forecasting equation as follows:

$$\pi^m_t - \pi^n_t = \alpha_{m,n} + \beta_{m,n}\,(i^m_t - i^n_t) + \eta^{m,n}_t$$

where

$\alpha_{m,n} = r_t^n - r_t^m$. $\beta_{m,n}=1$

$\eta^{m,n}_t = \varepsilon_t^m - \varepsilon_t^n - (u_t^m - u_t^n)$

$u_t^m = r_t^m - \bar{r}^m$

$u_t^n = r_t^n - \bar{r}^n$

Tests of the statistical significance of the $\beta_{m,n}$ coefficient and whether it differs from one reveal how much information there is in the slope of the term structure about future changes in inflation. We must point out the fact that the above equation is quite restrictive, because of its theoretical framework. The real interest rate is assumed to remain constant over time and as a result the nominal interest rate of any given term is only useful for predicting inflation of the same term. So it is important to recognize that using the regression equation above to test for whether there is information in the term structure about the future path of inflation implies that the word "information" is being used quite narrowly.

If we assume that expectations are rational (REHTS[3]) and the slope of the real term structure remains constant over time, then ordinary least squares (OLS) estimates of the forecasting equation produce a consistent estimate of β_n^m. We see this by first recognizing that rational expectations implies that the forecast errors of inflation ε_t^m and ε_t^n are orthogonal to the right hand regressors because under rational expectations $E_t\ \varepsilon_t^m = E_t\ \varepsilon_t^n = 0$. Constancy of the real term structure then makes OLS estimates consistent because the $(u_t^m - u_t^n)$ term disappears, leaving an error term for the forecasting equation of $(\varepsilon_t^m - \varepsilon_t^n)$ which is orthogonal to the right-hand-side regressors under rational expectations. If the slope of the real term structure is not constant, then the nominal term structure can still contain information about future changes in the inflation rate, but it is no longer an optimal predictor because $(u_t^m - u_t^n)$ is no longer equal to zero.

2.2.3 DATA AND EMPIRICAL RESULTS

Mishkin used various datasets in his empirical analyses. In one of his papers (1990) he used monthly data on inflation rates and one- to twelve-month US Treasury bills for the period February 1964 to December 1986.

As we can extrapolate from Mishkin's estimations, at the shortest end of the term structure (maturities six months or less) the results for the inflation change equations

[3] The REHTS (Rational Expectations Hypothesis on Term Structure) is based on the arbitrage condition that after adjusting for risk the expected return from holding for one period a bond that has n-periods to maturity is the same as the certain return from a one period bond. In other words, the expected excess holding period return is equal $E_tY(n,t+1) - Y_{1t} = \varphi(n,t)$, where the holding period return is the capital gain from holding an n-period bond for one-period and then selling it i.e. $Y(n,t+1) = \ln P(n-1) - \ln P(n,t)$, $P(n,t)$ is the price at time t of a discount bond with face value \$1 and n-periods to maturity, $Y_{n,t}$ is the certain one period return, and $\varphi(n,t)$ is the risk premium perceived at time t. Under risk neutrality the risk premium is assumed to be equal to zero, $\varphi(n,t) = 0$.

are strikingly different than that found for inflation level regressions of the type popularized by Fama (1975). In contrast to Fama-type regressions, which find that the level short term interest rates contain a great deal of information about the future level of inflation, at least before 1979, there is little ability of the slope of the term structure to forecast future changes in inflation. Apparently, the term structure for maturities of six months or less contains almost no information about the path of future inflation. The results for the longer maturities are quite different, however. The β coefficients are quite close and significantly different from zero at the 1% level.

Mishkin conducted the same study using monthly data from 1953 to 1987 for inflation rates and interest rates on one-through five-year Treasury bonds. His results indicate that there is a great deal of information in the longer maturity term structure about the future path of inflation.

Mishkin finds two principal results. First, the coefficient b is significantly greater than zero, showing that the term structure does contain information useful for predicting the path of the future inflation rate. Second β is significantly less than one. This is a rejection of the null hypothesis that real interest rates are constant.

This finding (the non-constancy of the real interest rate), casts doubt on the results of the approaches that consider the real interest rate to be constant over time. In the next paragraphs, (2.3, 2.4), we report some generally accepted facts about the real interest rate and we describe a series of empirical studies that uncover the non-constancy of the real interest rate.

2.3 STYLIZED FACTS ABOUT REAL INTEREST RATES

Various papers (*Fama, Nelson and Schwert, Mishkin (1981b, 1984a,b), Fama and Gibbons, Summers, Huizinga and Mishkin (1984, 1986), and Cumby and Mishkin,* have uncovered a number of basic facts about the behavior of real interest rates in USA and other countries:

a) The hypothesis that there is a constant real interest rate is strongly rejected for most sample periods.

b) The real interest rate is negatively correlated with expected inflation.

c) Increased money supply growth is associated with a decline in real interest rates. However, little evidence has been found that money growth affects real interest rates other through its effect on inflation, although this result could be due to the low power of the statistical tests.

d) In the postwar period in the United States, fluctuations in nominal interest rates have typically reflected changes in expected inflation rather than in real interest rates.

e) Other countries often display a different relation between nominal interest rates with either expected inflation or real interest rates than is found in the United States. Even before 1979, several other countries had only a very weak Fisher Effect, with low correlation between nominal interest rates and expected inflation yet had a strong positive correlation between nominal and real rates.

f) Real interest rates were extremely high during the contraction phase of the Great Depression. From the perspective of real rates, money was extremely tight in this period. In 1950s and 1960s real rates were positive and turned negative in the mid- to late 1970s, jumped up dramatically in the early 1980s to levels comparable to those during the phase of Great Depression, and although having declined somewhat have continued to remain at levels above those in 1950s and 1960s.

g) Real interest rates in other developed countries tend to move in tandem with those in the United States. Just as in the United States, real interest rates in these countries decline from the 1960s to the 1970s and then rose to levels unprecedented in the postwar period.

2.4 FRANKEL AND LOWN

Frankel and Lown apply a simple existing theoretical framework, which allows the real interest rate to vary in the short run but converge to a constant in the long run, to the problem of predicting the inflation spread. They show that the appropriate indicator of expected inflation can make use of the entire length of the yield curve, in particular by estimating the steepness of a specific nonlinear transformation of the curve, rather than being restricted to a spread between two points.

2.4.1 THE THEORETICAL FRAMEWORK THAT ALLOWS THE REAL INTEREST RATE TO VARY OVER TIME

Frankel and Lown assume that the real interest rate, though not constant, is expected by market to converge to a constant in the long run in the absence of future disturbances. Specifically they assume that the short-term nominal interest rate is expected to adjust to steady-state inflation rate according to a first-order differential equation

$$\frac{dit}{dt} = -\delta \, (i_t - \pi^e_0 - r) \qquad (1)$$

where i_t is the instantaneously short-term interest rate, π^e_0 is the exogenous long-run inflation rate expected at time 0, and r is the constant long-run real interest rate, all of which are not directly observable.

This equation is a way of specifying, as simply as possible, the notion that the real interest rate is not constant in the short run but has a tendency in that direction in the long run.

2.4.2 METHODOLOGY

The above equation implies that at time 0, market participants expect the short term nominal interest rate at time t to be a weighted average of the long run nominal interest rate ($\pi^e_0 + r$) and the current short-term nominal interest rate i_0:

$i_t = [\,1 - \exp(-\delta t)]\,(\pi^e_0 + r) + [\exp(-\delta t)]\,i_0$ (2)

By assuming that i^τ_0 (the interest rate on τα-maturity bonds issued at time 0) is the average of the expected instantaneously short-term interest rates between time 0 and time τ, plus a possible liquidity premium term k_t, Frankel and Lown derive the following equation:

$i^\tau_0 = k_t + i_0 + (\pi^e_0 + r - i_0)(1 - w_t)$,

where w_t the weights given by $w_t = \dfrac{1-\exp(-\delta\tau)}{\delta\tau}$.

For any given point in time they observe the Treasure security rate for two or more maturities. Then (given estimate of δ) they solve for the unknown ($\pi^e_0 + r - i_0$), which is the appropriately calculated slope of the yield curve at that point in time.

For each monthly observation over the period January 1960 to December 1988, they estimate the steepness of the yield curve at that point in time by running the regression

$$i^{\tau}_t = B0_t + B1_t \left[1- \frac{1-\exp(-\delta\tau)}{\delta\tau}\right];$$

the coefficient $B1_t$ is the appropriate estimate of the measure of steepness. Then they see if the time series $B1_t$, can help forecast the difference between the future 12-month inflation rate and the future 3-month inflation rate. They also subtract k_t premiums estimated as the average term structure over time from the interest rates (in log form), before running the regressions at each point in time to estimate $B1_t$.

2.4.3 DATA AND EMPIRICAL RESULTS

Frankel and Lown use in their paper US interest rates for the period 1960-1988 which are monthly averages of daily figures. The inflation rates used are constructed from the seasonally adjusted consumer price index.

The t-statistic is significant at the 99% level, indicating that the steepness of the yield curve does contain useful information regarding the future path of inflation. The results of this approach indicate that the measure extracted from the interest rate yield curve from it's entire length is a considerably better indicator than the simple term spread used by Mishkin.

2.5 OTHER EMPIRICAL STUDIES

Carlson (1977) contradicts Fama's assumptions according to which a) there is a constant expected real rate of interest, and b) all relevant information about future inflation is fully incorporated in the expected-inflation component of the market rate of interest. Carlson supports that the first assumption does not hold up when an expected-real-rate series is constructed from survey data on inflation expectations. The second assumption is challenged on Fama's own ground by showing that significant information about subsequent inflation has not been fully reflected in nominal interest rates and by arguing that the common trend in the data gives rise to the statistical illusion that variations in interest rates on Treasury Bills are good predictors of variations in inflation. **Carlson concludes that real interest rates do have notable variation**, rejecting Fama's assumption. According to his empirical results real interest rates fall during recessions. Moreover he points out that information about inflation that is not fully reflected in nominal interest rates is provided by an additional variable, the ratio of employment to population. These evidence results in the rejection of Fama's joint hypothesis that: (i) expected real returns on US Treasury Bills are constant through time and (ii) The Treasury Bill market is efficient in the limited sense that interest rates on bills are based on assessments of expected future inflation rates that properly use the information in past inflation rates.

Jorion and Mishkin (1991) work on a multi country comparison of term structure forecasts at long horizons. The results for countries like Britain, Germany and Switzerland are similar to those found for the U.S. The evidence of their empirical analysis suggest that for these countries the term structure can be used to help assess long-run inflationary pressures: a steepening of the slope of the longer maturity term structure indicates that the inflation rate will rise several years in the future, and conversely, a negative slope indicates reduced inflationary pressures.

Wickens and Tzavalis (1995) provide a derivation of the connection between the term spread and future inflation based on the n-period Fisher equation. They show that the forecasting ability of the spread regressions is very poor and alternatively they propose a model that makes fuller use of the information contained in the term structure. It is an error correction model and is based on all of the cointegrating vectors implied by the REHTS (Rational Expectations Hypothesis on Term Structure), the one-period Fisher equation, together with the short run dynamics. This model is proved to provide more reliable forecasts. Wickens and Tzavalis conclude that basing a forecasting equation for future inflation solely on the term spreads is not the best approach and that more information about inflation is contained in the real return.

Shmuel Kandel, Aharon R. Ofer, and Oded Sarig (1996) develop a method of measuring ex-ante real interest rates using prices of index and nominal bonds. Employing this method on Israeli data they directly test the Fisher hypothesis that the real rate of interest is independent of inflation expectations. Their empirical results indicate that a negative correlation between ex-ante real interest rates and expected inflation exists.

Mishkin (1992) reexamines the widely accepted view that there is a strong Fisher effect in postwar US data. Recognition that the level of inflation and interest rates may contain stochastic trends suggests that the ability of short-term interest rates to forecast inflation in the postwar United States is spurious. Mishkin's finding that the forecasting relationship between inflation and short term interest rates might be spurious suggests that there might be no short run Fisher effect. However, the absence of a short-run Fisher effect does not rule out the possible existence of a long-run Fisher effect in which inflation and interest rates trend together in the long run when they exhibit trends. Cointegration tests for a common trend in interest rates and inflation provide support for the existence of a long-run Fisher effect. Mishkin also gives an explanation why the Fisher effect appears to strong only for particular sample periods, but not for others. The conclusion that there is a long-run Fisher effect implies that when inflation and interest rates exhibit trends these two series will trend together and thus there will be a strong correlation between inflation and interest rates.

Hutchison and Keeley demonstrate that a change in the stochastic process generating money can alter the relationships between money and inflation and between inflation and interest rates. The extent to which inflation is forecastable is shown to depend significantly on the extent to which money is forecastable.

Martin Evans and Paul Wachtel (1992) use modern asset pricing theory to examine the behavior of short-term nominal interest rates. Their analysis investigates whether variation in the stochastic behavior of consumption growth and inflation can explain movements in the rate of interest. Their results reveal that much of the month-to-month movement in nominal interest rates reflects changes in the real rate and in the risk premia in addition to inflationary expectations. They conclude that there is little correlation between changes in nominal interest rates and expected inflation with quarterly or monthly data.

Malliaropulos (2000) finds empirical evidence that inflation, nominal, and real interest rates in the US are trend-stationary with a structural break in both the unconditional mean and the drift rate of a deterministic trend, which occurs shortly after the change in operating procedures of the Fed in September 1979. This finding casts some doubts on cointegration tests of the long-run Fisher effect conducted in recent studies, since the results of these tests can be affected by the existence of common structural breaks in the series. He proposes an alternative test of the Fisher effect, based on a VAR representation in appropriately detrended variables. He finds strong support for the Fisher effect both in the medium term and in the long term.

3 MOTIVATION FOR OUR EMPIRICAL STUDY

From the empirical studies that are reported in the previous paragraphs, it is easily understood that in the short-run, the real interest rate cannot be assumed constant over time. *Mishkin (1984), Carlson (1977), Frankel and Lown (1991) Wickens and Tzavalis (1995)* among others, support such a result.

But despite the fact that the real interest rate seems to vary over time, there is to our knowledge, no empirical study that examines the impact of a time varying real interest rate on the future path of inflation.

In this section we attempt to establish the theoretical relationship between real interest rate and inflation based on a standard macroeconomic model which combines simple equilibrium relationships of the real sector, the monetary sector and agents' expectations.

3.1 THE MODEL: DERIVATION

Our model is derived from the following relationships:

RELATIONSHIPS

$y_t - y^* = -\gamma\,(i_t - \pi^e_{t+1} - r)$

(1)

$m_t - p_t = \varphi\, y_t - \lambda i_t$

(2)

RELATIONSHIPS

$\pi_t = \rho(\,y_{t-1} - y^*) + \pi^e_T$

(3)

$\Delta m_t = \mu + u_t$

(4)

where

$m_t \sim$ RW with drift μ

$u_t \sim N\,(0,\sigma^2_u)$

y_t: the log of output at time t

y^* : the log of normal or potential output

i_t: the short-term nominal interest rate at time t

π^e: the expected long run inflation rate

p_t: the log of price level at time t

r: constant long run real interest rate in the absence of disturbances

EXPLANATION OF THE MACROECONOMIC RELATIONSHIPS

(1) is an IS relationship; it says that the output gap is related to the current real interest rate through investment demand. In the long run when $y_t = y^*$, we have $i = \pi^e + r$.

(2) is an LM relationship which shows all combinations of interest rates and level of income such that the demand for real balances is equal to the supply. The demand for real balances increases with income and decreases with the interest rate, the cost of holding money than other assets. Real money demand depends positively on income, with an elasticity φ, and negatively on the interest rate, with a semielasticity of λ.

(3) is a Lucas supply function: it says that the rate of price change is given by the sum of an excess demand term and the expected steady state inflation rate.

An increase in the money supply reduces the interest rate, increases investment spending and aggregate demand, and thus increases output. This happens because an increase in money supply generates portfolio disequilibrium. That is, at the prevailing interest rate and level of income, people are holding more money than they want. This causes portfolio holders to attempt to reduce their money by buying other assets. In the process asset prices increase and yields decline. The decrease of the interest rate leads to an increase in output.

In the short-run if the nominal money stock is increased, at each price level interest rates are lower and the demand for output rises. For the price level before the money supply has risen there is excess demand for goods. Firms find that their inventories are running down and accordingly hire more labor and raise output until the economy reaches a short run equilibrium, where both prices and output are higher. The rise in prices is due to the increased labor costs as production and employment rise. As a result, increased money growth will reduce the real interest rate.

In the long run, as long as output is above normal, wages are rising. Because wages are rising, firms experience cost increases and, as a result, output will be declining and prices will keep rising. So in the long run, once wages and prices have time to adjust fully, an increase in the money stock has no real effects. After all adjustments have taken place the real interest rate will return to its initial level. Since the nominal interest rate equals the real interest rate plus the inflation rate, once the real rate has returned to its initial level, the nominal interest rate will have increased. In other words there are only nominal effects in the long run caused by an increased money supply. This is called money-neutrality and holds only for the long run. In the short-run the adjustment of wages and prices are in fact slow due to long-term contracts and "menu costs".

From (2) we have

(2΄): $\Delta m_t - \Delta p_t = \varphi \Delta y_t - \lambda \Delta i_t$; Note $\Delta \pi_T = \pi_t$

From (1) we have

(1΄): $\Delta y_t = -\gamma\, (\Delta i_t - \Delta \pi^e_{t+1})$: Assuming $\Delta y^* = \Delta r = 0$

If we substitute (3), (4), (1΄) in (2΄) we get

(2΄΄): $\mu + u_t - \rho(y_{t-1} - y^*) - \pi^e_T = -\varphi\gamma(\Delta i_t - \Delta \pi^e_{t+1}) - \lambda \Delta i_t$

Then we substitute (1) in (2΄΄) for $(y_{t-1} - y^*)$ and we get

(2΄΄΄): $\mu + u_t + \rho\gamma\, (i_{t-1} - \pi^e_t - r) - \pi^e_t = -(\lambda + \varphi\gamma)\Delta i_t + \varphi\gamma \Delta \pi^e_{t+1}$

Assuming that the long-run expected rate of inflation equals any rate of money supply, $\pi^e_t = \mu$, we obtain from (2΄΄΄)

$$(5)\ \Delta i_t = -\frac{\rho\gamma}{\lambda + \phi\gamma}(i_{t-1} - \pi^e_t - r) + \frac{\phi\gamma}{\lambda + \phi\gamma}\Delta \pi^e_{t+1} - \frac{1}{\lambda + \phi\gamma}u_t$$

Rearranging (5):

$$\Delta \pi^e_{t+1} = -\frac{\rho}{\phi}r + \frac{\lambda + \phi\gamma}{\phi\gamma}\Delta i_t + \frac{\rho}{\phi}(i_{t-1} - \pi^e_t) + \frac{1}{\phi\gamma}u_t$$

$\Delta\pi^{e}_{t+1}$ is not directly observable. So we have to make the assumption of rational expectations

RE: $\pi^{e}_{t+1}=E_t(\pi_{t+1})= \pi_{t+1} + \varepsilon_t$ where $\varepsilon_t \sim N(0,\sigma^2_\varepsilon)$

Then

$$6.\Delta\pi_{t+1} = -\frac{\rho}{\phi}r + \frac{\lambda+\phi\gamma}{\phi\gamma}\Delta i_t + \frac{\rho}{\phi}(i_{t-1} - \pi_t) + \eta_{t+1}$$

where $\eta_{t+1}= \frac{1}{\phi\gamma} u_t +(1-\frac{\rho}{\phi}) \varepsilon_t - \varepsilon_{t+1}$

Since agents have rational expectations, ε_{t+1} will not be correlated with lagged information. Furthermore, ε_{t+1} will not be correlated with π_{t+1} or $\Delta\pi_{t+1}$.

$\Delta\pi_{t+1}$ depends on a constant (r), Δi_t and the time varying short term real interest rate $(i_{t-1}-\pi_t)$.

In the long run $\Delta\pi_{t+1} = \Delta i_t = \eta_{t+1} = 0$, and the above equation becomes:

$\frac{\rho}{\phi}r = \frac{\rho}{\phi}(i_{t-1} - \pi_t) \Rightarrow r = i_{t-1} - \pi_t$ which is the Fisher hypothesis.

Our study will focus on the estimation of the model above (equation 6). First we will examine the properties of our time series and test their stationarity. Augmented

Dickey-Fuller tests and Andrews and Zivot tests for a unit root will be used for this part of our analysis.

Then we will estimate equation 6. We must point out the fact that the constant long run real interest rate r and the structural parameters ρ, φ, λ, and γ cannot be estimated from equation 6. In order to estimate their values we need one more relationship, since we have five unknown parameters and four equations. The modeling of the residuals shows that they contain a moving average part that must be taken into account in the estimation. We will use the non-linear least squares method allowing for ma error terms. Regressions will be estimated in E-Views 3.1.

4 DATA DESCRIPTION

Data[4] are monthly observations for the US and cover the period 1970:01 to 1986:12. We use the same data set as Mishkin (1990a), namely zero coupon yields constructed by McCulloch and reported in the appendix of Shiller (1990). We use the same data set with many other researchers, in order to obtain comparable results. Moreover we do not update this data base, since it involves a quite complicated method, which is described in appendix A. Data are monthly observations of the nominal yields of 1,3,6,9,12 month United States government bonds, continuously compounded and expressed as annual percentage rates over the period 1970:01-1986:12. The inflation rate is based on the CPI and is the same series with that used by Mishkin.

Specifically

- the one-month inflation rate of period t is $\Pi t = \ln \frac{CPI(t+1)}{CPI(t)} * 12$.
- the quarterly inflation rate of period t is $\Pi t = \ln \frac{CPI(t+3)}{CPI(t)} * 4$.
- the six months inflation rate of period t is $\Pi t = \ln \frac{CPI(t+6)}{CPI(t)} * 2$.
- the nine months inflation rate of period t is

  $\Pi t = \ln \frac{CPI(t+9)}{CPI(t)} * 1{,}33$.
- the annual inflation rate of period t is $\Pi t = \ln \frac{CPI(t+12)}{CPI(t)}$.

[4] Descriptive Statistics of our series are presented in Appendix B.

The ex-post real rate is computed from the Fisher Equation as,

$$r(t,t+j) = i(t,t+j) - \Pi(t,t+j) \text{ where } \Pi(t,t+j) = \ln\frac{CPI(t+j)}{CPI(t)}, \; j=1,3,6,9,12 \text{ m}$$

as we have made the assumption of rational expectations that is:

RE: $\pi^{e}_{t+1}=E_t(\pi_{t+1})= \pi_{t+1} + \varepsilon_t$ where $\varepsilon_t \sim N(0,\sigma^2_{\varepsilon})$.

According to *Huizinga and Mishkin (1986)*, this sample period covers three different monetary regimes:

1. In the first period, 1970:01 – 1979:09, the Fed targeted nominal interest rates.

2. In the second period, from 1979:10 – 1982:09, the Fed used money base control and the nominal interest rates were allowed to fluctuate freely.

3. Thereafter (1982:10 – 1986:12) the Fed returned to a system in which movements in the nominal interest rates were restricted.

However inspection of the figures of real one, three, six, nine and twelve month returns, suggest that two only regimes exist for the real rate, with the break occurring somewhere after the middle of 1980.

An important puzzle is why real interest rates increased so sharply in the 1980s.The different behavior of the 1970's from the 1980's is often attributed to the large budget deficits of the 1980's combined with tight (disinflationary) policy which has been followed under Reagan's administration, see *Bosner-Neal (1990)*. But closer scrutiny of the data does not give strong support to this view.[5]

Instead, research by *Huizinga and Mishkin (1986)* points to monetary policy as the source of high real interest rates. Monetary theory (Lucas) suggests that regime changes have an important impact on the stochastic process of many economic variables. With the change in the monetary regime in October 1979, the Federal Reserve Board changed the method of conducting monetary policy in order to reverse the inflationary monetary policy of the 1970s. The basic question is whether this disinflationary, monetary regime change is associated with a shift in the stochastic process of real interest rates, which resulted in the high real rates in the 1980s.

The answer appears to be yes. When the Federal Reserve Board altered its behavior in October 1979, there was a statistically significant shift in the process of real interest rates.[6] In addition, if one asks when the shift in the stochastic process of real rates actually occurred, statistical evidence indicates that it corresponded exactly to the October 1979 change in the monetary policy regime. These results point at Paul Volkcker's change to a disinflationary monetary policy regime as a major factor causing the high level of real rates.

[5] See Blanchard and Summers, Evans

[6] See Malliaropulos (2000)

Truly convincing evidence that the Federal Reserve's monetary policy regime change led to high real interest rates must involve examination of a similar case. Examination of another episode[7] of a monetary regime shift in 1920 has many similarities to the October 1979 shift. At the beginning of 1920, the pursuit of the real bills doctrine by the Federal Reserve led to rapid money supply growth, a sustained high level of inflation with double-digit levels similar to that of 1970's, and a weak dollar. In January and June 1920, the Federal Reserve decided to reverse its inflationary monetary policy by raising the discount rate sharply. In the early years of the Federal Reserve System, changing the discount rate was the main tool of monetary policy, and it was particularly potent at this time because the total amount of member bank borrowing from the Federal Reserve exceeded the amount of non-borrowed reserves. This disinflation was similar to the one in the early 1980's.

The empirical analysis of the period surrounding 1920 reveals a significant shift in the stochastic process of real interest rates, which has many similarities to the experience of 1980s. Not only is the shift significant, but also the dating of the shift is coincident with the Federal Reserve Board's taking action to raise the discount rate in June 1920. Furthermore, the 1920 monetary regime change and the subsequent disinflation are associated with a strengthening of the correlation of nominal interest rates and real interest rates and a shift to a sustained higher level off real interest rates.[8] According to Mishkin the striking correspondence between the impact of monetary regime shifts on real interest rates in 1920 and 1979 provides strong support for the view the shift in real rate behavior after October 1979 is a monetary phenomenon. Particularly important in this regard is that high budget deficits were not a feature of the

[7] See Mishkin (1988)

[8] The 1920 regime shift is also associated with a weakening of the Fisher effect, just as occurred after October 1979.

1920's[9] thus suggesting that monetary factors are more important than budget deficits to the behavior of real interest rates.

An alternative explanation is offered in *Shoven et. al. (1991)* based on a decade long crisis in the savings and loan markets and the Federal Governments efforts to handle this.

Whatever the exogenous events associated with the changes in the real interest rate, we treat the break point as an unknown parameter, which is identified, endogenously, from the data by using the sequential unit-root testing endogenizing the break point (see Zivot and Andrews (1992))

In the next pages we give the charts of our data and then we analyze the methodology being used for testing unit roots in our series.

[9] Although the federal government ran substantial budget deficits in the years 1917-1919 as a result of World War I, there were budget surpluses in every year from 1920 to 1929.

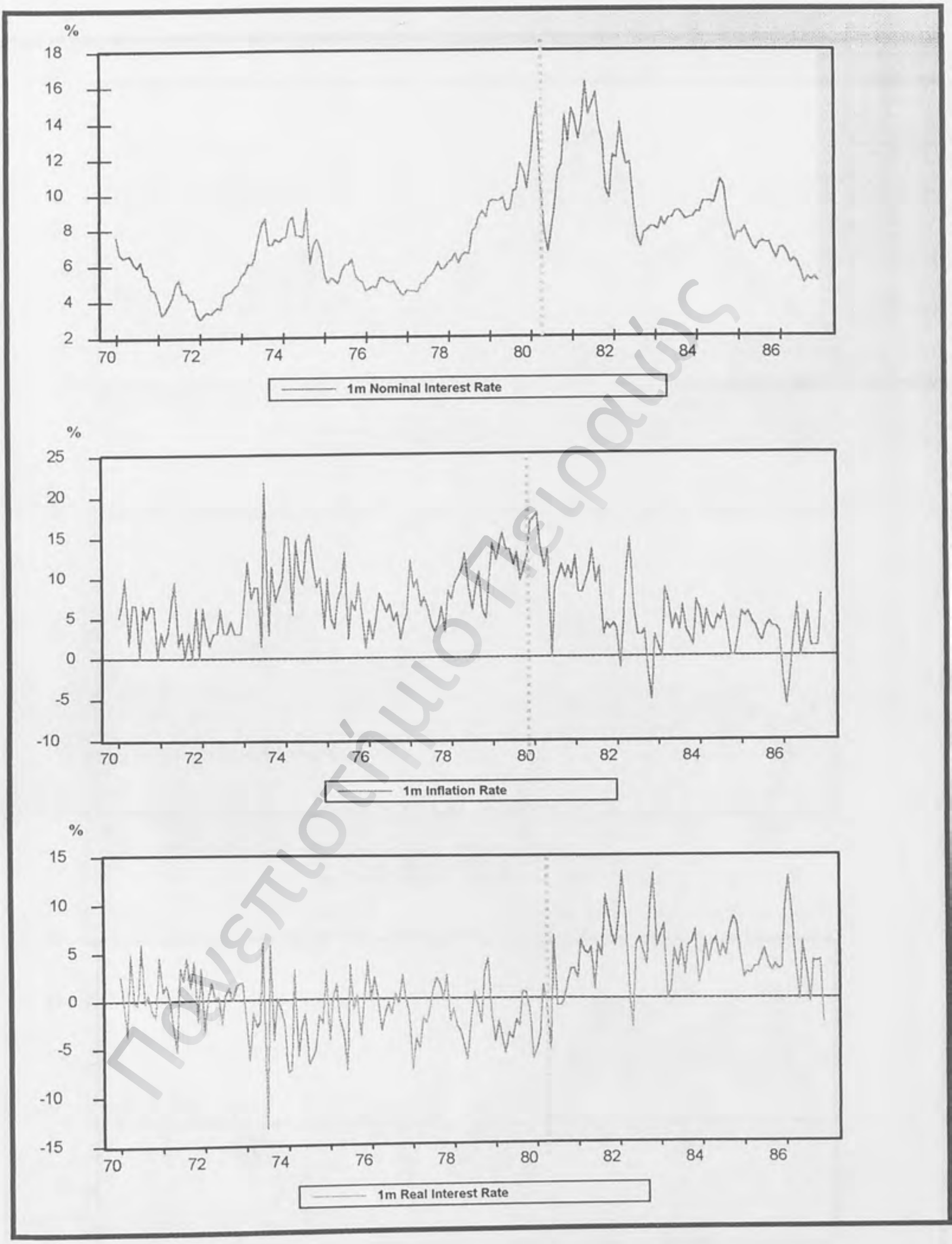


CHART 1: One-Month data

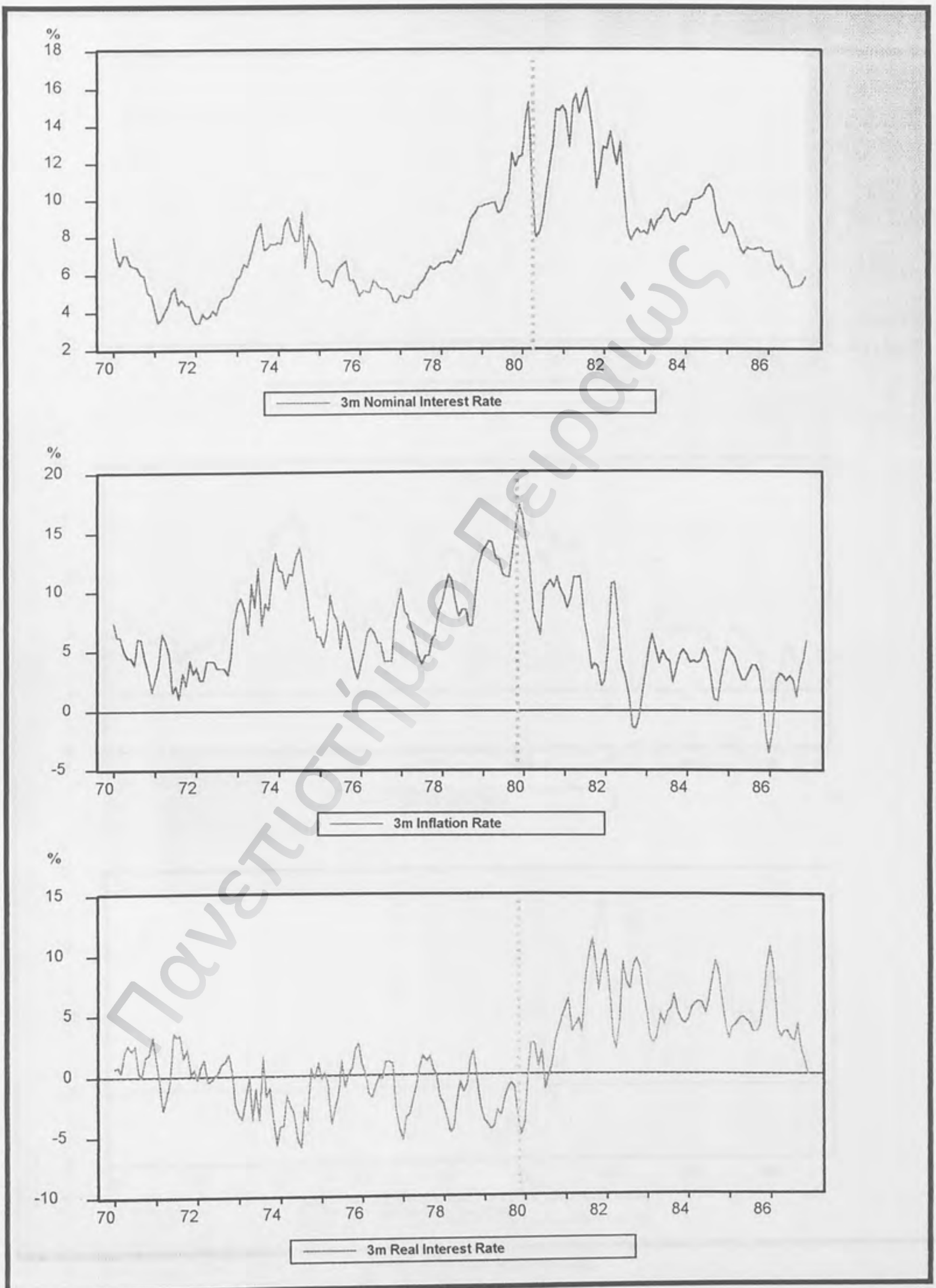


CHART 2: Three-Month data

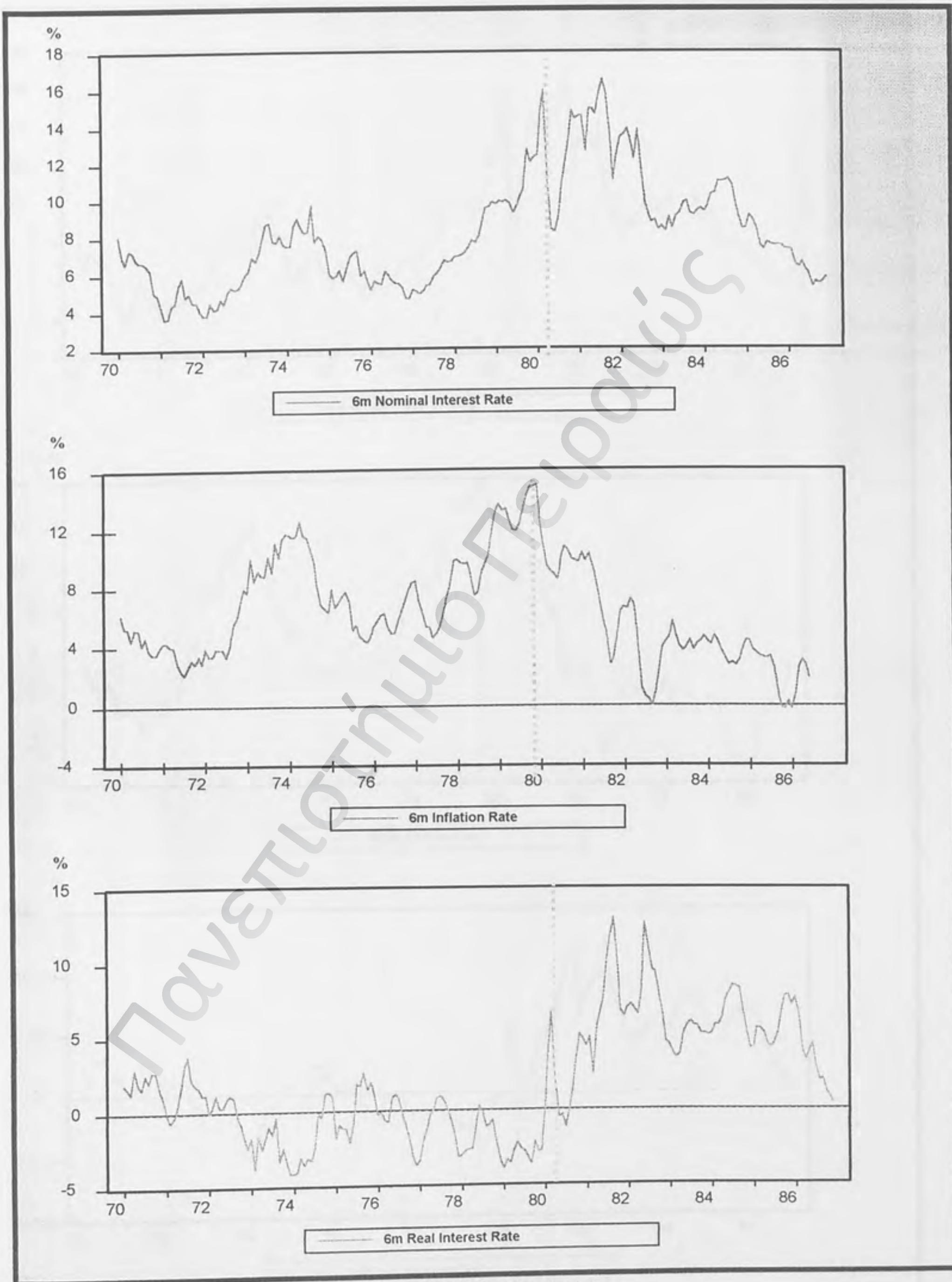


CHART 3: Six-Month data

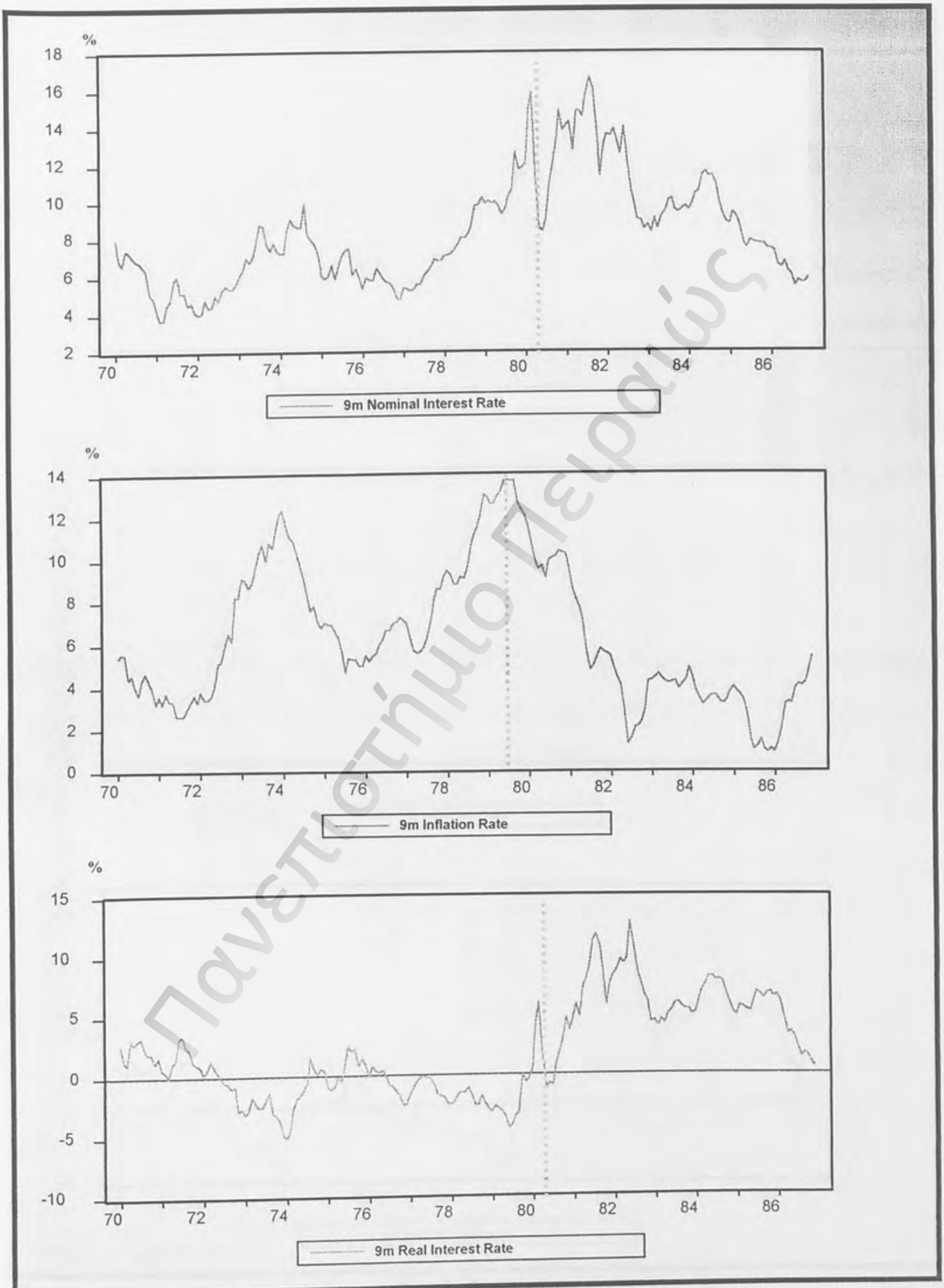


CHART 4: Nine-Month data

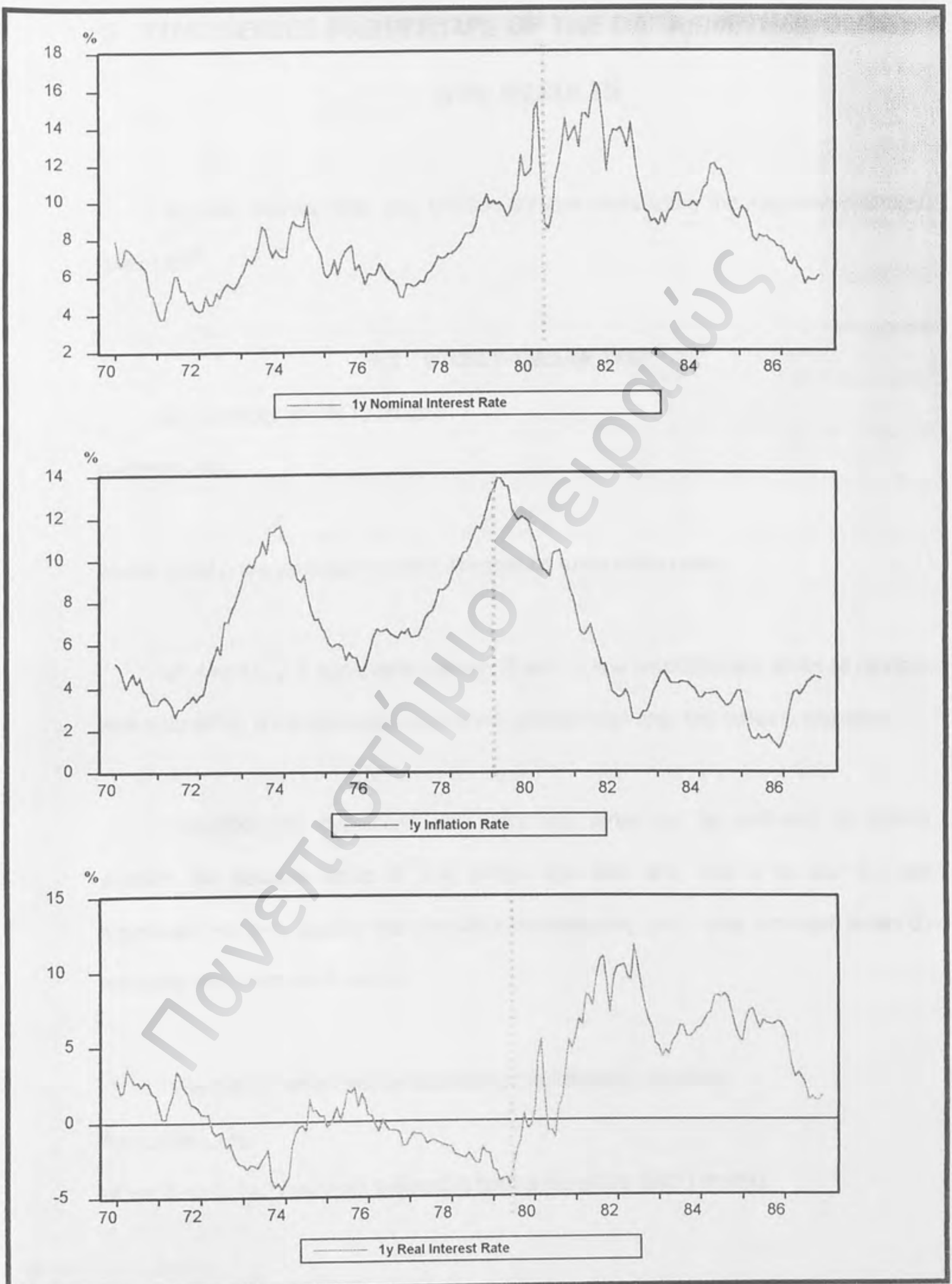


CHART 5: One-Year data

5 TIME SERIES PROPERTIES OF THE DATA: METHODOLOGY AND RESULTS

We test interest rates and inflation for unit roots using the Augmented Dickey Fuller test[10].

5.1 DICKEY-FULLER TEST

Let's consider an AR(1) process:

$y_t=\mu+\rho y_{t-1}+e_t$,

where μ and ρ are parameters and e_t is assumed to be white noise.

If $-1<\rho<1$, y is a stationary series. If $\rho=1$, y is a nonstationary series (a random walk with drift). If the absolute value of ρ is greater than one, the series is explosive.

Therefore, the hypothesis of a stationary series can be evaluated by testing whether the absolute value of ρ is strictly less than one, that is to test the null hypothesis Ho: $\rho=1$, against the one-sided alternative H_1: $\rho<1$, since explosive series do not make much economic sense.

The test is carried out by estimating the following equation:

$\Delta y_t=\mu+\beta y_{t-1}+e_t$,

where $\beta=\rho-1$, (y_{t-1} has been subtracted both sides of the AR(1) model)

[10] See Dickey-Fuller(1979)

The null and alternative hypotheses are:

Ho: $\beta=0$

H_1: $\beta<0$.

This test cannot be carried out by performing a t-test on the estimated β, because the t-statistic under the null hypothesis of a unit root does not have the conventional t-distribution. Dickey and Fuller (1979) showed that the distribution under the null hypothesis is non-standard, and simulated the critical values for selected sample sizes. More recently, MacKinnon (1991) has implemented a much larger set of simulations than those tabulated by Dickey and Fuller. In addition, MacKinnon estimates the response surface using the simulation results, permitting the calculation of Dickey-Fuller critical values for any sample size and for any number of right-hand variables. E-Views reports these MacKinnon critical values for unit root tests.

5.2 AUGMENTED DICKEY-FULLER TEST

The simple unit root test described above is valid only if the series is an AR(1) process. If the series is correlated at higher order lags, the assumption of white noise disturbances is violated. The ADF test makes a parametric correction for higher-order correlation by assuming that the y series follows an AR(p) process and adjusting the test methodology.

The ADF approach controls for higher-order correlation by adding lagged difference terms of the dependent variable y to the right-hand side of the regression:

$$\Delta y_t=\mu+\beta y_{t-1}+ \delta_1\Delta y_{t-1}+ \delta_2\Delta y_{t-2} + \ldots + \delta_{p-1}\Delta y_{t-p+1} + e_t.$$

This augmented specification is then used to test:

H_o: $\beta=0$

H_1: $\beta<0$

in this regression. An important result obtained by Fuller is that the asymptotic distribution of the t-statistic on β is independent of the number of lagged first differences included in the ADF regression. Moreover, while the parametric assumption that y follows an auto-regressive process may seem restrictive, Said and Dickey (1984) demonstrate that the ADF test remains valid even when the series has a moving average (MA) component, provided that enough lagged difference terms are added to the regression. Apart from specifying the number of lagged differenced terms, we also have to decide whether to include a constant, a trend or both in the augmented regression. The null hypothesis of a unit root is rejected against the one-sided alternative if the t-statistic is less than (lies to the left of) the critical value.

5.2.1 CHOICE OF OPTIMAL LAG-LENGTH

We will use the Akaike[11] Information Criterion (AIC) in order to determine the optimal lag-length of the ADF regression equation. The model we will select will be the one with the smallest information criterion.

[11] See Akaike (1970)

5.3 ADF RESULTS OF THE TIME SERIES

Augmented Dickey-Fuller tests for unit roots are reported in table 5.1.3.1. The results of these test suggest that all the series are I(1) processes. The ADF t-statistics drive us to the conclusion that the nominal interest rate for different maturities, the real interest rate and the CPI inflation rate in USA for the examined period contain a unit root and have zero drift.

VARIABLE	t_{μ}*	t_{τ}	t	LAGS**
i1m	0,93923	-0,28004	-0,9556	0
i3m	-2,2370	-2,38907	-0,61708	17
i6m	-2,06149	--2,31316	-0,59782	17
i9m	-1,20493	-1,48016	-0,53063	8
i1y	-1,40303	-1,7886	-0,46505	11
π1m	-1,88491	-1,73604	-0,91521	8
π3m	-3,08402	-2,85083	-0,93509	19
π6m	-3,03893	-2,39309	-1,02634	19
π9m	-2,69937	-2,66025	-0,70242	20
π12m	-2,64103	-2,25561	-0,91118	13
r1m	-2,29213	-1,62539	-1,51439	8
r3m	-2,34431	-1,65412	-1,52572	6
r6m	-1,90700	-1,40549	-1,29114	6
r9m	-2,13280	-1,72088	-1,5783	11
r1y	-2,28945	-1,77759	-1,61033	5
5% cv	**-3.4212**	**-2.8679**	**-1.9401**	

Table 5.3.1

Notes

*Augmented Dickey-Fuller tests are shown in the columns headed t_{μ}(with constant and trend), t_{τ} (with constant only), t (no constant, no trend).

** The optimal number of lags is indicated by Akaike Criterion (AIC)

However, the evidence that the real return of different periods(1m, 3m, 6m, 9m, 1y) *is stochastic non-stationary may lack an economic interpretation.* Economic theory suggests (for instance Rose (1988)) that the one-period real interest rate equals the consumption growth rate which by implication is stationary since the level of consumption is an I(1) process. The nonstationarity of the real interest rate could lead to the rejection of some equilibrium asset pricing models such as the consumption CAPM. Also, the widely used Black-Scholes formula for pricing options is based on the assumption of a constant ex-ante real interest rate. *Garcia and Perron (1996)* provided a statistical description of the ex-post real interest rate that allows nonstationarity in the form of infrequent changes in mean and variance. Such structural changes are important in characterizing the ex-post real interest rate, as noted by *Perron (1990)*. Their results support Fama's original characterization of the ex-ante real interest rate as essentially constant with, however, the crucial difference that the mean of the series is subject to occasional shifts.

The finding of non-stationary real return[12] may be due to the fact that the unit root tests do not consider a changing drift component of the real interest rate related to the monetary regime changes and thus they are biased towards nonrejection. To remove this type of bias, Perron suggested to include in the tests dummy variables conditional on structural changes at a known point in time. Although it seems reasonable to regard the date of the announcement by the Fed of the new operating procedures, in October 1979, as known, it may appear more plausible to treat the break point as unknown. If one takes the view that economic agents need time to learn about the new operating procedures there will be a learning transmission period in which the agents form expectations about the future variables weighting still the past[13].

[12] See Perron (1990)
[13] See Lewis (1989)

A class of unit-root tests endogenizing the break point selection procedure has been suggested by *Zivot and Andrews (1992)*. These tests are described in the next paragraph (5.4).

5.4 SEQUENTIAL UNIT ROOT TESTS OF ZIVOT & ANDREWS

Next, we apply the sequential ADF test of Zivot and Andrews (1992), which accounts for structural breaks in the data with endogenous timing. Under the null, these tests consider a unit-root process with a mean or drift or mean and drift that excludes any structural change. Under the alternative, they assume a stationary process with a one-time shift in the mean or the drift parameter or both mean and drift of unknown date. The selection procedure of the break point is based on the minimum value, over all possible break points, of the relevant, unit root t-statistic, $Inft_a(\lambda)$, where λ denotes the break function of the sample. This test can be conducted using three variants of the alternative hypothesis:

Model A: The series is trend-stationary with a break in the mean.

Formally, the sequential ADF test equation of model A reads:

$$y_t = b_0 + b_1 t + b_2 D_{t,\tau} + a y_{t1} + \Sigma c_i \Delta y_{ti} + u_t$$

where

- $D_{t,T}$ is a dummy variable that takes the value of zero (0) when $t < T\tau$ and one(1) else,
- T is the number of observations and
- $\tau \in (0,1)$ is the relative timing of the structural break.

Model B: The series is trend-stationary with a break in the drift rate of the trend.

Formally, the sequential ADF test equation of model b reads:

$y_t = b_0 + b_1 t + b_3 D_{t,T} t + a y_{t1} + \Sigma c_i \Delta y_{ti} + u_t$

Model C: The series is trend-stationary with a break in both the mean and the drift rate of the trend.

Formally, the sequential ADF test equation of model C reads:

$y_t = b_0 + b_1 t + b_2 D_{t,T} + b_3 D_{t,T} t + a y_{t1} + \Sigma c_i \Delta y_{ti} + u_t$

It is obvious that model A (the case considered in Evans and Lewis) assumes $b_3=0$ and model B assumes $b_2=0$, i.e both models are nested within model C. The null hypothesis in all three models is that the series has a unit root, i.e $a=0$.

5.5 ZIVOT- ANDREWS TEST: Results

The unit root hypothesis can not be rejected at the 5% level against the alternative of stationarity with a shift in both mean and drift for one, three, six, nine and twelve months inflation, nominal and real interest rates. Based on the lowest value of the t-statistic the examined series are trend-stationary with a break in both the mean and the drift rate of the trend. Nine-month inflation rate was found to have a shift only in the drift in 1980:12. The corresponding t- statistic is -4.455710, and the c.v is –4,42. The 9m real interest rate is found to be stationary with a shift in the mean, with the break occurring in 1980:07. The computed t-statistic is -5.01746 (the 5% c.v. is –4.80). The nominal interest rates of 1m and 9m are found to be non-stationary stochastic processes. The relative timing for the structural break for our series and the lowest t-statistics (Model C) are reported in the following table (5.4.1), and then the figures of sequential unit root tests are given.

	t-statistic	Tt
π(1m)	-5,18766*	1981:08
r(1m)	-5,92128*	1980:09
π(3m)	-6,71173*	1981:06
i(3m)	-5,23917*	1980:06
r(3m)	-5,23984*	1981:08
π(6m)	-8,47884*	1980:11
i(6m)	-5,17649*	1980:03
r(6m)	-5,63199*	1981:03
π(1y)	-6,22368*	1981:02
i(1y)	-8,19897*	1980:03
r(1y)	-12,2865*	1980:11
Critical value	*-5,08*	

Table 5.5.1

Note

* significant at the 5% and 10% significance level

SEQUENTIAL UNIT ROOT TESTS: FIGURES

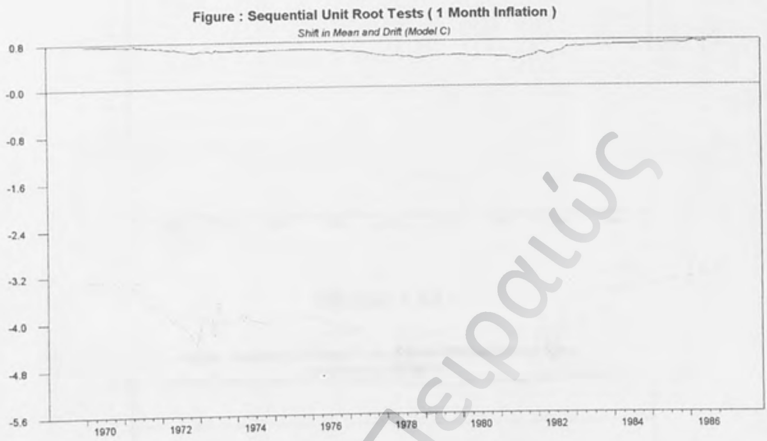


FIGURE 5.5.1

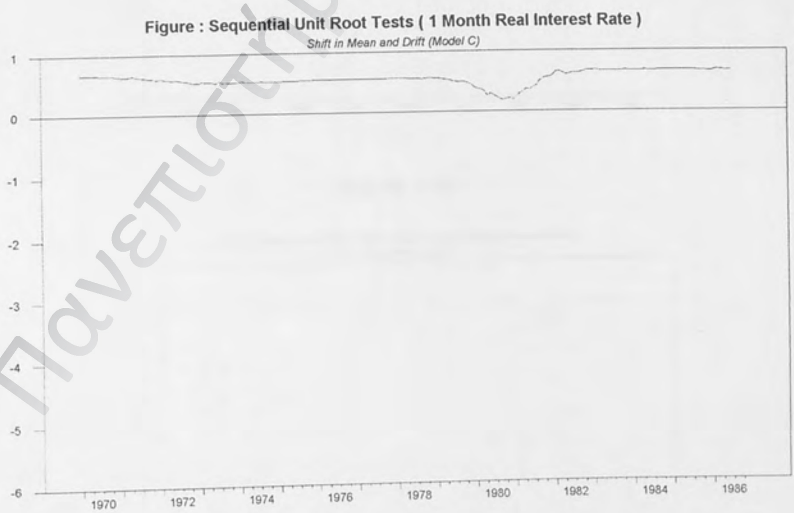


FIGURE 5.5.2

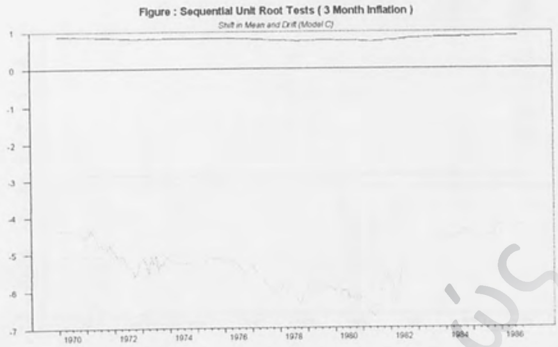


FIGURE 5.5.3

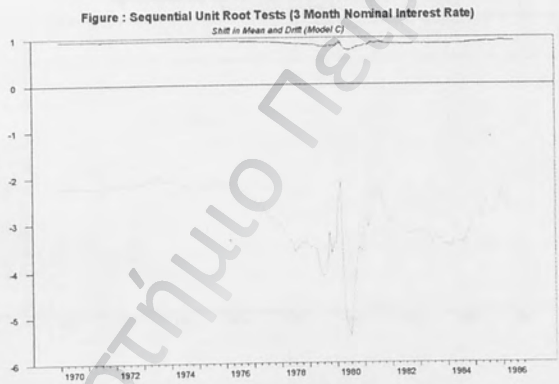


FIGURE 5.5.4

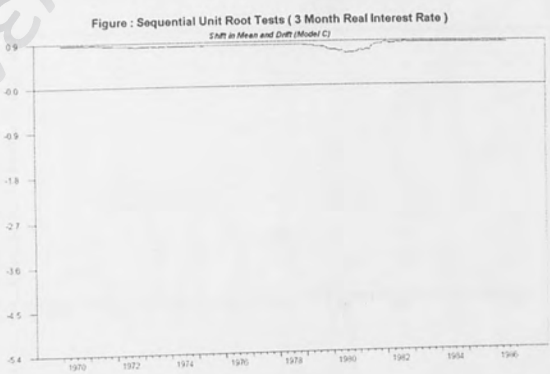


FIGURE 5.5.5

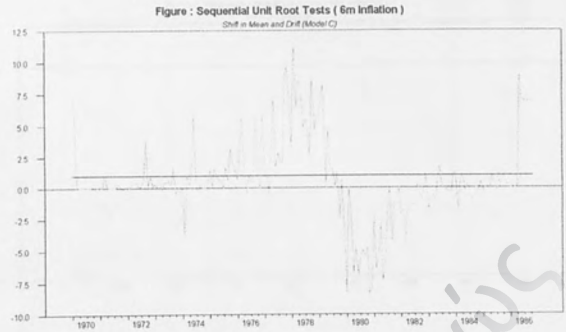


FIGURE 5.5.6

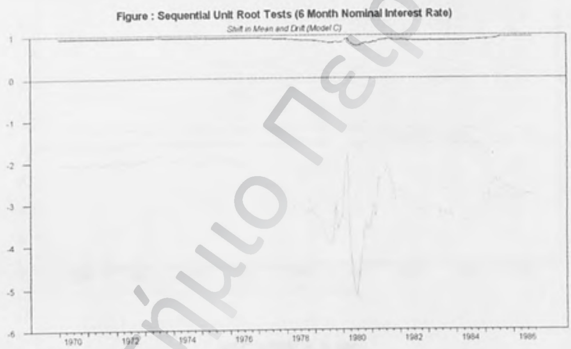


FIGURE 5.5.7

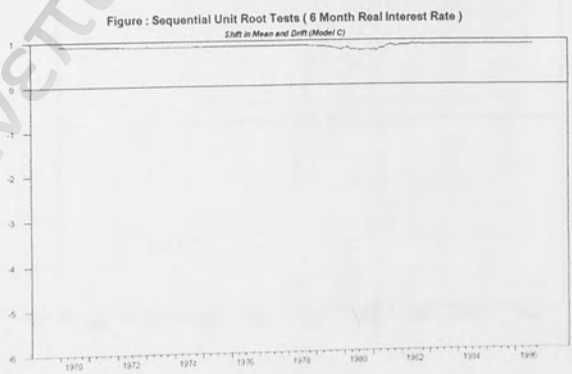


FIGURE 5.5.8

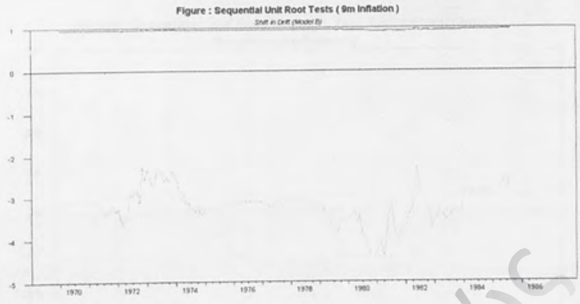


FIGURE 5.5.9

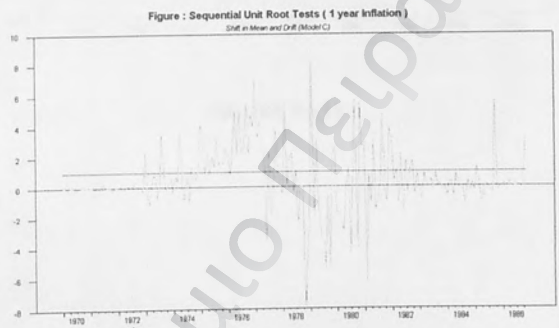


FIGURE 5.5.10

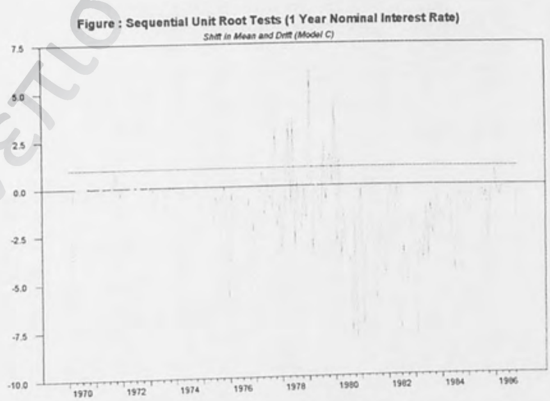


FIGURE 5.5.11

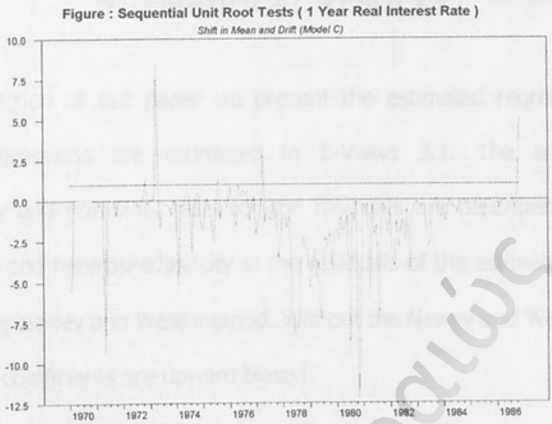


Figure : Sequential Unit Root Tests (1 Year Real Interest Rate)
Shift in Mean and Drift (Model C)
10.0
7.5
5.0
2.5
0.0
-2.5
-5.0
-7.5
-10.0
-12.5
1970
1972
1974
1976
1978
1980
1982
1984
1986

FIGURE 5.5.12

6 EMPIRICAL RESULTS

In this section of our paper we present the estimated regressions and their results. The regressions are estimated in E-Views 3.1. The serial correlation, heteroskedasticity and normality tests for the residuals are described in Appendix D. Serial correlation and heteroskedasticity in the residuals of the estimated regressions is corrected with the Newey and West method. Without the Newey and West correction the t-statistics of the coefficients are upward biased.

6.1 1 MONTH DATA: EMPIRICAL RESULTS

The estimated regression for one-month data is

$$\Delta\pi_t = \alpha + \beta_0 D + \beta_1 \Delta i_{t-1} + \beta_2 r_{t-1} + u_t$$

where $\Delta\pi_t$: the change of the CPI inflation rate from period t-1 to period t,

Δi_{t-1}: the change of one-month nominal interest rate from period t-2 to period t-1, which is known,

r_{t-1}: the one-month real interest rate of period t-1[14]

D: a dummy variable that takes the value of zero from 1970:01 to 1981:07, and the value of 1 from 1981:08 to 1986:12 (1981:08 is the relative timing of the structural break in both the mean and the drift of the one-month inflation rate).

The estimated residuals exhibit heteroskedasticity (Arch Effects) and serial correlation, and are corrected with Newey and West (see Newey and West (1987). They are also non-normal. We observe that there is a strong positive relationship between the lagged real interest rate and the change of the inflation rate. The estimated coefficient of r(t-1) is positive, substantially high (0,746041) and statistically significant at the 1% level. This is in line with our model and confirms that a time varying short-term real interest rate reflects inflation fluctuations. The R^2 is quite high (35,24%) and indicates the close relationship of these two variables. The change in the one-month nominal rate seems to have no explanatory force in predicting inflation and is statistically insignificant. The coefficient of the estimated dummy variable is significant at the 1% level, showing

[14] The lag of real interest rate that is implied from our model is lag (t-2) for 1-month data, (t-5) for 3-month data, (t-8) for 6-month data, (t-11) and (t-14) for 9 and 12-month data respectively. Our research uncovered that the relationship that exists and is investigated is between $\Delta\pi_t$ and the lag (t-1), (t-3), (t-6), (t-9), (t-12) of real interest rate for 1,3,6, 9,12 month data respectively. Although the real interest rate and $\Delta\pi$ are not contemporaneous, they overlap partly, i.e. by one month.

the impact of the change of monetary regime in the fluctuations of inflation. Results of the above regression are reported in table 6.1.1.

Dependent Variable Δπ(t)	Independent variables			
	constant	D	Δi(t-1)	r(t-1)
coefficients	0.6213	-4.5296	-0.14139	0.74604
t-statistic	1.7960***	-5.0568*	-0.6111	5.9163*

*significant at the 1%, 5%, and 10% level ***significant at the 10% level

R*2	D.W	White Heteroskedasticity Test	ARCH(1)	Residuals Normality Test
0.352487	2.049431	Obs*R-squared 26,464 (0,000874)	Obs*R-squared 10,292 (0.000137)	20,491 (0,0003)

Note: The numbers in parentheses are probabilities.

Breusch-Godfrey Serial Correlation LM Test			
Lags	2	6	10
Obs*R-squared	6.494896	7.634921	15.61384
Prob.	0.038873	0.266088	0.111233
RESIDUALS AUTOCORRELATIONS			
LAG	AUTOCORRELATION	Q-Stat	Prob
1	-0.038	0.2885	0.591
2	0.174	6.5247	0.038
3	0.03	6.7163	0.082
4	0.033	6.9472	0.139
5	0.043	7.3329	0.197
6	-0.042	7.6963	0.261
7	0.123	10.877	0.144

Table 6.1.1

We have also regressed the change of the inflation rate on the spread of ten-year nominal interest rate[15] and one-month nominal interest rate at lag (t-1). Our results

[15] We have used a 10 year US Government Composite Bond. This series has been downloaded from Datastream and its code is USOCLNG%.

(table 6.1.2) suggest that the coefficient of this spread (0,10438) is not statistically significant. The ADF tests prove that this spread is stationary[16]. Autocorrelation in the residuals and heteroskedasticity (Arch Effects) are corrected non-parametrically with Newey and West.

Dependent Variable Δπ(t)	Independent variables				
	constant	D	Δi(t-1)	r(t-1)	i10y-i1m (t-1)
coefficients	0.6786	-4.3296	-0.1865	0.745255	0.1043
t-statistic	1.6441	-4.8478*	-0.7641	5.8340*	0.47811

*Significant at 1%, 5% and 10% level

R*2	D.W	White Heteroskedasticity Test	ARCH(1)	Residuals Normality Test
0.353567	2.041393	Obs*R-squared 34,38 (0,001053)	12,20579 (0.000587)	18,428 (0,0001)

Note: The numbers in parentheses are probabilities.

Breusch-Godfrey Serial Correlation LM Test			
Lags	2	6	10
Obs*R-squared	6.49768	7.825046	15.81864
Prob.	0.038819	0.251203	0.104943

RESIDUALS AUTOCORRELATIONS

LAG	AC	Q-Stat	Prob
1	-0.034	0.2306	0.631
2	0.175	6.5105	0.039
3	0.032	6.7251	0.081
4	0.034	6.9612	0.138
5	0.044	7.3694	0.195
6	-0.043	7.7670	0.256
7	0.122	10.938	0.141
8	0.015	10.987	0.202
9	0.181	17.987	0.035
10	-0.008	18.000	0.055

Table 6.1.2

[16] ADF tests for the nominal interest rate spreads are given in Appendix C.

6.2 3 MONTH DATA: EMPIRICAL RESULTS

The estimated regression for three-month data is

$$\Delta\pi_t = \alpha + \beta_0 D + \beta_1 \Delta i_{t-1} + \beta_2 r_{t-3} + \beta_3 ma(1) + \beta_4 ma(2) + u_t$$

Where $\Delta\pi_t$: the change of the CPI inflation rate from period t-1 to period t,

Δi_{t-1}: the change of three-month nominal interest rate from period t-2 to period t-1,

r_{t-3}: the three-month real interest rate of period t-3,

ma (i) : moving average terms of order i, where i=1 to 2

D: a dummy variable that takes the value of zero from 1970:01 to 1981:05, and the value of 1 from 1981:06 to 1986:12 (1981:06 is the relative timing of the structural break in both the mean and the drift of the quarterly inflation rate).

The exhibition of serial correlation in the residuals is corrected with Newey and West, in order to obtain the appropriate t-statistics of the coefficients. We also observe that the estimated residuals are non-normal, and that there is no heteroskedaticity. Results of the above regression are reported in table 6.2.1.

The estimated results suggest that there is a strong relationship between the real interest rate and the change in the inflation rate. The estimated coefficient of r_{t-3} is 0,25569 and is significantly different from zero at the 1% significance level. So we can say that the lagged real return is an important variable to forecast inflation. The higher the real return, the larger the predicted increase in the inflation rate. Moreover the R^2 (21%) is quite satisfactory if we take into account the fact that the model is expressed in first differences, but is relatively lower than in one-month data. The included dummy variable, which denotes the change in the monetary regime, is significant, but the change of the nominal interest rate seems not to be a serious determinant of the future change in the CPI inflation rate.

Dependent Variable Δπ(t)	Independent variables					
	constant	D	Δi(t-1)	r(t-3)	MA(1)	MA(2)
coefficients	0.216428	-1.68202	0.210407	0.25569	0.261998	0.239382
t-statistic	1.453277	-3.9677*	1.73374**	4.5792*	2.5970**	2.4815**

*significant at the 1%, 5%, and 10% level ** significant at the 5% and 10% level
***significant at the 10% level

R^*2	D.W	White Heteroskedasticity Test	ARCH(1)	Residuals Normality Test
0.206846	2,199688	Obs*R-squared 9.1332 (0.3311)	Obs*R-squared 7,842 (0,005)	6.32252 (0,035093)

Note: The numbers in parentheses are probabilities.

Breusch-Godfrey Serial Correlation LM Test			
Lags	2	6	10
Obs*R-squared	35.702	53.44442	56.61507
Prob.	0	0	0

RESIDUALS AUTOCORRELATIONS

LAGS	AC	Q-Stat	Prob
1	-0.103	2.0904	
2	-0.065	2.9148	
3	-0.312	22.248	0.000
4	0.198	30.015	0.000
5	0.151	34.580	0.000
6	-0.185	41.450	0.000
7	0.034	41.685	0.000
8	0.023	41.789	0.000
9	0.140	45.790	0.000
10	-0.031	45.988	0.000

Table 6.2.1

We have also regressed the change of the inflation rate on the spread of ten-year nominal interest rate and three-month nominal interest rate at lag (t-3). According to the ADF tests this spread is stationary.

Our results suggest that this spread is statistically significant only at the 10% level and does not determine the future path of inflation rate in the extent that the real interest rate does. The dummy variable is again significant at the 1% significance level and the change of the nominal interest rate at lag (t-1) contains no predictive power for the future inflation. The ma terms are all significant. Autocorrelation and Arch effects are taken into account by correcting non-parametrically with Newey and West.

These findings support our model and is line with Wickens and Tzavalis (1995) who pointed out that more information about inflation is contained in the real return and that the term structure is less useful for forecasting inflation than information about the dynamic adjustment of real interest rates to current and past monetary shocks.

The results of the above regression are reported in table 6.2.2.

Dependent Variable Δπ(t)	Independent variables						
	constant	D	Δi(t-1)	r(t-3)	i10y-i3m (t-3)	MA(1)	MA(2)
coefficients	0.179803	-1.984089	0.183966	0.264865	0.161030	0.242838	0.215698
t-statistic	1.120388	-4.9384*	1.51288	5.16315*	1.69821***	2.40372**	2.18748**

*significant at the 1%, 5%, and 10% level ** significant at the 5% and 10% level
***significant at 10% level

R*2	D.W	White Heteroskedasticity Test	ARCH (1)	Residuals Normality Test
0.218384	2.180065	Obs*R-squared 0.2682 (0,000)	Obs*R-squared 7.950924 (0.0048)	5.901632 (0,52297)

Note: The numbers in parentheses are probabilities.

Breusch-Godfrey Serial Correlation LM Test			
Lags	2	6	10
Obs*R-squared	36.1156	55.32981	58.73281
Prob.	0	0	0

RESIDUALS AUTOCORRELATIONS

LAGS	AC	Q-Stat	Prob
1	-0.094	1.7281	
2	-0.059	2.4077	
3	-0.316	22.247	0.000
4	0.209	30.932	0.000
5	0.167	36.544	0.000
6	-0.167	42.134	0.000
7	0.044	42.520	0.000
8	0.011	42.546	0.000
9	0.137	46.404	0.000
10	-0.044	46.799	0.000

Table 6.2.2

6.3 6 MONTH DATA: EMPIRICAL RESULTS

The estimated regression for six-month data is

$$\Delta\pi_t = \alpha + \beta_0 D + \beta_1 \Delta i_{t-1} + \beta_2 r_{t-6} + \beta_3 ma(1) + \beta_4 m\alpha(2) + \beta_5 ma(3) + \beta_6 ma(4) + \beta_7 ma(5) + u_t$$

where

$\Delta\pi_t$: the change of the CPI inflation rate from period t-1 to period t,

Δi_{t-1}: the change of the six-month nominal interest rate from period t-2 to period t-1,

r_{t-6}: the six-month real interest rate of period t-6,

ma(i) : moving average terms of order i, where i=1 to 5

D: a dummy variable that takes the value of zero from 1970:01 to 1980:10, and the value of 1 from 1980:11 to 1986:12 (1980:11 is the relative timing of the structural break in both the mean and the drift of the six-month inflation rate).

The exhibition of autocorrelation and ARCH effects in the residuals is corrected non-parametrically with Newey and West HAC, in order to obtain the appropriate t-statistics of the estimators of the independent variables.

The estimated results suggest that there is no significant positive relationship between the real interest rate of the lag (t-6) and the change in the six-month inflation rate. The estimated coefficient of r_{t-6} (-0.00958) is not significantly different from zero. Moreover it is not positive as indicated by our model. We see that these results are quite different with results for short term interest rates and inflation (1m, 3m) and that the real interest rates do not contain substantial information for the future path of inflation rate when we examine data with longer maturities. Results of the above regression are reported in the following table (6.3.1).

Dependent Variable Δπ(t)	Independent variables								
	constant	D	Δi(t-1)	r(t-6)	MA(1)	MA(2)	MA(3)	MA(4)	MA(5)
coefficients	0.389	-0.278	-0.0330	-0.00958	0.494	0.862	0.535	0.883	0.49
t-statistic	1.848***	-1.010	-0.5806	-0.2505	24.771*	44.295*	15.392*	24.860*	12.54*

*significant at the 1%, 5%, and 10% level
***significant at the 10% level

R*2	D.W	White Heteroskedasticity Test	ARCH (1)	Residuals Normality Test
0.354129	2.259945	Obs*R-squared 1,03 (0,9980)	Obs*R-squared 6,168 (0,013)	18.72589 (0,0001)

Note: The numbers in parentheses are probabilities.

Breusch-Godfrey Serial Correlation LM Test			
Lags	2	6	10
Obs*R-squared	14.5418	41.29444	44.02627
Prob.	0.000695	0	0.000003

RESIDUALS AUTOCORRELATIONS

LAGS	AC	Q-Stat	Prob
1	-0.154	4.6097	
2	-0.174	10.566	
3	-0.032	10.765	
4	-0.088	12.310	
5	0.045	12.708	
6	-0.240	24.206	0.000
7	0.100	26.204	0.000
8	-0.014	26.245	0.000
9	0.100	28.295	0.000
10	-0.011	28.321	0.000

Table 6.3.1

We have also regressed the change of the inflation rate on the spread of ten-year nominal interest rate and six-month nominal interest rate at lag t-6. We must point out the fact that this spread is proved to be a non-stationary stochastic process. Our results suggest that this spread is statistically significant and does determine the future path of inflation rate. The estimated coefficient is 0.086660 and it is statistically different from zero at the 10% significance level.

Moreover we see that the real interest rate when it is regressed with the spread is statistically significant and has important information for the future path of the inflation rate. The estimated coefficient is 0,053343 and is lower than the coefficients estimated for one and three month data.

Results of the above regression are presented in the following table (6.3.2).

Dependent Variable Δπ(t)	Independent variables									
	constant	D	Δi(t-1)	r(t-6)	i10y-i6m (t-6)	MA(1)	MA(2)	MA(3)	MA(4)	MA(5)
coefficients	0.0672	-0.578	0.0075	0.0533	0.086	0.241	0.341	-0.506	0.076	-0.0906
t-statistic	0.898	-2.721*	0.108	1.854***	1.948***	2.601**	5.320*	-10.90*	0.861	-1.516

*significant at the 1%, 5%, and 10% level
** significant at the 5% and 10% level
***significant at 10% level

R*2	D.W	White Heteroskedasticity Test	ARCH (1)	Residuals Normality Test
0.264359	1.897047	Obs*R-squared 14,751 (0,323)	Obs*R-squared 8,029 (0,004)	4,138 (0,126)

Note: The numbers in parentheses are probabilities.

Breusch-Godfrey Serial Correlation LM Test			
Lags	2	6	10
Obs*R-squared	9.43087	54.23001	63.72168
Prob.	0.008956	0	0

RESIDUALS AUTOCORRELATIONS

LAGS	AC	Q-Stat	Prob
1	0.050	0.4787	
2	-0.037	0.7525	
3	0.170	6.4314	
4	-0.076	7.5832	
5	-0.133	11.083	
6	-0.379	39.810	0.000
7	-0.025	39.938	0.000
8	0.052	40.494	0.000
9	0.083	41.897	0.000
10	0.076	43.068	0.000

Table 6.3.2

6.4 9 MONTH DATA: EMPIRICAL RESULTS

The estimated regression for nine-month data is

$\Delta\pi_t = \alpha + \beta_0 D + \beta_1 \Delta i_{t-1} + \beta_2 r_{t-9} + \beta_3 ma(1) + \beta_4 ma(2) + \beta_5 ma(3) + \beta_6 ma(4) + \beta_7 ma(5) + \beta_8 ma(6) + \beta_9 ma(7) + \beta_{10} ma(8) + u_t$

where

$\Delta\pi_t$: the change of the CPI inflation rate from period t-1 to period t,

Δi_{t-1}: the change of nine-month nominal interest rate from period t-2 to period t-1,

r_{t-9}: the nine-month real interest rate of period t-9,

ma(i) : moving average terms of order i, where i=1 to 8

D: a dummy variable that takes the value of zero from 1970:01 to 1980:11, and the value of 1 from 1980:12 to 1986:12 (1980:12 is the relative timing of the structural break in both the mean and the drift of the nine-month inflation rate).

The exhibition of autocorrelation and Arch Effects in the residuals is corrected with Newey and West HAC, in order to obtain the appropriate t-statistics of the estimators of the independent variables. The estimated results suggest that there is no significant positive relationship between the real interest rate of the lag (t-9) and the change in the nine-month inflation rate. The estimated coefficient of r_{t-9} is 0,30333 and is not significantly different from zero, showing that for longer maturities data real interest rate does not contain great deal of information for future inflation fluctuations.

Results of the above regression are reported in table 6.4.1.

Dependent Variable Δπ(t)	Independent variables											
	constant	D	Δi(t-1)	r(t-9)	MA(1)	MA(2)	MA(3)	MA(4)	MA(5)	MA(6)	MA(7)	MA(8)
coefficients	0.080	-0.339	-0.0107	0.03033	0.347	0.326	0.016	0.219	0.170	-0.098	0.111	0.048
t-statistic	0.959	-1.752	-0.2289	0.99	3.504	5.025*	0.215	3.275*	2.783*	-1.174	1.184	0.721

*significant at the 1%, 5%, and 10% level
** significant at the 5% and 10% level
***significant at the 10% level

R*2	D.W	White Heteroskedasticity Test	ARCH (1)	Residuals Normality Test
0.230846	2.024803	Obs*R-squared 5,653 (0,685)	Obs*R-squared 4,645 (0,031)	17.36767 (0,000169)

Note: The numbers in parentheses are probabilities.

Breusch-Godfrey Serial Correlation LM Test			
Lags	2	6	10
Obs*R-squared	7.47544	13.26568	19.63765
Prob.	0.023808	0.039006	0.032872

RESIDUALS AUTOCORRELATIONS

LAGS	AC	Q-Stat	Prob
1	-0.015	0.0475	
2	-0.028	0.2029	
3	0.024	0.3228	
4	-0.052	0.8708	
5	-0.035	1.1231	
6	0.014	1.1612	
7	-0.070	2.1501	
8	-0.012	2.1801	
9	-0.207	11.060	0.001
10	0.146	15.466	0.000
11	0.019	15.542	0.001
12	0.160	20.903	0.000

Table 6.4.1

We have also regressed the change of the inflation rate on the spread of ten-year nominal interest rate and nine-month nominal interest rate at lag t-9.

This spread, as in the case of the six-month data is non-stationary. Our results suggest that this spread is statistically significant and does determine the future path of inflation rate. The estimated coefficient is 0,009942 and it is statistically different from zero at the 1% significance level. This finding is different from results found for one and three month data, where the term structure does not play any significant role in the fluctuation of the change on the inflation rate. Moreover this is line with Mishkin whose results suggest that the term structure has predictive ability for inflation at longer horizons.

We see that the real interest rate, when it is regressed with the spread is statistically significant and seems to have important information for the future path of the inflation rate. The estimated coefficient is equal to 0,0409 and is substantially lower to those found for one month and three month data.

Results of this regression are reported in the following table (6.4.2). Serial correlation in the residuals is corrected with Newey and West in order to obtain the appropriate t-statistics of the estimators.

Dependent Variable Δπ(t)	Independent variables												
	constant	D	Δi(t-1)	r(t-9)	i10y-i9m(t-9)	MA(1)	MA(2)	MA(3)	MA(4)	MA(5)	MA(6)	MA(7)	MA(8)
coefficients	0.0711	-0.459	0.011	0.0409	0.09942	0.272	0.216	-0.127	0.109	0.059	-0.24	0.013	-0.001
t-statistic	1.282	-3.1168**	0.2313	1.971***	2.891*	2.92*	2.89*	-1.84***	1.61	1.14	-2.69*	0.127	-0.02

*significant at the 1%, 5%, and 10% level
** significant at the 5% and 10% level
***significant at the 10% level

R*2	D.W	White Heteroskedasticity Test	ARCH (1)	Residuals Normality Test
0,253830	1.988530	Obs*R-squared 10,479 (0,654)	Obs*R-squared 3,102 (0,078)	18.28 (0,000107)

Note: The numbers in parentheses are probabilities.

Breusch-Godfrey Serial Correlation LM Test			
Lags	2	6	10
Obs*R-squared	7.11293	13.61242	19.4316
Prob.	0.02854	0.034278	0.035111

RESIDUALS AUTOCORRELATIONS

LAGS	AC	Q-Stat	Prob
1	0.001	1.E-04	
2	-0.004	0.0039	
3	0.056	0.6370	
4	-0.039	0.9430	
5	-0.021	1.0362	
6	0.043	1.4192	
7	-0.069	2.3981	
8	-0.032	2.6031	
9	-0.223	12.918	0.000
10	0.122	16.021	0.000
11	-0.011	16.047	0.001
12	0.141	20.199	0.000

Table 6.4.2

6.5 1 YEAR DATA: EMPIRICAL RESULTS

The estimated regression for one-year data is

$$\Delta\pi_t = \alpha + \beta_0 D + \beta_1 \Delta i_{t-1} + \beta_2 r_{t-12} + \beta_3 ma(1) + \beta_4 ma(2) + \beta_5 ma(3) + \beta_6 ma(4) + \beta_7 ma(5) + \beta_8 ma(6) + \beta_9 ma(7) + \beta_{10} ma(8) + \beta_{11} ma(9) + \beta_{12} ma(10) + \beta_{13} ma(11) + u_t$$

where

$\Delta\pi_t$: the change of the CPI inflation rate from period t-1 to period t,

Δi_{t-1}: the change of one-year nominal interest rate from period t-2 to period t-1,

r_{t-12}: the one-year real interest rate of period t-12,

ma (i) : moving average terms of order i, where i=1 to 11

D: a dummy variable that takes the value of zero from 1970:01 to 1980:08, and the value of 1 from 1980:09 to 1986:12 (1980:09 is the relative timing of the structural break in both the mean and the drift of the annual inflation rate).

The exhibition of autocorrelation in the residuals is corrected with Newey and West method, in order to obtain the appropriate t-statistics of the estimators of the independent variables. The estimated results suggest that there is no significant positive relationship between the real interest rate of the lag (t-12) and the change in the annual inflation rate. The estimated coefficient of r_{t-12} is 0,001984 and is not significantly different from zero. The R^2 is quite high (44%) and may be due to the fact that the ma terms are significant. We see that these results are quite different with results for short term interest rates and inflation (1m, 3m) and that real interest rates do not contain substantial information for the future path of inflation rate when we examine data with longer maturities. Results of the above regression are reported in the following table (6.5.1).

Dependent Variable Δn(t)	Independent variables														
	constant	D	Δi(t-1)	r(t-12)	MA(1)	MA(2)	MA(3)	MA(4)	MA(5)	MA(6)	MA(7)	MA(8)	MA(9)	MA(10)	MA(11)
coefficients	0.040	-0.167755	-0.038	0.001984	0.318	0.522	0.283	0.087	0.139	-0.143	0.151	-0.138	0.334	0.426	0.331
t-statistic	0.603	-1.370063	-1.41	0.117595	3.875*	7.980*	4.129*	1.556	2.282**	-2.590**	2.449**	-2.664*	5.257*	6.679*	4.756*

*significant at the 1%, 5%, and 10% level
** significant at the 5% and 10% level
***significant at the 10% level

R*2	D.W	White Heteroskedasticity Test	ARCH (1)	Residuals Normality Test
0.448413	2.007372	Obs*R-squared 1,87211 (0,9847)	Obs*R-squared 0,939 (0,333)	38.45928 (0,000)

Note: The numbers in parentheses are probabilities.

Breusch-Godfrey Serial Correlation LM Test			
Lags	2	6	10
Obs*R-squared	10.0122	17.31217	24.30842
Prob.	0.006697	0.008202	0.006823

RESIDUALS AUTOCORRELATIONS

LAGS	AC	Q-Stat	Prob
1	-0.006	0.0082	
2	-0.051	0.5184	
3	-0.046	0.9428	
4	0.031	1.1334	
5	0.050	1.6233	
6	0.082	2.9550	
7	-0.020	3.0316	
8	-0.038	3.3261	
9	-0.017	3.3826	
10	-0.041	3.7203	
11	-0.006	3.7284	
12	-0.228	14.492	0.000
13	0.058	15.183	0.001
14	-0.002	15.184	0.002

Table 6.5.1

We have also regressed the change of the inflation rate on the spread of ten-year nominal interest rate and one-year nominal interest rate at lag (t-12). This spread is also non-stationary, as in the case of six and nine month data.

Our results suggest that this spread is statistically significant (the value of the coefficient is 0,106) and does determine the future path of inflation rate, confirming Mishkin's empirical results.

In addition in this regression the lagged real interest rate (0,0033) is proved to be statistically significant at the 5% level and seems to contain useful information for the future path of the inflation rate. The R^2 is relatively high (47%), and slightly higher than when the term structure is not included in the regression. The included dummy variable is significant at the 1% significance level, showing that the regime change in the monetary policy affected the path of inflation. The change of the nominal interest rate at (t-1) seems to have no information for the inflation fluctuations. Most of the ma terms are also statistically significant.

Results of the above regression are reported in table 6.5.2.

Dependent Variable Δπ(t)	Independent variables															
	constant	D	Δi(t-1)	r(t-12)	i10y-i1y (t-12)	MA(1)	MA(2)	MA(3)	MA(4)	MA(5)	MA(6)	MA(7)	MA(8)	MA(9)	MA(10)	MA(11)
coefficients	0.062	-0.358	-0.022	0.033	0.106	0.261	0.469	0.21	-0.03	0.05	-0.27	0.071	-0.23	0.32	0.40	0.26
t-statistic	0.987	-3.664*	-0.8387	2.239**	3.661*	3.414*	7.685*	3.18*	-0.83	0.85	-4.68*	1.126	-4.32*	4.98*	7.55*	3.86*

*significant at the 1%, 5%, and 10% level
** significant at the 5% and 10% level
***significant at the 10% level

R*2	D.W	White Heteroskedasticity Test	ARCH (1)	Residuals Normality Test
0.477504	2.016946	Obs*R-squared 2.436030 (0,999)	Obs*R-squared 0,557 (0,455)	38,46 (0,000)

Note: The numbers in parentheses are probabilities.

Breusch-Godfrey Serial Correlation LM Test			
Lags	2	6	10
Obs*R-squared	11.2674	18.92419	23.4381
Prob.	0.003575	0.004294	0.00924

RESIDUALS AUTOCORRELATIONS

LAGS	AC	Q-Stat	Prob
1	-0.011	0.0237	
2	-0.072	1.0442	
3	-0.063	1.8299	
4	0.040	2.1407	
5	0.027	2.2871	
6	0.079	3.5458	
7	-0.028	3.7023	
8	-0.048	4.1662	
9	-0.065	5.0363	
10	-0.095	6.8711	
11	-0.034	7.1036	
12	-0.245	19.542	0.000
13	0.048	20.020	0.000
14	-0.006	20.027	0.000
15	0.049	20.539	0.000

Table 6.5.2

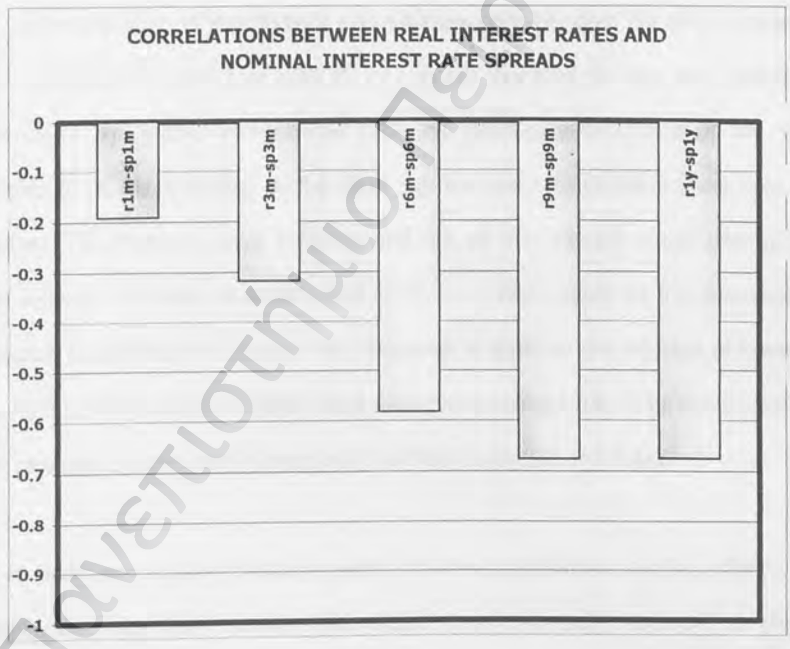
7 INTERPRETATION

From our empirical results, which are presented in section 6, it is obvious that the lagged real interest rate has predictive power for the future inflation for short maturities (1, 3 month), whereas for longer maturities the term structure is proved to contain important information for future inflation. An interesting question is: «Why does the real interest rate predict inflation for short maturities?»

In short periods, shocks affect the nominal interest rates more than inflation. If the nominal money stock is increased, at each price level interest rates are lower and the demand for output rises. The change in the inflation rate, Δπ, changes with a lag due to price stickiness (short-run non-neutrality of money). In the short-run the adjustment of wages and prices are in fact slow due to long-term contracts and "menu costs". As a result a shock will reduce the real interest rate. To the contrary, the term structure does not change much over short maturities, and consequently it has no forecasting power for inflation.

On the other hand, for longer maturities, (6m, 9m, 12m) the term structure is found to contain a great deal of information for inflation. This finding is in line with Mishkin's (1990) empirical results. In the long run, socks affect predominantly the term structure. Over longer periods the change of the slope of the term structure reflects inflation expectations. If the nominal interest rate i reflects inflation expectations, then the real interest rate is constant in the long run. As a result the real rate has no predictive power for future inflation.

The fact that for longer maturities data the lagged real interest rate seems to have predictive power for the future path of inflation only when the term structure is included in the regression of our model must be attributed to the high correlation between the lagged real interest rate and the nominal interest rate spread that exists in the longer maturities data. The correlation between these two variables is low for short maturities, (1m, 3m), but increases substantially for longer maturities (6m, 9m, 12m). The correlations are presented in the following graph.



Notes

- **r1m-sp1m**: the correlation between the real interest rate of one-month data and the nominal interest rate spread i_{10y}-i_{1m}.
- **r3m-sp3m**: the correlation between the real interest rate of three-month data and the nominal interest rate spread i_{10y}-i_{3m}.
- **r6m-sp6m**: the correlation between the real interest rate of six-month data and the nominal interest rate spread i_{10y}-i_{16}.
- **r9m-sp9m**: the correlation between the real interest rate of nine-month data and the nominal interest rate spread i_{10y}-i_{9m}.
- **r12m-sp12m**: the correlation between the real interest rate of twelve-month data and the nominal interest rate spread i_{10y}-i_{12m}.

8 CONCLUSIONS

In this paper we have provided a formal derivation of the connection between the change in the inflation rate and the lagged real interest rate. Using US data for the period 1970:01 to 1986:12 we found that the inflation, the nominal and the real interest rate in the US are trend-stationary series with a break in both the mean and the drift rate of the trend[17].

Our results suggest that there is a strong connection between the future inflation rate and real rate for short time periods. In one and three-month data the estimated coefficients of the lagged real interest rate are statistically significant at the 1% significance level, showing that in the short run the real rate contains great deal of information. This finding is new evidence and can be very helpful in the attempt to forecast inflation for short maturities, and of course it casts doubt on the approaches that assume a constant real interest rate. Moreover it confirms the intuition of Frankel and Lown, and Wickens and Tzavalis, who have already pointed out the fact that the real interest rate may contain useful information for the future path of inflation.

In addition, the term structure seems to have no ability to predict inflation in short maturities data. This is in line with Mishkin's empirical results, according to which we cannot use the nominal term structure to assess future fluctuations of inflation.

[17] The only exceptions were the nine-month data. The 9m inflation was found to be trend stationary with a shift only in the drift rate of the trend, the 9m real interest rate stationary with a shift in the mean, and the 9m nominal interest rate was found to contain a unit root.

In contrast, for longer maturities our results suggest that the term structure does have significant ability to forecast changes in inflation. This finding is in line with Jorion and Mishkin, whose empirical study (1991) for the U.S., Britain, Germany, and Switzerland resulted in a similar conclusion. On the other hand, the real interest rate seems to have no predictive power for future inflation changes, and its statistical significance only when the term structure is included in the regressions may be attributed to the high correlation between these two series (the real interest rate and the nominal interest rate spread) for long maturities data. The inability of the real interest rate to predict future fluctuations of inflation for long periods is attributed to the fact that the real interest rate converges to a constant in the long run.

An interesting proposal for further research would be the investigation of the validity of our model by using an updated dataset, which would include data from the 1990s. We also believe that it would be interesting to test the predictive power of real interest rates for future inflation by using data from other countries.

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10 APPENDIX A: DATA CONSTRUCTION

The spot rates for the US market are taken from Shiller and McCulloch (1987), until December 1986. The method of their construction involves fitting splines to the discount function for each month of the sample.

First, a functional form for the discount function is postulated. The discount function is formally defined as the present value of a dollar paid at time t, and will be used to price any bond. The discount function[18] is assumed to take the shape of a cubic spline:

$D(t) = a_{i0} + a_{i1}t + a_{i2}t^2 + a_{i3}t^3$, $t_i < t < t_{i+1}$,

where the values of the parameter a_{ij} are allowed to differ across knot points t_i, i=1,.....,N. To ensure continuity and smoothness of the spline curve, the function value and its first and second derivatives are restricted to be the same at the knot points. In addition the discount function is set equal to one at time zero. For N knot points, this yields 3+N parameters to estimate.

Second, the theoretical prices of the sample of selected bonds are computed from this discount function. For instance, if a bond pays the cash flows c_k, k=1....., K, at a respective times t_k, the model price is set at the present value of the future cash flows,

$$P^* = \sum_k c_k D(t_k),$$

[18] See Jorion and Mishkin (1990).

which can be written as a linear function of the spline parameters,

$$P^* = a_0 f_0(c_k) + a_1 f_1(c_k) + a_2 f_2(c_k) + \dots$$

The parameter {a} are estimated by comparing the theoretical prices P* with the market prices P for the selected bonds. Minimizing the square of the disperancies amounts to running a simple regression of the market prices on the model prices expressed as a linear function of the parameters of interest.

Once the discount function is found from market data, annualized spot rates can be derived as D(t) = exp(-i(t) * t), where t is expressed in years.

11 APPENDIX B: DESCRIPTIVE STATISTICS

1m Nominal Interest Rate		
Mean	7.45711	(%)
Median	6.97	(%)
Maximum	16.21	(%)
Minimum	3.02	(%)
Std. Dev.	2.85192	(%)
Skewness	0.93501	
Kurtosis	3.46677	
Jarque-Bera	31.5765	

1m Inflation Rate		
Mean	6.34427	(%)
Median	5.725	(%)
Maximum	21.67	(%)
Minimum	-5.92	(%)
Std. Dev.	4.54168	(%)
Skewness	0.38963	
Kurtosis	3.13621	
Jarque-Bera	5.31927	

1m Real Interest Rate		
Mean	1.1126	(%)
Median	0.895	(%)
Maximum	13.35	(%)
Minimum	-13.39	(%)
Std. Dev.	4.38255	(%)
Skewness	0.05516	
Kurtosis	3.13074	
Jarque-Bera	0.24874	

3m Nominal Interest Rate		
Mean	7.830784	(%)
Median	7.295	(%)
Maximum	16	(%)
Minimum	3.38	(%)
Std. Dev.	2.928828	(%)
Skewness	0.891652	
Kurtosis	3.272885	
Jarque-Bera	27.66445	

3m Inflation Rate		
Mean	6.338922	(%)
Median	5.63	(%)
Maximum	17.29	(%)
Minimum	-3.55	(%)
Std. Dev.	3.809542	(%)
Skewness	0.397848	
Kurtosis	2.791266	
Jarque-Bera	5.751958	

3mReal Interest Rate		
Mean	1.491127	(%)
Median	1.1	(%)
Maximum	11.31	(%)
Minimum	-5.96	(%)
Std. Dev.	3.822674	(%)
Skewness	0.360366	
Kurtosis	2.623564	
Jarque-Bera	5.619841	

6m Nominal Interest Rate		
Mean	8.13279	(%)
Median	7.53	(%)
Maximum	16.51	(%)
Minimum	3.61	(%)
Std. Dev.	2.92253	(%)
Skewness	0.85625	
Kurtosis	3.14767	
Jarque-Bera	25.1129	

6m Inflation Rate		
Mean	6.39682	(%)
Median	5.505	(%)
Maximum	14.99	(%)
Minimum	-0.2	(%)
Std. Dev.	3.48937	(%)
Skewness	0.43808	
Kurtosis	2.43731	
Jarque-Bera	8.94509	

6m Real Interest Rate		
Mean	1.8026	(%)
Median	1.12	(%)
Maximum	12.9	(%)
Minimum	-4.19	(%)
Std. Dev.	3.81773	(%)
Skewness	0.52976	
Kurtosis	2.63296	
Jarque-Bera	10.6869	

9m Nominal Interest Rate		
Mean	8.270392	(%)
Median	7.675	(%)
Maximum	16.64	(%)
Minimum	3.69	(%)
Std. Dev.	2.894297	(%)
Skewness	0.834873	
Kurtosis	3.095937	
Jarque-Bera	23.77665	

9m Inflation Rate		
Mean	6.325637	(%)
Median	5.485	(%)
Maximum	13.69	(%)
Minimum	0.85	(%)
Std. Dev.	3.280722	(%)
Skewness	0.519676	
Kurtosis	2.284746	
Jarque-Bera	13.53064	

9m Real Interest Rate		
Mean	1.944314	(%)
Median	1.045	(%)
Maximum	12.78	(%)
Minimum	-5.11	(%)
Std. Dev.	3.870354	(%)
Skewness	0.497754	
Kurtosis	2.409679	
Jarque-Bera	11.38587	

1y Nominal Interest Rate		
Mean	8.35706	(%)
Median	7.685	(%)
Maximum	16.35	(%)
Minimum	3.75	(%)
Std. Dev.	2.83799	(%)
Skewness	0.78684	
Kurtosis	2.94552	
Jarque-Bera	21.0753	
1y Inflation Rate		
Mean	6.31686	(%)
Median	5.69	(%)
Maximum	13.81	(%)
Minimum	1.08	(%)
Std. Dev.	3.19542	(%)
Skewness	0.53904	
Kurtosis	2.21876	
Jarque-Bera	15.0668	
1y Real Interest Rate		
Mean	2.03995	(%)
Median	1.25	(%)
Maximum	11.69	(%)
Minimum	-4.52	(%)
Std. Dev.	3.98622	(%)
Skewness	0.43783	
Kurtosis	2.16966	
Jarque-Bera	12.3781	

12 APPENDIX C: UNIT ROOT TESTS FOR THE USED INTEREST RATE SPREADS.

VARIABLE	t_{μ}*	t_{v}	t	LAGS**
i(10y) – i (1m)	-6,598	-3,857	-1,122	0
i(10y) – i (3m)	1,660	-15,638	-3,135	0
i(10y) – i (6m)	-1,664	-2,187	-0,849	0
i(10y) – i (9m)	-1,957	-1,741	-1,608	8
i(10y) – i (1y)	-1,763	-1,539	-1,434	9
5% cv	**-3.4212**	**-2.8679**	**-1.9401**	

Notes

*Augmented Dickey-Fuller tests are shown in the columns headed t_{μ} (with constant and trend), t_{v} (with constant only), t (no constant, no trend).

** The optimal number of lags is indicated by Akaike Criterion (AIC)

13 APPENDIX D: RESIDUALS TESTS

13.1 AUTOCORRELATIONS (AC)

The autocorrelation of a series y at lag k is estimated by:

$$r_k = \frac{\sum_{t=k+1}^{T}(y_t-\bar{y})(y_t-k)}{\sum_{t=1}^{T}(y_t-\bar{y})^2}$$

where $\bar{y}$ is the sample mean of y. This is the correlation coefficient for values of the series k periods apart. If r_1 is nonzero, it means that the series is first order serially correlated. If r_k dies off more or less geometrically with increasing lag k, it is a sign that the series obeys a low-order autoregressive (AR) process. If r_k drops to zero after a small number of lags, it is a sign that the series obeys a low-order moving-average (MA) process.

13.1.1 Q-STATISTICS

The last two columns reported in the correlogram are the Ljung-Box Q-statistics and their p-values. The Q-statistic at lag k is a test statistic for the null hypothesis that there is no autocorrelation up to order k and is computed as

$$Q_{LB} = T(t+2)\sum_{j=1}^{k}\frac{r^2_j}{T-j}$$

where r_j is the j-th autocorrelation and T is the number of observations. If the series is not based upon the results of ARIMA estimation, then under the null hypothesis, Q is asymptotically distributed as a x^2 with degrees of freedom equal to the number of autocorrelations. If the series represents the residuals from ARIMA estimation, the appropriate degrees of freedom should be adjusted to represent the number of autocorrelations less the number of AR and MA terms previously estimated. Note also that some care should be taken in interpreting the results of a Ljung-Box test applied to the residuals from an ARMAX specification (see Dezhbaksh, 1990, for simulation evidence on the finite sample performance of the test in this setting).

The Q-statistic is often used as a test of whether the series is white noise. There remains the practical problem of choosing the order of lag to use for the test. If you choose too small a lag, the test may not detect serial correlation at high-order lags. However, if you choose too large a lag, the test may have low power since the significant correlation at one lag may be diluted by insignificant correlations at other lags. For further discussion, see Ljung and Box (1979) and Harvey (1990, 1993).

13.2 SERIAL CORRELATION LM TEST

This test is an alternative to the Q-statistics for testing serial correlation. The test belongs to the class of asymptotic (large sample) tests known as Lagrange multiplier (LM) tests.

Unlike the Durbin-Watson statistic for AR (1) errors, the LM test may be used to test for higher order ARMA errors, and is applicable whether or not there are lagged dependent variables. Therefore, we recommend its use whenever you are concerned with the possibility that your errors exhibit autocorrelation.

The null hypothesis of the LM test is that there is no serial correlation up to lag order p, where p is a pre-specified integer. The local alternative is ARMA(r,q) errors, where the number of lag terms p = max{r,q}. Note that the alternative includes both AR (p) and MA (p) error processes, and that the test may have power against a variety of autocorrelation structures. See Godfrey (1988) for a discussion.

The test statistic is computed by an auxiliary regression as follows: suppose you have estimated the regression

$Y_t = X_t b + e_t$,

where e are the residuals.

The test statistic for lag order p is based on the regression

$e_t = X_t \gamma + \alpha_1 e_{t-1} + \alpha_2 e_{t-2} + \ldots\ldots + \alpha_p e_{t-p} + u_t$

This is a regression of the residuals on the original regressors X and lagged residuals up to order p. E-Views reports two test statistics from this test regression. The F-statistic is an omitted variable test for the joint significance of all lagged residuals. Because the omitted variables are residuals and not independent variables, the exact finite sample distribution of the F-statistic under H_0 is not known, but E-Views still presents the F-statistic for comparison purposes.

The Obs*R-squared statistic is the Breusch-Godfrey LM test statistic. This LM statistic is computed as the number of observations, times the (uncentered) R^2 from the test regression. Under quite general conditions, the LM test statistic is asymptotically distributed as a $x^2(p)$. The original regression may include AR and MA terms, in which case the test regression will be modified to take account of the ARMA terms.

13.3 ARCH LM TEST

This is a Lagrange multiplier (LM) test for autoregressive conditional heteroskedasticity (ARCH) in the residuals (Engle 1982). This particular specification of heteroskedasticity was motivated by the observation that in many financial time series, the magnitude of residuals appeared to be related to the magnitude of recent residuals. ARCH in itself does not invalidate standard LS inference. However, ignoring ARCH effects may result in loss of efficiency.

The ARCH LM test statistic is computed from an auxiliary test regression. To test the null hypothesis that there is no ARCH up to order q in the residuals, E-VIEWS runs the regression

$$e_t = b_0 + b_1 e^2_{t-1} + ... + b_q e^2_{t-q} + u_t,$$

where e is the residual. This is a regression of the squared residuals on a constant and lagged squared residuals up to order q. E-Views reports two test statistics from this test regression. The F-statistic is an omitted variable test for the joint significance of all lagged squared residuals. The Obs*R-squared statistic is Engle's LM test statistic, computed as the number of observations times R^2 from the test regression. The exact finite sample distribution of the F-statistic under H_0 is not known but the LM test statistic is asymptotically distributed as $x^2(q)$ under quite general conditions.

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13.4 WHITE'S HETEROSKEDASTICITY TEST

This is a test for heteroskedasticity in the residuals from a least squares regression (White, 1980). Ordinary least squares estimates are consistent in the presence heteroskedasticity, but the conventional computed standard errors are no longer valid. If evidence of heteroskedasticity is found, someone should either choose the robust standard errors option to correct the standard errors (see HAC) or should model the heteroskedasticity to obtain more efficient estimates using weighted least squares.

White's test is a test of the null hypothesis of no heteroskedasticity against heteroskedasticity of some unknown general form. The test statistic is computed by an auxiliary regression, where E-Views regresses the squared residuals on all possible (non redundant) cross products of the regressors. For example, suppose the following regression is estimated:

$$Y_t = b_1 + b_2 X_t + b_3 z_t + e_t$$

The test statistic is then based on the auxiliary regression:

$$e^2_t = a_0 + a_1 x_t + a_2 z_t + a_3 x^2_t + a_4 z^2_t + a_5 x_t z_t$$

E-Views reports two test statistics from the test regression. The F-statistic is an omitted variable test for the joint significance of all cross products, excluding the constant. It is presented for comparison purposes.

The Obs*R-squared statistic is White's test statistic, computed as the number of observations times the centered R^2 from the test regression. The exact finite sample distribution of the F-statistic under H_0 is not known, but White's test statistic is asymptotically distributed as a x^2 with degrees of freedom equal to the number of slope coefficients (excluding the constant) in the test regression.

White also describes this approach as a general test for model misspecification, since the null hypothesis underlying the test assumes that the errors are both homoskedastic and independent of the regressors, and that the linear specification of the model is correct. Failure of any one of these conditions could lead to a significant test statistic. Conversely, a non-significant test statistic implies that none of the three conditions is violated.

When there are redundant cross products, E-Views automatically drops them from the test regression. For example, the square of a dummy variable is the dummy variable itself, so that E-Views drops the squared term to avoid perfect collinearity.

E-Views has two options for the test: cross terms and no cross terms. The cross terms version of the test is the original version of White's test that includes all of the cross product terms. However, with many right-hand side variables in the regression, the number of possible cross product terms becomes very large so that it may not be practical to include all of them. The no cross terms option runs the test regression using only squares of the regressors.

13.5 JARQUE-BERA NORMALITY TEST

Jarque-Bera is a test statistic for testing whether the series is normally distributed. The test statistic measures the difference of the skewness and kurtosis of the series with those from the normal distribution. The statistic is computed as:

$$JB = (N-k)/6^* \{S^2 + (1/4)^*(K-3)\}^2$$

where S is the skewness, K is the kurtosis, and k represents the number of estimated coefficients used to create the series.

Under the null hypothesis of a normal distribution, the Jarque-Bera statistic is distributed as x^2 with 2 degrees of freedom. The reported Probability is the probability that a Jarque-Bera statistic exceeds (in absolute value) the observed value under the null—a small probability value leads to the rejection of the null hypothesis of a normal distribution.