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DIRECTION GÉNÉRALE DES ÉTUDES ET DES RELATIONS INTERNATIONALES

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Term Structure Anomalies : Term Premium or Peso problem?

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Résumé

L'objectif de cet article est de tester l'importance relative des hypothèses de "prime de terme variable dans le temps" et de "l'effet peso" pour le rejet empirique de la théorie des anticipations de la structure par terme des taux (TAST). Le raisonnement est fondé sur l'étude d'un modèle linéaire de la structure par terme autorisant la présence de ces deux phénomènes de façon simultanée. Sous l'hypothèse que seul un régime est observé ex-post, il est possible de quantifier les distorsions engendrées par ces deux hypothèses. Par ailleurs il est également possible de tester la présence d'un "effet peso" dans les données. Les résultats sont les suivants : l'effet peso pourrait expliquer le rejet de la théorie des anticipations de la structure par terme en Allemagne et au Royaume Uni ; en revanche, cette hypothèse ne peut expliquer le rejet de la théorie aux Etats-Unis.

Mots clés : Théorie des anticipations de la structure par terme, effet peso, prime de terme variable.

Classification JEL : E43, E52, C22.

Abstract

The goal of this paper is to develop a test for the relative importance of the time-varying term premium and the peso-problem for rejection of the Expectation Hypothesis of the Term Structure (EHTS). Our reasoning is based on a term structure model that allows for both phenomena simultaneously. If we assume that only one regime is observed ex-post, we can estimate all the information we need to evaluate distortions generated by both hypotheses. We can also test the presence of a peso-problem. Firstly we find that a peso-problem might explain rejection of the EHTS in Germany and the United Kingdom after the European exchange rate crisis. Secondly, we show that this explanation appears inappropriate to explain the EHTS failure in the United States.

Keywords: Expectation theory of the term structure, peso problem, time varying term premium.

JEL Classification : E43, E52, C22.

Résumé non technique

Le modèle le plus répandu pour expliquer les mouvements de la courbe des taux repose sur la théorie des anticipations de la structure par terme. Une version simple stipule que le rendement d'un titre de maturité N est égal à la moyenne des taux courts futurs anticipés.

Une telle représentation a fait l'objet de multiples tentatives de validation empiriques. La plupart des tests réalisés sur données américaines sont défavorables à la théorie des anticipations. Elle est également rejetée sur données allemandes.

Deux possibilités sont évoquées pour expliquer ce rejet. La première revient à considérer une prime de terme variable dans le temps ; la seconde met en cause les erreurs de prévision des agents. Plus recemment, tout un pan de la littérature s'est développé autour de l'effet de ces erreurs de prévisions lorsqu'elles résultent de l'anticipation (rationnelle) d'évènements non observés ex-post. On fait alors référence à un ensemble de résultats d'inférence sur petit échantillon connus sous l'appellation "d'effet peso".

Un "effet peso" a lieu quand la possibilité d'un changement de régime dans la distribution des chocs futurs affecte les anticipations (rationnelles) des agents. Généralement, cette situation est observée quand la distribution du processus générateur de données inclut un évènement rare, entraînant une très forte désutilité pour les agents économiques. Parce que cet état est rare, il est peu probable qu'il soit observé sur un échantillon de taille habituellement considéré convenable pour l'inférence statistique. Parce qu'il est considéré comme "catastrophique", la simple éventualité qu'il puisse se réaliser affecte significativement les anticipations des agents, et donc les prix d'équilibre. Les conséquences d'un "effet peso" sur l'inférence sont alors immédiates : si la distribution ex-post, c'est à dire celle observée, est différente de celle ex-ante, c'est à dire celle à partir de laquelle les agents fondent leurs anticipations, alors les moments calculés à partir de l'échantillon ont peu de chance de coïncider avec leurs contreparties théoriques.

Peu d'études ont cherché à déterminer lequel de ces deux effets, "prime de terme" et "effet peso", prédomine dans le rejet de la théorie des anticipations de la structure par terme. Cela tient notamment à deux difficultés. Premièrement, la prime de terme et les erreurs de prévision ne sont pas directement observables et il est donc difficile d'identifier correctement leurs

effets respectifs. Deuxièmement, la présence d'un "effet peso" induit des distorsions au niveau de l'inférence statistique impliquant qu'il ne peut être estimé par les méthodes économétriques habituelles (modèles à changement de régime par exemple).

L'exercice mené dans cet article permet de lever ces difficultés et propose une évaluation de l'influence respective de ces deux phénomènes dans le rejet de la théorie des anticipations de la structure par terme. Le raisonnement repose sur l'étude d'un modèle linéaire des taux. La prime de terme est supposée suivre un processus autoregressif. L'effet peso résulte de l'hypothèse suivante : on suppose que les agents anticipent ex-ante un changement de régime du taux court qui ne se réalise pas ex-post. On montre alors que l'estimation de ce modèle nous donne toute l'information nécessaire pour évaluer les biais engendrés par la prime de terme et par "l'effet peso".

Cette approche est appliquée aux taux de maturités 3 et 6 mois pour l'Allemagne, le Royaume-Uni et les Etats-Unis. Les tests menés font ressortir la présence d'un "effet peso" en Allemagne et au Royaume Uni après la crise européenne du change de 1992. En revanche, les tests ne permettent pas de conclure à la présence d'un "effet peso" aux Etats-Unis.

1 Introduction

One of the most interesting empirical paradoxes that emerges when one focuses on the yield curve is the rejection of the Expectation Hypothesis of the Term Structure (hereafter EHTS). Recent versions[1] of this theory suggest that the long-term interest rate is a weighted average of current and expected future short-term interest rates, plus a constant term premium. To what extent this theory is a good representation of the observed yield curve is a crucial question for investors and financial market participants. Actually, a direct consequence of this hypothesis is that the interest rate spread, defined as the difference between a long rate and a short rate, should be a good leading indicator of changes in interest rates. Unfortunately, most of the tests performed on US data reject the EHTS (see Campbell and Shiller (1987,1991), Hardouvelis (1994), Kugler (1990), Evans and Lewis (1994)). Studies concerning other countries lead generally to the same conclusions (see Hardouvelis (1994), Gerlach (1996), Gerlach and Smets (1997), Jondeau (1997)).

Two main alternative explanations for this rejection have been proposed. The first one assumes that the information in the spread is composite information about the variation of both expected future rates and time-varying term premia. Consequently, changes in spreads that result from changes in term premia, provide no information regarding future short rate variations. However, as noted by Mankiw and Miron (1986), while stories of highly variable term premia might explain the failure of the EHTS for very long-term yields, such stories seem less plausible applied to the markets for three-month and six-month bills. On the other hand, Evans and Lewis (1994) show that a stationary time-varying, even highly variable, is not sufficient to generate all the empirical characteristics of the US term structure. The second explanation draws attention to the forecast errors made by agents when they predict future changes in interest rates. The key element of this reasoning is that the short rate is hard to predict when monetary authorities use it as a policy instrument. Actually, potential changes in monetary policy regime should influence rational expectations made by the agents. Hamilton (1988) and Kugler (1996) point to the empirical relevance of regime uncertainty for

[1] In contrast with the original version of EHTS in which the term premium is zero.

the US term structure. A more recent strand of literature has investigated consequences of expected regime changes when they are not materialized ex-post. Indeed an increase (decrease) in the spread might result from the expectation of a monetary policy regime with higher (lower) short rates. If this regime is not observed on the sample of data, behaviour of the spread appears inconsistent with short term interest rates (and the EHTS). In this case, ex-post market forecasts of the future short rates are biased. This phenomenon falls into a particular issue in small sample inference known as peso problem.
A peso problem occurs when the potential for discrete shifts in the distribution of future shocks to the economy affects the rational expectations of market participants. It is generally the case when the distribution of the data generating process includes a low probability, usually catastrophic, state that generates extreme disutility to economic agents. Because this state has a low probability, it is unlikely to be observed in a given small sample. Because the state is catastrophic the possibility that it may occur substantially affects agents' decisions and hence equilibrium prices. Consequences of a peso-problem for inference are then immediate : if the ex-post distribution, that is the observed one, is different from the true ex-ante distribution, then the sample moments calculated from the available data do not coincide with their theoretical counterparts. Lewis (1994) shows that this phenomenon can explain some empirical anomalies in US term structure, when a spread increase (decrease) is not followed by a long term interest rate increase (decrease). Bekeart, Hodrick and Marshall (2001) incorporate a peso-problem and a low volatility term premium into the EHTS model. They show that this model is largely consistent with term structure data from the United States, the United Kingdom and Germany. Furthermore, their paper outlines the importance of taking into account both explanations, a time-varying term premium and a peso problem to reproduce the properties of the term structure of interest rates.

To our knowledge, no papers have investigated which of these two hypotheses dominate in explaining the rejection of the EHTS. Indeed, answering this question raises difficult methodological problems. Firstly term premia and forecast errors are not observable. Therefore they have to be estimated from the sample of data. This point leads to a second empirical implementation issue: how can one do inference from the sample distribution underlying a peso-problem when, by definition, a peso-problem only exists when there are insufficient data to estimate that population distribution? Bekeart and al. (2001) overcome this small-sample problem by pooling short-rate data from several different countries. In so doing, they make the strong assumption that the data generating process is the same for all the countries considered. This assumption seems far-fetched given the empirical international heterogeneity of interest rate properties. Therefore, although the paper of Bekeart and al. (2001) incontestably provides an interesting first approximation of the peso-problem phenomenon, we think that alternative solutions to the small-sample issue have to be examined.

The goal of this paper is to develop an alternative test for the relative importance of a time varying term premium or the peso-problem for rejection of the EHTS. Our reasoning is based on a term structure model that allows for both phenomena simultaneously. The long rate is defined as the average of current and expected future short-rates, plus a time varying term premium. We assume that changes in the short rate can switch between two processes. In order to generate a peso-problem, we assume that only one of these regimes is observed ex-post. Under this assumption, we can estimate a reduced form of the model that provides all the information we need to distinguish and evaluate econometric distorsions generated by the time varying term premium on one hand, and by the peso-problem on the other hand. We also propose a procedure that allows us to test the null hypothesis of "no peso-problem" against the alternative hypothesis of "possibility of a peso-problem".

Our approach is applied to three-month and six-month maturity eurorates for Germany, the United Kingdom and the United States. Our findings are as follows. First we divided each sample into sub-samples characterizing periods of economic stability and we show that the EHTS is rejected for some of this sub-period. Secondly, we test whether these rejections are attributed to a peso-problem. We show that we are unable to reject the null hypothesis of no peso-problem in Germany and the United Kingdom before the European exchange rate crisis of 1992 and we reject this hypothesis after this date. We also show that this hypothesis is not rejected in the United States for all the periods considered. In addition, with regard to Germany and the United Kingdom, we show that distorsions generated by a time-varying term premium are quantitatively more important than those generated by a peso problem.

The paper is organized as follows. In the following section we test the EHTS for Germany, the United Kingdom and the United States. In section 3 we present the model. In section 4 we show how to evaluate biases generated by the time-varying term premium and the peso-problem. We also present a procedure that allows us to test the presence of a peso-problem. In section 5 we report empirical results. Section 6 provides a conclusion.

2 Evidence on the EHTS in Germany, the United Kingdom and the United States

2.1 The Campbell-Shiller regressions

The EHTS describes how a longer-term N-period interest rate R_t^N is related to a shorter-term one, say a one-period interest rate i_t. More precisely, it states that:

$$R_t^N = \frac{1}{N} \sum_{k=0}^{N-1} E[i_{t+k}|\Omega_t] + \overline{\Phi} \tag{1}$$

where Ω_t denotes the information set available in t, and $\overline{\Phi}$ a constant parameter. Therefore, the N-period interest rate is a constant, plus a simple average of the current and expected future one-period rates up to N-1 periods in the future. The parameter $\overline{\Phi}$ reflects a term premium, that is a predictable excess return on the N-period bond over the one-period bond. According to the EHTS the term premium is allowed to vary with N, but is assumed to be constant through time.

Equation (1) can be obtained directly if one assumes that expected continuously compounded yields to maturity on all discount bonds are equal, up to a constant (Fama (1984)). It can also be derived as a linear approximation to any several different non linear expectations theories of the term structure. According to Campbell and Shiller (1991) the approximation is quite adequate for most purposes.

Let us focus our attention on the behaviour through time of the spread between the N-period rate and the one-period rate, $S_t^N = R_t^N - i_t$. The EHTS implies that the spread is a constant term premium, plus an optimal forecast of changes in future interest rates. For instance, if $N = 2$ equation (1) reduces to:

$$R_t = \frac{1}{2} i_t + \frac{1}{2} E[i_{t+1}|\Omega_t] + \overline{\Phi} \tag{2}$$

where R_t is the two-period interest rate.
If $S_t = R_t - i_t$ is the spread between the two-period and one-period interest rate, and $\Delta i_{t+1} = i_{t+1} - i_t$ the one-period changes in the short term interest rate, rearrangement of equation (2) gives:

$$E[\Delta i_{t+1}|\Omega_t] = 2S_t - \overline{\Phi} \tag{3}$$

This expression reflects the fact that under the EHTS the current value of the spread should help to predict the one-period changes in the short term interest rate. Therefore, following Campbell and Shiller (1991), we can test the model by regressing the one-period changes of i_t onto the spread and testing whether the coefficient equals 2.

To be more precise, we consider the following regression:

$$\Delta i_{t+1} = b_0 + b_1 S_t + u_{t+1} \tag{4}$$

Coefficients of this equation can be estimated by ordinary least square (OLS) methods and one can test whether estimation of the slope coefficient, that is b_1, is consistent with the EHTS. In other words we test the hypothesis $H_0 : b_1 = 2$. Rejection of H_0 leads us to reject the EHTS.
In the next paragraph this test is performed on three OECD countries: Germany, the United Kingdom and the United States.

2.2 The case of Germany, the United Kingdom and the United States

In this paragraph we investigate whether the EHTS gives a good representation of the short-end of the euro-rate term structure in Germany, the United Kingdom and the United States[2]. For that purpose, we apply the

[2]The choice of Euro-rates guards us against heterogeneity such as differences in duration, in calculation of yield etc. The interest rates from different countries are directly comparable (See also Kugler (1990) and Gerlach and Smets (1997)).

Campbell-Shiller linear regression test presented above to the 3-month (for i_t) and the 6-month (for R_t) Euro-rates[3].

For each country, the whole period of observation is divided into sub-periods (see table 5) which are historically identified as different regimes of stability[4].

[Insert Table 5 about here]

Tables 6, 7 and 8 report OLS results of equation (4) for Germany, the United Kingdom and the United States respectively. For each country we provide results obtained for each sub-period described in table 5. The first column reports the b_1 coefficient estimate. In the second column we report the Student statistic associated with the H_0 hypothesis $b_1 = 0$ (the number in brackets indicates the corresponding p-value). The third column presents the Student statistic associated with the H_0 hypothesis $b_1 = 2$ and the corresponding p-value. The EHTS is rejected at the 5% level when the hypothesis $b_1 = 2$ is rejected at the 5% level (the p-value is smaller than 5%). The last column reports conclusions regarding the EHTS, i.e whether this theory is rejected or not.

[Insert Table 6 about here]

[Insert Table 7 about here]

[Insert Table 8 about here]

We are unable to reject the EHTS in Germany during the periods immediately following the first and the second oil shock, up to 1984. On the other hand, results of the estimate indicate that the theory is rejected after this date.
The EHTS is not rejected in the United Kingdom until the exchange rate crisis in Europe that occurred in September 1992. After this date, the 3-month and 6-month interest rate data reject the theory.
In the United States, the EHTS is not rejected for two sub-periods. First, during the short period of the non-borrowed-reserves operating procedure

[3]Data are described in the appendix.
[4]See Clarida and Gertler (1996) for Germany, Tootley (2002), Vila Wetherilt (2002) for the United Kingdom, Walsh (1998) and Mishkin (1995) for the United States

(1979 to 1982) and during Alan Greenspan's term as Fed chairman. From 1973 to the date of the appointment of Paul Volcker as Fed chairman, the EHTS is rejected. The theory is also rejected from november 1982 to august 1987, that is in the period preceding the appointment of Alan Greenspan as Fed chairman.

In conclusion, for the three countries considered we can identify sub-periods for which the 3-month and 6-month interest rate data reject the EHTS. In the remainder of the paper we investigate whether this rejection is due to a "time-varying term premium" or to a peso-problem.

3 The Model

In the literature we mainly find two possible explanations for the rejection of the expectation hypothesis. First, the term premium may vary with time. This explanation is the most natural one. Actually, many studies support the existence of a time varying term premium (Evans Lewis (1994), Cochrane and Piazzesi (2002)). In what follows, we introduce this possibility assuming that the term premium follows an autoregressive process of order p.
The rejection of the expectation hypothesis may also occur when the expectation errors are not orthogonal to the period t information set. This phenomenon may be generated by a small sample inference problem known as the peso-problem. This is the case when agents anticipate a regime switch that is not materialized $ex-post$. In order to formalize this idea, we assume that changes in the short term interest rate can switch between two regimes.

3.1 The time-varying term premium

We introduce a time varying stationary term premium, Φ_t. In order to simplify calculations, we assume that :

$$\Phi_t = \overline{\Phi} + \eta_t \tag{5}$$

where η_t is an autoregressive process of order p: $\Theta(L)\eta_t = \omega_t$. L is the lag operator, $\Theta(z)$ is a polynomial function of order p, $E(\omega_t) = 0$, $E(\omega_t\omega_t) = \sigma^2_\omega$ and $E(\omega_t\omega_{t+k}) = 0$, $\forall k \neq t$.

3.2 Changes in the short term interest rate

We assume that changes in short term interest rate Δi_{t+1} can switch between two processes. We also assume that switches in the process are indicated by changes in a discrete-valued variable, $z_t = \{0, 1\}$, and let $\Delta i_{t+1}(z)$ denote the realized change in short term interest rates in regime $z_{t+1} = z$[5].

3.2.1 "within-regime", ex-ante and ex-post forecast errors

We first focus our attention on the forecast error that agents make when they know that the $t+1$ regime is regime z. Let :

$$\varepsilon_{t+1}(z) = \Delta i_{t+1}(z) - E(\Delta i_{t+1}(z)|\Omega_t) \tag{6}$$

designate this forecast error, for $z = \{0, 1\}$

We refer to this as the $within - regime$ forecast error. To be more precise, $\varepsilon_{t+1}(0)$ and $\varepsilon_{t+1}(1)$ are the forecast errors made by agents within the regime 0 and 1 respectively. In both cases, this forecast error inherits the properties of conventional rational expectations forecast errors :
$E(\varepsilon_{t+1}(z)|\Omega_t) = 0$, for $z = \{0, 1\}$.
The value of z is not observed by market participants.
Given the definition of the *within-regime* forecast error, we can break down actual changes in interest rate into the conditionally expected change in regime z, $E(\Delta i_{t+1}(z)|\Omega_t)$, and a residual ε_{t+1} :

$$\Delta i_{t+1} = E(\Delta i_{t+1}(0)|\Omega_t) + z_{t+1}(E(\Delta i_{t+1}(1)|\Omega_t) - E(\Delta i_{t+1}(0)|\Omega_t)) + \varepsilon_{t+1} \tag{7}$$

where $\varepsilon_{t+1} = \varepsilon_{t+1}(0) + z_{t+1}(\varepsilon_{t+1}(0) - \varepsilon_{t+1}(1))$

We now focus on the forecast error that agents make when they predict Δi_{t+1} without knowing the $t+1$ regime. We express this forecast error as :

$$e_{t+1} = \Delta i_{t+1} - E(\Delta i_{t+1}|\Omega_t) \tag{8}$$

We refer to it as the *ex-ante* forecast error in contrast to the *ex-post* forecast error, $e_{t+1}(z)$, that is the actual value of e_{t+1} when $z_{t+1} = z$. More precisely, the *ex-post* forecast error is given by :

$$e_{t+1}(z) = \Delta i_{t+1}(z) - E(\Delta i_{t+1}|\Omega_t) \tag{9}$$

[5] In the following, we will denote $X_{t+1}(z)$ the realized value of the random variable X_{t+1} when $z_{t+1} = z$ for $z = 0, 1$.

The $ex-ante$ forecast error always inherits conventional rational expectation properties. In contrast, we can show that the $ex-post$ forecast errors may appear biased and correlated with $ex-ante$ information when the market participants expect a regime switch that does not occur in the sample period of observation. In this case, we will say that there is a peso-problem. By way of illustration, let us rewrite (9) using equation (7). We obtain the result that the $ex-post$ forecast error does not coincide with the $within-regime$ forecast error as defined in (6). Indeed, we have :

$$e_{t+1}(z) = \varepsilon_{t+1} + (z - E(z_{t+1}|\Omega_t))(E(\Delta i_{t+1}(1)|\Omega_t) - E(\Delta i_{t+1}(0)|\Omega_t)) \quad (10)$$

Hence, under the assumption that z is constant, say equal to zero, during the period of observation, the $ex-post$ residual is given by:

$$e_{t+1}(0) = \varepsilon_{t+1} - Pr(z_{t+1} = 1|\Omega_t)(E(\Delta i_{t+1}(1)|\Omega_t) - E(\Delta i_{t+1}(0)|\Omega_t)) \quad (11)$$

Consequently, as soon as the market believes that regime 1 is possible, i.e that $Pr(z_{t+1} = 1|\Omega_t) > 0$, the $ex-post$ forecast error is biased and correlated with $ex-ante$ information. Thus, when market participants expect a regime switch which does not occur in the sample period observation, a peso-problem occurs[6].

3.2.2 Specification of each regime

To go into more detail, we assume that :

$$\Delta i_{t+1}(0) = \mu_0 + \rho_0 \Delta i_t + \varepsilon_{t+1}(0) \quad (12)$$

and

$$\Delta i_{t+1}(1) = \mu_1 + \rho_1 \Delta i_t + \varepsilon_{t+1}(1) \quad (13)$$

We also suppose that the variable z_t follows a Markov process of order 1, that is, we have $Pr(z_{t+1}|\Omega_t) = Pr(z_{t+1}|z_t)$.
Let us designate :

$$P_{00} = Pr(z_{t+1} = 0|z_t = 0)$$

[6]More generally, this kind of distorsion occurs as soon as the number of shifts in the sample observation period is not representative of the underlying distribution of z_{t+1}. For details, we refer the reader to the excellent exposition of Evans (1995).

and

$$P_{11} = Pr(z_{t+1} = 1 | z_t = 1)$$

Then, based on these assumptions the *ex-post* forecast error is given by:

$$e_{t+1}(z) = \varepsilon_{t+1} + (z - E(z_{t+1}|\Omega_t))((\mu_1 - \mu_0) + (\rho_1 - \rho_0)\Delta i_t) \tag{14}$$

and if only regime 0 occurs $ex - post$, we have:

$$e_{t+1}(0) = \varepsilon_{t+1}(0) - (1 - P_{00})((\mu_1 - \mu_0) + (\rho_1 - \rho_0)\Delta i_t(0)) \tag{15}$$

To summarize, we assume that the data generating process is a two-regime switching model as in (12) and (13) and that the market participants make their forecasts according to this DGP. In what follows, we also assume that during the sample period, only regime 0 occurred. These two assumptions generate a Peso-problem.

3.3 The interest rate spread

We assume that the two-period interest rate, R_t is given by :

$$R_t = \frac{1}{2} i_t + \frac{1}{2} E(i_{t+1}|\Omega_t) + \Phi_t \tag{16}$$

Therefore, the spread between the two-period and the one-period interest rate, S_t, is given by:

$$S_t = \frac{1}{2} E(\Delta i_{t+1}|\Omega_t) + \Phi_t \tag{17}$$

Using (5) and (8), we have :

$$S_t = \overline{\Phi} + \frac{1}{2}\Delta i_{t+1} - \frac{1}{2} e_{t+1} + \eta_t \tag{18}$$

4 How can we explain the rejection of the EHTS

In this section we show how a mixture of a time-varying term premium and a peso-problem produce biaises on estimates of the Campbell-Shiller regressions. If we assume that only one regime is observed ex-post, we show these biases can be distinguished and evaluated. Furthermore, we also develop a procedure that allows us to test the null hypothesis of "no peso-problem" against the alternative hypothesis of "possibility of a peso-problem".

4.1 Consequences for the Campbell-Shiller regressions : "term premium" and peso-problem biases

If we consider the regression :

$$\Delta i_{t+1} = b_0 + b_1 S_t + u_{t+1}$$

The least square estimate of b_1 is given by :

$$\widehat{b}_1 = \frac{cov_T(S_t, \Delta i_{t+1})}{var_T(S_t)} \tag{19}$$

where cov_T and var_T denote the sample covariance and variance, and T is the number of observations.
Hence we have :

$$\widehat{b}_1 \rightarrow_{T \rightarrow \infty} \frac{cov(S_t, \Delta i_{t+1})}{var(S_t)} \tag{20}$$

A rearrangement of equation (18) leads to:

$$\Delta i_{t+1} = -2\overline{\Phi} + 2S_t + e_{t+1} - 2\eta_t \tag{21}$$

Hence :

$$\frac{cov(S_t, \Delta i_{t+1})}{var(S_t)} = 2 - 2\frac{cov(S_t, \eta_t)}{var(S_t)} + \frac{cov(S_t, e_{t+1})}{var(S_t)} \tag{22}$$

If T is large enough we have:

$$\widehat{b}_1 \simeq 2 - 2\frac{cov(S_t, \eta_t)}{var(S_t)} + \frac{cov(S_t, e_{t+1})}{var(S_t)} \tag{23}$$

Equation (23) indicates how the point estimate of b_1 can deviate from 2. The bias can be broken down into two parts. The first one, $-2\frac{cov(S_t,\eta_t)}{var(S_t)}$, is the bias generated by the time varying term premium. This term is zero when the term premium is constant, that is $\sigma_\eta = 0$ (where σ_η denotes standard deviation of η_t).
The second part, $\frac{cov(S_t,e_{t+1})}{var(S_t)}$, vanishes as soon as the market forecast error e_{t+1} inherits conventional rational expectation error properties. Indeed, in this case, e_{t+1} is not correlated with the information set in t, and we have $cov(S_t, e_{t+1}) = 0$. In a sample period where z is always equal to zero, the residual is given by (15) and is correlated with $ex - ante$ information. The econometrician who ignores values of z and applies the standard regression procedures will get a biased result. We will call this bias the peso-problem bias.

In practice it is difficult to evaluate the "term premium" and the peso-problem biases, principally because they are not observable. However, if we assume that only one regime, say regime 0, occurs ex-post, these biases can be distinguished and evaluated. This is what we do in the subsequent paragraph.

4.2 Evaluating biases

We now assume that the only regime that occurs $ex-post$ is regime 0. Consequently, the forecast error that agents make when they predict changes in the short term interest rate, $e_{t+1}(0)$, is given by (15). Hence, using equation (18) and after rearranging expression, we obtain:

$$S_t(0) = \alpha_0 + \alpha_1 \Delta i_t(0) + \eta_t \tag{24}$$

where

$$\alpha_0 = \overline{\Phi} + \frac{\mu_0 + (1 - P_{00})(\mu_1 - \mu_0)}{2} \tag{25}$$

$$\alpha_1 = \frac{\rho_0 + (1 - P_{00})(\rho_1 - \rho_0)}{2} \tag{26}$$

If we could estimate parameters $\overline{\Phi}$, μ_0, μ_1, ρ_0, ρ_1 and P_{00}, we should estimate equation (24) and test the restrictions (25) and (26). This test should provide evidences for or against the EHTS with the presence of a peso-problem. Some authors, Hamilton (1988) and Kugler (1996)for example, propose testing such restrictions in order to test the EHTS in a first-order Markov switching model. Unfortunately we argue that this procedure is not appropriate when one wants to test the EHTS in the presence of a peso-problem. Indeed, if agents expect a regime switch that is not observed on the ex-post sample of observations, estimates of $\overline{\Phi}$, μ_0, μ_1, ρ_0, ρ_1 and P_{00} are very likely to be biased.

In what follows, we show that estimation of equation (24) is nevertheless informative because it allows us to evaluate biases generated by the term premium and the peso-problem, and to test the presence of a peso-problem (next section).

Using equations (15) and (24), we obtain[7]:

$$\frac{cov(S_t(0), e_{t+1}(0))}{var(S_t(0))} = -\frac{(2\alpha_1 - \rho_0)\alpha_1 \sigma_\varepsilon^2}{\alpha_1^2 \sigma_\varepsilon^2 + \sigma_\eta^2 (1 - \rho_0^2)} \tag{27}$$

[7]See appendix for details

and

$$\frac{cov(S_t(0),\eta_t)}{var(S_t(0))} = \frac{\sigma_\eta^2(1-\rho_0^2)}{\alpha_1^2\sigma_\varepsilon^2 + \sigma_\eta^2(1-\rho_0^2)} \tag{28}$$

where $\sigma_\varepsilon^2 = var(\varepsilon_t(0))$.
Therefore, we have:

$$\widehat{b}_1 \simeq 2 - 2\frac{\sigma_\eta^2(1-\rho_0^2)}{\alpha_1^2\sigma_\varepsilon^2 + \sigma_\eta^2(1-\rho_0^2)} - \frac{(2\alpha_1-\rho_0)\alpha_1\sigma_\varepsilon^2}{\alpha_1^2\sigma_\varepsilon^2 + \sigma_\eta^2(1-\rho_0^2)} \tag{29}$$

The second term of the right-hand is the bias generated by the time varying term premium. The last term of the right member is the bias generated by the "Peso effect". This term is zero when $P_{00} = 1$ or $\rho_0 = \rho_1$. In the first case, market participants forecast no regime switching, and then their forecast errors inherit properties of conventional rational expectation errors. In the second case, a regime switching may be expected $ex-post$, but when it occurs, it only modifies the unconditional mean of the process Δi_{t+1}. In this case, market forecast errors are not correlated with the information set in t, but they remain biased.
In the following, we denote TPB and PPB the term premium and peso-problem biases respectively[8]

$$TPB = -2\frac{\sigma_\eta^2(1-\rho_0^2)}{\alpha_1^2\sigma_\varepsilon^2 + \sigma_\eta^2(1-\rho_0^2)}$$

$$PPB = -\frac{(2\alpha_1-\rho_0)\alpha_1\sigma_\varepsilon^2}{\alpha_1^2\sigma_\varepsilon^2 + \sigma_\eta^2(1-\rho_0^2)}$$

Where μ_0, ρ_0, σ_ε^2, α_0, α_1 and σ_η^2 are parameters the of equations :

$$\Delta i_{t+1}(0) = \mu_0 + \rho_0\Delta i_t(0) + \varepsilon_{t+1}(0) \tag{30}$$

and

$$S_t(0) = \alpha_0 + \alpha_1\Delta i_t(0) + \eta_t \tag{31}$$

Therefore, all the information we need to evaluate TPB and PPB (that is μ_0, ρ_0, σ_ε^2, α_0, α_1 and σ_η^2) can be obtained by estimating equations (30) and (31). These equations are estimated on sub-periods for which the EHTS is rejected.
We designate the parameter estimates $\widehat{\mu}_0$, $\widehat{\rho}_0$, $\widehat{\sigma}_\varepsilon^2$, $\widehat{\alpha}_0$, $\widehat{\alpha}_1$ and $\widehat{\sigma}_\eta^2$. Of course

[8]Under the same hypotheses, we can calculate the expression of the Student statistics for the hypotheses $H_0 : b_1 = 0$ and $H_0 : b_1 = 2$. See appendix for details

this approach does not allow us to identify P_{00}, P_{11}, ρ_0, ρ_1, μ_0, μ_1 and $\overline{\Phi}$, but this information is not necessary to simulate the model.

In the following we designate estimates of the "term premium bias" $\widehat{TPB}$, and of the "peso problem bias" $\widehat{PPB}$:

- $\widehat{TPB} = -2\frac{\widehat{\sigma}_\eta^2(1-\widehat{\rho}_0^2)}{\widehat{\alpha}_1^2\widehat{\sigma}_\varepsilon^2+\widehat{\sigma}_\eta^2(1-\widehat{\rho}_0^2)}$
- $\widehat{PPB} = -\frac{(2\widehat{\alpha}_1-\widehat{\rho}_0)\widehat{\alpha}_1\widehat{\sigma}_\varepsilon^2}{\widehat{\alpha}_1^2\widehat{\sigma}_\varepsilon^2+\widehat{\sigma}_\eta^2(1-\widehat{\rho}_0^2)}$

4.3 Testing the possibility of a peso-problem

Our task is to test whether the rejection of EHTS is generated by a time-varying term premium, a peso-problem or both. As noted earlier addressing this issue is not straightforward. First, term premium and forecast errors are not observable. Second, when there is a peso-problem, estimation can produce biased results.
In this paragraph we argue that both difficulties can be circumvented in our model if we assume that only one regime is observed *ex-post*. Hence, we can propose an unbiased procedure that allows us to reject or not reject the possibility of a peso-problem.

In what follows we consider the following hypotheses:

- H_0 : A peso-problem cannot occur.
- H_1 : A peso-problem may occur.

Our aim is to test hypothesis H_0 against hypothesis H_1. In fact, this test can be reduced to a parameter restriction test. A peso problem may occur as soon as the market expects a regime switch. In our model, that means that the transition probability P_{00} is different from one. Conversely, a peso-problem cannot occur when the market expects no regime switching. In this case, the transition probability equals 1. According to equation (26), this implies the following parameter restriction $\alpha_1 = \rho_0/2$. Consequently the previous hypothesis is reduced to:

- H_0 : $\alpha_1 = \rho_0/2$.
- H_1 : $\alpha_1 \neq \rho_0/2$.

Once again, this restriction test can be performed using equations (30) and (31)[9].

5 Empirical results

For each country we consider sub-periods for which the EHTS is rejected and investigate whether this rejection may be generated by a peso-problem. For each country, the sub-periods considered are historically identified as periods of stability. Consequently, we can reasonably hope that the hypothesis of no regime switching during the considered sub-period is borne out.Therefore, we can apply the test procedure proposed in the previous section. In other words we test the hypothesis $H_0 : \alpha_1 = \rho_0/2$. Rejection of the H_0 hypothesis means that agents expect a regime switching. In this case a peso-problem is possible. If H_0 is not rejected, agents expect no regime switching during the observation period. Consequently, there is no peso-problem.
We also report estimates of the "term premium bias" ($\widehat{TPB}$), and, if any, the "peso problem bias" ($\widehat{PPB}$). Results are reported in tables 6, 7 and 8.

In Germany and the United Kingdom, the period following the European exchange rate crisis of 1992 is associated with market expectation of a regime switch. As indicated in tables 9 and 10 the H_0 hypothesis is rejected (at the 5% level) after 1992. A consequence of this result is that a peso-problem is possible during this period. This phenomenon could explain in part the rejection of the EHTS experienced for this period. Evaluations of TPB and PPB give more information about contributions of a time-varying term premium and a peso-problem to the rejection of the EHTS hypothesis. In both countries, PPB is dominated by TPB, indicating that distorsions generated by a time-varying term premium are more important than those generated by a peso-problem.
In the United States (see table 11), the peso-problem hypothesis is always rejected. Indeed, the H_0 hypothesis is not rejected, indicating that we can reasonably consider (at the 5% level) that the market expect no regime switch during the periods in question. Consequently, the rejection of the EHTS is only due to a time-varying term premium.

[insert Table 9 about here]

[9] See appendix

[insert Table 10 about here]

[insert Table 11 about here]

6 Conclusion

In this paper, we attempt to explain the empirical rejection of the Expectation Hypothesis of the term structure for three-month and six-month Eurorates for Germany, the United Kingdom and the United States. To do so, we estimate a model that incorporate a time-varying term premium and a peso problem. A restriction test is performed in order to test the possibility of a peso-problem during the observation period. Furthermore, we estimate and compare biases generated by a time varying term premium and a peso-problem.

Our findings are as follows. On the first hand we show that the immediate period following the European exchange rate crisis that occurred in 1992 is associated with the market expectation of a regime switch in Germany and the United Kingdom. We therefore argue that a peso-problem may have occurred in these countries during this period. This phenomenon could explain in part the EHTS rejection experienced in Germany and United Kingdom after 1992. On one hand, we cannot reject the "no peso problem" hypothesis in the United States. We therefore conclude that the rejection of the EHTS in the United States is only generated by a time-varying term premium.

A Data

The short-term interest rate is the 3-month Euro-rate, and the long -term interest rate is the 6-month Euro-rate.
We use monthly data from the Bank for International Settlements.

	Germany, 1970-01 to 1998-12	
	mean	standard deviation
3-month	5.90	2.57
6-month	6.01	2.47
	the United Kingdom, 1970-01 to 1998-12	
	mean	standard deviation
3-month	10.09	3.56
6-month	10.13	3.43
	the United States, 1970-01 to 2000-10	
	mean	standard deviation
3-month	7.49	3.15
6-month	7.64	3.11

Table 1: Euro-Rates: descriptive statistics

B peso-problem test

A combination of equations (30) and (31) leads to:

$$S_t(0) - \frac{1}{2}\Delta i_{t+1}(0) = \alpha_0 - \frac{1}{2}\mu_0 + (\alpha_1 - \frac{1}{2}\rho_0)\Delta i_t(0) + \eta_t - \frac{1}{2}\varepsilon_{t+1}(0) \quad (32)$$

Consider the regression:

$$S_t(0) - \frac{1}{2}\Delta i_{t+1}(0) = \gamma_0 + \gamma_1 \Delta i_t(0) + v_{t+1} \quad (33)$$

Hence, we test $H_0 : \gamma_1 = 0$ ($\Leftrightarrow H_0 : \alpha_1 = \rho_0/2$)

Germany					
	$\widehat{\rho}_0$	$\widehat{\alpha}_1$	$\widehat{\gamma}_1$	$\widehat{t}_{\widehat{\gamma}_1=0}$ (*P-value*)	H_0
1984-01 to 1989-12	0.06	0.04	0.02	0.38 (*0.70*)	NOT REJECTED
1990-01 to 1992-09	0.01	0.07	-0.05	-0.43 (*0.66*)	NOT REJECTED
1992-10 to 1998-12	0.43	0.09	0.09	2.21 (*0.03*)	REJECTED

Table 2: "Peso problem" test - Germany

the United Kingdom					
	$\widehat{\rho}_0$	$\widehat{\alpha}_1$	$\widehat{\gamma}_1$	$\widehat{t}_{\widehat{\gamma}_1=0}$ (*P-value*)	H_0
1992-10 to 1998-12	0.24	0.32	0.14	3.39 (*0.00*)	REJECTED

Table 3: "Peso problem" test - the United Kingdom

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

the United States					
	$\widehat{\rho}_0$	$\widehat{\alpha}_1$	$\widehat{\gamma}_1$	$\widehat{t}_{\widehat{\gamma}_1=0}$ (*P-value*)	H_0
1973-01 to 1979-09	0.10	-0.01	-0.05	-0.67 (*0.50*)	NOT REJECTED
1982-11 to 1987-08	0.01	0.03	0.02	0.33 (*0.74*)	NOT REJECTED

Table 4: "peso problem" test - the United States
* significant at the 10% level,** significant at the 5% level ,*** significant at the 1% level

C Term premium and Peso-problem biases

When we only observe regime 0 $ex-post$, changes in the short interest rate evolves according to:

$$\Delta i_{t+1}(0) = \frac{\mu_0}{1-\rho_0} + \frac{1}{1-\rho_0 L}\varepsilon_{t+1}(0) \tag{34}$$

Hence, using (34) and (18) we have:

$$S_t(0) = \alpha_0 + \frac{\alpha_1\mu_0}{1-\rho_0} + \frac{\alpha_1}{1-\rho_0 L}\varepsilon_t(0) + \eta_t \tag{35}$$

Given equations (15) and (25) we have:

$$\begin{aligned} cov(S_t(0), e_{t+1}(0)) &= -(1-P_{00})(\rho_1-\rho_0)\alpha_1 var(\Delta i_t(0)) \\ &= -(1-P_{00})(\rho_1-\rho_0)\alpha_1\frac{\sigma_\varepsilon^2}{1-\rho_0^2} \end{aligned} \tag{36}$$

Using (25) and (26), we have:

$$-(1-P_{00})(\mu_1-\mu_0) = -(2\alpha_0 - 2\overline{\Phi} - \mu_0) \tag{37}$$

and

$$-(1-P_{00})(\rho_1-\rho_0) = -(2\alpha_1-\rho_0) \tag{38}$$

Hence, we can rewrite $cov(S_t(0), e_{t+1}(0))$ as:

$$cov(e_{t+1}(0), S_t(0)) = -\alpha_1(2\alpha_1-\rho_0)\sigma_\varepsilon^2\frac{1}{1-\rho_0^2}$$

Furthermore, from (35), we have:

$$cov(S_t(0), \eta_t) = \sigma_\eta^2 \tag{39}$$

and

$$var(S_t(0)) = \frac{\alpha_1^2 \sigma_\varepsilon^2}{1-\rho_0^2} + \sigma_\eta^2$$

We can deduce b_1:

$$b_1 = 2 - 2\frac{\sigma_\eta^2(1-\rho_0^2)}{\alpha_1^2\sigma_\varepsilon^2 + \sigma_\eta^2(1-\rho_0^2)} - \frac{(2\alpha_1 - \rho_0)\alpha_1\sigma_\varepsilon^2}{\alpha_1^2\sigma_\varepsilon^2 + \sigma_\eta^2(1-\rho_0^2)} \tag{40}$$

Student statistics are given by expressions:

$$t_{b_1=0} = \widehat{b}_1\sqrt{\frac{Tvar(S_t(0))}{var(u_t)}}$$

$$t_{b_1=2} = (\widehat{b}_1 - 2)\sqrt{\frac{Tvar(S_t(0))}{var(u_t)}}$$

where $u_t = e_{t+1}(0) - 2\eta_t$.

$$\begin{aligned} e_{t+1}(0) &= -(1-P_{00})(\mu_1-\mu_0) - (1-P_{00})(\rho_1-\rho_0)\frac{\mu_0}{1-\rho_0} + \left(1 - \frac{(1-P_{00})(\rho_1-\rho_0)L}{1-\rho_0 L}\right)\varepsilon_{t+1}(0) \\ &= -(2\alpha_0 - 2\overline{\Phi} - \mu_0) - (2\alpha_1 - \rho_0)\frac{\mu_0}{1-\rho_0} + \left(1 - \frac{(2\alpha_1-\rho_0)L}{1-\rho_0 L}\right)\varepsilon_{t+1} \end{aligned} \tag{41}$$

Using (35) and (41) we have:

$$t_{b_1=0} = \widehat{b}_1\sqrt{\frac{T(\alpha_1^2\sigma_\varepsilon^2 + (1-\rho_0^2)\sigma_\eta^2}{\sigma_\varepsilon^2(1-\rho_0^2 + (2\alpha_1-\rho_0)^2) + 4(1-\rho_0^2)\sigma_\eta^2}} \tag{42}$$

and

$$t_{b_1=2} = (\widehat{b}_1 - 2)\sqrt{\frac{T(\alpha_1^2\sigma_\varepsilon^2 + (1-\rho_0^2)\sigma_\eta^2}{\sigma_\varepsilon^2(1-\rho_0^2 + (2\alpha_1-\rho_0)^2) + 4(1-\rho_0^2)\sigma_\eta^2}} \tag{43}$$

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Germany	
1973-01 to 1978-12	Period immediately after the end of the Bretton Woods system and the first major oil shock
1979-01 to 1983-12	Period following the second oil shock
1984-01 to 1989-12	Era of stagnation and late recovery in West Germany
1990-01 to 1992-09	Early years of reunification
1992-10 to 1998-12	Period following the exchange rate crisis in Europe
United Kingdom	
1973-01 to 1978-12	Immediate period after the end of the Bretton Woods system and the first major oil shock
1979-01 to 1992-09	Period following the second oil shock
1992-10 to 1998-12	Period following the exchange rate crisis in Europe
United States	
1973-01 to 1979-09	Period immediately after the end of the Bretton Woods system. Period of federal fund rate operating procedure and use of monetary aggregate as intermediate target.
1979-10 to 1982-10	Appointment of Paul Volcker as Fed chairman Period of non-borrowed-reserves operating procedure known as “the Volcker’s experience”
1982-11 to 1987-08	End of the “Volcker experiment” beginning of a borrowed-reserves operating procedure and return to a policy of smoothing interest rates
1987-09 to 1998-12	Alan Greenspan’s appointment as Fed chairman Period following the 1987 stock market crash

Table 5:

Germany				
	$\widehat{b}_1$	$\widehat{t}_{\widehat{b}_1=0}$ (*P-value*)	$\widehat{t}_{\widehat{b}_1=2}$ (*P-value*)	EHTS
1973-01 to 1978-12	1.82***	3.35	-0.32	NOT REJECTED
		(*0.00*)	(*0.74*)	
1979-01 to 1983-12	1.30**	2.50	-1.34	NOT REJECTED
		(*0.01*)	(*0.18*)	
1984-01 to 1989-12	0.59	0.81	-1.94	REJECTED
		(*0.41*)	(*0.05*)	
1990-01 to 1992-09	0.36	0.68	-3.08	REJECTED
		(*0.49*)	(*0.00*)	
1992-10 to 1998-12	1.29***	5.18	-2.81	REJECTED
		(*0.00*)	(*0.00*)	

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

Table 6: EHTS test - Germany

United Kingdom				
	$\widehat{b}_1$	$\widehat{t}_{\widehat{b}_1=0}$ (P-*value*)	$\widehat{t}_{\widehat{b}_1=2}$ (P-*value*)	EHTS
1973-01 to 1978-12	1.75***	2.98	-0.41	NOT REJECTED
		(*0.00*)	(*0.68*)	
1979-01 to 1992-09	1.35***	2.93	-1.41	NOT REJECTED
		(*0.00*)	(*0.16*)	
1992-10 to 1998-12	1.06***	7.74	-2.17	REJECTED
		(*0.00*)	(*0.03*)	

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

Table 7: EHTS test - the United Kingdom

United States				
	$\widehat{b}_1$	$t_{\widehat{b}_1=0}$ (*P-value*)	$t_{\widehat{b}_1=2}$ (*P-value*)	EHTS
1973-01 to 1979-09	0.09	0.10	-2.20	REJECTED
		(*0.91*)	(*0.03*)	
1979-10 to 1982-10	1.53*	1.67	-0.48	NOT REJECTED
		(*0.09*)	(*0.63*)	
1982-11 to 1987-08	-0.70	-0.90	-3.45	REJECTED
		(*0.37*)	(*0.00*)	
1987-09 to 1998-12	1.34***	2.74	-1.32	NOT REJECTED
		(*0.00*)	(*0.18*)	

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

Table 8: EHTS test - the United States

Germany			
	H_0	TPB	PPB
1984-01 to 1989-12	NOT REJECTED	-1.85	0
	no peso-problem	[-1.99 -1.56]	
1990-01 to 1992-09	NOT REJECTED	-1.93	0
	no peso-problem	[-1.98 -1.29]	
1992-10 to 1998-12	REJECTED	-1.06	-0.31
	possibility of a peso-problem	[-1.59 -0.61]	[-0.69 0.23]

Table 9: peso problem test - Germany

the United Kingdom			
	H_0	TPB	PPB
1992-10 to 1998-12	REJECTED	-1.02	-0.62
	possibility of a peso-problem	[-1.42 -0.64]	[-0.99 -0.06]

Table 10: peso problem test - United Kingdom

the United States			
	H_0	TPB	PPB
1973-01 to 1979-09	NOT REJECTED	-1.96	0
	no peso-problem	[-1.98 -1.78]	
1982-11 to 1987-08	NOT REJECTED	-1.84	0
	no peso-problem	[-1.99 -1.50]	

Table 11: peso problem test - United States

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