The Interest Rate Term Structure in the Greek Money Market

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Abstract

Using monthly data on Greek money market rates, we provide several tests of the Expectations Hypothesis (EH) with constant term premia. The empirical analysis draws on cointegration techniques, perfect foresight spread (PFS) regressions and the Campbell-Shiller VAR approach. On the basis of cointegration analysis, PFS regressions and VAR approach, the results are unfavourable to the EH. Spread stationarity and weak exogeneity tests appear to support the theory. We present some tentative explanations of these results.

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Introduction

The relationship between interest rates on equivalently rated debt instruments with differing dates to maturity, that is the term structure of interest rates, is recognised as an important tool in facilitating macroeconomic modelling and monetary policy evaluation (Andersen and Risager (1988), Mankiw, Miron and Weil (1987)). One aspect of term structure determination that has received much attention is the predictive power of the yield spread between long and short-term rates. Specifically, the expectations hypothesis (EH) of the interest rate term structure implies that the yield spread is an optimal predictor of future changes in the short rates over the life of the long-term instrument.

There is a great deal of evidence on the EH in the USA. Campbell and Shiller (1987) develop a Vector Autoregression (VAR) model to test the EH and find that the restrictions imposed by the theory are statistically rejected. In a more comprehensive study, Campbell and Shiller (1991) find again that the predictions of the EH for long rates are comfortably rejected in all instances. Nevertheless, they also conclude that future short-term rates move in the direction predicted by the theory. Other studies investigating the US interest rate term structure and utilising multivariate cointegration techniques include Hall, Anderson and Granger (1992), Shea (1992) and Engsted and Tanggaard (1994a). In brief, all aforementioned studies provide some support for the cointegration implications of the EH either at the short-end (Hall et al. (1992)), or at the medium term (Engsted and Tanggaard (1994a)), or at the long-end of the maturity spectrum (Shea (1992)).

The VAR methodology and cointegration analysis have been extensively employed to test the EH on UK data (MacDonald and Speight (1988, 1991), Taylor (1992), Cuthbertson (1996) and Cuthbertson, Hayes and Nitzsche (1996)). To summarise the findings, the evidence is supportive for the validity of the EH in the money market (Cuthbertson (1996), Cuthbertson et al. (1996)), whereas it is rather powerless for longer maturity yields (MacDonald and Speight (1988, 1991), Taylor (1992)).

The EH has not been extensively tested for developed countries of comparable size with that of Greece. Engsted (1996) and Engsted and Tanggaard (1994b), using Danish data, find support for the EH, especially during periods of high volatility in interest rates. Similar findings are obtained by da Fonseca (2002) who investigates the Portuguese Treasury bill market for a period (1990-1998) characterised by increased uncertainty and turbulence in interest rates. The cointegration test results indicate the presence of a unique cointegration vector in the bivariate models examined, hence confirming the stable relationship between the Treasury bill rates. These findings are broadly consistent with the *Mankiw-Miron* (1986) hypothesis which states that the EH is likely to perform better empirically under a policy of monetary targeting rather than interest rate smoothing.

In the present paper we explore the validity of the EH using short-term interest rates for a market that has not been investigated before, namely the Greek money market. The econometric methodologies employed are drawn on cointegration techniques, perfect foresight spread (PFS) regressions and the VAR approach of Campbell and Shiller (1987, 1991). Most previous research papers have employed these techniques using data on Treasury bill yields and/or government bonds. Instead we use money market interest rates which are closer linked to monetary policy, and are therefore the appropriate rates to use when assessing the usefulness of the term structure in conducting monetary policy.

The Greek money market became fully active in 1996 when it started offering a wide variety of maturities, ranging from 1- to 12-months. The liberalization of capital movements during the 1990s, the adoption by the Greek Central Bank of a low inflation targeting policy, and the stabilization of the Greek drachma/Ecu exchange rate in 1997, led to the enhancement of the role of the Greek money market as a vehicle of macroeconomic policy implementation. Our sample period, which runs from January 1996 to October 2001, is characterized by highly volatile interest rates due to speculation attacks against the Greek drachma and liquidity considerations in the Greek capital market.[i] Therefore, a study on the validity of the EH in the Greek money market allows us to ascertain the efficiency of this market as a tool of macroeconomic modelling for small developed economies. This is of crucial importance not only for the case of Greece which opted for a monetary targeting policy in order to successfully meet the requirements for participation in Stage III of EMU, but also for a number of other small European emerging economies (Poland, Hungary, Czech Republic) which may adopt similar policies during their preparation for membership in the euro area.

Finally, it should be noted that the utilization of the EH as the equilibrium model of the relationship between longer and shorter rates allows us to investigate the Efficient Market Hypothesis (EMH) from a broader perspective. Should the EH holds, economic agents would appear to process information in a way which is consistent with the concept of market efficiency.

Our findings provide limited evidence for the validity of the EH. On the one hand, the unit root tests in interest rate spreads and weak exogeneity tests are generally supportive of the hypothesis. On the other hand, the cointegration analysis, the PFS regressions and VAR cross-equation restrictions suggest that agents appear to process information in a way which is inconsistent with the EH.

The remainder of the paper is organized as follows. Section 1 provides a brief restatement of the EH, and outlines the methodology employed in the present study. Section 2 describes the data used in the empirical analysis. Section 3 presents the empirical findings. Finally, section 4 summarizes and concludes the paper.

1. The expectations theory of the term structure

The EH of the term structure of interest rates is a relationship between a longer-term n-period interest rate, $R_{t,n}$, and a shorter-term m-period interest rate, $R_{t,m}$, where n/m is an integer. In the case of money market rates, as with our data here, this is:

$$Rt,n = \left(\frac{1}{k}\right)\sum_{i=0}^{k-1} EtRt + mi,m + Cn,m \qquad \text{k=n/m} \qquad (1)$$

Equation (1) states that the n-period money market rate can be expressed as a simple average of the expected future m-period rates, $E_t R_{t+im,m}$, over the holding period (n-m periods), plus a constant term premium, $C_{n,m}$.

Subtracting $R_{t,m}$ from both sides of equation (1) and suppressing the constant risk premium implies:

$$R_{t,n} - R_{t,m} = \left(\frac{1}{k}\right)\sum_{i=1}^{k-1}\sum_{j=1}^{i} E[\Delta_m R_{t+jm,m}] = E_t\sum_{i=1}^{k-1}(1-\frac{i}{k})\Delta_m R_{t+im,m} \qquad (2)$$

The left hand side of equation (2) is the yield spread $S_{t,(n,m)}\left[= R_{t,n} - R_{t,m}\right]$. Δ_m is a difference operator measured over m periods, so that $\Delta_m R_t = R_t - R_{t-m}$. The EH implies that the spread is an optimal forecast of changes in future interest rates. The term on the right hand side of equation (2), $\sum_{i=1}^{k-1}(1-\frac{i}{k})\Delta mRt + im,m$, is termed the *perfect-foresight spread*, $S^*_{t,(n,m)}$, since it is the estimated spread that would be obtained by the model if agents had perfect foresight about future interest rates (Campbell and Shiller (1991)).

Suppose now that we want to test the assumption of rational expectations of the term structure (RETS), that is, whether the market's expectation is correct on average. We may write the actual future short rate as the sum of the expectation and a forecast error:

$$R_{t+im,m} = E_t R_{t+im,m} + \eta_{t+im,m} \tag{3}$$

Rearranging equation (2) using equation (3) yields a single equation test of the EH:

$$S^*_{t,(n,m)} = S_{t,(n,m)} + \varepsilon_t \tag{4}$$

The EH implies that the actual spread is a forecast of changes in future short period rates, or the perfect-foresight spread, $S^*_{t,(n,m)}$. The single equation test of the null of RETS is given by equation (5):

$$S^*_{t,(n,m)} = \alpha + \beta S_{t,(n,m)} + \gamma F_t + \varepsilon_t \tag{5}$$

where the null hypothesis is that α=γ=0, β=1 (Cuthbertson (1996)). If α≠0, then equation (5) suggests that there is a constant term premium. F_t represents the full information set available at time *t*, and ε_t is a moving average process which requires a Newey and West (1987) correction of order (*n-m-1*).

1.1 Bivariate autoregression tests of the expectations hypothesis

As Campbell and Shiller (1987) note, if the short rate is an I(1) process, then equation (2) implies that the spread is stationary, or alternatively, short and long-term rates are cointegrated with a cointegrating vector (-1, 1). The latter suggests that the system of interest rates possesses a long run equilibrium, even though random shocks push the system away from equilibrium in the short run. The error correction mechanism, or otherwise the spread, identifies such disequilibria and guides the interest rates back to the long run path. It follows that a *weak test* of the EH is that the spread *Granger causes* future changes in short and long rates.

Given the above considerations it follows that the vector $Z_t \equiv \left(\Delta R_{t,m}, S_{t,(n,m)}\right)$ is stationary and may be approximated by a VAR of order ρ which in companion form is:

$$Z_t = \sum_{i=1}^{\rho} A Z_{t-i} + u_t \tag{6}$$

where A is the *companion matrix* of coefficients and Z_t has 2ρ elements, namely $\Delta R_{t,m}$ and ρ-1 lags and $S_{t,(n,m)}$ and ρ-1 lags. The vector Z_t effectively summarises the whole history of $\Delta R_{t,m}$ and $S_{t,(n,m)}$. Using $h'Z_t = \Delta R_{t,m}$ and $g'Z_t = S_{t,(n,m)}$, Campbell and Shiller (1991) show that the *theoretical* spread can be formally expressed as:

$$S'_{t,(n,m)} \equiv h'A\left[I - \left(\frac{m}{n}\right)(I - A^n)(I - A^m)^{-1}\right](I - A)^{-1} Z_t \tag{7}$$

If the EH is true, then the theoretical should equal the actual spread. Formally, we write:

$$S'_{t,(n,m)} = g'Z_t = S_{t,(n,m)} \tag{8}$$

Equation (8) states that the actual spread, $S_{t,(n,m)}$ must equal the optimal prediction of future changes in $R_{t,m}$ based on ($S_{t,(n,m)}, \Delta R_{t,m}$) using the VAR. The resultant non-linear cross equation restrictions on the VAR can be tested by means of a Wald test (Campbell and Shiller (1991)).

Campbell and Shiller (1991) and Taylor (1992) argue that formal tests of equation (8) may lead to rejection of the rationality implications of the EH, even though the deviations from the null hypothesis are generated by economically uninteresting factors, such as minor data imperfections and the use of linearisations. Hence, we suggest an additional way of evaluating the model's performance. Specifically, we suggest computing the theoretical spread without imposing the VAR restrictions and testing whether the variance of the theoretical spread is equal to the variance of the actual spread. If the null hypothesis of equality in variances cannot be accepted, then the volatility of the actual spread is considered to be significantly different from that imposed by the EH.

1.2 Multivariate cointegration tests of the expectations hypothesis

As mentioned above, Campbell and Shiller (1987) show that if $R_{t,m}$ is an I(1) process, then equation (2) implies that $R_{t,n}$ and $R_{t,m}$ are cointegrated with a cointegrating vector (-1,1). Given a set of *r* yield variables, equation (2) suggests that each yield is cointegrated with all other yields, and hence there should be *r-1* cointegration vectors. Then each of the *r-1* linearly independent spread vectors (-1,1,0, ... ,0), (-1,0,1, ... ,0), etc. should span the cointegrating space. Johansen (1988) and Johansen and Juselius (1990) provide two tests to determine the number of cointegrating vectors in a multivariate setting, as well as likelihood ratio statistics to test whether the cointegrating vectors can be expressed in terms of the spreads (see Hall et al. (1992) and Engsted and Tanggaard (1994a, 1994b)).

2. Discussion of the data

The data analysis is conducted on Greek money market rates (ATHIBOR/EURIBOR as from 1/1/2001). This data set, provided by the Bank of Greece, has not been analysed before in the literature. The file contains six rates with different maturities, namely, the 1-, 2-, 3-, 6-, 9-, and, 12-month rate. The full sample period consists of 70 monthly observations for each series, dating from January 1996 until October 2001. The limited sample period is dictated by the lack of data for all six rates prior 1996. However, the data are particularly appropriate for a preliminary investigation of the underlying mechanisms of the Greek money market. The use of discount data allows us to circumvent the problem of having to use (arbitrary) approximations to zero-coupon yields that may bias the subsequent statistical analysis. The plots of the 1- and the 12-month money market rate in Figure 1 and Figure 2, respectively, are representative of the money market rates at both the short and long-end of the maturity spectrum. They clearly reveal the sharp increase of money market rates in autumn 1997 (when speculation attacks against the Greek drachma took place) and their subsequent decline due to their convergence to those in the euro-zone. However, the visual evidence also indicates that for most of the time covered by this research, the negative slope of the interest rates is maintained, hence providing no evidence for structural instability.

[Figures 1 & 2]

3. Empirical findings

3.1 Time series properties of individual money market rates

Before proceeding with our analysis, we briefly discuss the time series properties of the Greek money market rates. By means of ADF (Dickey and Fuller (1981)) and Phillips-Perron (PP) (1988) unit root tests we establish that each of the six money market rates contains a unit root at levels, but not in first differences.

[Table 1]

3.2 Cointegration analysis

The EH, as expressed in equation (2), implies that if the short interest rate is an I(1) process, then interest rate spreads should be stationary. This assumption is tested in Table 2. The 1-month rate is the basic short rate relative to which spreads are calculated. The ADF and PP tests are all significant at at least 10% significance level, confirming the predictions of the EH under the assumption of a constant or I(0) term premium. However, this finding should be carefully interpreted since the spread unit root tests constitute a relatively *naïve* approach in testing the EH.

[Table 2]

We now consider the hypothesis that the money market rates are cointegrated with the spread vectors corresponding to the cointegrating vectors. First, it is necessary to formulate the vector error correction (VEC) model:

$$\Delta Z_t = c + \Gamma_1 \Delta Z_{t-1} + \ldots + \Gamma_{k-1} \Delta Z_{t-k-1} + \Pi Z_{t-k} + u_t \qquad (9)$$

On the basis of both the Johansen's maximal eigenvalue test ($\lambda_{\max}$) and the trace statistic (λ_{trace}) we accept the restriction that the rank of the cointegrating space is not more than three. Given the set of six money market rates, this finding suggests that there are less cointegrating vectors than the ones predicted by the theory. Alternatively stated, there are more than one common trends in the system of six money market rates. Hence we conclude that interest rates in the Greek money market, are only partially cointegrated. All subsequent tests are conditional on the rank of the cointegrating space, $r=3$. The estimates of α and β obtained from applying the Johansen technique are presented in Table 3. Here, β is a (3x6) transposed, normalized by the first element matrix whose rows are the cointegrating vectors, and α is the associated matrix of error correction parameters, which measures the influence of the error correction term in each of the equations for the money market rate. The EH implies that the error correction mechanism of the rates (i.e., the spread) should adjust from disequilibrium and bring money market rates back to the long run equilibrium path. Stated alternatively, the EH holds the speed of adjustment coefficients, α_s, should be statistically significant. The reported estimates indicate statistically insignificant error correction terms mainly at the short-end of the maturity spectrum (up to 3-months). This finding suggests that the short-term money market rates may be weakly exogenous with respect to the long run parameters in this model. To further investigate the issue of weak exogeneity, we conduct a test by placing appropriate restrictions on α_s. Likelihood ratio test results are also reported in Table 3 and indicate that none of the money market rates can be assumed to be weakly exogenous. Furthermore, the adjustment in the money market rates is in the right direction, as the strongly negative values for α_s in the second cointegrating vector suggest. This finding supports one of the central implications of the EH, namely that the spreads should be able to predict changes in short term rates.

Next, we examine whether the cointegrating vectors tell us anything about the structural relationship underlying the long run model. The estimates of β in

Table 3 are not necessarily unique, and therefore, we impose restrictions motivated by rational arguments and then test whether the columns of β are identified. Table 3 provides a likelihood ratio test that the cointegrating vectors are of the form (-1, 1, 0, … , 0). A clear pattern emerges from these results. In all instances, there is a strong rejection of the null that the individual spread vectors form a basis for the cointegrating space.

Tests based on the Johansen procedure assume serially uncorrelated and homoscedastic disturbances. Table 3 reports a number of multivariate and univariate diagnostic tests for the presence of serial correlation (of fourth order), the presence of autoregressive conditional heteroscedasticity (of second order) and a test for normality in the residuals. The chi-square tests for serial correlation and ARCH effects reveal no insample evidence of misspecification. Non-normal residuals, however, seem to be a significant problem in all equations. Eitrheim (1992) and Gonzalo (1994) explore the finite sample properties of the maximum likelihood estimators of the error correction mechanism under deviations from the normality assumption. According to the results of these simulation studies, the estimated coefficients of the α- and β-matrix seem tenable, in the sense that the Johansen method is not outperformed by any other method proposed in the cointegration analysis literature.

The overall conclusion is that the cointegration results are inconsistent with the EH. The money market rates cannot be described as fully integrated processes which cointegrate into stationary spreads. These findings contrast those obtained in previous studies where cointegration analysis is at least partially supportive of the EH (da Fonseca (2002), Hall et al. (1992), Engsted and Tanggaard (1994a, 1996b), Cuthbertson (1996)). The rejection of the statistical implications of the EH may be attributed to the presence of a small I(1) time-varying risk premium which distorts the cointegrating vectors (Evans and Lewis (1994)). As noted in the Introduction, the sample period is not one of interest rate smoothing, hence possibly giving rise to nonstationary risk premia. Therefore, the formulation of the EH with constant term premia may be unable to capture the dynamics of the Greek money market yield curve variation.

[Table 3]

3.3 Perfect foresight regressions

The regression results of the perfect foresight spread, $S^{*}_{t,(n,m)}$ on the actual spread, $S_{t,(n,m)}$ and the information subset I_t (consisting of four lags of $S_{t,(n,m)}$ and $\Delta_m R_{t,m}$) are presented in Table 4. The method of estimation is the Generalized Method of Moments (GMM) to correct the covariance matrix for moving average errors and possible heteroscedasticity (Newey and West (1987)). According to RETS, $S'_{t,(n,m)}$ should be unpredictable given information at time *t* or earlier, and the slope coefficient β in equation (5) should be equal to one. In general, the results are unfavorable for RETS. In all cases we reject the null that information available at time *t* or earlier does not incrementally add to the predictions of future interest rates (Ho:γ=0). Furthermore, we note that the statistically significant, and correctly signed, regression coefficient of the PFS on the actual spread is in the region of 1.10-2.13, rather than close to unity. The latter could be assessed as an underreaction of the yield spread. In other words, the spread between the two rates is smaller than it can be justified by rational expectations of future short rate changes.[ii] This ascertainment is further reinforced when the null H_1:β=1 (given

γ=0) is tested and strongly rejected in all cases. Similar results are obtained when the null H_2:α=0, β=1 (given γ=0) is tested. These findings are consistent with those obtained from the cointegration analysis and suggest that agents do not optimally utilize all available information in forecasting Greek money market rates. Generally speaking, our results are broadly consistent with those of Campbell and Shiller (1991) who also find little or no support of the EH at maturities less than one year from the regression of the PFS on the actual spread. On the contrary, our findings are different from those obtained by Engsted (1996), Cuthbertson (1996), Cuthbertson et al. (1996) who provide more support for the EH within the PFS regression framework.

[Table 4]

3.4 VAR methodology

Section 1 indicates that one implication of the bivariate VAR representation of the term structure is that the spread must Granger-cause $\Delta_m R_{t,m}$. If agents have more information, other than the history of $\Delta_m R_{t,m}$, about the future course of $R_{t,n}$, it follows from equation (2) that the spread between $R_{t,n}$ and $R_{t,m}$ should have additional explanatory power for forecasting future changes in the short rate. The Granger causality test results are reported in Table 5. The null hypothesis of no Granger causality from spreads to short rate changes is rejected at the 5% significance level in only two cases, namely, the 1- to 2-month spread, and the 1- to 3-month spread. In all other instances, the null hypothesis cannot be rejected. The latter contradicts the weak exogeneity test results reported in Table 3 which provide evidence for the statistical significance of the error-correction coefficients (α_s) in the cointegrating vectors. The standard Granger causality test (when cointegration is not taken into account) assumes that the information for the prediction of the variables is contained only in the time series data of these variables. The VEC model, however, improves upon standard Granger causality tests by allowing for long run information in the data. Hence, we place more emphasis on the results obtained from the cointegration analysis.

[Table 5]

Table 6 reports the VAR cross equation restrictions as outlined in Campbell and Shiller (1991) and McDonald and Speight (1988, 1991). The aforementioned researchers show that the non-linear VAR cross equation restrictions can be simplified, so that they become linear and more easily interpreted. In particular, they show that the variable $V_t\left[= S_t - 1\left(\frac{1}{\alpha}\right)S_{t-1} + \Delta_m R_{t,m}\right]$ should be unpredictable given lagged $\Delta_m R_{t,m}$ and $S_{t,(n,m)}$. The heteroscedasticity robust Wald tests indicate that the VAR restrictions are rejected in all cases. These rejections imply that the information subset at time *t* or earlier (other than $S_{t,(n,m)}$) influences future changes in short rates. This explanation is in accordance with the findings in Table 4 where the single equation PFS regressions reject the null that the limited information set F_t available at time *t* or earlier does not add to the predictions of future interest rates. Furthermore, the rejection of the VAR restrictions may originate from the low frequency of our data set. In a money market, one would expect participants to formulate their decisions utilizing intraday observations. Therefore, forecasts based on monthly data may not adequately mimic such behavior.

Taylor (1992) suggests that formal testing of equation (8) may often lead to rejections of the RETS restrictions because of statistically significant, but economically unimportant factors, such as minor data imperfections. Hence, Table 6 also reports a test for equality of the variance of the actual spread against the variance of the theoretical spread. The idea here is that even if the VAR restrictions do not hold, the behavior of the theoretical and actual spread allows us to estimate the deviation from the EH. Under the EH, the null hypothesis of equal variances should not be rejected. However, the F-statistic suggests that the null is strongly rejected in all instances, hence confirming the rejections of the VAR restrictions.

[Table 6]

4. Conclusions

The presence of well-developed money market instruments is a prerequisite for the proper functioning of the Greek capital market. In this paper, we investigate the structure of the Greek money market and assess its operational efficiency by testing the validity of the EH with constant term premia. Conditional on money market rates being I(1), the Johansen cointegration analysis and spread restrictions are generally inconsistent with the EH over the sample period. Furthermore, the PFS regressions and the Campbell-Shiller VAR approach suggest that agents do not appear to utilise all information available to them in a way that is consistent with the EH. It is only the spread unit root tests and weak exogeneity tests that provide limited evidence in favour of the theory.

The rejection of the validity of the EH in the Greek money market implies either (i) the limited efficiency of this market as a vehicle of macroeconomic policy implementation or, (ii) the misspecification of the EH with time independent term premia. Should the first possible justification holds, market participants would appear to underreact to the arrival of new information, hence violating the concept of market efficiency.[iii] Therefore, for the sample period that we examine, the Greek money market failed to accomplish its role as a means of formulating market expectations in accordance with those of monetary policy makers. Alternatively, our unfavourable findings may be explained by the initial specification of the EH with constant term premia. The sample period in the present study can be characterised as a period of high interest rate volatility. In 1997 there was a shift in the Greek foreign exchange policy with the stabilisation of the Greek drachma/Ecu exchange rate. This crucial political decision may have enhanced the credibility of the Greek authorities' policy of convergence towards the Maastricht criteria, but it also led to an increased volatility of short rates, hence possibly resulting in time varying term premia. Explicit test of this conjecture are left in the agenda for future research.

References

Andersen, T. M. and O. Risager (1988), 'Stabilization policies, credibility, and interest rate determination in a small open economy', *European Economic Review*, Papers and Proceedings, Vol. 32, pp. 669-679.

Campbell, J. and R. J. Shiller (1987), 'Cointegration and test of present-value models', *Journal of Political Economy*, Vol. 95, pp. 1063-1088.

Campbell, J. and R. J. Shiller (1991), 'Yield spreads and interest rate movements: a bird's eye view', *Review of Economic Studies*, Vol. 58, pp. 495-514.

Cuthbertson, K. (1996), 'The expectations hypothesis of the term structure: The UK interbank market', *The Economic Journal*, Vol. 106, pp. 578-592.

Cuthbertson, K., S. Hayes and D. Nitzsche (1996), 'The behaviour of certificate of deposit rates in the UK', O*xford Economic Papers*, Vol. 48, pp. 397-414.

Dickey, D. A. and W. A. Fuller (1981), 'Likelihood Ratio Statistics for Autoregressive Time-Series with a Unit Root', *Econometrica*, Vol. 49, pp. 1057-1072.

Dockery, E. and M. G. Kavussanos (1996), 'Testing the Efficient Market Hypothesis Using Panel Data, with Application to the Athens Stock Market', *Applied Economics Letters,* Vol. 3, pp. 121-123.

Eitrheim, O. (1992), 'Inference in small cointegrated systems: some Monte Carlo results', Department of Economics, University of California, Working paper (92-31).

Engsted, T. (1996), 'The predictive power of the money market term structure', *International Journal of Forecasting*, Vol. 12, pp. 289-295.

Engsted, T. and C. Tanggaard (1994a), 'Cointegration and the US term structure', *Journal of Banking and Finance*, Vol. 18, pp. 167-181.

——————— (1994b), 'A cointegration analysis of Danish zero-coupon bond yields', *Applied Financial Economics*, Vol. 24, pp. 265-278.

Evans, M. D. D. and Lewis, K. K. (1994), 'Do stationary risk premia explain it all?', *Journal of Monetary Economics*, Vol. 33, pp. 285-318.

da Fonseca, J. S. (2002), 'The risk premiums in the Portuguese treasury Bills interest rates', *European Review of Economics and Finance*, Vol. 1, pp. 69-82,

Gonzalo, J. (1994), 'Five alternative methods of estimating long run equilibrium relationships', *Journal of Econometrics*, Vol. 60, pp. 203-233.

Hall, A. D., H. M. Anderson and C. W. J. Granger (1992), 'A cointegration analysis of treasury bill yields', *Review of Economics and Statistics*, Vol. 74, pp. 116-126.

Johansen, S. (1988), 'Statistical analysis of cointegration vectors', *Journal of Economic Dynamics and Control*, Vol. 12, pp. 231-254.

Johansen, S. and K. Juselius (1990), 'Maximum likelihood estimation and inference on cointegration - with applications to the demand for money', *Oxford Bulletin of Economics and Statistics*, Vol. 52, pp. 169-210.

Johansen, S. and Nielsen, B. (1993), 'Asymptotics for cointegration rank tests in the presence of intervention dummies', Institute of Mathematical Statistics, University of Copenhagen, (manuscript).

Kavussanos, M. G. and E. Dockery (2001), 'A Multivariate Test for Stock Market Efficiency: The Case of ASE', *Applied Financial Economics*, Vol. 11, pp. 573-579.

MacDonald, R. and A. E. H. Speight (1988), 'The term structure of interest rates in the UK', *Bulletin of Economic Research*, Vol. 40, pp. 287-299.

——————— (1991), 'The term structure of interest rates under rational expectations: some international evidence', *Applied Financial Economics*, Vol. 1, pp. 211-221.

Mankiw, N. G. and J. Miron (1986), 'The changing behavior of the term structure of interest rates', *Quarterly Journal of Economics*, Vol. CI, pp. 221-228.

Mankiw, N. G., J. Miron and D. Weil (1987), 'The adjustment of expectations to a change in regime: A study of the founding of federal reserve', *American Economic Review,* Vol. 77, pp. 358-374.

Newey, W. K. and K. D. West (1987), 'A simple, positive semi-definite, heteroskedasticity and autocorrelation consistent covariance matrix', *Econometrica*, Vol. 55, pp. 703-708.

Phillips, P.C.B. and P. Perron (1988), 'Testing for a Unit Root in Time Series Regression', *Biometrika*, Vol. 75, pp. 335–346.

Shea, G. S. (1992), 'Benchmarking the expectations hypothesis of the term structure: an analysis of cointegration vectors', *Journal of Business and Economic Statistics*, Vol. 10, pp. 347-365.

Siourounis, G. D. (2002), 'Modelling Volatility and Testing for Efficiency in Emerging Capital Markets: The Case of the Athens Stock Exchange', *Applied Financial Economics*, Vol. 12, pp. 47-55.

Stengos, T. and Panas E. (1992), 'Testing the Efficiency of the Athens Stock Exchange: Some Results from the Banking Sector', *Empirical Economics*, Vol. 17, pp. 239-252.

Taylor, M. P. (1992), 'Modelling the yield curve', *The Economic Journal*, Vol. 102, pp. 524-537.

Footnotes

[i] In October 1997 the Bank of Greece increased the discount rate to 170% p.a. in order to prevent capital outflows and to smooth pressures against the Greek drachma. The underlying reasons for this currency attack may be found in the Asian crisis, as well as, in the specific structure and weaknesses of the Greek economy at that time (high current account deficit and external debt, lack of competitiveness, and pessimistic expectations regarding the entrance of the drachma in the ERM). Subsequent financial crises (such as the crisis in Russia in 1998) also had an impact (albeit at a lesser extent) on Greece's capital market liquidity, resulting in highly volatile money market rates.

[ii] It should be noted, however, that the β coefficient approaches its theoretical value of unity as the maturity of money market rates, n, increases (for given m).

[iii] This is consistent with a number of other studies which investigate the issue of market efficiency employing solely data on stock prices listed in the Athens Stock Exchange (ASE) (Siourounis (2002), Kavussanos and Dockery (2001), Dockery and Kavussanos (1996)). The notable exception in this area of research is the study by Stengos and Panas (1992) which finds support for the weak and semi-strong form of efficiency using data on selected stocks from the Greek banking sector.

Appendix 1

Figure 1. The 1-month money market rate

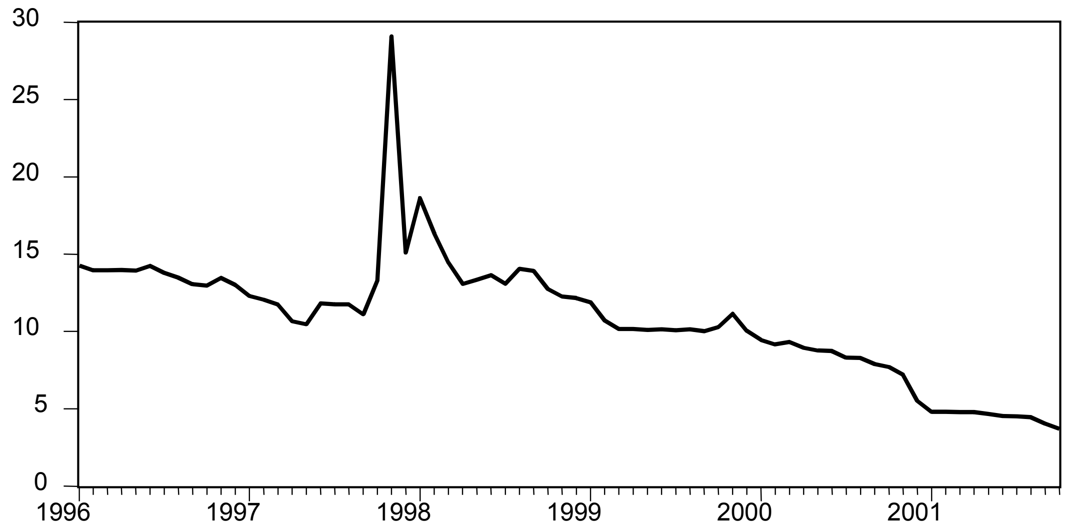
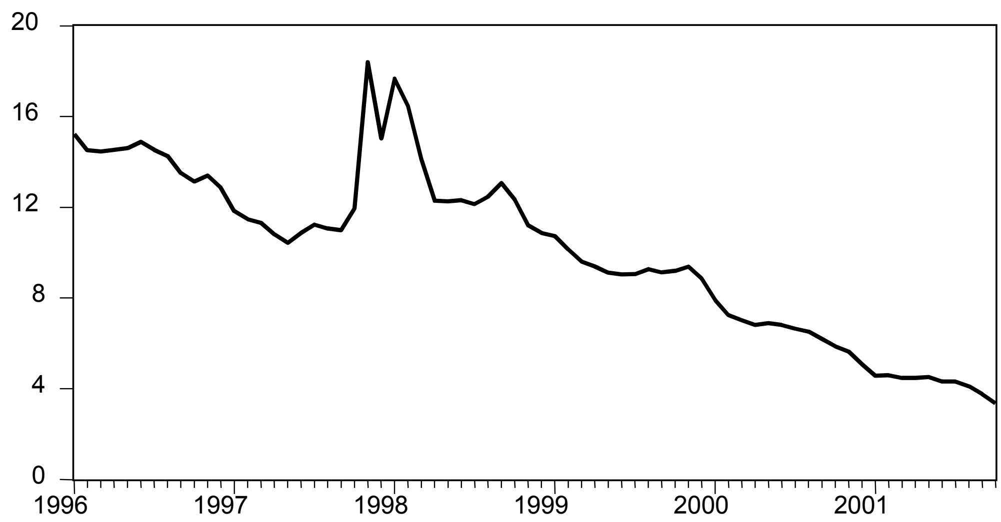


Figure 2. The 12-month money market rate



Appendix 2

Table 1. Test for unit roots in Greek money market rates

1996:1 - 2001:10	ADF(4)	PP(3)
Levels:		
1-month	-0.81	-2.27
2-months	-0.78	-1.72
3-months	-0.75	-1.32
6-months	-0.69	-0.92
9-months	-0.68	-0.82
12-months	-0.70	-0.82
First Differences:		
1-month	-4.37*	-15.09*
2-months	-4.23*	-14.09*
3-months	-4.22*	-13.03*
6-months	-4.43*	-11.55*
9-months	-4.52*	-10.62*
12-months	-4.65*	-10.03*

Notes: The table gives values of the *t*-statistic in the augmented Dickey-Fuller (ADF) regressions with 4 lags. The critical value at the 5% significance level is -2.90. The table also reports the *t*-statistic of Phillips-Perron's (PP(3)) test for a unit root. The truncation lag is set equal to 3 to ensure Newey-West heteroskedasticity and autocorrelation consistent estimates. The critical value of the *t*-statistic at the 5% significance level is –2.90.
* denotes rejection of the null hypothesis at the 5% significance level.

Table 2. Tests for unit roots in Greek money market spreads

1996:1 - 2001:10	ADF(4)	PP(3)
Levels:		
1-month/2-months	-3.89*	-7.85*
1-month/3-months	-3.77*	-7.85*
1-month/6-months	-3.26*	-7.10*
1-month/9-months	-2.93*	-6.82*
1-month/12-months	-2.70**	-6.50*

Notes: See Table 1.
* denotes rejection of the null hypothesis at the 5% significance level.
** denotes rejection of the null hypothesis at the 10% significance level.

Table 3. Cointegration Analysis

Test for the Number of Cointegrating Vectors

$H_0 = r$	k-r	λ_{max}	90% critical value	λ_{trace}	90% critical value
0	6	112.34	24.63	238.43	89.37
1	5	66.00	20.90	126.09	64.74
2	4	40.56	17.15	60.09	43.84
3	3	13.25	13.39	19.53	26.70

Estimated α- & β-matrices (based on three cointegrating vectors)

β-matrix (transposed and normalised by the first element)

1-month	2-months	3-months	6-months	9-months	12-months
1.00	-1.06	0.01	-1.10	2.32	-1.17
1.00	-0.80	-1.62	3.50	-3.04	0.97
1.00	-2.87	2.38	-0.34	-0.14	0.46

α-matrix

1-month	2-months	3-months	6-months	9-months	12-months
-0.05	0.71	1.29	1.77*	1.61*	1.62*
-8.06*	-6.40*	-5.27*	-4.56*	-3.89*	-3.47*
2.01	2.02	1.48	1.47	1.61	1.56

Restrictions on β-matrix: Testing for Stationary Spreads

	$\chi^2(3)$	ρ-value
1-month through 12-months	14.49	0.00
1-month & 2-months	25.07	0.00
1-month & 3-months	28.13	0.00
1-month & 6-months	30.65	0.00
1-month & 9-months	32.57	0.00
1-month & 12-months	33.98	0.00

Restrictions on α -matrix: Testing for Weak Exogeneity

	χ^2 (3)	ρ-value
H_0: $\alpha_1 = 0$	11.75	0.00
H_0: $\alpha_2 = 0$	11.04	0.01
H_0: $\alpha_3 = 0$	12.09	0.01
H_0: $\alpha_4 = 0$	20.32	0.00
H_0: $\alpha_5 = 0$	21.05	0.00
H_0: $\alpha_6 = 0$	21.64	0.00

Residual Analysis

Multivariate Statistics

$\chi^2(4)_{ar}$ 31.16 (0.70)

Univariate Statistics

	1-m	2-m	3-m	6-m	9-m	12-m
$\chi^2(2)_{arch}$	1.35	1.73	2.05	0.76	0.86	1.00
$\chi^2(2)_{nor}$	29.84**	28.67**	26.79**	24.05**	20.14**	16.42**

Notes: r denotes the cointegration rank. k-r denotes the number of common trends present in the system of yields. Critical values for cointegration tests are from Johansen and Nielsen (1993). ρ-value denotes the marginal significance levels in tests of stationary yield spreads, as well as in tests of weak exogeneity.

* denotes statistically significant (at the 5% significance level) error-correction terms.

** denotes rejection of the null hypothesis of normally distributed residuals.

Table 4. Regression of PFSt,(n,m) on St,(n,m)

	Coefficients		Hypothesis Tests (ρ-values)		
(n,m)	α (.) = s.e	β (.) = s.e	H_0: γ = 0	H_1: β=1	H_2: α = 0, β = 1
(2,1)	0.11 (0.16)	2.13 (0.32)	0.00	0.00	0.00
(3,1)	0.10 (0.22)	1.39 (0.17)	0.00	0.00	0.00
(6,1)	0.16 (0.37)	1.12 (0.16)	0.00	0.00	0.00
(9,1)	0.18 (0.47)	1.13 (0.17)	0.00	0.00	0.00
(12,1)	0.14 (0.56)	1.10 (0.19)	0.02	0.00	0.00

Notes: The regression coefficients are from regression of PFSt,(n,m) on St,(n,m) with γ=0 imposed (eq. (5)). The reported standard errors (s.e.) are corrected for heteroscedasticity and moving average errors of order (n-m-1), using Newey and West (1987) weights to guarantee positive semi-definiteness. For H_0: γ=0 the reported results are for an information set H_t which includes four lags of the change in short rates and of the spread. The null H_1: β=1, is conditional on γ=0 (and α≠0), while the null H_2: α=0, β=1 is also conditional on γ=0.

Table 5. Granger Causality Tests from St to ΔmRt,m

VAR lags	**F-statistics H_0: $S_{t,(n,m)}$ does not Granger-cause $\Delta_m R_{t,m}$** 1m/2m	1m/3m	1m/6m	1m/9m	1m/12m
4	4.98 (0.00)	4.93 (0.00)	1.25 (0.30)	1.23 (0.31)	1.17 (0.33)

Notes: Figures in parentheses denote marginal significance levels. Spreads are defined as the long money market rate minus the 1-month money market rate.

Table 6. Tests of the Rational Expectations of the Term Structure

(n,m)	**Wald Statistic** (.) = ρ-value	**F-statistics H_0: Equality of variances between $S'_{t,(n,m)}$ and $S_{t,(n,m)}$**
(2,1)	0.00	73.34 (0.00)
(3,1)	0.00	82.00 (0.00)
(6,1)	0.00	96.00 (0.00)
(9,1)	0.00	99.16 (0.00)
(12,1)	0.00	100.70 (0.00)

Notes: The numbers in the second column are the marginal significance levels of Wald statistics for testing the exclusion restrictions given in eq. (8) with the constant included. Wald statistics are heteroscedastic-robust. The third column provides an estimate of Var(S′t,(n,m)) = Var(St,(n,m)). Figures in parentheses denote marginal significance levels.