ΘΕΜΑ ΔΙΠΛΩΜΑΤΙΚΗΣ ΕΡΓΑΣΙΑΣ:

Εξέταση της ικανότητας της μακροοικονομίας στην πρόβλεψη μελλοντικών μεταβολών των επιτοκίων.

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1. ABSTRACT

Most of the recent empirical literature on the term structure of interest rates is concerned with the informative content of either the forward rates or the yields to maturity of long-term bonds. According to *the expectations hypothesis*, these rates are, up to a constant, an unbiased predictor of the future expected spot rate and the future expected average of the short-term bond yields, respectively. In the rate of change form and assuming that the market expectations are formed *rationally* previous hypotheses have been tested extensively and almost invariably rejected. The empirical failure is generally attributed either to systematic expectations errors or to shifts in the time-varying term premia. In fact, the empirical tests, based on the estimation of single-equation models, are not able to discriminate between these two hypotheses.

A recent strand of the macroeconomic literature has analyzed monetary policy by including the central bank reaction function in small empirical macro models. By simulating these models forward it is possible to derive the full forward path of short-term interest rates and hence to construct any long-term interest rate consistent with the expectations theory. Then, a direct test of the validity of the expectation models using macroeconomics, based on full information, can be immediately constructed by comparing observed long-term rates with the simulated ones and the associated 95% confidence interval.

Within our framework, we examine the predictive power of previous macro models in future changes of actual interest rates. We use regression models that take into account the simulated term structure path derived by these models rather than forward rates or yields to maturity of long-term bonds. It is crucial to mention that this procedure seems to be free of time-varying term premia which have been recognized in the literature as the main factor for the rejection of classical expectations theory. Our work is based on Favero (2001), who introduced the idea of using a structural macroeconomic model to test the validity of the REHTS. However, our work differs in an important respect from Favero as we use differences rather than levels of interest rates -that are nonstationary- allowing the application of standard statistical inference.

2. THEORETICAL FRAMEWORK ON EXPECTATIONS HYPOTHESIS

The expectation model of the term structure states that the yield to maturity of long-term bonds is equal to the average of expected future short-term bond yields.

A general form of this specification may be written as an equation that provides an implicit definition of the expected term premium and is the following:

$$R_t^{(n)} = E_t \theta_t (n) + \sum_{i=0}^{n-1} w_i E_t r_{t+i} \ , \qquad \sum_{i=0}^{n-1} w_i = 1 \quad (1)$$

where E_t denotes the expectations operator conditional on information available at the beginning of time *t*, $R_t^{(n)}$ is the yield on a bond with a maturity of *n* periods, r_{t+i} is the interest rate on a 1-period debt instrument at the beginning of time *t+i*, and w_i is a geometric declining weight that sums to 1. Specifically, if we take into account a zero coupon bond then:

$$w_i = \frac{1}{n}$$

while for a n-period annually coupon bond:

$$w_i = \left[\frac{1-g}{1-g^n} \right] g^i \quad \text{and} \quad g \equiv \left(1+\overline{R}\right)^{-1} \qquad i = 0,1,\ldots,n-1$$

But the expectations hypothesis imposes the restriction that the expected term premium is constant: $E_t \theta_{t+n}(n) = \theta(n)$ and thus (1) can be rewritten as:

$$R_t^{(n)} = \theta(n) + \sum_{i=0}^{n-1} w_i E_t r_{t+i} \ , \qquad \sum_{i=0}^{n-1} w_i = 1 \quad (2)$$

The expectations hypothesis is compatible with an arbitrage-pricing equilibrium and is less restrictive than the pure expectations hypothesis, which states that $E_t\theta_{t+n}(n)$=0. Cox, Ingersoll and Ross (1981) have criticized the different versions of the pure expectations hypothesis that exist in the literature as being incompatible with each other and with an arbitrage-pricing equilibrium. However, Campbell (1986) has shown that their criticisms do not carry over to the more general expectations hypothesis that I consider in this study.

Then supposing that market has rational expectations whereby the actual future spot rate is equal to the expected future spot rate augmented by an error term: $r_{t+n} = E_t r_{t+n} + w_{t+n}$ we can take from (2) the following linear regression equation:

$$\sum_{i=0}^{n-1} w_i r_{t+i} - r_t = a_n + b_n \left[R_t^{(n)} - r_t\right] + w_{t+n} \qquad (2^*)$$

where the test of H_0: $\hat{b}_n = 1$ on the term spread is a joint test of the null hypothesis that expectations are rationally formed and that arbitrage between short- and long-term rates holds as assumed by the expectations hypothesis.

For ease of exposition, in a simple two-period case (Lange 1999) the estimate of the $\hat{b}_n$ coefficient converges to:

$$\hat{b} \equiv \frac{\sigma^2(\mathrm{E}_t \Delta r_{t+1}/2) + 2\rho\sigma(\mathrm{E}_t \Delta r_{t+1}/2)\sigma(\theta_t)}{\sigma^2(\mathrm{E}_t \Delta r_{t+1}/2) + 2\rho\sigma(\mathrm{E}_t \Delta r_{t+1}/2)\sigma(\theta_t) + \sigma^2(\theta_t)} \qquad (3)$$

where, under the assumption of rational expectations, $E_t \Delta r_{t+1}$ is the expected change in the short-term rate, ρ is the correlation between the time-varying term premium θ_t and the expected change in the short rate, and σ is the standard deviation of θ_t.

Equation (2*) is the standard equation estimated in empirical work on the expectations hypothesis of the term structure. For example (Rudebusch 1995), let $r(1)_t$ and $r(2)_t$ be the yields on one- and two-period (zero-coupon) bonds, respectively. Then, the expectations hypothesis with a constant term premium, $\pi(2)$, implies (to the close approximation):

$$r(2)_t = 1/2\left[r(1)_t + E_t r(1)_{t+1}\right] + \pi(2) \qquad (4)$$

that is, the current two-period yield equals the average of the actual and expected one-period yields plus a term premium. Assuming rational expectations,

$$r(1)_{t+1} = E_t r(1)_{t+1} + \varepsilon_{t+1} \qquad (5)$$

where ε_{t+1} is a forecast error orthogonal to information available at time t. Substituting (5) into (4) and rearranging provides:

$$1/2\left[r(1)_{t+1} - r(1)_t\right] = \alpha + \beta\left[r(2)_t - r(1)_t\right] + u_{t+1} \qquad (6)$$

Under the expectations hypothesis of the term structure, $\beta = 1$ and $\alpha = -\pi(2)$, that is, after taking expectations of both sides of (6), one-half the optimal forecast of the change in the short rate should equal the spread between the long rate and short rate (minus a term premium). The error term is orthogonal to the right-hand-side regressors ($u_{t+1} = \varepsilon_{t+1}/2$), so OLS provides consistent coefficient estimates.

The term spread can also be used to express the market's forecast of a 1-period change in the long-term yield:

$$(n-1)(R_{t+1}^{n-1} - R_t^n) = a_0 + a_1(R_t^n - r_t) + \varepsilon_{t+1} \qquad (7)$$

where R_{t+1}^{n-1} is the next period's long-term yield and the left-hand-side is the 1-period change in the n-periods yield. If the expectations hypothesis holds and the 1-period term premium $\theta_{1,t+1}$ is constant, then changes in the long-term yield $R_{t+1}^{n-1} - R_t^n$ reflect changes in the term spread on the right-hand-side of (7). Intuitively, equation (7) states that if the n-periods yield is expected to rise next period (positive left-hand-side of (7)), which will result in a capital loss, then the n-period bond has to have a higher current yield than the 1-period short-term instrument (positive term spread) in order to equate expected returns over the next period.

In a similar way the one-period forward rate can always be decomposed into $E_t\theta_{t+n}(n)$ the expected time-varying risk premium, and $E_t r_{t+n}$ the expected rate on time *t+n* in the future. This can be written as follows:

$$f_t(n,1) = E_t r_{t+n} + E_t\theta_{t+n}(n) \qquad (8)$$

but expectations hypothesis imposes the restriction that the expected risk premium is constant: $E_t\theta_{t+n}(n)$ = $\theta(n)$ and thus:

$$f_t(n,1) = E_t r_{t+n} + \theta(n) \qquad (9)$$

Next, we assume that market expectations are formed rationally:

$$r_{t+n} = E_t r_{t+n} + w_{t+n} \qquad (10)$$

where forecast error w_{t+n} must be uncorrelated with variables in the information set at time *t*. Thus, w_{t+n} orthogonal with the forward rate $f_t(n)$ that market participants know at time *t*. Substituting (10) into (9) for $E_t r_{t+n}$, subtracting r_t from both sides and rearranging, we obtain:

$$r_{t+n} - r_t = -\theta(n) + (f_t(n) - r_t) + w_{t+n} \qquad (10^*)$$

Equation (10*) shows the relation between the change of the spot rate between *t* and *t+n* and the actual forward spread. It is a result of both expectations hypothesis (unbiased forward rate hypothesis) and rational expectations and it indicates that if $\theta_t(n)$ is a constant risk premium θ, then ($r_{t+n} - r_t$) and $(f_t(n) - r_t)$

move one-to-one. That is if $\theta_t(n) = \theta$, as the expectations hypothesis suggests, then the following linear regression:

$$r_{t+n} - r_t = a_n + b_n\left(f_t(n) - r_t\right) + w_{t+n} \qquad (11)$$

must give, in accordance with the previous results, an estimation about the coefficient b_n equals unity:

$$\hat{b}_n = 1 \quad \text{(Forward rate unbiased hypothesis [FRUH]).}$$

If $\hat{b}_n > 0$ but $\neq 1$, then the forward-spot spread $\left(f_t(n) - r_t\right)$ observed at time *t* has an obvious predictive power to forecast the change in the spot rate *n* years ahead but not a one-to-one relationship with $r_{t+n} - r_t$.

3. PREVIOUS EMPIRICAL EVIDENCES

Does the slope of the term structure-the yield spread between longer-term and shorter-term interest rates-predict future changes in interest rates? And if so, is the predictive power of the yield spread in accordance with the expectations theory of the term structure?

These questions are important, both for forecasting interest rates and for interpreting shifts in the yield curve. The most prevalent explanation of fluctuations in the yield curve is the expectations theory, which posits that the slope of the yield curve reflects the market expectation of the future change in interest rates. If the expectations theory is an adequate description of the term structure, then rational expectations of future interest rates are the dominant force determining current long-term interest rates. On the other hand, if the expectations theory is very far for accurate, then predictable changes in expected excess returns must be the main influence moving the term structure.

The literature on the term structure, however, presents evidence that the data are inconsistent with the joint hypothesis of the expectations theory and rational expectations. Almost all studies statistically reject the expectations theory of the term structure; but some studies suggest that the yield spread does predict interest rate movements in roughly the way one would expect if the expectations theory were true, while other studies reach the opposite conclusion. Different studies use different econometric methods, test different implications of the expectations theory, and look at different interest rate maturities.

The empirical failure is generally attributed either to systematic expectations errors, or to shifts in term premia. In fact, the empirical tests, based on the estimation of single-equation models, are not able to discriminate between these two hypotheses. Under these results, the simple theory that the slope of the term structure can be used to forecast the direction of future changes in interest rates seems worthless.

The bulk of contrary evidence shows that:

i) high yield spreads fare poorly in predicting short-run changes in long rates as equation (7) requires (see Campbell, 1995)

ii) changes in yields does not move one-to-one with the forward-spot-spread (see Fama and Bliss, 1987 and Hardouvelis, 1988)

However, as Hardouvelis (1988) said, the predictive power of forward rates is an ambiguous indicator of the market's ability to predict. This is because forward rates contain composite information on the market's expectations about both future spot rates and time-varying term premia. If, for example, a non-constant risk premium θ is correlated with the forward premium, then OLS on (11) gives an inconsistent ***b***-coefficient and the t-test about the unbiasness of the forward spread is not reliable. In another occasion ***b***-coefficient is inconsistent too, if the forecast error w_{t+n} in (11) is not orthogonal within the sample of all available information so it is not a white noise process $(E_t\left[w_{t+n}\right]=0)$ as the assumption of rational expectations requires.

Thus, formula (3) indicates that the size of ***b***-coefficient under rational expectations depends on three terms (Lange 1999): the variation of the expected changes in the short-term rate, the variation of the term premium $\sigma(\theta t)$, and the correlation ρ between the term premium and the expected changes in the short rate. The ***b***-coefficient is 0 if the short-term interest rate is not predictable ($E_t \Delta r_{t+1} = 0$) and is equal to 1 in the absence of a time-varying premium ($\sigma(\theta t)^2 = 0$). However, variations in the term premium $\sigma(\theta t)$ will bias downwards the coefficient on the term spread; the size of the bias depends on the variance of the expected change in the future short rate. The bias also depends on ρ, the correlation between the term premium and the expected changes in the short rate, so that ***b*** can be greater than 1 when ρ is sufficiently negative and the variation in expected changes in future short rate is low.

Much of the previous empirical literature has relied on regression models that can be used to test the predictive ability of the yield spread for changes in long- and short-term interest rates. This empirical approach bases on *ex post test of the rational expectations hypothesis* and contains the following two equations:

$$r_{t+m}^{(n-m)} - r_t^{(n)} = a_0 + a_1 \frac{m}{n-m}(r_t^{(n)} - r_t^{(m)}) + error\ term \qquad (3.1)$$

$$\frac{m}{n}\sum_{i=0}^{n/m-1} r_{t+mi}^{(m)} - r_t^{(m)} = b_0 + b_1(r_t^{(n)} - r_t^{(m)}) + error\ term \qquad (3.2)$$

Many researchers have tested this proposition on the spread, using postwar data, for almost any combination of maturities between one month and ten years. However other papers assess the expectations theory for the longer end of the term structure of interest rates using other empirical approaches. Generally, the main *three* of them are:

(i) *ex post test of the rational expectations hypothesis*

(ii) *cointegration tests of the long-run unbiased hypothesis*

The changes in future short-term rates on the left-hand side of equation (3.2) is an I(0) series, with the level of the short-rate being non-stationary and integrated of order 1. The right-hand side is a linear combination of two I(1) variables, $r_t^{(n)}$ and $r_t^{(m)}$, plus a term premium θ_t and a forecast error. If the expectations hypothesis holds, then the term premium and the forecast error are stationary. This implies that the term spread is stationary, and thus long- and short-term rates are cointegrated. The term spread is an unbiased predictor of changes in future short-term rates over the long run if there is a stable one-to-one relationship between short- and long-term interest rates. However, the existence of a cointegration relationship is a necessary condition for the expectations hypothesis to hold, though it is not an explicit test of the hypothesis itself.

The research on the existence of a long-run equilibrium relationship between Canadian interest rates as implied by the expectations hypothesis can be traced to Boothe (1991). Using residual-based tests, he found that cointegration between short- and long-term interest rates is always rejected at the 5 per cent level for the 1972-89 period. Furthermore, the coefficient on the short-rate was always significantly below the theoretical value of 1.0 that is required for the term spread to be unbiased predictor of changes in short rates over the long run. Lange (1999) re-estimated Boothe's equation for a longer sample period, from 1956 to 1998, and obtained:

$$R_t^n = 2.73 + 0.72 r_t,$$

where R_t^n is the 10-year-and-over government bond yield and r_t is the 90-day commercial paper rate. As in Boothe, the augmented Dickey-Fuller test statistic at –3.21 suggests that the null hypothesis of no cointegration cannot be rejected by the data at the 10 per cent level. The coefficient on the short rate is slightly larger than the 0.59 in Boothe's regression for the 1972-89 period, but still noticeably less than 1.

(iii) simulations of a theoretical long-term yield that is consistent with the expectations hypothesis

Specifically, in a seminal paper, Campbell and Shiller (1987) use the cointegration property of short- and long-term interest rates to specify a VAR model that can simulate the expected future changes in short-term rates. The approach tests for the expectations hypothesis by generating the VAR forecasts of changes in future short-term rates, and then comparing the counterfactual long-term yield that is consistent with the expectations hypothesis with the behavior of the actual long-term yield. The Campbell-Shiller methodology allows for multi-period forecasting of changes in short-term rates without having to both estimate regressions with overlapping errors and drop large portions of the estimation period in order to test for the *ex post* success of the term spread as a predictor of changes in future short-term rates.

Campbell and Shiller (1987,1991) specify a bivariate VAR model for two stationary variables, the term spread, $S_t \equiv R_t - r_t$, and the change in the short interest rates, Δr_t:

$$\begin{bmatrix} \Delta r_t \\ S_t \end{bmatrix} = \begin{bmatrix} \alpha(L) b(L) \\ c(L) d(L) \end{bmatrix} x \begin{bmatrix} \Delta r_{t-1} \\ S_{t-1} \end{bmatrix} + \begin{bmatrix} \mu_{rt} \\ \mu_{st} \end{bmatrix} \qquad (3.3)$$

where $a(L),..., d(L)$ are lagged polynomials. The model is estimated for demeaned values, which guarantees a non-varying component of the term premium. This constant component is accounted for by a non-zero difference of the unconditional means of the long- and short-term interest rates.

In the VAR model, changes in the short-term interest rate would only be dependent on past changes in the short rate if $b(L)$ were 0. On the other hand, if market participants have additional information beyond the history of past changes in the short rate (and therefore past S_t), then S_t will have incremental explanatory power. If agents do not have such information, then they form S_t as an exact linear function of current and lagged Δr_t.

The estimation methodology proceeds in three steps. First, a second-order VAR model is estimated for the change in the short-term rate, and the spread between long- and short-term interest rates is estimated as in equation (3.3). Second, the VAR framework is used as a model for a multi-period forecast of changes in future short-term rates. Assuming a constant term-premium, the predicted changes in short-term rates, along with a set of declining geometric weights, are then used to compute a theoretical or counterfactual long-term yield or term spread that is consistent with the expectations hypothesis. Third, the theoretical yield or spread is compared with the historical behavior of the actual series in order to access how well expectations hypothesis explains movements in long-term yields or the term spread over time.

Campbell and Shiller show that, although the restriction implied by the expectations theory can easily be rejected on U.S. data, long- and short-term interest rates computed under the assumption of the expectations hypothesis evolve over time much as actual term spreads do.

As far as the long end of the maturity spectrum (data examined maturity) is concerned, the literature is not unanimous about long-term changes in short-rates, and it is plausible that the reported rejections follow from applying inappropriate econometric methods (Lanne (1999)), but for shorter maturity spectrum the evidence against the expectations hypothesis seems to be strong, especially with U.S data (Campbell and Shiller (1991), and Evans and Lewis (1994)). Interestingly, Campbell (1995), Mishkin (1988) and Hardouvelis (1994) find that there is much more truth in the proposition that high yield spreads should forecast long-term changes in short-rates, especially at very short and very long maturities.

About forward rates that according to the expectations hypothesis are, up to a constant, unbiased predictors of future realized spot rates, the expectations hypothesis has been tested extensively and rejected. However, rejection of the expectations hypothesis does not imply that there is no information in the term structure. Fama (1984) examines one- to six-month Treasury bills from 1959 through 1982 and, although he rejects the expectations hypothesis, finds predictive power in forward rates that lasts about three to five months during the first half of his sample and one month during the second half of his sample. Hardouvelis (1988) found strong evidence that forward rates have predictive power. Until October 1979, forward rates were only able to predict changes in interest rates that would occur one week later. However, when the Fed allowed interest rates to fluctuate relatively freely during the period October 1979 through October 1982, predictive power increased substantially lasting for at least six week into the future. (There is also predictive power from fourteen to twenty-one weeks ahead). After October 1982, when the Fed returned to partial interest rate targeting, predictive power remained strong, lasting nine weeks into the future.

The failure of the expectation model to predict short-run changes in long yields and the (partial) success in the prediction of long-run changes in short yields is explained first by the role of measurement errors. In fact, in the regression (3.1) of long rate changes onto the yield spread, changing rational expectations about excess long bond returns act like a measurement error that appears positively in the regressor and negatively in the dependent variable. Conversely, in the regression (3.2) of short-rate changes onto the yield spread, changing rational expectations about excess long-bond returns act like a measurement error that appears only in the regressor. In the first case a small measurement error can change the sign of the relevant regression coefficient, while in the second case the measurement error biases the coefficient toward zero but cannot affect its sign. In previous situations, period excess returns on long-term bond are predictable using the information of yield spread, while both forms of the expectations hypothesis require being unforecastable, in order not to have biased forecast errors.

But, is there a puzzle here? As we have already said, an extensive literature documents that the spread between long- and short-term interest rates can predict, at least, the correct direction of future changes in short rates [Campbell and Shiller (1987), Fama (1984,1990), Fama and Bliss (1987), Hardouvelis (1988), Mankiw and Minor (1986), Mishkin (1988)]. When the long rate rises relative to the short rate, future short rates tend to increase. Such predictive power is consistent with the expectations hypothesis of the term structure, which claims that long rates are weighted averages of current and expected future short rates. According to the expectations hypothesis, a rise in the long rate relative to the short rate is due to the expectation of higher short rates in the future. Thus, if the market makes correct predictions on average, future short rates would subsequently tend to rise, generating a positive correlation of the change in short rates with the earlier spread.

On the other hand, Mankiw and Summers (1984), Mankiw (1986), and Campbell and Shiller (1991) observed that the spread predicts the wrong direction in the subsequent change of the long rate: A rise of the current long rate relative to the current short rate is followed by a subsequent decline, rather than a rise, in the long rate next period. *This behavior is puzzling: How can the movement of future cumulative short rates obey the overall direction predicted by the expectations hypothesis but at the same time the short-run movement of long rates does not?*

Two main alternative explanations to the puzzle have been proposed: The first, and what appears to be the most popular one, claims that movements in current long rates do obey the general direction predicted by the expectations hypothesis, but those movements are sluggish relative to the movements of the current short rates: Long rates underreact to current short rates (or overreact to future short rates). This explanation assumes that term premia are constant and that the spread between long and short rates correctly incorporates the information about expectations of future interest rates, but that the market's expectations themselves violate the strict definition of rational expectations. Suppose, for example, that a policy announcement by the Fed increases the market's expectation of future short rates but, since the policy will be implemented in the future, leaves the current short rates intact. The hypothesis claims that markets would overreact to the announcement, raising their expectations of future spot rates by more than is warranted. The current long rate would thus increase by more than warranted, making the spread between long and short rates larger than it should be. During the next month or quarter, long rates would fall somewhat correcting the previous overreaction, thus generating a negative correlation between the change in long rates and the previous spread. Short rates, on the other hand, would begin their predicted rise, generating a positive correlation between the change in short rates and the earlier spread.

The second explanation assumes that the market's expectations are rational but the information in the spread is composite information about the variation of both expected future rates and term premia. This explanation requires, however, that the time-varying term premium has a very special structure. Consider the earlier example in which the Fed announcement increases the expected level of future short rates. For the term premium explanation about the observed correlations in the data, the term premium ought not to show a large response at the time of the announcement in order for the expectations hypothesis to approximately holds, yet one to three month later the term premium ought to decline so that the long-term bond yield also declines. Froot (1989) uses U.S. survey data on short-term and long-term interest rates and is able to distinguish between the two competing hypothesis. He finds that the negative correlation between changes in long rates and previous spread is not due to time-varying term premium, but is due to a violation of the rational expectations assumption, namely, an overreaction of the spread. When buying long-term bonds, market participants would do better to place more weight on the contemporaneous short rate and less weight on the expected future short rates.

Hardouvelis (1994) takes a fresh look at the puzzle by examining the relation between the spread and the future evolution of long and short rates internationally. He uses post-war data on an approximately ten-year yield and a three-month yield of each country that belongs of the Group of Seven (G7). The paper finds that, curiously, the puzzle is manifested primarily in the United States, the country with the most sophisticated and liquid financial markets. In France and Italy, long rates move in the correct direction. In Canada, UK, Germany and Japan, long rates move in the opposite direction, but this movement is apparently due to a white noise error that does not materially affect the information in the term structure. Specifically, a simple white noise deviation of long rates from their theoretically correct value, predicted by the expectations hypothesis, be responsible for the presence of the estimated negative correlation between changes in long rates and the earlier term structure spread. Such a deviation could be due to temporary mistakes that the market makes or it could be due to a simple econometric measurement error. The use of instrumental variables reverses the negative regression sign. In the U.S., a white noise error on long rates cannot explain the puzzle. The use of instrumental variables results in equally sharp rejections of the expectations hypothesis. Furthermore, time-varying term premia cannot provide an adequate explanation to the puzzle: Holding premia vary way too much relative to the variability of expected changes in long rates to be able to accommodate regression estimates that are different from zero.

Mankiw and Miron (1986) provided an alternative explanation for the term structure evidence that was consistent with rational expectations. In contrast to the earlier results based on postwar data, they showed that the three-month and six-month yield spread did significantly help to predict future changes in the three-month rate from 1890 to 1914, a period that predated the founding of the Federal Reserve System. Mankiw and Miron argued that the negligible predictive power of the spread after the founding of the Fed did not reflect a failure of the expectations theory. Instead, they suggested that the Fed “stabilized” short-term rates, such as the three-month rate, by including a random walk behaviour that eliminated any predictable variation. Thus, to a first approximation, expected future short rates have equalled current short rates since the founding of the Fed. In such a situation, even if the rational expectations theory holds, supporting empirical evidence cannot be obtained from the forecasting ability of the slope of the yield curve because there is no predictable variation in future short rates to incorporate into yield spreads. In essence, Mankiw and Miron argued that the absence of predictive information in the term structure for future short rates reflects the manner in which the Fed controls interest rates and is not a rejection of the rational expectations theory of the term structure.

Another explanation that has recently been offered for the rejection of the expectations hypothesis by Lewis (1991), Evans and Lewis (1994), and Bekaert, Hodrick and Marshall (1997a), is the presence of so called peso effects that influence the distribution of the typically used test statistics. By peso effects these authors refer to potential regime shifts in the process of the short-term rate that occur less frequently in the actual sample than they should according to the probability distribution of the process. Whereas classical regression-based tests indicate rejection, tests based on a new model allowing for potential-but unrealized-regime shifts provide support for the expectations hypothesis. Even if there were not a single regime shift in the observed data, the fact that these shifts have a positive probability, affects the expectations that the market forms of the future short-term rates, and thus the data seems to be irreconcilable with the expectations hypothesis. The estimation results suggest that potential regime shift had an effect on expectations concerning the longer-term interest rates only for a short while in the early phase of the sample period, when interest rates were at their highest.

4. CENTRAL BANK REACTION FUNCTION WITHIN SMALL MACRO MODELS

The claim of very low predictability of policy rates contradicts a growing body of empirical literature, which has established interest rate rules as a convenient way to model and interpret monetary policy. This macroeconomic literature has analyzed monetary policy by including the central bank reaction function in small empirical macro models of inflation and output.

Two main factors underlie the interest about these macro models. First, after a long period of near exclusive focus on the role of no monetary factors in the business cycle, a stream of work beginning in the late 1980s has made the case that monetary policy significantly influences the short-term course of the real economy. The precise amount remains open to debate. On the other hand, there now seems to be broad agreement that the choice of how to conduct monetary policy has important consequences for aggregate activity. It is no longer an issue to downplay.

Second, there has been considerable improvement in the underlying theoretical framework used for policy analysis. To provide theoretical underpinnings, the literature has incorporated the techniques of dynamic general equilibrium theory pioneered in real business cycle analysis. A key point of departure from real business cycle theory is the explicit incorporation of frictions such as nominal price rigidities that are needed to make the framework suitable for evaluation of monetary policy.

Taylor(1993), from whom this literature originates, simply postulates that the central bank should base the setting of the short-term interest rates on the current situation with regard to inflation and the business cycle:

Taylor interest rates= real equilibrium interest rates

+(expected)inflation rate

+a$_p$ * output gap

+a$_I$ * inflation gap

In another form the Taylor rule is given by:

$$i_t = r^{eq} + \pi_t + 0.5\{y_t + (\pi_t - \pi^{ob})\}$$

where $i_t, r^{eq}, \pi_t, \pi^{ob}$ *and* y_t denote the nominal interest rate, the equilibrium real interest rate(assumed constant), the rate of inflation over the past year, policymakers' inflation objective, and the output gap, respectively.

The use of the equilibrium real rate in the Taylor rule emphasizes that real rates play a central role in formulating monetary policy. Although the nominal federal funds rate is identified as the instrument that policymakers adjust, the real interest rate is what affects real economic activity. In particular, the rules clarify that real interest rates will be increased above equilibrium when inflation is above target or output is above its potential.

The output gap is the relative difference between the actual and the potential output level, the inflation gap is the difference between the measured inflation rate and the rate of inflation that the central bank aims for. Both variables are included in the Taylor interest rate with positive weighting of a$_p$ and a$_I$, respectively. This reflects the idea that an excessive price rise and an overutilisation of production should be counteracted by a higher short-term interest rate and vice versa. Accordingly, given full use of capacity and realization of the envisaged rate of inflation, the "real equilibrium interest rate" is the level of the real rate of interest at which the long-term equilibrium is not changed by monetary policy. The (expected) inflation rate is added to the sum of these three components to make the Taylor interest rate comparable with the relevant nominal interest rate.

In his original paper, Taylor applied the concept subsequently named after him to US monetary policy from 1987 to 1992. For his deliberately simple calculation, he selected the following approximations of the non-observable variables: he substituted the realized inflation rate of the preceding four quarters for the expected inflation rate over the same period, assumed a 2.2% annual rate of growth for the production potential (which corresponds to the trend growth of real income in the United States between 1984 and 1992), set the equilibrium real short-term rate of interest at 2%, and calculated the inflation gap as the difference between the current inflation rate and the inflation target which is a constant 2%. He gave equally high weighting to the inflation and output gaps, at 0.5 each. Measured by its simplicity, the Taylor interest rate thus calculated captures the behavior of the US Federal Funds Rate in the period reviewed quite well.

Calculating and using the Taylor interest rate appear at first glance to be very simple. In actual fact, however, it raises a number of practical and theoretical problems. First of all, for example, the weightings of the output and inflation gaps have to be determined. The weighting scheme used by Taylor is not necessarily appropriate. The central bank's orientation and the structure of the economy have to be taken into consideration when determining the coefficients. The weights are to be estimated and are thus method-dependent. Depending on the relative weight, however, there may be considerable differences in the Taylor interest rate at different periods resulting in a correspondingly varied assessment of current monetary policy.

The original work by Taylor has been refined in a number of papers by Clarida, Gali and Gertler (1998,1999,2000), who have shown that a forward looking version of the Taylor-rule with some interest rates persistence not only tracks the data well but is also capable of explaining the high inflation in the seventies in terms of an accommodating behavior towards inflation in the pre-Volcker era.

Specifically, 1998 paper by Clarida et al. estimate the monetary policy reaction functions for two sets of countries: the G3 (Germany, Japan and the US) and the E3 (UK, France and Italy). They found that since 1979 each of the G3 central banks has pursued an implicit form of inflation targeting, which might account for the broad success of monetary policy in those countries over this time period. The net effect was transition from a global environment where inflation seemed a virtually intractable problem to the current era where the major economies of the world enjoy relative price stability.

As for the E3, even prior to the emergence of the 'hard ERM', the E3 central banks were heavily influenced by German monetary policy. Further, using the Bundesbank's policy rule as a benchmark, they found that at the time of the EMS collapse, interest rates in each of the E3 countries were much higher than domestic macroeconomic conditions warranted. Taken all together, the results lend support to the view that some form of inflation targeting may be superior to fixing exchange rates, as a means to gain a nominal anchor for monetary policy.

The evidence also suggests that these central banks have been forward looking: they respond to anticipated inflation as opposed to lagged inflation. It first presents estimable policy reaction function. The baseline specification has a central bank adjust the nominal short-term interest rate in response to the gaps between expected inflation and output and their respective targets. It is essentially a forward-looking version of the simple backward looking reaction function popularized by Taylor.

Given this background scenario, policy reaction works as follows: They assume that within each operating period the central bank has a target for the nominal short term interest rate, r_t^*, that is based on the state of the economy. In the baseline case, we assume that the target depends on both expected inflation and output. Specifically,

$$r_t^* = \bar{r} + \beta(\mathrm{E}[\pi_{t+n} / \Omega_t] - \pi^*) + \gamma(\mathrm{E}[y_t / \Omega_t] - y_t^*) \qquad (4.1)$$

where $\bar{r}$ is the long-run equilibrium nominal rate, π_{t+n} is the rate of inflation between periods t and t+n, y_t is real output, and π^*, y_t^* are respective bliss points for inflation and output. We assume that y_t^* is given by potential output, defined as the level that would arise if wages and prices were perfectly flexible. In addition, E is the expectation operator and Ω_t is the information available to the central bank at the time it sets interest rates. It is highly possible that when choosing the target interest rate, the central bank may not have direct information about the current values of either output or the price level. Their specification allows for this possibility.

It is instructive to consider the implied target for the ex ante real interest rate, $rr_t \equiv r_t - E[\pi_{t+n} / \Omega_t]$. Rearranging Eq. (4.1) yields:

$$rr_t^* = \overline{rr} + (\beta - 1)(\mathrm{E}[\pi_{t+n} / \Omega_t] - \pi^*) + \gamma(\mathrm{E}[y_t / \Omega_t] - y_t^*) \qquad (4.2)$$

where $\overline{rr}$ is the long-run equilibrium real rate of interest. Given the economic environment we are presuming, purely real factors determine $\overline{rr}$. According to Eq. (4.2), the target real rate adjusts relative to its natural rate in response to departures of either expected inflation or output from their respective targets. A straightforward but critical point is that the magnitude of the parameter β is key. If β>1, the target real rate adjusts to stabilize inflation, as well as output (given $\gamma > 0$). With β<1, it instead moves to accommodate changes in inflation: Though the central bank raises the nominal rate in response to an expected rise in inflation, for example, it does not increase it sufficiently to keep the real rate from declining. In this 'accommodative' regime, self-fulfilling bursts of inflation and output may be possible (Bernanke and Woodford, 1996; Clarida et al.,1997). The estimated magnitude of the parameter β thus provides an important yardstick for evaluating a central bank's policy rule.

Clarida et al. (1998) then proceed to an empirical specification. An immediate concern is that a simple rule like Eq. (4.1) cannot capture the tendency of central banks to smooth changes in interest rates (Goodfriend, 1991). Traditional explanations for smoothing interest rate changes include: fear of disrupting capital markets, loss of credibility from sudden large policy reversals, the need for consensus building to support a policy change, etc. Explicitly capturing these factors is obviously quite difficult. Instead, they simply assume that the actual rate partially adjusts to the target, as follows:

$$r_t = (1 - \rho) r^* + \rho r_{t-1} + \upsilon_t \qquad (4.3)$$

where the parameter $\rho \in [0,1]$ captures the degree of interest rate smoothing.

The specification also includes an exogenous random shock to the interest rate, υ_t. Importantly, they assume that υ_t is i.i.d. Several interpretations are possible here. First, υ_t could reflect a pure random component to policy, of the type stressed in the recent identified VAR literature on monetary policy. Second, it could arise because the central bank imperfectly forecasts idiosyncratic reserve demand and, for some reason, does not instantly supply reserves to offset the shock. Under this scenario, the interest rate jumps in response to an unexpected movement in inflation and output.

To obtain an estimable equation, they first define $\alpha \equiv \bar{r} - \beta\pi^*$ and $x_t \equiv y_t - y_t^*$. They then rewrite Eq. (4.1) as:

$$r_t^* = \alpha + \beta \mathrm{E}[\pi_{t+n} / \Omega_t] + \gamma \mathrm{E}[x_t / \Omega_t]. \qquad (4.4)$$

Combining the target model (4.4) with the partial adjustment mechanism (4.3) yields:

$$r_t = (1-\rho)(\alpha + \beta \mathrm{E}[\pi_{t+n} / \Omega_t] + \gamma \mathrm{E}[x_t / \Omega_t]) + \rho r_{t-1} + \upsilon_t \qquad (4.5)$$

Finally, they eliminate the unobserved forecast variables from the expression by rewriting the policy rule in terms of realized variables as follows:

$$r_t = (1-\rho)\alpha + (1-\rho)\beta\pi_{t+n} + (1-\rho)\gamma x_t + \rho r_{t-1} + \varepsilon_t \qquad (4.6)$$

where the error term $\varepsilon_t \equiv -(1-\rho)\{\beta(\pi_{t+n} - \mathrm{E}[\pi_{t+n} / \Omega_t]) + \gamma(x_t - \mathrm{E}[x_t / \Omega_t])\} + \upsilon_t$ is a linear combination of the forecast errors of inflation and output and the exogenous disturbance υ_t.

Their econometric approach relies on the assumption that, within out short samples, short-term interest rates and inflation are I(0). Standard Dickey-Fuller tests of the null that inflation in the G3 countries is I(1) is rejected in favor of the alternative of stationarity. Also for Germany, they reject that the short-term interest rate is I(1). For the US and Japan there is less evidence against the null that short-term interest rates are I(1). However, they know that the Dickey-Fuller test has low power against the alternative of stationarity for the short sample they are studying.

Finally, they set u_t to be a vector of variables within the central bank's information set at the time it chooses the interest rate (i.e., $u_t \in \Omega_t$) that are orthogonal to ε_t. Possible elements of u_t include any lagged variables that help forecast inflation and output, as well as any contemporaneous variables that are uncorrelated with the current interest rate shock υ_t. Then since $\mathrm{E}\left[\varepsilon_t / u_t\right] = 0$, Eq. (4.6) implies the following set of orthogonal conditions that they exploit for estimation:

$$\mathrm{E}\left[r_t - (1-\rho)a - (1-\rho)\beta\pi_{t+n} - (1-\rho)\gamma\hat{y}_t - \rho r_{t-1} / u_t\right] = 0 \qquad (4.7)$$

In another work, Clarida et al. (1999) summarized what they had learned from the recent research on monetary policy. They reviewed the process that had been made and also identified the central questions that remain. To organize the discussion, they exposited the monetary policy design problem in a simple theoretical model. They started with a stripped-down baseline model in order to characterize a number of broad principles that underlie optimal policy management. They then considered the implications of adding various real world complications. Finally, they showed how the predictions from theory square with policy-making in practice. From their extensive work, Clarida et al. concluded in a number of key results:

1. To the extent cost-push inflation is present there exists a short run trade-off between inflation and output variability.
2. The optimal policy incorporates inflation targeting in the sense that it requires to aim for convergence of inflation to its target over time. Extreme inflation targeting, however, *i.e.*, adjusting policy to immediately reach an inflation target, is optimal under only one of two circumstances: (1) cost push inflation is absent, or (2) there is no concern for output deviations (*i.e.*, α=0).
3. Under the optimal policy, in response to a rise in expected inflation, nominal rates should rise sufficiently to increase real rates. Put differently,

in the optimal rule for the nominal rate, the coefficient on expected inflation should exceed unity.

4. The optimal policy calls for adjusting the interest rate to perfectly off-set demand shocks, g_t, but perfectly accommodate shocks to potential output, z_t, by keeping the nominal rate constant.
5. If the central bank desires to push output above potential, then under discretion a suboptimal equilibrium may emerge with inflation persistently above target, and no gain in output.
6. If price setting depends on expectations of future economic conditions, then a central bank that can credibly commit to a rule faces an improved short-run trade-off between inflation and output. This gain from commitment rises even if the central bank does not prefer to have output above potential.
7. The globally optimal policy rule under commitment has the central bank partially adjust demand in response to inflationary pressures. The idea is to exploit the dependence of current inflation on expected future demand.
8. With imperfect information, stemming either from data problems or lags in the effect of policy, the optimal policy rules are the certainty equivalent versions of the perfect information case. Policy rules must be expressed in terms of the forecast of target variables as opposed to the ex post behavior. Using observable intermediate targets, such as broad money aggregates is a possibility, but experience suggests that these indirect indicators are generally too unstable to use in practice.
9. Large unobservable shocks to money demand produce high volatility of interest rates when a monetary aggregate is used as the policy instrument. It is largely for this reason that an interest rate instrument may be preferable.

Finally, Clarida et al. (2000), estimate a forward-looking monetary policy reaction function for the postwar United States economy, before and after Volcker's appointment as Fed chairman in 1979. Their results point to substantial differences in the estimated rule across periods. In particular, interest rate policy in the Volcker-Greenspan period appears to have been much more sensitive to changes in expected inflation than in the pre-Volcker period. They then compare some of the implications of the estimated rules for the equilibrium properties of inflation and output, using a simple macroeconomic model, and show that the Volcker-Greenspan rule is stabilizing.

Specifically, from the late 1960s through the early 1980s, the United States economy experienced high and volatile inflation along with several severe recessions. Since the early 1980s, however, inflation has remained steadily low, while output growth has been relatively stable. Many economists cite supply shocks-and oil price shocks, in particular-as the main force underlying the instability of the early period. It is unlikely, however, that supply shocks alone could account for the observed differences between the two eras. In this paper, it is explored the role of monetary policy. Authors go to argue that this difference could be an important source of the shift in the way monetary policy was conducted pre- and post-1979.

First of all, they identify how monetary policy differed before and after Volcker came to office by estimating policy rules for each era. Authors use the same policy rule specification, as their 1998 paper [eq. (4.1)-(4.5)], and estimate a general type of rule that treats the Federal Funds rate as the instrument of monetary policy. The rule assumes forward-looking behavior on the part of the central bank and calls for adjustment of the Funds rate to the gaps between *expected* inflation and output and their respective target levels.

The key difference in the estimated policy rules across time involves the response to expected inflation. They find that the Federal Reserve was highly “accommodative” in the pre-Volcker years: on average, it let short-term interest rates decline as anticipated inflation rose. While it raises nominal rates, it typically did so by less than the increase in expected inflation. On the other hand, during the Volcker-Greenspan era the Federal Reserve adopted a proactive stance toward controlling inflation: it systematically raised real as well as nominal short-term interest rates in response to higher expected inflation. Their results thus lend quantitative support to the popular view that not until Volcker took office did controlling inflation become the organizing focus of monetary policy.

The second part of the paper presents a theoretical model designed to fresh out how the observed changes in the policy rules account for the change in macroeconomic performance. After log-linearization around a zero inflation steady state, the model's equilibrium conditions are summarized by the following equations (ignoring uninteresting constants):

$$\pi_t = \delta E\{\pi_{t+1} / \Omega_t\} + \lambda(y_t - z_t) \qquad (4.8)$$

$$y_t = E[y_{t+1} / \Omega_t] - (1/\sigma)(r_t - E[\pi_{t+1} / \Omega_t]) + g_t \qquad (4.9)$$

$$r_t^* = \beta E[\pi_{t+1} / \Omega_t] + \gamma x_t \qquad (4.10)$$

$$r_t = \rho r_{t-1} + (1-\rho) r_t^* \qquad (4.11)$$

Equation (4.8) describes the change in the aggregate price level as a function of expected future inflation and the deviation of (log) output y_t from its natural rate z_t, where the latter is defined as the level of output that would obtain under fully flexible prices. Eq. (4.9) combines a standard Euler equation for consumption with a market clearing condition, determining the current output gap as a function of the ex ante real rate and expected future output. Equation (4.10) and (4.11) specify the policy rule [like eq. (4.1)-(4.5)].

By this model, they show that the estimated rule for the pre-Volcker period permits greater macroeconomic instability than does the Volcker-Greenspan rule. It does so in two distinct respects.

First, the pre-Volcker rule leaves open the possibility of bursts of inflation and output that result from self-fulfilling changes in expectations. These sunspot fluctuations may arise under this rule because individuals (correctly) anticipate that the Federal Reserve will accommodate a rise in expected inflation by letting short-term real interest rates decline (which in turn stimulates the rise in aggregate demand and inflation). On the other hand, self-fulfilling fluctuations cannot occur under the estimated rule for the Volcker-Greenspan era since, within this regime, the Federal Reserve adjusts interest rates sufficiently to stabilize any changes in expected inflation.

Second, the pre-Volcker rule is less effective than the Volcker-Greenspan rule at mitigating the impact of fundamental shocks to the economy. That is, holding constant the volatility of exogenous fundamental shocks, the economy exhibits greater stability under the post-1979 rule than under a rule that closely approximates monetary policy pre-1979.

5. FAVERO'S DISCUSSION PAPER

Carlo Favero with a recent discussion paper (June 2001), which will be my benchmark, tries to answer the following question: "Does macroeconomics help us to understand the term structure of interest rates? ".

In this paper, he wants to provide an assessment of the expectations model validity of the term structure of interest rates using information from monetary policy by including the central bank reaction function in small macro models. His approach differs from the limited information approach taken commonly in the literature because it allows testing directly the prediction of the relevant model using the information generated by this. The future path of policy rates is derived consistently with the adopted macro models rather than using the assumption of rational expectations. He believes that his empirical results strikingly contradict the previous empirical results. The whole term structure of US policy rates behaves in a way that is statistically consistent with the expectations theory.

Specifically, these macro models are closed and by stimulating them forward it is possible for someone to derive the full forward path of short-term interest rates and hence to construct any long-term interest rate consistent with the expectations model (pure expectations hypothesis). Comparing then the observed long-term rates with the simulated ones in any confidence interval can immediately perform a direct test of the expectational model, based on full information. Importantly is that the presence of term premia does not affect this procedure since the macro model used to derive directly the full path of future policy rates is risk-free. Therefore, if the expectations theory is true then the term premium can be possibly measured as the difference between actual and simulated yields on long-term bonds.

The discussion paper by Favero tests this expectations model within a backward looking small macro model for US and German economy. His empirical analysis of the US economy based on the following estimated model built-in 4 equations:

$$\pi_t = \underset{(0.054)}{1.38}\,\pi_{t-1} + \left(1 - \underset{(0.054)}{1.38}\right)\pi_{t-2} + \underset{(0.01)}{0.03}\,y_{t-1} + u_{1t}$$

$$S.E. = 0.29$$

$$y_t = \underset{(0.06)}{0.12} + \underset{(0.02)}{0.92}\, y_{t-1} - \underset{(0.156)}{0.39}\left(\bar{r}_{t-3} - \bar{\pi}_{t-3}\right) + u_{2t}$$

$$S.E. = 0.63$$

$$r_{t,t+1} = \left(1 - \underset{(0.01)}{0.92}\right)\left[\underset{(0.54)}{4.58} + \underset{(0.5)}{1.12}\, E_t\left(\pi_{t+12} - \pi^*\right) + \underset{(0.46)}{1.43}\, E_t y_t\right] + \underset{(0.01)}{0.92}\, r_{t-1,t} + u_{3t}$$

$$S.E. = 0.36$$

$$r_{t,t+n} = \frac{1}{n}\sum_{i=0}^{n-1} E_t r_{t+i,t+i+1}$$

where $r_{t,t+i}$ is the continuously compounded interest rate observed at month t on a bond with maturity at month $t+i$ in an annual base, π_t is annual CPI inflation at time t, $\bar{r}_t = \frac{1}{3}\left(r_{t,t+1} + r_{t-1,t} + r_{t-2,t-1}\right)$, $\bar{\pi}_t = \frac{1}{3}\left(\pi_{t,t+1} + \pi_{t-1,t} + \pi_{t-2,t-1}\right)$ and y_t is the output gap at time t, defined as $\ln(Y_t) - (ln(Y_t))^*$, Y_t is the seasonally adjusted industrial and $(ln(Y_t))^*$ is Hodrick-Prescott filtered log of industrial productions, giving a stochastic trend. Finally, π^* is the target level of inflation, which he set to an annual rate of 2%.

The first two equations, estimated individually by OLS over the sample 1976:1-1999:12 and offer a simple supply-demand empirical model of US output and inflation, stable over time, and capable of matching the properties of more articulated models such as the FRB/US macro econometric model. Though it has been noted that a forward-looking version of this model is more justifiable in terms of macroeconomic foundations (Mc Callum and Nelson, 1999a), Favero's estimated supply equation is backward looking. In practice, he considers that as his empirical evidence will be based on forward simulation of the small macro econometric models, using a backward or a forward version of the supply function does not alter substantially the conclusions.

The third equation in the model is a standard forward-looking Taylor rule with interest rate smoothing, proposed by Clarida, Gali and Gertler (1998,1999,2000), and successfully implemented to model monetary policy in a number of industrialized countries. The error term u_{3t} can be rationalized by considering that central bank applies a partial adjustment model, around a target interest rates, described by the following two equations (like Clarida et al. (1998)):

$$r_{t,t+1}^{*} = \bar{r} + a_1 \mathrm{E}_t(\pi_{t+12} - \pi^{*}) + a_2 \mathrm{E}_t(y_t - y_t^{*})$$

$$r_{t,t+1} = (1-\rho) r_{t,t+1}^{*} + \rho r_{t-1,t} + u_{1t}$$

where r^* is the target interest rate and $\bar{r}$ is the equilibrium value for r^*.

Note that, if the central bank is setting interest rates optimally, then a_1, a_2 are convolutions of parameters describing the structure of the economy and the preferences of the monetary policy maker. He estimated the reaction function over the sample 1984:1-1999:12 because an interest rate rule cannot describe effectively the behavior of the Federal Reserve from 1979 to 1982, when the operating procedure followed by the Fed is better by reserves targeting rather than interest rate targeting. Moreover, Clarida et al. (1999) have concluded that monetary was far less aggressive in fighting inflation in the pre-Volcker era than in the later period, since it was relatively more focused on output stabilization and it allowed real interest rates to decrease in presence of inflationary shocks. Favero says that such policy, described by a value of a_1 smaller than one would prevent his small macro model to converge when simulated, therefore it would not allow deriving model consistent long-term interest rates.

Favero estimates his monetary policy rule considering the one-month rate as policy determined, fully in line with results already available in the empirical literature (see Clarida et al., 1999). Similarly he has chosen to include a one-year ahead horizon for the expected inflation.

The fourth equation generates interest rates for any maturity, consistently with the pure expectations theory. Note that the equation does not contain any estimated parameter and does not feature feedback with the other three equations in the system. Therefore, at any point in time the small macro model consisting of demand equation, supply equation and the interest rate rule can be solved forward to derive the full path of all the monthly future policy rates. Given future policy rates, long-term interest rates at all maturities can be derived using the last equation. Stochastic simulation of the model generates confidence intervals around the point estimates. A direct test of the expectation model is then feasible by comparing actual long-term rates with those simulated by the model and its confidence intervals. Within this framework, under the null hypothesis that agents form their expectations for future policy rates, we are able to measure directly expectations and to provide a direct test for the expectations model.

As we have already said, Favero's empirical evidence is based on solving forward recursively and stochastically his simple model, to derive the full path of expected policy rates and their associated standard errors. To mimic at any point in time the decision of agents who form expectations on the basis of the available data, he re-simulated the model as he move forward along the sample.

The figures show clearly that the expectation model is never rejected at all sample points and at all frequencies (from 3-month to 10-year period forward with the other maturities to be 6-month, 1, 2, 3, 5 and 7-year). Moreover, there is a clear tendency for the observed interest rates to commove with the simulated interest rates at the mid-point of the 95% confidence interval. Clearly, the 95% confidence intervals get larger as the maturity of the relevant interest rate get longer, but the observed interest rates get nowhere near the upper and the lower band, with some few exceptions only for the shorter maturity, three-month.

His empirical test based on the use of a very simple small macro model to generate expectations of future policy rates, contradicts all the available evidence based on limited-information procedures. Under the null hypothesis that agents use his three-equation models to generate expected future policy rates, the expectation model of the term structure delivers confidence interval for long-term rates at all frequencies which always contain the observed long-term rates. It is interesting for his point of view to consider that the time-series behavior of the difference between actual and simulated series, which, under the null of validity of the expectation model, can be naturally interpreted as the term premium. The series have a positive mean, monotonically increasing with maturity. The correlation pattern is such that correlation among term premia at any two different maturities decreases with the distance between maturities.

The following regression for 10-year period forward is an example of this study:

$$r_{t,t+120} = \underset{(0.47)}{2.73} + \underset{(0.07)}{0.78}\, r^{*}_{t,t+120} + u_t \qquad (5.1)$$

$$R^2 = 0.38, \sigma = 1.5, S.D. of.dep.\mathrm{var.} = 7.67$$

$$r^{*}_{t,t+120} = \frac{1}{120}\sum_{i=0}^{119} E_t r_{t+i,t+i+1}$$

The regression of the actual redemption yield on 10-year bonds on the redemption yield derived using model-based simulated future policy rates delivers strongly significant results. The results of the estimation of (5.1) are much more in line with the expectations theory than those on the estimation of the equation:

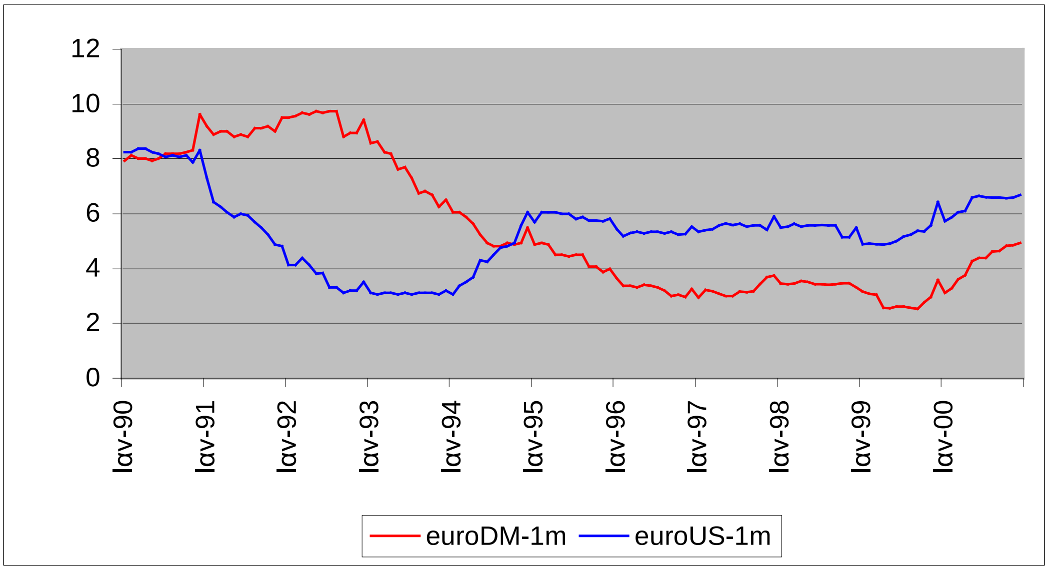
$$r_{t+1,t+121} - r_{t,t+120} = -\underset{(0.036)}{0.022} - \underset{(2.48)}{0.54}\frac{1}{120}\left(r_{t,t+120} - r_{t,t+1}\right) + u_{t+1} \qquad (5.2)$$

$$R^2 = 0.0002, \sigma = 0.34.$$

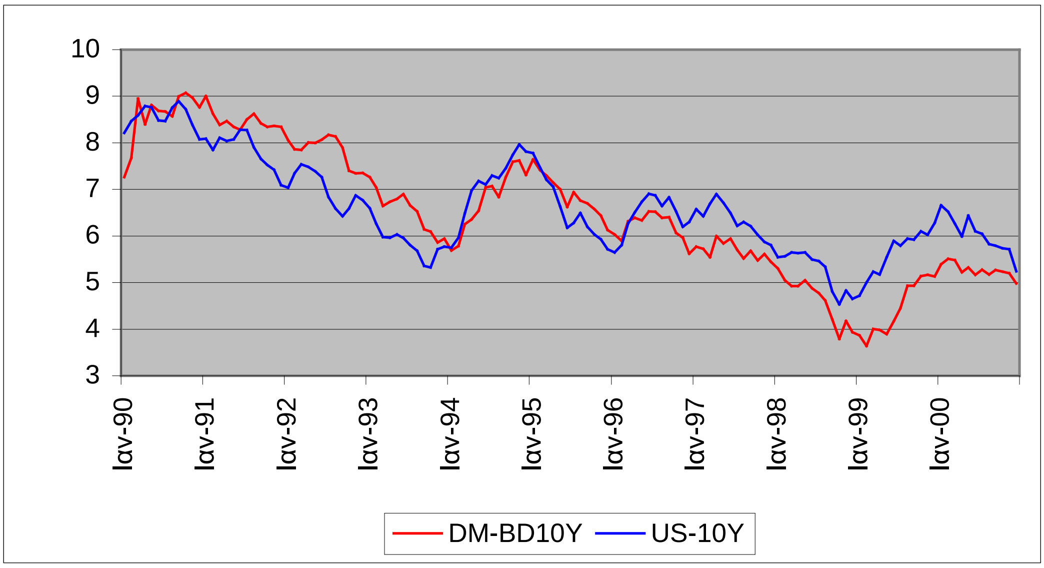
Favero estimates equation (5.2) in order to show that the traditional test based on the regression of the short-term change in the long-term yields on the lagged yield spread fares very poorly in predicting changes in long-term yields leading to the rejection of the expectations theory. For that purpose, he considers the same sample used for his simulation-based results, 1984:1-1999:12.

At (5.1) the coefficient on $r^{*}_{t,t+120}$ is significantly different from zero and as high as 0.78, but it is also significantly different from one. This result can be explained by the existence of a term-premium negatively and moderately correlated with the future path of interest rates. He noticed that the apparent contradiction between the weak rejection of the theory in the regression based results and the lack of rejection of the theory in the simulation results can be explained considering that the uncertainty surrounding the long-term yields consistent with the expectations theory is not considered in the regression results. In fact, equation (5.1) is estimated using as a $r^{*}_{t,t+120}$ just the average of the stochastically simulated long-term yields.

Then, Favero extends the model with some international evidence, which on the behavior of long-term and short-term interest rates in the course of the nineties poses some interesting questions for the expectation model. He reports the one-month rates on Eurodollars and Euro-D.Mark, his measures of the policy rates, and the redemption yields on ten-year benchmark Treasury bonds denominated in the two currencies. The figures clearly show that there is a strong tendency for the long-term rates to co-move, even in presence of remarkable asymmetry in the stance of monetary policy.



US and German policy rates



US and German 10-year bond yields

The evidence from 1994 is particularly interesting because the increase in the US long-term bond yields generated by the typical episode of "inflation scare" was entirely shared by German long-term bond yields, although the stances of German and US monetary policy were completely different in the course of that year. In fact, German bond yields began to decrease only towards the end of the year when the restrictive US monetary policy became successful in curbing the inflation scare. The symmetric behavior of long-term bond yields in presence of asymmetric monetary policy poses some serious challenge for a limited information approach to the expectation model. In fact, the validity of the model in both countries requires that in at least one of the two the current monetary policy stance is not informative on the stance of future monetary policy. The information generated by a small macro model might help in understanding the behavior of bond yields. To assess this possibility he extended his four-equation, closed economy model for the US, to the following eight-equation model for the US and Germany:

$$\pi_t^{US} = \underset{(0.054)}{1.38}\,\pi_{t-1}^{US} + \left(1 - \underset{(0.054)}{1.38}\right)\pi_{t-2}^{US} + \underset{(0.01)}{0.03}\,y_{t-1}^{US} + u_{1t}$$

$$S.E. = 0.29$$

$$y_t^{US} = \underset{(0.06)}{0.12} + \underset{(0.02)}{0.92}\, y_{t-1}^{US} - \underset{(0.156)}{0.39}\left(\bar{r}_{t-3}^{US} - \bar{\pi}_{t-3}^{US}\right) + u_{2t}$$

$$S.E. = 0.63$$

$$r_{t,t+1}^{US} = \left(1 - \underset{(0.01)}{0.92}\right)\left[\underset{(0.54)}{4.58} + \underset{(0.5)}{1.12}\, E_t\left(\pi_{t+12}^{US} - \pi^*\right) + \underset{(0.46)}{1.43}\, E_t y_t^{US}\right] + \underset{(0.01)}{0.92}\, r_{t-1,t}^{US} + u_{3t}$$

$$S.E. = 0.36$$

$$r_{t,t+n}^{US} = \frac{1}{n}\sum_{i=0}^{n-1} E_t r_{t+i,t+i+1}^{US}$$

$$\pi_t^{GER} = \underset{(0.016)}{0.95}\, \pi_{t-1}^{GER} + \underset{(0.008)}{0.02}\, \pi_{t-9}^{US} + \underset{(0.01)}{0.028}\, y_{t-1}^{US} + \underset{(0.0009)}{0.018}\, y_{t-11}^{GER} + u_{4t}$$

$$S.E. = 0.29$$

$$y_t^{GER} = \underset{(0.15)}{0.18} + \underset{(0.045)}{0.60}\, y_{t-1}^{GER} - \underset{(0.04)}{0.05}(r_{t-10}^{GER} - \pi_{t-10}^{GER}) + \underset{(0.05)}{0.11}\, y_{t-3}^{US} + u_{5t}$$

$$S.E. = 1.4$$

$$r_{t,t+1}^{GER} = \left(1 - \underset{(0.03)}{0.92}\right)\left[\underset{(0.29)}{5.10} + \underset{(0.41)}{1.61}\, E_t\left(\pi_{t+12}^{GER} - \pi^*\right) + \underset{(0.18)}{0.32}\, E_t y_t^{GER}\right] + \underset{(0.03)}{0.92}\, r_{t-1,t}^{GER} + u_{6t}$$

$$S.E. = 0.48$$

$$r_{t,t+n}^{GER} = \frac{1}{n}\sum_{i=0}^{n-1} E_t r_{t+i,t+i+1}^{GER}$$

As we can see, equations for the US economy are unaltered, while equation for the German economy have been specified so that US output gap and inflation have a potential role as a leading indicators for the German macroeconomic variables. Only local expected inflation and output gap enter the reaction function estimated by the Bundesbank. The estimation results point towards a significant role of US macroeconomic variables in determining European variables, moreover the weights attached by the German central bank to inflation and output gap give more prominence to inflation than in the US case. This may reflect either different central bank preferences or a different structure of the US and German economy.

By this extension, Favero concentrates on the capability of the expectation model to predict the behavior of long-term bond yields applying the stochastic simulation framework proposed in the previous section. He first assess comparatively the behavior of actual and simulated German ten-year benchmark bond-yields and as for US long-term bond-yields, the expectation model is not rejected by the data. In fact, the simulated series tracks very well the observed one, which gets nowhere near the upper and lower bounds of the 95% confidence interval. Then, he reports the difference for actual and simulated series on US and German data, which, under the null of the validity of the expectations model, is interpreted as the term premium. The figure shows a strong correlation between the two series, which stands at 0.72, although the German variable shows a higher mean than the US one, 1.25 versus 0.43.

The conclusion is that the observed behavior of US and German long-term bond yields is compatible with the expectations model of the term structure when expectations are formed by simulating a very simple macro model of the US and German economy. Favero mentions that these results are important because the use of these small macro models which try to define the information set relevant to agents changes completes the outcome of traditional testing procedures based on a limited information approach. He concludes that testing the expectations theory by deriving the future path of policy rates consistently with a macro model delivers empirical results drastically different from the regression based test of the same theory where future expected rates are simply substituted by *ex-post* observed rates.

6. THESIS EMPIRICAL FRAMEWORK

Within this thesis study, I am going to examine the predictive power of the interest rate structure Favero's macro models suggest, using the traditional test of the expectation hypothesis. This set of macro models seems to be free of time-varying term premia and it is interesting for someone to test whether the amount of information included in their yield curve can explain better future changes in short rates than actual long-term bonds. Therefore, we are looking from macro models results much closer to expectations hypothesis than these of yields to maturity of long-term bonds.

The great difference between my work and Favero's approach is that the simulated term structure proposed by his macro models does not examined for its effectiveness to predict *future* changes of interest rates. Favero works on levels and he just compares simulated rates with observed long-term bonds. This is a different approach of the validity of pure expectations hypothesis, which, unfortunately, does not answer some crucial questions, both for forecasting and interpreting shifts in the yield curve. Since the most prevalent explanation of fluctuations in the yield curve is the expectations theory, which posits that the slope of the yield curve reflects the market expectation of the future change in interest rates, we always try to find the dominant force determining current long-term interest rates.

Favero simulates forward the macro models and derives the full forward path of short-term interest rates, in order to construct any long-term interest rate consistent with the expectations model (pure expectations hypothesis). Then, he just compares the observed long-term rates with the simulated ones in the 95% confidence interval for a direct test of the expectational model, based on full information:

$$r^{*}_{t,t+n} = \frac{1}{n}\sum_{i=0}^{n-1} E_t r_{t+i,t+i+1} \quad (\textit{pure expectations hypothesis})$$

$$r_{t,t+n} = \alpha + \beta r^{*}_{t,t+n} + u_t \qquad (6.1)$$

so he tests the prediction of the relevant model using the information generated by this model, while the future path of policy rates is derived consistently with the adopted macro models rather than using the assumption of rational expectations. While the output figures show clearly that the expectation model is never rejected at all sample points and at all frequencies, he noticed that the apparent contradiction between the weak rejection of the theory in the regression based results and the lack of rejection in the simulation results can be explained considering that the uncertainty surrounding the long-term yields consistent with the expectations theory is not considered in the regression results. In fact, equation (6.1) is estimated using as a $r^{*}_{t,t+n}$ just the average output of the stochastically simulated long- term yields without considering the upper and lower bounds of the confidence interval.

What about the future path of actual short-term rates, though? Favero's approach has some disadvantages, indeed. No sign about the predictive power of these models for future movements of actual interest rates appears. In addition, equation (6.1), which examines, on levels, the relation between actual and simulated long-term bonds, could be a long-run cointegration relationship, since no prerequisite of the stationarity of these long-term yields is secured.

On the other hand, while the assessment of the validity of the expectation models using macroeconomics gives us results that cannot reject the expectations model, we can use these macro models to our following framework. Specifically, we can use the simulated term structure of interest rates in comparison with the observed one on a classical empirical approach and examine by *ex post tests of expectations hypothesis* its predictive ability for long-run changes in short rates. For this matter, we use regressions models that take into account the full term structure path derived by macroeconomics rather than this of yields to maturity of long-term bonds. Thus, I'II focus in classical regressions with the general form of:

$$\sum_{i=0}^{n-1} w_i r_{t+i} - r_t = a_n + b_n \left[R_t^{(n)} - r_t \right] + w_{t+n} \qquad (6.2)$$

while, $R_t^{(n)}$ is now the simulated long rate $r_{t,t+n} = \frac{1}{n}\sum_{i=0}^{n-1} E_t r_{t+i,t+i+1}$, $i = 1,2,...,120$ (i=month) according to the pure expectations hypothesis that Favero uses to construct the theoretical term path. Thus, the $R_t^{(n)}$ is exactly the simulated output suggested by Favero, which is tested for its predictive ability on differences (6.2) rather than on levels (6.1). On the other hand, $\sum_{i=0}^{n-1} w_i r_{t+i}$ is the *actual* geometric declining weight of interest rates on a ***short-term*** debt instrument until period *t+n* in the future. We subtract from this sum the actual r_t in order to construct the long-run changes of short interest rates, which will be the dependent variable on our examined regression.

The advantage on those regressions is that the presence of term premia does not affect the whole procedure since the macro model that is used to derive directly the full path of future policy rates is risk-free. Thus, the coefficient a_n does not imply any risk assumption, so under the pure expectations hypothesis simulated $R_t^{(n)}$ is the following:

$$R_t^{(n)} = \sum_{i=0}^{n-1} w_i E_t r_{t+i} \ , \qquad \sum_{i=0}^{n-1} w_i = 1$$

Moreover, we have mentioned that the main problem in the following formula:

$$\hat{b} \equiv \frac{\sigma^2(\mathrm{E}_t \Delta r_{t+1}/2) + 2\rho\sigma(\mathrm{E}_t \Delta r_{t+1}/2)\sigma(\theta_t)}{\sigma^2(\mathrm{E}_t \Delta r_{t+1}/2) + 2\rho\sigma(\mathrm{E}_t \Delta r_{t+1}/2)\sigma(\theta_t) + \sigma^2(\theta_t)}$$

which estimates $\hat{b}_n$ coefficient, is that $\hat{b}_n$ depends on the variations in the term premium *σ(θt)* that give downwards bias on the coefficient; the size of the bias depends on the variance of the expected change in the future short rate. In addition, $\hat{b}_n$ is equal to 1 in the absence of a time-varying premium (*σ(θt)* ^{2} = 0). In our approach, variations of the term premium *σ(θt)* are almost zero by construction.

Our previous regression models are obvious different from those of Favero's since we focus on the rate of change form of the expectations theory rather than on levels.

I examine models for both America and German economy, and the range of short-term debt instrument maturity is between 1-month and 1-year, while the simulated long-term of interest rates between 1-year and 10-years.

At first, we re-calculate Favero's empirical macro model to take the appropriate simulated long-term interest rates for our following work.

7. FAVERO'S MACRO MODEL RE-ESTIMATION

Favero's empirical macro model re-estimated by OLS over the sample 1977:1-2001:8, using the same data sources (All the macroeconomic time-series are taken from DATASTREAM, while USA interest rates at all maturities are taken from the FRED database at the website of the Federal Bank of St.Louis). Our empirical output based on the following estimated model and it is in full accordance with Favero's results:

$$\pi_t^{US} = \underset{(0.053)}{1.42}\,\pi_{t-1}^{US} + \left(1 - \underset{(0.053)}{1.42}\right)\pi_{t-2}^{US} + \underset{(0.01)}{0.02}\,y_{t-1}^{US} + u_{1t}$$

$$S.E. = 0.29$$

$$y_t^{US} = \underset{(0.06)}{0.08} + \underset{(0.02)}{0.94}\,y_{t-1}^{US} - \underset{(0.19)}{0.37}\left(\overline{r}_{t-3}^{US} - \overline{\pi}_{t-3}^{US}\right) + u_{2t}$$

$$S.E. = 0.64$$

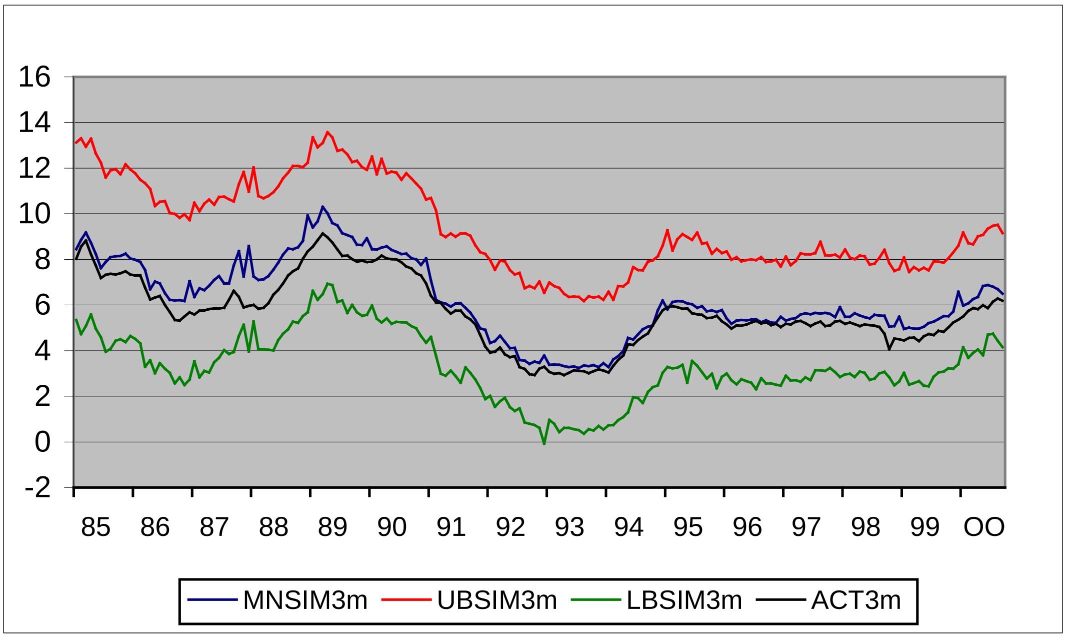
$$r_{t,t+1}^{US} = \left(1 - \underset{(0.017)}{0.92}\right)\left[\underset{(0.8)}{5.35} + \underset{(0.26)}{0.97}\,E_t\left(\pi_{t+12}^{US} - \pi^*\right) + \underset{(0.49)}{1.29}\,E_t y_t^{US}\right] + \underset{(0.017)}{0.92}\,r_{t-1,t}^{US} + u_{3t}$$

$$S.E. = 0.52$$

$$r_{t,t+n}^{US} = \frac{1}{n}\sum_{i=0}^{n-1} E_t r_{t+i,t+i+1}^{US}$$

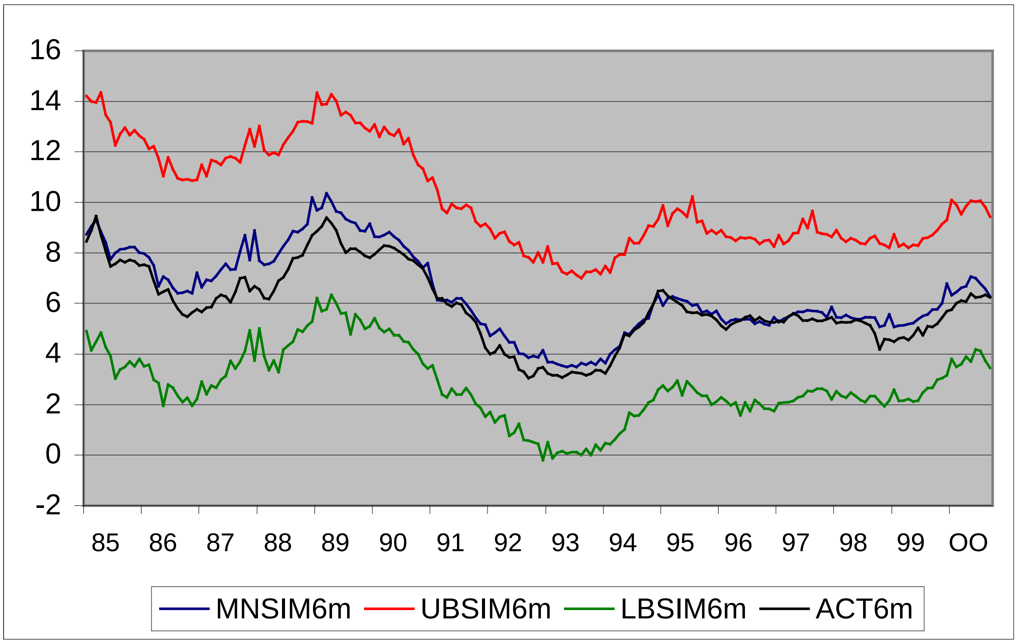
Then, we solve forward recursively and stochastically our new model, to derive the full path of expected policy rates and their associated standard errors. We implement our simulation exercise from 1985:1 to 2000:12. To mimic at any point in time the decision of agents who form expectations on the basis of the available data, we re-simulate the model as we move forward along the sample. Stochastic simulation of the model generates confidence intervals around the point estimates.

The results are reported in the following figures:

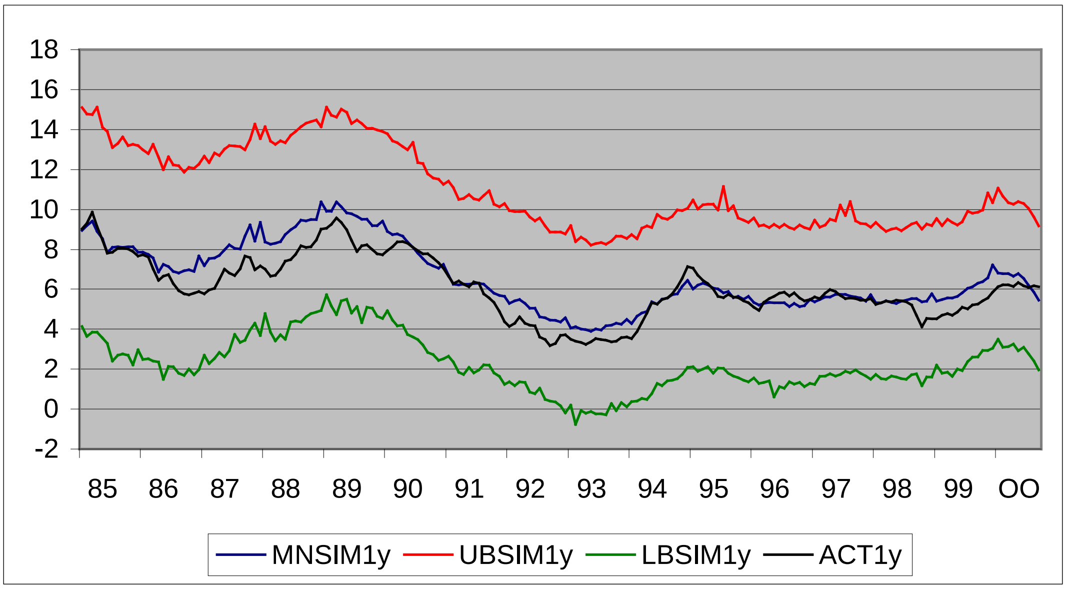


3-month

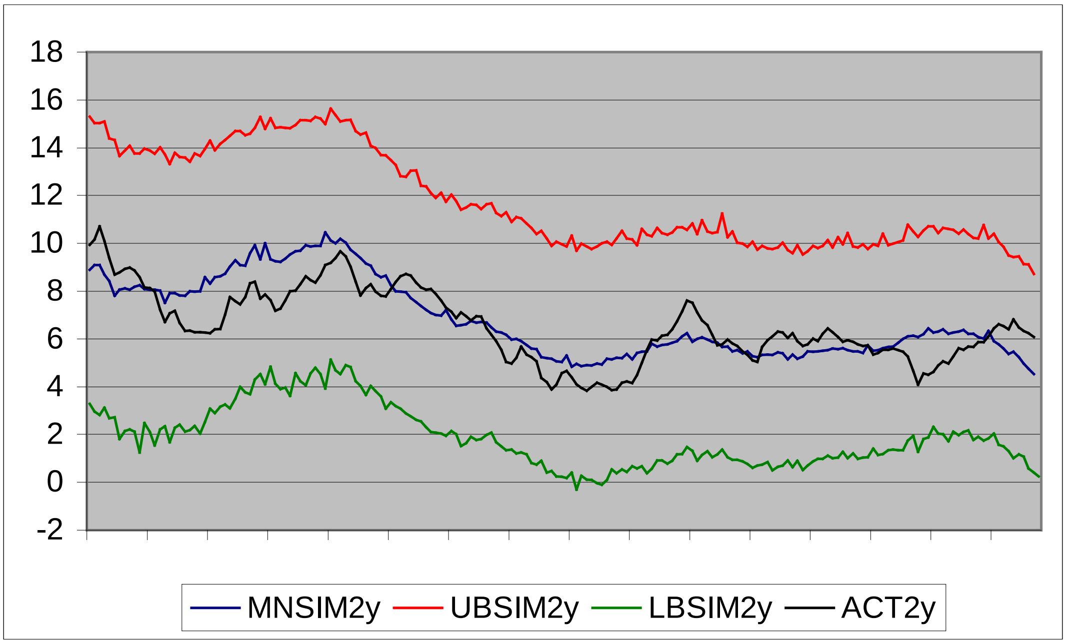
Observed and simulated interest rates for the 3-month and 6-month maturities. Observed rates are labelled ACT3m, ACT6m, simulated rates are instead labelled MNSIM3m, MNSIM6m. Simulation are stochastic so average simulated rates are reported with upper (UBSIM3m, UBSIM6m) and lower bounds (LBSIM3m, LBSIM6m) of their ninety-five per cent confidence interval



6-month

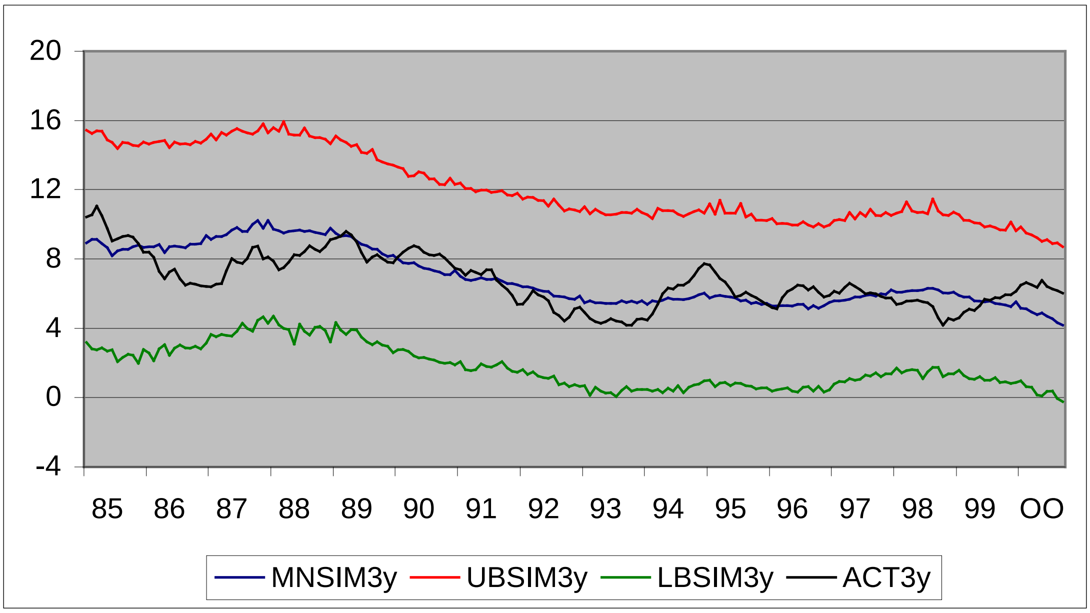


1-year

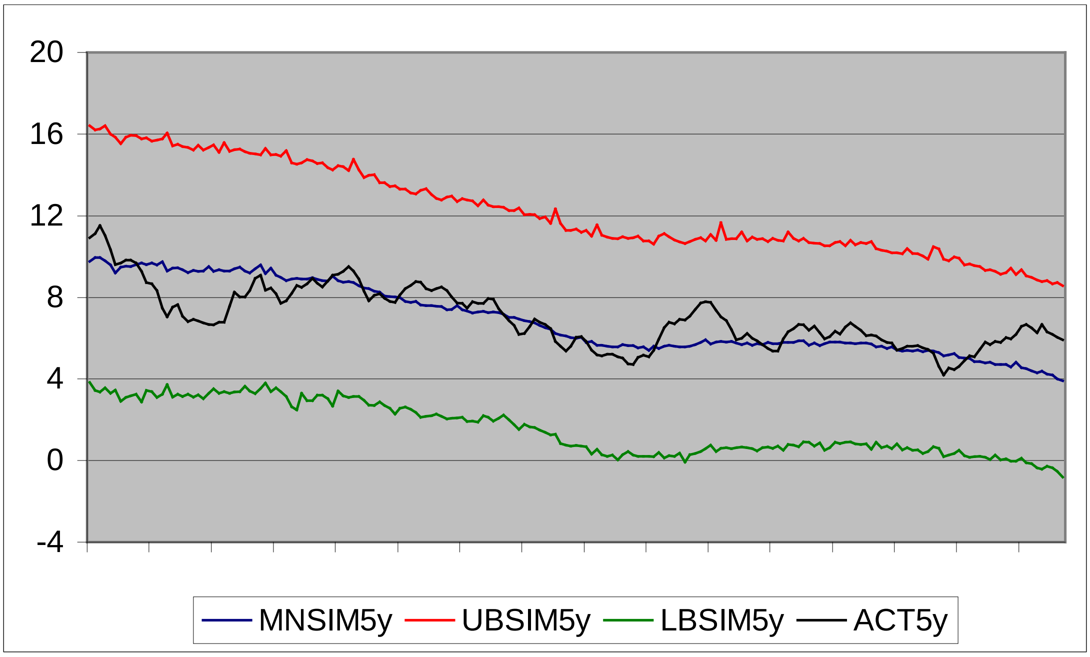


2-year

Observed and simulated interest rates for the 1-year and 2-year maturities. Observed rates are labelled ACT1y, ACT2y, simulated rates are instead labelled MNSIM1y, MNSIM2y. Simulation are stochastic so average simulated rates are reported with upper (UBSIM1y, UBSIM2y) and lower bounds (LBSIM1y, LBSIM2y) of their ninety-five per cent confidence interval

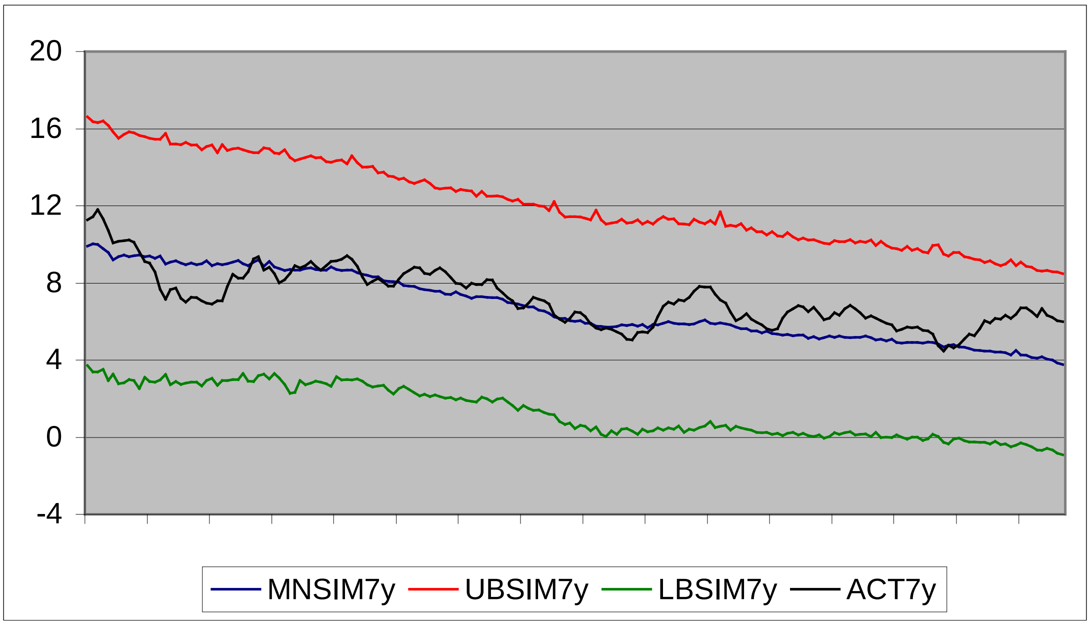


3-year

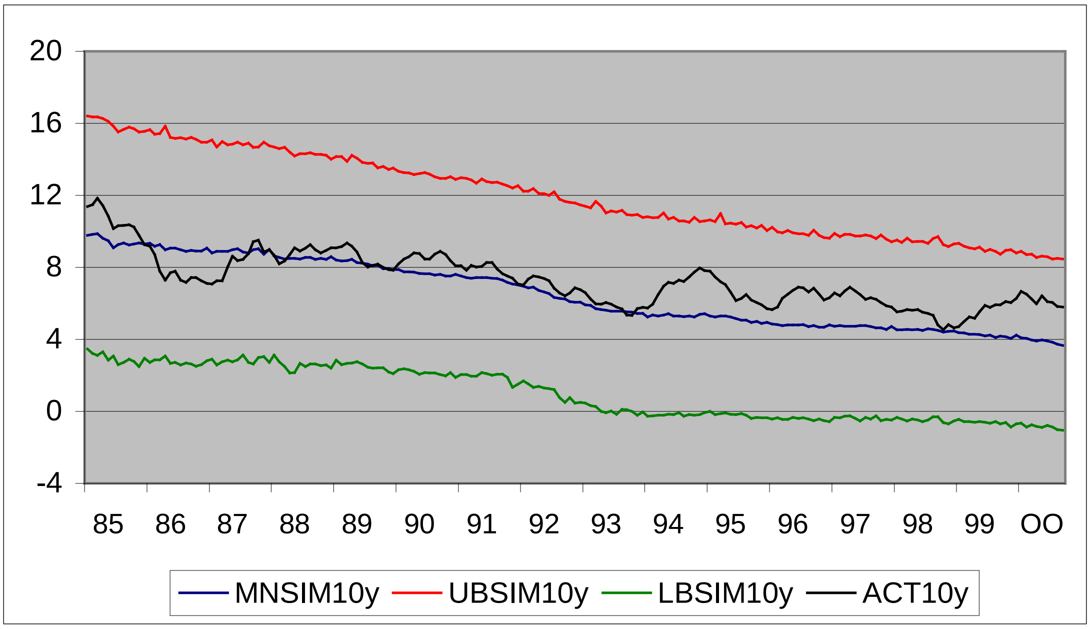


5-year

Observed and simulated interest rates for the 3-year and 5-year maturities. Observed rates are labelled ACT3y, ACT5y, simulated rates are instead labelled MNSIM3y, MNSIM5y. Simulation are stochastic so average simulated rates are reported with upper (UBSIM3y, UBSIM5y) and lower bounds (LBSIM3y, LBSIM5y) of their ninety-five per cent confidence interval



7-year



10-year

Observed and simulated interest rates for the 7-year and 10-year maturities. Observed rates are labelled ACT7y, ACT10y, simulated rates are instead labelled MNSIM7y, MNSIM10y. Simulation are stochastic so average simulated rates are reported with upper (UBSIM7y, UBSIM10y) and lower bounds (LBSIM7y, LBSIM10y) of their ninety-five per cent confidence interval

The figures show clearly that the expectation model is never rejected at all sample points and at all frequencies from 3-, 6-month to 1,2,3,5,7 and 10-years. Moreover, there is a clear tendency for the observed interest rates to commove with the simulated interest rates at the *mid-point* of the 95% confidence interval. Clearly, the 95% confidence intervals get larger as the maturity of the relevant interest rate get longer, but the observed interest rates get nowhere near the upper and the lower band. Under the null hypothesis that agents use this three-equation models to generate expected future policy rates, the expectation model of the term structure delivers confidence interval for long-term rates at all frequencies which always contain the observed long-term rates.

The following table gives the coefficients on regressions $r_{t,t+n} = \alpha + \beta r^{*}_{t,t+n} + u_t$ for all frequencies, like Favero's one for 10-years maturity:

Regression Coefficients

Regression equation $r_{t,t+n} = \alpha + \beta r^{*}_{t,t+n} + u_t$ (7.1)

		Simulated long bond $r^{*}_{t,t+n}$ maturity (months)					
Dependent variable		**3**	**12**	**24**	**36**	**60**	**120**
Actual long bond $r_{t,t+n}$ maturity	***α***	0.53 (0.09)	0.6 (0.11)	1.29 (0.29)	1.95 (0.32)	2.32 (0.27)	2.93 (0.21)
	β	0.87 (0.03)	0.88 (0.03)	0.77 (0.05)	0.69 (0.05)	0.68 (0.05)	0.67 (0.04)

The regression of the actual redemption yields on the redemption yield derived by simulated future policy rates, delivers strongly significant results.

The β coefficients on $r^{*}_{t,t+n}$ are significantly different from *zero*, but there are also significantly different from one. This result can be explained by the existence of a term-premium negatively and moderately correlated with the future path of interest rates. This apparent contradiction, as Favero notices, between the weak rejection of the theory in the regression based results and the lack of rejection in the simulation results can be explained considering that the uncertainty surrounding the long-term yields consistent with the expectations theory is not considered in the regression results. In fact, equation (7.1) is estimated using as a $r^{*}_{t,t+n}$ just the ***mean*** output of the stochastically simulated long- term yields with 95% confident intervals.

Finally, the re-calculated extension of the closed economy model to the Germany economy is the following:

$$\pi_t^{GER} = \underset{(0.016)}{0.94}\,\pi_{t-1}^{GER} + \underset{(0.01)}{0.03}\,\pi_{t-9}^{US} + \underset{(0.013)}{0.036}\,y_{t-1}^{US} + \underset{(0.01)}{0.012}\,y_{t-11}^{GER} + u_{4t}$$

$$S.E. = 0.36$$

$$y_t^{GER} = \underset{(0.2)}{0.17} + \underset{(0.05)}{0.57}\,y_{t-1}^{GER} - \underset{(0.06)}{0.05}(r_{t-10}^{GER} - \pi_{t-10}^{GER}) + \underset{(0.06)}{0.18}\,y_{t-3}^{US} + u_{5t}$$

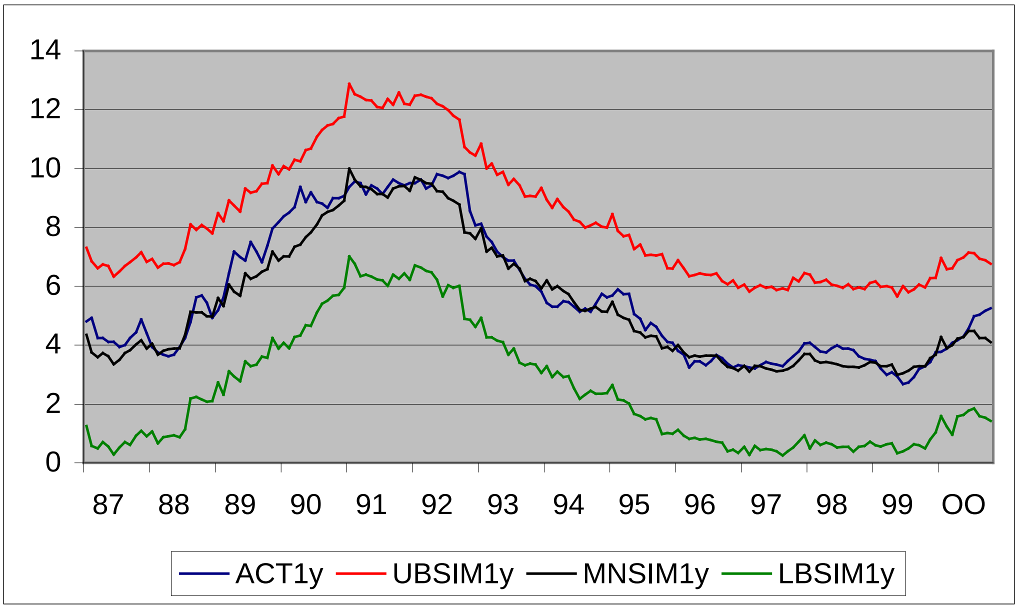
$$S.E. = 1.5$$

$$r_{t,t+1}^{GER} = \left(1 - \underset{(0.03)}{0.95}\right)\left[\underset{(0.55)}{4.85} + \underset{(0.34)}{1.47}\,E_t\left(\pi_{t+12}^{GER} - \pi^{*}\right) + \underset{(0.31)}{0.84}\,E_t y_t^{GER}\right] + \underset{(0.03)}{0.95}\,r_{t-1,t}^{GER} + u_{6t}$$

$$S.E. = 0.4$$

$$r_{t,t+n}^{GER} = \frac{1}{n}\sum_{i=0}^{n-1} E_t r_{t+i,t+i+1}^{GER}$$

Germany's economy results are near to those of US economy macro model, showing clearly that the expectation model is never rejected at all maturities. An example figure of the 1-year maturity is the following:



Observed and simulated interest rates for the 1-year maturity. Observed rate is labelled ACT1y, simulated rate is instead labelled MNSIM1y. Simulation is stochastic so average simulated rate is reported with upper UBSIM1y and lower bound LBSIM1y of its ninety-five per cent confidence interval.

8. EMPIRICAL RESULTS

We now proceed our empirical study by testing the informative content of simulated yields previous macro model suggest, full in accordance with the classic empirical literature on the term structure of interest rates. As we have analyzed, empirical research has focused upon the 'expectations theory' that relates the yield on long-term bonds to expected future short rates. This theory has frequently been tested and rejected using regression tests of the slope of the yield curve. Explanation for the rejections have focused on the presence of either time-varying term premia or biased forecast errors in the regression residuals. Either explanation implies, contrary to expectations theory requirements, that excess bond returns are predictable, since they are the sum of risk premia and forecast errors by definition. Indeed,

$$R_t^{(n)} = \theta(n) + \sum_{i=0}^{n-1} w_i E_t r_{t+i} \quad \text{(expectations hypothesis with constant term premium)}$$

$$r_{t+n} = E_t r_{t+n} + w_{t+n} \qquad \text{(rationally formed expectations)}$$

Using the same theoretical framework by regression-based tests, we expect a better performance of yields derived by macroeconomics to predict future changes in spot short-interest rates. As we have already noticed, these stochastic simulated yields are free of term premia, so in our procedure the greatest problem that causes the rejection of expectations hypothesis seems to be absent.

Specifically, we use the following regression equation to test the validity of expectations hypothesis:

$$\sum_{i=0}^{n-1} w_i r_{t+i} - r_t = a_n + b_n \left[R_t^{(n)} - r_t\right] + w_{t+n} \qquad (8.1)$$

where $R_t^{(n)}$ is the yield on a bond with a maturity of period *n*, r_{t+i} is always the actual interest rate on a 1-period debt instrument, and w_i is a declining weight that sums to 1.

Long-term maturity instrument in the right side of the equation will be in our work, either the observed long-term bond or the corresponding simulated one, in order to find which of the two has better predictive ability about the future path of expected short term interest rates.

Therefore, if arbitrage between short- and long- term rates holds as assumed by the expectations hypothesis, previous linear regression equation must give estimation about the coefficient b_n equals unity:

$$H_0: \quad \hat{b}_n = 1 \quad \text{(Expectations Hypothesis)}$$

under the alternative, $H_1: \quad \hat{b}_n \neq 1$ (for values of $\hat{b}_n$<1 with $t_{n,0.025} = 1.6$)

-or- $H_1: \quad \hat{b}_n > 1$ (for values of $\hat{b}_n$>1 with $t_{n,0.05} = 2$)

The following tables report our empirical outputs for the slope coefficients b_n (with standard errors underneath in parentheses), from a series of regressions of long-run changes of actual short rates on a constant and the long-short yield spread for U.S data at different maturities.

Table1 **Regression Coefficients** (***1-month*** short rate yields)

(Each row shows a regression coefficient b_n, with the standard error below in parentheses)

		Long bond maturity (months)						
Dependent variable		**6**	**12**	**24**	**36**	**60**	**84**	**120**
Long-run changes in short yields	Actual Long bond rates	0.41 (0.08)	0.57 (0.1)	0.55 (0.13)	0.87 (0.14)	1.21 (0.11)	1.22 (0.09)	0.68 (0.11)
	R^2	0.14	0.15	0.11	0.20	0.47	0.63	0.37
	Simulated long bond rates	0.75 (0.09)	0.79 (0.07)	0.88 (0.06)	0.99 (0.06)	1.23 (0.06)	1.33 (0.04)	0.91 (0.02)
	R^2	0.28	0.41	0.54	0.62	0.76	0.91	0.97

Table 2

Regression Coefficients (***3-month*** short rate yields)

Dependent variable		*Long bond maturity (months)* 6	12	24	36	60	84	120
Long-run changes in short yields	Actual Long bond rates	0.44 (0.11)	0.45 (0.12)	0.57 (0.14)	0.81 (0.16)	1.17 (0.13)	1.18 (0.11)	0.69 (0.11)
	R^2	0.09	0.13	0.09	0.18	0.39	0.53	0.37
	Simulated long bond rates	0.58 (0.04)	0.63 (0.05)	0.69 (0.05)	0.91 (0.06)	1.01 (0.07)	1.19 (0.06)	0.87 (0.02)
	R^2	0.12	0.32	0.50	0.57	0.65	0.81	0.96

As we observe for almost all frequencies, long-short yield spread, using simulated rates, gives us strong results and seems to be a better predictor of cumulative changes in short rates than actual long bond rates. As the forecasting horizon increases, both the size of the slope coefficient and the regression R^2 increase. The difference between actual and simulated regressors on the R^2 is great and becomes bigger as the horizon increase. At 10-years maturity horizon, R^2 is as high as 0.97 and 0.96 for simulated long bond rates and 0.37 for actual rates, respectively. This is an obvious sign of how better long-run changes of short rates can be explained by the simulated long-short yield spread than the observed one for U.S data at almost all maturities.

As a matter of fact, previous conclusion is fully justified. As we have already said, Favero tests the prediction of the relevant model using the information generated by this model, while the future path of policy rates is derived consistently with the adopted macro models rather than using the assumption of rational expectations. Conversely, when actual long bond rates are used, in the regression (8.1) of short-rate changes onto the yield spread, changing rational expectations about excess long-bond returns act like a measurement error that appears only in the regressor. In this case, a small measurement error, while cannot affect its sign, biases the coefficient toward zero and this biased behaviour gives smaller R^2 estimation. On the other hand, when simulated long rates are used, we do not have any rational expectations assumption so those measurement errors are minimized. This lack of measurement errors gives the opportunity for a better explanation of the movement of future short rates and a bigger R^2 estimation as well.

Many of $\hat{b}_n$ in tables 1 and 2 are less than one but, indeed, all of them are significantly different from zero. There is a clear dependence of the size and statistical significance of the $\hat{b}_n$ on the maturity of the long securities. As the forecast horizon increases, the predictive power in both cases becomes significantly bigger with the estimated coefficient b_n to approach unity. Especially, when we use simulated long bond rates, $\hat{b}_n$ is closer to unity than with actual bond rates.

However, while estimated slope coefficients b_n show clearly that the simulated yields have better predictability, since $\hat{b}_n$ is significantly nearer to unity than the actual situation, statistical tests about the expectations hypothesis H_0: $\hat{b}_n = 1$ reject H_0 at many sample points. Indeed, classic ***t-tests*** for H_0: $t_n = \dfrac{\hat{b}_n - 1}{\hat{\sigma}_{b_n}}$ reject the theory at many frequencies due to small standard errors at simulation cases. The reason of these small standard errors is that simulated series of interest rates, as we can see from figures on pages 41-44, are calm series without important fluctuations, especially for longer end maturities. As a result, standard errors of $\hat{b}_n$ tend to decrease and the lower value is observed at maturity of 10 years (see second row of tables 1-2).

Taking previous point into consideration, statistical t-tests of H_0: $\hat{b}_n = 1$, for both cases, are the following:

(***1-month*** short rate yields)

t-tests of $H_0: \hat{b}_n = 1$	Long bond maturity (months)							
		6	**12**	**24**	**36**	**60**	**84**	**120**
reject H_0: **REJ** *cannot reject* H_0: **N-REJ**	Actual Long bond rates	REJ	REJ	REJ	N-REJ	N-REJ	REJ	REJ
	Simulated long bond rates	REJ	REJ	N-REJ	N-REJ	REJ	REJ	REJ

(***3-month*** short rate yields)

t-tests of $H_0: \hat{b}_n = 1$	Long bond maturity (months)							
		6	**12**	**24**	**36**	**60**	**84**	**120**
reject H_0: **REJ** *cannot reject* H_0: **N-REJ**	Actual Long bond rates	REJ	REJ	REJ	N-REJ	N-REJ	N-REJ	REJ
	Simulated long bond rates	REJ	REJ	REJ	N-REJ	N-REJ	REJ	REJ

As we can see, previous statistical tests of H_0 give approximately same results for both cases. Either the weak rejection or the lack of rejection of the theory in the regression-based results appears at same points of maturity. However, when simulated long bond rates are used (second row), results can be explained considering that the uncertainty surrounding the long-term yields consistent with the expectations theory is not considered in the regression results. In fact, equation

(8.1) is estimated with the $R_t^{(n)}$ in the difference $\left[R_t^{(n)} - r_t\right]$ to be just the average output of the stochastically simulated long- term yields without considering the upper and lower bounds of the 95% confidence interval. This point is going to be discussed more detailed later.

The following tables are same empirical outputs for different short rates maturities of U.S. data and give us results full in accordance with previous ones:

Table 3

Regression Coefficients (***6-month*** short rate yields)

Dependent variable		*Long bond maturity (months)* 12	24	36	60	84	120
Long-run changes in short yields	Actual Long bond rates	0.33 (0.17)	0.36 (0.17)	0.77 (0.18)	1.2 (0.14)	1.29 (0.11)	0.71 (0.12)
	R^2	0.06	0.03	0.13	0.37	0.54	0.36
	Simulated long bond rates	0.56 (0.04)	0.64 (0.05)	0.86 (0.05)	0.98 (0.06)	1.23 (0.06)	0.88 (0.02)
	R^2	0.25	0.48	0.56	0.64	0.80	0.95

Table 4

Regression Coefficients (***12-month*** short rate yields)

Dependent variable		*Long bond maturity (months)* 24	36	60	84	120
Long-run changes in short yields	Actual Long bond rates	− 0.04 (0.19)	0.58 (0.19)	1.23 (0.15)	1.36 (0.11)	0.71 (0.13)
	R^2	0.0002	0.05	0.34	0.55	0.27
	Simulated long bond rates	0.48 (0.04)	0.66 (0.05)	0.90 (0.06)	1.16 (0.06)	0.88 (0.03)
	R^2	0.43	0.54	0.61	0.74	0.93

With these two tables, we continue to examine whether high yield spreads are associated with future changes of short yields under the expectations hypothesis.

Table 3-4 show that beyond one year the coefficients increase, and at 7 years coefficients are even significantly greater than one, giving strong forecasting power for short rate movements. Around one year, however, yield spread variation seems less related to subsequent movements in short rates.

It is obvious, though, for all examined maturities that simulated yields have better predictability, since $\hat{b}_n$ is significantly nearer to unity than the actual situation. The main reason for these results is that the simulation procedure of macro models is not affected by the presence of term premia. Macro models do not need any risk assumption to derive the full path of future policy rates. Therefore, if the expectations theory is true, the term premium can be possibly measured as the difference between actual and simulated yields on long-term bonds.

In order to see this point clearer, we have just to remember the following formula, in a simple two-period case:

$$\hat{b} \equiv \frac{\sigma^2(\mathrm{E}_t \Delta r_{t+1}/2) + 2\rho\sigma(\mathrm{E}_t \Delta r_{t+1}/2)\sigma(\theta_t)}{\sigma^2(\mathrm{E}_t \Delta r_{t+1}/2) + 2\rho\sigma(\mathrm{E}_t \Delta r_{t+1}/2)\sigma(\theta_t) + \sigma^2(\theta_t)}$$

In this formula, estimation of $\hat{b}_n$ depends on the variations in the term premium *σ(θt)* that give downwards bias on the coefficient. The bigger is the variance *σ(θt)* ^{2} of the term premium, the smaller is the estimation output of $\hat{b}_n$. The size of the bias also depends on the variance of the expected change in the future short rate. It is obvious, though, that $\hat{b}_n$ is equal to 1 in the absence of a time-varying premium (*σ(θt)* ^{2} = 0).

Finally, the corresponding statistical t-tests of $H_0: \hat{b}_n = 1$ for the last two maturities, are the following:

(***6-month*** short rate yields)

t-tests of $H_0: \hat{b}_n = 1$		*Long bond maturity (months)*					
		12	**24**	**36**	**60**	**84**	**120**
reject H_0: **REJ** *cannot reject* H_0: **N-REJ**	Actual Long bond rates	REJ	REJ	N-REJ	N-REJ	REJ	REJ
	Simulated long bond rates	REJ	REJ	REJ	N-REJ	REJ	REJ

As we can see from the three tables of statistical t-tests of $H_0: \hat{b}_n = 1$ that have been examined so far, there are some specific maturities (3 and 5 years of long bond maturity) that give statistically significant results according to expectations hypothesis. Those maturities have strong predictive ability about the future path of expected short-term interest rates, since we cannot reject H_0 for the slope coefficient on regression equation (8.1) to be one ($H_0: \hat{b}_n = 1$). Occasionally, there are also some others maturities around them that can give us same results for both cases (actual and simulated).

(***12-month*** short rate yields)

t-tests of $H_0: \hat{b}_n = 1$	*Long bond maturity (months)*					
		24	**36**	**60**	**84**	**120**
reject H_0: **REJ** *cannot reject* H_0: **N-REJ**	Actual Long bond rates	REJ	REJ	N-REJ	REJ	REJ
	Simulated long bond rates	REJ	REJ	N-REJ	REJ	REJ

The final point that we have to analyze more is the following: we have said that statistical t-tests $t_n = \dfrac{\hat{b}_n - 1}{\hat{\sigma}_{b_n}}$ about the expectations hypothesis $H_0: \hat{b}_n = 1$ reject H_0 at many sample points due to small standard errors in simulation cases, while, on the other hand, the output figures of the regression results show clearly that the simulated yields have better predictability, since $\hat{b}_n$ is significantly nearer to unity than the actual situation.

However, the fact that on regression equation (8.1) is used as $R_t^{(n)}$ just the average output of the stochastically simulated long-term yields is something required by the construction of those models under the pure expectations hypothesis. Thus, for the regression-based results the mean difference $\left[R_t^{(n)} - r_t\right]$ is used not by chance but as a result of our assumptions (8.2).

$$R_t^{(n)} = \frac{1}{n}\sum_{i=0}^{n-1} E_t r_{t+i,t+i+1} \quad (\textit{pure expectations hypothesis}) \qquad (8.2)$$

$$\sum_{i=0}^{n-1} w_i r_{t+i} - r_t = a_n + b_n\left[R_t^{(n)} - r_t\right] + w_{t+n} \qquad (8.1)$$

When Favero compares on his approach the observed long-term rates with the simulated ones in the 95% confidence interval for a direct test of the expectational model, he noticed that the apparent contradiction between the weak rejection of the theory in the regression based results and the lack of rejection in the simulation results can be explained considering that the uncertainty surrounding the long-term yields consistent with the expectations theory is not considered in the regression results.

In fact, equation (8.3) is estimated using as a $r^*_{t,t+n}$ just the average output of the stochastically simulated long- term yields without considering the upper and lower bounds of the 95% confidence interval.

$$r^*_{t,t+n} = \frac{1}{n}\sum_{i=0}^{n-1} E_t r_{t+i,t+i+1} \quad (\textit{pure expectations hypothesis})$$

$$r_{t,t+n} = \alpha + \beta r^*_{t,t+n} + u_t \qquad (8.3)$$

This stochastic *mean* output, though, is required to be used in the regression-based results by definition. While the prediction of the relevant macro model is examined under the *pure expectations hypothesis,* the average output of the stochastically simulated long- term yields must be concerned.

On the other hand, when Favero compares observed long-term rates with the simulated ones in the 95% confidence interval for a direct test of the expectational model, he does not really compare the observed long-term rates with the mean simulated ones, but with the whole stochastic simulated output of interest rates and its confidence interval. This seems to be a deviation of what *pure expectations* *hypothesis* requires, but it is just the manner for him to show the stochastic character of those simulated interest rates that have an average output and an upper and lower bound to reflect the standard error of them.

In our approach, the regressor of mean differences $\left[R_t^{(n)} - r_t\right]$ has its standard error that OLS estimator gives for $\hat{b}_n$ in regression (8.1). This st.error is small enough to give us statistical rejections of H_0: $\hat{b}_n = 1$ at many frequencies. However, in accordance with Favero procedure, we can also compare observed long-run changes of short-rates with the simulated differences $\left[R_t^{(n)} - r_t\right]$ in the 95% confidence interval for a direct test of the expectational model. It is obvious that results from this comparison would be absolutely similar to results of pages 41-44 that take into account the whole stochastic character of the simulated interest rates.

Finally, I give some empirical results for German economy:

Table 1

Regression Coefficients (1-month short rate yields)

Dependent variable		*Long bond maturity (months)* **6**	**12**	**24**	**36**	**60**	**120**
Long-run changes in short yields	Actual Long bond rates	0.8 (0.05)	0.83 (0.09)	0.63 (0.12)	0.85 (0.11)	0.91 (0.14)	1.51 (0.13)
	R^2	0.49	0.32	0.19	0.33	0.34	0.74
	Simulated long bond rates	0.84 (0.08)	0.87 (0.09)	1.07 (0.11)	1.13 (0.06)	1.09 (0.04)	1.33 (0.04)
	R^2	0.70	0.37	0.43	0.85	0.95	0.99

Table 2

Regression Coefficients (12-month short rate yields)

Dependent variable		24	36	60	120
		Long bond maturity (months)			
Long-run changes in short yields	Actual Long bond rates	0.17 (0.15)	0.41 (0.14)	0.76 (0.07)	1.73 (0.11)
	R^2	0.01	0.07	0.56	0.82
	Simulated long bond rates	0.39 (0.10)	1.05 (0.11)	1.03 (0.06)	1.30 (0.04)
	R^2	0.12	0.45	0.81	0.99

As we can see for German economy, as well as U.S., at all examined maturities, simulated yields have better predictability, with $\hat{b}_n$ and R^2 to be significantly nearer to unity than the actual one. In both cases, beyond one year and two years the coefficients increase, and at 10 years coefficients are even significantly greater than one, give strong forecasting power for short rate movements.

8. CONCLUSIONS

Within our framework, we examined the predictive power of small empirical macro models that a recent strand of the macroeconomic literature has analyzed by including the central bank reaction function in those models. Our regression equations used, take into account the simulated term structure path derived by these models rather than forward rates or yields to maturity of long-term bonds for a better prediction in future changes of interest rates. As this procedure is free of time-varying term premia, which recognized as the main factor for the rejection of classical expectations theory, our empirical results are much more in the line of the expectations theory, giving outputs for the slope of the regression nearer unity than classical situation of actual long bond rates.

The forecasting validity of macroeconomics by this set of theoretical models give us an important theoretical tool when future path of expected interest rates is concerned.

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