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Testing the Expectations Hypothesis of the Term Structure of Interest Rates in the Presence of a Potential Regime Shift

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Abstract

The expectations hypothesis of the term structure of interest rates is tested using monthly Eurodollar deposit rates for maturities 1, 3 and 6 months covering the period 1983:1–1996:6. 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. The peso problem is modelled by means of a threshold autoregression. The estimation results suggest that potential regime shift had an effect on expectations concerning the longer-term interest rate only for a short while in the early phase of the sample period, when interest rates were at their highest.

Key words: peso problem, TAR models, term structure of interest rates

Korkojen aikarakenteen testaaminen mahdollisten regiimimuutosten tapauksessa

Suomen Pankin keskustelualoitteita 20/99

Markku Lanne
Tutkimusosasto

Tiivistelmä

Korkojen aikarakenteen odotushypoteesia testataan tässä tutkimuksessa käyttäen kuukausihavaintoja 1, 3 ja 6 kuukauden dollarin eurokoroista aikajaksolta 1983:1–1999:6. Klassiset regressiotestit eivät tue odotushypoteesia. Sen sijaan hypoteesia ei voitu hylätä mallissa, joka sallii mahdolliset mutta havaintoaineistossa toteutumattomat regiiminmuutokset. Nämä pesovaikutukset mallinnettiin autoregressiivisen kynnysmallin avulla. Mallin estimointitulosten mukaan regiimimuutoksen mahdollisuus on vaikuttanut pitempää korkoa koskeviin odotuksiin vain lyhyen aikaa otoksen alussa, jolloin korot olivat korkeimmillaan.

Asiasanat: peso-ongelma, autoregressiiviset kynnysmallit, korkojen aikarakenne

Contents

1 Introduction

The classical expectations hypothesis of the term structure of interest rates is typically rejected, especially with U.S. data (see e.g. Campbell and Shiller (1991), and Evans and Lewis (1994)). As far as the long end of the maturity spectrum is concerned, the literature is not unanimous, and it is plausible that the reported rejections follow from applying inappropriate econometric methods (Lanne (1999)), but for the shorter-term interest rates the evidence against the expectations hypothesis seems to be strong. One 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. 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. For a survey of the literature on peso problems, see Evans (1996).

Previous term structure literature has mainly attempted to take the effect of regime switches into account by testing the rational expectations restrictions within models with more than one regime. Among others Hamilton (1988), Sola and Driffill (1994), and Kugler (1996) have employed the so called Markov switching model. This approach is useful only if a regime shift has actually taken place several times in the sample which indeed was the case in these papers since they examined U.S. interest rates in periods including the three years of the new operating procedures of the Federal Reserve. The remedy to the peso problem that Bekaert et al. (1997a) suggest, is to estimate the data generating process (DGP) and by simulation obtain the appropriate finite sample distributions. To this end they pool data from seven countries (Australia, Germany, Italy, Japan, Sweden, the U.K., and the U.S.) making the bold assumption that all these data follow the same process, and use the consequent finite sample distribution to conduct inference in U.S. term structure data. The main idea is that in the pooled data the occurrence of the different regimes is expected to come closer to the true frequencies than in the U.S. data alone, and hence the model estimated with this data approximates well the true DGP. A potential problem with this approach is that it relies on the assumption of a common DGP across the countries.

A closely related literature attempts to resolve the rejection of the expectations hypothesis with models where the expectations are affected by infrequent but only partially predictable target interest rate changes. See e.g. Balduzzi et al. (1997, 1998), and Jääskelä and Vilmunen (1999).

In this paper we test the expectations hypothesis in the short-term Eurodollar deposit rates from the period after the new operating procedures of the Federal Reserve, 1983–1999, which has been little studied in the previous literature. Thus our a priori assumption is that no regime shifts have occurred in the sample period, and to take the potential peso effects into account, we employ a version of the model recently introduced by Saikkonen and Ripatti (1999). The central idea in the model is that the expectations of the shorter-term interest rate are computed from a threshold autoregressive (TAR) model which in the actual sample reduces to a linear autoregressive (AR) model due to the absence of regime shifts. If there indeed are peso effects and the model is correctly specified, the discrepancy between the expectations based on the TAR and AR models should then be able to reconcile the data with the expectations hypothesis. The empirical results indicate that this is the case. While the expectations hypothesis is rejected when the assumption of no peso effects is made, it cannot be rejected when peso effects are allowed for.

The plan of the paper is as follows. Section 2 introduces the model and discusses estimation and inference. In Section 3 the empirical results are presented. Finally, Section 4 concludes.

2 The model

In the previous literature the peso problem has mainly been studied in the framework of a Markov switching model (see e.g. Hamilton (1993)), the simplest version of which, in the case of the expectations hypothesis of the term structure, can be written as follows.

$$r_t^{(m)} = \mu_{s_t} + \xi_t \tag{1}$$

$$r_t^{(n)} = \alpha + \gamma \frac{m}{n} \sum_{j=0}^{n/m-1} E_t r_{t+jm}^{(m)}, \tag{2}$$

where ξ_t is a zero mean stationary autoregressive process, s_t is a homogenous two-state Markov chain independent of ξ_t, and $r_t^{(m)}$ and $r_t^{(n)}$ are m and n-period interest rates, respectively, and $n > m$. According to equation (1) the shorter-term interest rate follows an AR process with a mean value changing between two regimes according to a Markov chain, while equation (2) gives the relationship between the shorter-term and longer-term interest rates. Under the pure expectations hypothesis $\alpha = 0$ and $\gamma = 1$, so that the n-period interest rate is the arithmetic average of the present and expected future m-period interest rates over the life of the n-period rate. In this paper we follow recent empirical literature and consider the less stringent version of the expectations hypothesis and allow α to deviate from zero, i.e. allow for a constant term premium.

The choice of the Markov switching model (1) is motivated by the fact that typically there have been a priori reasons to believe that regime shifts have occurred in the sample period. In this case the peso problem is present if the number of shifts in the sample is not representative of the underlying distribution of the Markov switching process. In contrast to this case we consider a situation where no regime shifts have actually occurred, but the model that the market uses to form expectations, includes the possibility of regime shifts. As pointed out by Saikkonen and Ripatti (1999), the Markov switching model cannot be used to estimate the peso effect in the absence of actual regime shifts. The reason is that the conditional expectations in equation (2) are then computed conditional on the known regime, and their sum is the same that would obtain if $r_t^{(m)}$ followed a linear AR process save an additive constant (see Hamilton (1993)). Although this additive constant adheres to the peso problem, it cannot be separated from the constant term α in equation (2) and thus the peso effect is not identified. One possibility to identify it might be to make the transition probabilities of the Markov chain dependent on some observable variables. Although this kind of a Markov switching model has been considered in some different contexts in the literature, it seems unnecessarily complicated, and hence we follow the congenial but simpler approach of Saikkonen and Ripatti (1999) in replacing equation (1) in the model by a pth order TAR process

$$r_t^{(m)} = \beta_0 + \beta_1 r_{t-1}^{(m)} + \cdots + \beta_p r_{t-p}^{(m)} + \delta I(z_{t-d} \geq c) + \eta_t, \tag{3}$$

where the indicator function $I(\cdot)$ equals unity when the threshold variable z_{t-d} is greater than or equal to the threshold value c and zero otherwise, and η_t is a martingale difference sequence satisfying $E(\eta_t | r_s^{(m)}, s < t) = 0$ and having constant variance. Moreover we assume that the roots of the lag polynomial $\beta(L) = 1 - \beta_1 L - \cdots - \beta_p L^p$ lie outside the unit circle. Since we are assuming that all the observations come from one regime, $z_t < c$, $t = 1, \ldots, T$, and the TAR model (3) reduces to a linear AR model in the sample. However, if $\delta > 0$ a jump to a higher level is possible at each point in time although it does not occur in the observed realization. When the expectations in (2) are formed using process (3) with positive δ, there is a peso effect stemming from the fact that the upper regime does not occur in the observed time series.

Saikkonen and Ripatti (1999) considered only a model where the threshold variable is the lagged level, here a lag of the m-period interest rate $r_t^{(m)}$. Also other predetermined variables, such as functions of the level, could be considered but then the interpretation of the model, of course, somewhat changes. In this paper we consider lagged differences of the m-period interest rate $r_t^{(m)}$ computed over different spans in addition to the lagged level as threshold variables. The choice of the threshold variable typically has to be based on statistical criteria, and following Teräsvirta (1994) we suggest selecting the model that minimizes the p value of the linearity test. The reason is that if the threshold variable is incorrectly selected, the model is misspecified and the power of the linearity test against a misspecified nonlinear model can hardly be expected to exceed the power against the correctly specified model. Thus the linearity test against the correct model should have the lowest p value, i.e. strongest rejection. Because linearity testing is very complicated in this model (see below), the selection rule must in practice be modified as follows. *Assuming* that the null hypothesis of linearity is rejected, the model with the largest value of the log-likelihood function is selected.

To make the model operational, we augment equation (2) with an error term ε_t and assume that the two error terms η_t and ε_t are independent. The error term in equation (2) contains the forecasting errors due to the estimation errors in (3). Thus our model consists of equation (3), the following term structure equation

$$r_t^{(n)} = \alpha + \gamma \frac{m}{n} \sum_{j=0}^{n/m-1} E_t r_{t+jm}^{(m)} + \varepsilon_t \tag{4}$$

and the restriction that $z_{t-d} < c$, $t = 1, \ldots, T$. Provided the conditional expectations in (4) can be computed given some values of the parameters of (3), the model can be estimated by the method of (quasi) maximum likelihood. Assuming that the independent error terms η_t and ε_t have a joint normal distribution, the conditional log-likelihood function (conditional on the initial values $r_{-p+1}^{(m)}, \ldots, r_0^{(m)}$) can be written as

$$l_T = -\frac{T}{2}\log\sigma_\varepsilon^2 - \frac{1}{2\sigma_\varepsilon^2}\sum_{t=1}^{T}(r_t^{(n)} - \alpha - \gamma f_{t-1}(\beta_0, \beta_1, \ldots, \beta_p, \delta, c, \sigma_\eta^2))^2 \quad (5)$$

$$-\frac{T}{2}\log\sigma_\eta^2 - \frac{1}{2\sigma_\eta^2}\sum_{t=1}^{T}(r_t^{(m)} - \beta_0 - \beta_1 r_{t-1}^{(m)} - \cdots - \beta_p r_{t-p}^{(m)})^2$$

where $f_{t-1}(\cdot)$ denotes the average of the expected short-term rates, $(m/n)\sum_{j=0}^{n/m-1} E_t r_{t+jm}^{(m)}$, and the notation aims at expressing the complicated cross equation restrictions in the model. Obviously the maximum likelihood estimation requires iterative numerical methods, and furthermore a way of evaluating the function $f_{t-1}(\cdot)$ is needed. To this end Clements and Smith (1997) recommend a simulation method where (in our application), given some fixed values of the parameters and initial values $r_{t-p+1}^{(m)}, \ldots, r_t^{(m)}$, a large number of realizations of model (3) are simulated $m(n/m-1)$ periods ahead, and the conditional expectations are obtained as averages over these realizations at each horizon. This is computationally burdensome (especially if n/m is large) since the simulation is required in the estimation at each iteration and for all observations $t = 1, \ldots, T$.[1]

Saikkonen and Ripatti (1999) consider the identifiability of the parameters and the asymptotic properties of their maximum likelihood estimators in a model similar to ours. Their conclusions imply that the parameters of our model are identified provided δ is not equal to zero which has already been ruled out by assuming it to be positive. Supposing this identifiability result holds, the standard asymptotic distribution theory can be applied. Obviously this is also true in the restricted model where the peso effects are ruled out by replacing equation (3) by the corresponding linear AR model with the term involving the indicator function removed. Thus, by the asymptotic normality of the maximum likelihood estimator in these two models, Wald tests can be constructed in the usual way, making use of the numerical Hessian matrix of the log-likelihood function. The only exception is the hypothesis $\delta = 0$ which has the interpretation of the absence of peso effects. Because the threshold parameter c is not identified under this null hypothesis, standard asymptotic results do not apply and, say, a likelihood ratio test comparing the values of the log-likelihood functions of the restricted and unrestricted models does not have the usual asymptotic χ^2 distribution. There is by now a relatively large literature on testing in a situation like this where a parameter is not identified under the null hypothesis. Unfortunately, the results typically concern certain special cases, none of which is directly applicable in our model (see e.g. Hansen (1996) and references therein). Hansen (1996, 1999) consideres linearity tests in TAR models and suggested both a method of computing the asymptotic distribution and a bootstrap method to compute the finite-sample distribution of a test statistic. Although it would be rather straightforward to extend the bootstrap method to our model, the computational burden makes its application infeasible in practice. Furthermore, Hansen's (1999) results imply that the differences between the asymptotic and finite-sample distributions based on different sets of assumptions tend to be very large, and hence more research is required to establish the properties of these methods.

[1] The estimates in Section 3 are based on 1,000 simulated realizations.

3 Empirical results

In this section we present the estimation results of the ’peso’ model. We concentrate on U.S. interest rates, mainly because the expectations hypothesis has most clearly been rejected in the U.S. market whereas there is much more support for it in the other markets (see e.g. Gerlach and Smets ((1997)). Following much of the previous literature on peso problems in the term structure cited in the Introduction we study the short end of the maturity spectrum where the different tests typically unanimously reject the expectations hypothesis with U.S. data. In part this choice of maturities is also dictated by the computational complexity of our model which increases with the maturity of the longer-term interest rate, and the availability of data. Because we are assuming that no regime switches have occurred in the sample, only the period after the abandonment of the ’new operating procedures’ of the Federal Reserve is considered. The data set was obtained from the Federal Reserve Statistical Release, and it consists of the monthly Eurodollar deposit rates for maturities 1, 3 and 6 months (the last Friday of each month) covering the period 1983:1–1999:6 (198 observations). The data are depicted in Figure 1. The yield curve seems to have been upward sloping in almost the entire sample period. The interest rates have also undergone some rather abrupt level shifts which suggests that, provided peso effects are detected, their effect on the expectations (in our model) probably turns out to have been significant only at the beginning of the sample where the interest rates are at a high level.

As a starting point for the analysis we first present the results of the commonly used regression-based tests of the expectations hypothesis. Much of the previous empirical literature has relied on the following two regression models that can be used to test the predictive ability of the yield spread for changes in the long-term and short-term interest rate, respectively (see Campbell and Shiller (1991)).

$$r_{t+m}^{(n-m)} - r_t^{(n)} = a_0 + a_1 \frac{m}{n-m}(r_t^{(n)} - r_t^{(m)}) + error\ term \quad (6)$$

$$\frac{m}{n}\sum_{i=0}^{n/m-1} r_{t+mi}^{(m)} - r_t^{(m)} = b_0 + b_1(r_t^{(n)} - r_t^{(m)}) + error\ term \quad (7)$$

When the expectations hypothesis holds, the slope coefficient in each of these regressions equals unity. These regression models involve two econometric problems. First, due to overlap the error term in (7) follows a moving average process of order $n - m - 1$ which has to be corrected for in computing the standard errors. Second, since there are no observations for Eurodollar deposit rates with maturities two and five months, we must follow the common practice of approximating $r_{t+m}^{(n-m)}$ by $r_{t+m}^{(n)}$ in (6) which leads to a positively biased estimate of a_1, as was recently shown by Bekaert, Hodrick and Marshall (1997b). The results are presented in Table 1. Throughout the empirical analysis, the one-month rate is taken as the short-term rate. In spite of the positive bias the point estimate of the slope coefficient in (6) falls short of unity in both cases, contrary to the implication of the expectations hypothesis. The null hypothesis of a unit coefficient cannot, however, be rejected at any reasonable significance level. As far as regression (7) is concerned, the expectations hypothesis is rejected at the 7% level for the three-month rate and at the 5% level for the six-month rate. Hence the result is somewhat inconclusive.

Next we consider the estimation of the term structure model of Section 2. The first step is the choice of the threshold variable, and Table 2 gives the value of the maximized log-likelihood function for different specifications. The four threshold variable candidates are lagged one-month rate and its six-, twelve- and eighteen-month lagged differences. In models with any of the lagged differences as a threshold variable there is a high probability of a regime shift near the three local peaks of the one-month rate, whereas in the model with the lagged level as a threshold variable this probability is high only near the maximum of the series. According to the decision rule discussed above, the lagged level of the one-month rate is clearly selected which indicates that the potential peso problem is active only near the maximum of the one-month rate. Because the data are monthly, only the first lags were considered. The estimation results of the models involving the peso effects with the lagged level as a threshold variable and the restricted models are presented in Table 3. The order of the autoregressive polynomial was chosen to be one based on error autocorrelation tests for a linear AR model. The threshold parameter c was estimated using a grid search, and therefore no standard errors are available. In both the three- and six-month case c was estimated somewhat above the maximum of the one-month rate in the sample (11.810) as was to be expected.

Given the high persistence of the one-month rate the estimated values of δ indicate large jumps in both cases. It is, however, interesting to note that the estimate of δ is much lower in the model involving the six-month rate compared to that in the model involving the three-month rate despite the fact that the estimates of c are almost equal. Taken at face value this would indicate that the market uses two different models to form expectations of the one-month rate. The discrepancy is, however, just a consequence of the fact that with the longer forecast horizon, the regime shift becomes more probable, and this has not explicitly been taken into account in formulating the model. Thus to reconcile the model with the data, the higher probability must be compensated by a smaller value of the jump. The equality of the jumps in the models for the different maturities would indicate that the peso problem gets worse with the length of the forecast horizon, and there is no a priori reason to expect that to happen. The values of the likelihood ratio statistic for the null hypothesis of no peso effects, i.e. $\delta = 0$, are large in both cases (178.1, and 139.8, for the three and six-month rates, respectively), but as discussed in Section 2, the standard critical values from the χ^2 distribution are not valid for this test.

The test of the null hypothesis $\gamma = 1$ provides a tests of the expectations hypothesis. Because it is natural to assume that the presence of peso effects would manifest itself as a positive value of δ, values greater than unity of γ are expected to be found if the peso effects are erroneously set to zero. This follows because the expectations from the linear AR model cannot be greater than those from the nonlinear TAR model where $\delta > 0$, and if peso effects really prevail, a larger value of γ is required to reconcile the expectations of the one-month rate base on the AR model with the longer-term rate. Therefore, it seems natural use a one-tailed t test on γ. In the restricted models the hypothesis that $\gamma = 1$ is rejected at the 5% level (the observed levels of significance are 1.7% and 1.5% for the three and six-month rates, respectively), whereas in the unrestricted models it cannot be rejected (the observed significance levels are 10.2% and 11.2%, respectively). Thus the peso effects seem to explain the rejection of the expectations hypothesis. To further confirm that the nonrejection in the presence of peso effects was not just caused by low power, we also calculated the values of γ implied by the inverse power function (see Andrews (1989)) for 50% power. The values for the three and six-month rates were 1.033 and 1.081, respectively, indicating that for any γ less than these values the probability of rejection is less than 50%. Thus the test seems to have fair power against alternatives relatively close to the value given by the null hypothesis, and in particular, each value deviates less from unity than the estimate in the corresponding restricted model.

The derivation of diagnostic tests is complicated by the fact that the function $f_{t-1}(\cdot)$ in the likelihood function cannot be written in closed form. However, to give some idea, the residuals of the unrestricted models are plotted in Figures 2 and 3. Especially the residuals from the equation for the one-month rate seem to be somewhat heteroskedastic with lower variance in the latter half of the sample.[2] There also seem to be some outliers in all the residual series. The residuals of the first equation depicted in the upper panel of Figures 2 and 3 seem to be serially uncorrelated, whereas the residuals of the term structure equation are clearly autocorrelated. The first-order autocorrelation coefficients in the models for the three- and six-month rates are 0.34 and 0.58, respectively, and the higher-order autocorrelation coefficients die out relatively slowly. This is to be expected in small samples in this kind of models where forecasts computed using one equation are fed into another, and it need not indicate model misspecification. The probable reason are the estimation errors of the equation for the one-month rate that enter the error term of the term structure equation multiplied by lags of the persistent one-month rate. For instance, assuming that the lag length equals one in the linear model of the one-month rate, the error term of the equation for the three-month rate becomes

$$\widehat{\alpha}-\alpha+\frac{1}{3}\left(\widehat{\gamma}-\gamma\right)r_t^{(1)}+\frac{1}{3}\left(\widehat{\gamma}(\widehat{\beta}_1+\widehat{\beta}_1^2)-\gamma(\beta_1+\beta_1^2)\right)r_{t-1}^{(1)}+\varepsilon_t.$$

Obviously the high first-order autocorrelation coefficient of the one-month rate is a component in the first-order autocorrelation coefficient of the residual, and thus the residuals can be strongly autocorrelated even if the estimation error is small.

Interpretationally interesting is Figure 4 that shows the differences of the fitted values of the longer-term rate from the unrestricted and restricted models. The unrestricted model implies a somewhat higher overall level of the longer-term rate in both cases, but it is only in the period 1984:4–1984:7 that the peso effect is visible. This was, of course, the period when the one-month rate was closest to the threshold values and when the probability of a regime shift, according to the model used to form expectations, was highest.

4 Conclusion

In this paper we examine whether the data on U.S. Eurodollar deposit rates are reconcilable with the expectations hypothesis once the presence of a potential regime shift in the process of the one-month rate is allowed for. The main idea is that in forming expectations of the future one-month rate the market gives a positive probability to a regime shift, but in the sample period all the observations belong to a single regime. Hence, when the rationality of expectations is tested assuming that

[2] The heteroskedasticity does not seem to follow from restricting the error variance to be constant in the TAR model (3). We also attempted to estimate models with variance differing accross the regimes, but virtually no improvement in the likelihood function was achieved. The identification problems discussed above, make the formal testing of the constancy of the error variance cumbersome.

no regime shift was not even expected, the so called peso problem arises. The extent of the peso effect is measured by the discrepancy between the expectations based on the model allowing for a regime shift and the model having just a single regime. If this difference is considerable, the rationality of expectations can be rejected if the possibility of a regime shift is not taken into account in testing. Our approach differs from the previous literature in two interconnected respects. First, instead of the commonly applied Markov switching model we employ a version of the model due to Saikkonen and Ripatti (1999) in which the regime shift is modelled by means of a TAR model. Second, we assume that no regime shift has actually occurred in the sample period contrary to most of the previous literature that has a regime switching model as a starting point and claims the peso effect to be a consequence of having too few regime shifts in the observed sample. As Saikkonen and Ripatti (1999) pointed out, the peso effect is not identified in a homogeneous Markov switching model if no regime shift has occurred which motivates the TAR model as a viable alternative specification.

To be able to plausibly make the a priori assumption of no realized regime shift, monthly data from the period after the abandonment of the new operating procedures of the Federal Reserve (1983:1–1999:8) were used. The empirical results lend support to the expectations hypothesis at the short end of the maturity spectrum of U.S. Eurodollar deposit rates once a potential regime shift is allowed for. Furthermore, it turns out that the peso effect was really working only in an early period in the sample when interest rates were at a high level. Because this may be a feature of our particular model specification, alternative specifications were also considered to find out whether peso effects really were absent later in the sample period. The first model, however, outperform the other specifications. Yet another possibility that was not considered at all might be to model the one-month rate as a Markov switching model with transition probabilities dependent on some observable variables (see e.g. Diebold et al. (1994)) instead of a threshold autoregression. This possibility was also hinted at by Saikkonen and Ripatti (1999), but as they noted, this kind of a model would be rather complicated in practice.

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Table 1. **Estimates of the slope coefficients in the regressions of changes in the long-term (a_1) and short-term (b_1) interest rates on the yield spread, respectively.**

n	3	6
a_1	0.851 (0.227)	0.685 (0.372)
b_1	0.721 (0.153)	0.489 (0.257)

The shorter-term interest rate is the one-month rate and n is the maturity of the longer-term interest rate. The figures in parentheses are standard errors. The standard errors of b_1 are computed by the method of Newey and West (1987).

Table 2. **Values of the maximized log-likelihood function in term structure models with different threshold variables.**

Threshold variable	n = 3	n = 6
–	227.399	148.162
$r_{t-1}^{(1)}$	316.426	218.073
$\Delta^6 r_{t-1}^{(1)}$	308.496	214.313
$\Delta^{12} r_{t-1}^{(1)}$	295.456	205.433
$\Delta^{18} r_{t-1}^{(1)}$	298.410	148.162

n is the maturity of the longer-term rate.

Table 3. **Estimation results of the term structure models.**

n	3		6	
	unrestricted	restricted	unrestricted	restricted
c	11.819		11.813	
δ	1.674 (2.012)		0.806 (0.381)	
β_0	0.108 (0.134)	0.186 (0.092)	0.109 (0.137)	0.169 (0.092)
β_1	0.980 (0.017)	0.968 (0.013)	0.980 (0.017)	0.971 (0.013)
α	-0.011 (0.158)	-0.134 (0.121)	-0.067 (0.396)	-0.304 (0.272)
γ	1.025 (0.020)	1.037 (0.018)	1.060 (0.050)	1.086 (0.040)
σ_η	0.386 (0.011)	0.387 (0.020)	0.386 (0.011)	0.387 (0.020)
σ_ε	0.191 (0.005)	0.300 (0.015)	0.315 (0.011)	0.449 (0.023)

The shorter-term interest rate is the one-month rate and n is the maturity of the longer-term interest rate. The threshold variable is $r_{t-1}^{(1)}$. The figures in parentheses are standard errors computed by the inverse of the final Hessian. The unrestricted models allow for peso effects while these effects are absent in the restricted models.

Figure 1. **Monthly Eurodollar deposit rates 1983:1–1999:8 (solid line = one-month rate, dashes = three-month rate, dots = six-month rate).**

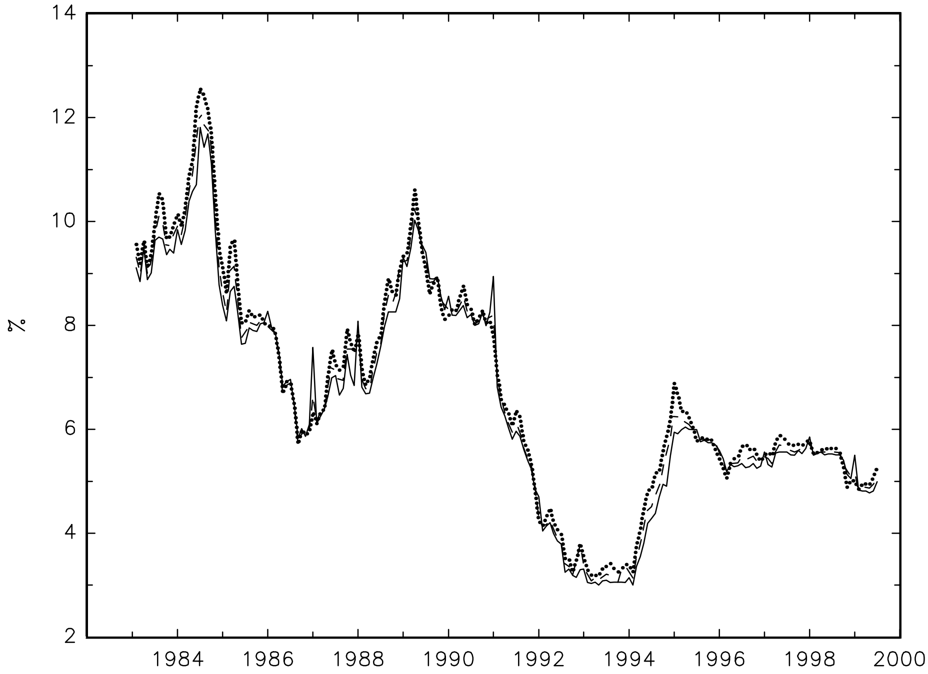

Figure 2. **Residuals of the model for the one-month (upper panel) and three-month (lower panel) rates.**

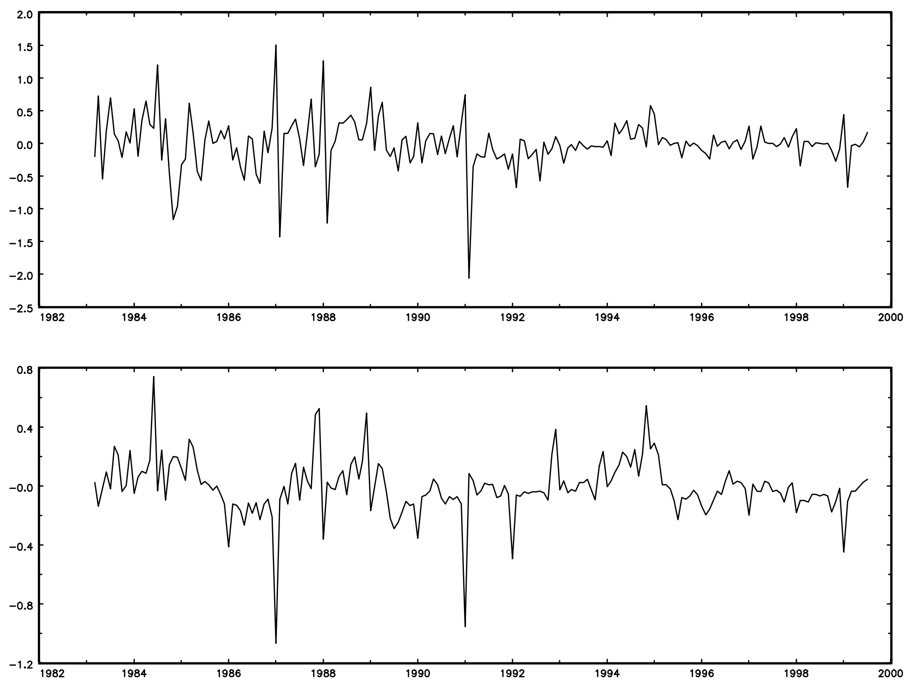

Figure 3. **Residuals of the model for the one-month (upper panel) and six-month (lower panel) rates.**

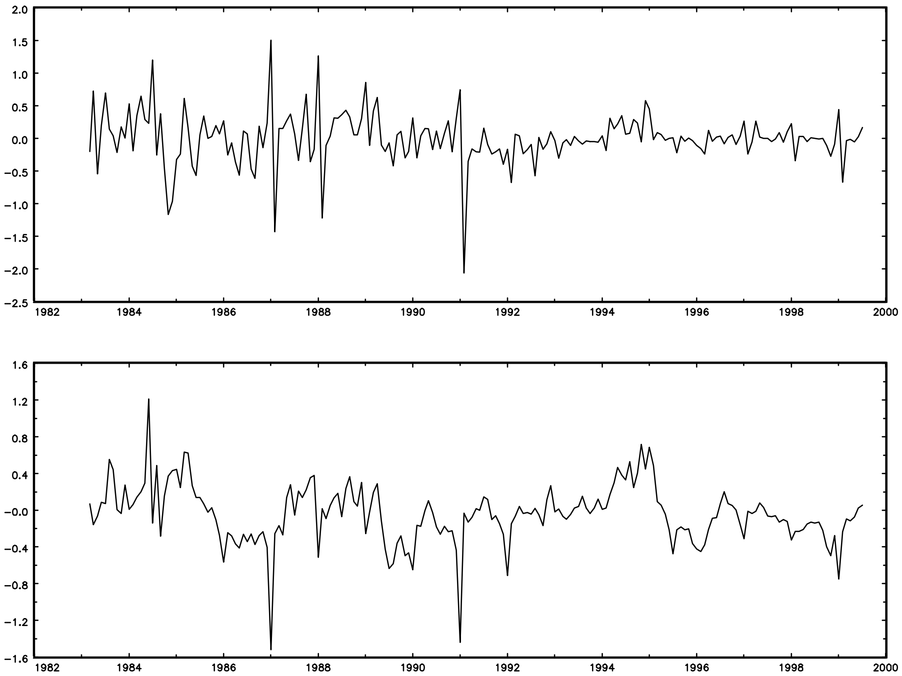

Figure 4. **Differences of the fitted values of the longer-term rate from the unrestricted and restricted models for the three-month (solid line) and six-month (dashes) rates.**