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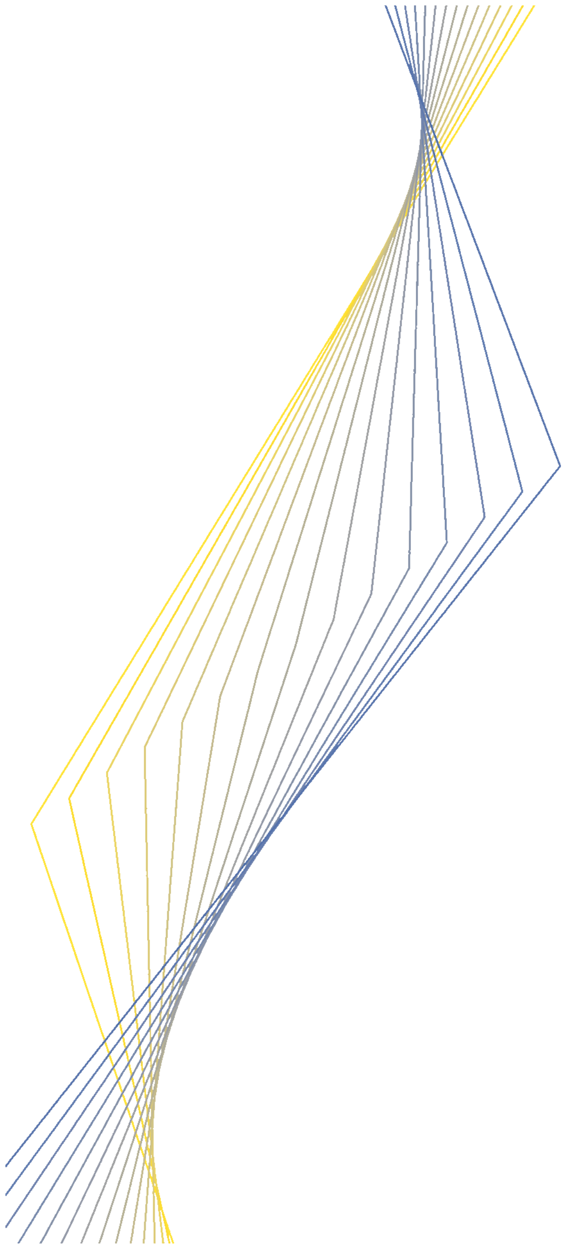

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ESTIMATING RISK PREMIA IN MONEY MARKET RATES[1]

BY ALAIN DURRÉ[2], SNORRE EVJEN[3] AND RASMUS PILEGAARD[4]

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Contents

Abstract

This paper empirically tests the expectations hypothesis on both daily EONIA swap rates and monthly EURIBOR rates extended backwards with German LIBOR rates. In addition, we quantify the size of the risk premia in the money market at maturities of one, three, six and nine months. Using implied forward and spot rates in a cointegrated VAR model, we find that the data support the expectations hypothesis in the euro area and in Germany prior to 1999. We find that risk premia are relatively limited at the shorter maturities but more significant at maturities of six and nine months. Furthermore, the results on LIBOR/EURIBOR rates tentatively indicate a downward shift in the structure of the risk premia after the introduction of the euro.

JEL classification: E43,C32

Keywords: Term structure of interest rates; Expectations hypothesis; Cointegrated VAR models

Non-technical summary

The aim of this paper is to estimate the size of the risk premia embedded in the implied forward interest rates derived from euro area money market interest rates. For this purpose, we estimate a cointegrating vector autoregressive model (CIVAR) using monthly and daily data for money market interest rates and implied forward rates.

In order to find estimates of risk premia, we first test whether the expectations hypothesis theory (EH) of the term structure of interest rates holds true. On German LIBOR data from 1989 until 1998 with monthly frequency we find that a "weak version" of this theory cannot be rejected on a horizon from one up to nine months. The weak version of the EH incorporates a constant term premium which increases with maturity. These estimations suggest risk premia of around 5 basis points at the one-month maturity. The indicated risk premia increase to around 10 basis points at the three-month maturity, reaching 15 and 27 basis points at the six- and nine-month maturities respectively.

When estimating on a longer sample period, incorporating the period from January 1999 until late 2001 using EURIBOR rates, we no longer find support to the expectations hypothesis. We have two potential explanations for this. On the one hand, the rejection of the EH may stem from the existence of time-varying risk premia following the launch of the European single currency. On the other hand, the results could be explained by a shift in the level of risk premia at different horizons.

We propose to test the latter potential explanation mainly for two reasons. Firstly, according to the convergence of the different national monetary policies towards the German approach, it would be surprising that time-varying risk premia should appear with the launch of the European single currency, given the indications that the risk premia in German money markets before this period were time-independent. Secondly, according to the results the rejection of the EH seems to be due to only a minor change in the estimates.

Therefore, we tested for the existence of a structural change in the level of the constant risk premium. We found that this change probably appeared in January 1999. By re-estimating the models for the whole sample period, we found that the introduction of EMU has entailed a decrease of risk premia of around 2, 5, 9 and 14 basis points respectively, at the 1-, 3-, 6- and 9-month horizons compared to the estimated risk premia in German data.

We also tested the expectations hypothesis (using a cointegrating VAR-model) on daily data for the euro area, using EONIA swap rates. The results seem to support a weak version of the expectations hypothesis as well, i.e. with a constant term premium. Using two different samples, the estimated risk premia are in ranges of 0-1, 2, 4-6 and 10-13 basis points at the horizons of one, three, six and nine months respectively. Although the estimated risk premia for the EONIA market should be interpreted with caution given the limited sample period, they are fairly close to the results obtained from the analysis on German LIBOR/EURIBOR data. However, one should keep in mind that the results from the analyses using EONIA swap rates and LIBOR/EURIBOR rates are not 100 per cent comparable as the credit risks in EONIA swap rates are lower than in LIBOR/EURIBOR rates.

The empirical evidence from other studies is mixed and few studies give clear indications about the size of risk premia in money market rates. However, our results seem to be broadly in line with other comparable studies on European data.

1. Introduction

Short-term interest rates contain information about market participants' expectations about the stance of monetary policy in the near future. An assessment of these expectations can prove useful in many ways. Firstly, knowledge of such expectations helps the central banks to predict whether a particular policy decision is likely to surprise market participants, and what their short-term response is likely to be to a given decision. Secondly, measures of interest rate expectations can also be useful to evaluate the central banks' communication with financial markets (Goodfriend, 1998 and Fisher, 2001) and, ex post, to assess whether monetary policy was predictable (Rudebusch, 1995, 1998 and Taylor, 2001).

If there were no uncertainty about the path of future interest rates, forward rates would equal expected future interest rates. As future interest rates are not known with certainty, risk averse investors will then require a risk premium to bear this interest risk. Hence, the existence of risk premia, arising from interest rate risk and investor risk aversion, implies that there will be a wedge between implied forward rates and the expected future interest rate which normally increases with time. Because the sizes of such risk premia are unobservable, the "true" market expectations about future interest rates are not known with certainty.

Forward rates in the euro area can be derived from the term structures of both EONIA swap rates and EURIBOR rates.[1] EURIBOR rates are offered rates on unsecured loans in the interbank market. Consequently, the forward rates derived from such rates will also include a credit risk premium. The existence of credit risk premia widens the wedge between the implied forward interest rates and the expected rates. Just as for term premia, credit risk considerations are likely to increase with maturity. Credit risk in EONIA swaps is more limited than in EURIBOR since the swaps do not involve an exchange of principal amounts. The breakdown of risk premia into different categories of risk is beyond the scope of this analysis, which aims at providing a quantification of the overall risk premia in money market rates.

In the economic literature, two distinct approaches have been developed for extracting information about interest rates expectations and risk premia from the term structure of the yield curve. The expectations hypothesis (EH) theory literature focuses on the time-series properties of interest rates to analyse the relationship between short and long-term rates.[2] The second approach is a market-based "no-arbitrage" approach, which tries to identify a factor structure affecting the shape of the term structure. This latter approach assumes stationary stochastic processes for economic fundamentals driving interest rates

[1] There exist many other financial instruments from which market expectations can be extracted, for example FRA rates (forward rate agreements) and EURIBOR futures, see the article "The information content of interest rates and their derivatives for monetary policy" page 37-55 in the ECB Monthly Bulletin of May 2000. See also Svensson and Soderlind (1997) for a detailed description of techniques to extract market expectations from financial instruments and for instance Carlson et al. (1995) for a description of American financial instruments used for this purpose.

dynamics, and fundamentals are represented as factors determining the decomposition of interest rates into expectations and risk premia.[3]

Following the first approach, the spreads between interest rates at different maturities, or the implied forward rates extracted from the yield curve, reflect the path of future short-term rates.

To give a sneak preview of our results, the expectations hypothesis was tested in both a single-equation framework (using the Phillips-Hansen methodology) and a cointegrating vector autoregressive model (CIVAR). The analysis using a CIVAR-model indicates that the "pure" version of the EH (with no risk premium) does not hold true in the German-LIBOR market for most of the maturities up to 9 months. However, the "weak" version of the EH (i.e. with a risk premium constant across time) is supported by data. The risk premia are found to be increasing with the maturity of the forward rates. Table 1 below provides a summary of measures of risk premia indicated from this analysis and other studies.

Moreover, adding EURIBOR data from January 1999 to the German LIBOR data implicated that the EH, using the Phillips-Hansen methodology, was slightly rejected at maturities of three months and beyond. Among potential explanations, this paper explicitly test the hypothesis of a level shift in the constant risk premia according to the Gregory and Hansen (1996) methodology. The results suggest that the constant risk premia decreased for all maturities after the launch of the European single currency in 1999. More specifically, the risk premia amounted at 2, 5, 9 and 14 basis points respectively for the one-, three-, six- and nine-months for the whole sample (see Table 1).

Our results are broadly in line with other comparable studies on data for European countries (e.g. Gerlach et.al., 1997, Cassola et.al., 2001, and Cuthbertson et. al., 2000, Brooke et. al., 2000). However, most studies focus on whether the expectations hypothesis holds true. The number of other studies quantifying the size of risk premia in interest rates is relatively limited. Cassola et.al. (2001) find evidence that the term premium in German interest rates from 1972 to 1998 is around zero basis points for the one-month maturity, increasing to around 40-50 basis points at the twelve-month maturity. Brooke et. al. (2000) estimate the size of term premia by comparing implied two-week interbank forward rates derived from the UK money market instruments with actual outturns of the two-week repo rate. The average biases suggested that term premia embodied in interbank forward rates are significant beyond a six-month horizon. They tentatively indicate that the average biases over the period from January 1993 to September 2002 for six-month, one-year and two-year maturities were 23, 45 and 109 basis points respectively. They

[2] Among others, see Shiller et al., 1983, and Campbell and Shiller, 1991.

[3] For more details, see for instance Balduzzi, Das, Foresi and Sundaram, 1996.

believe that credit risk considerations may account for 20-25 basis points, on average for these maturities.[4]

The note is organised as follows. Section 2 presents briefly the data. In Section 3, we present some other empirical evidence. In Section 4 we present our empirical results. Section 5 concludes.

Table 1. Indicated risk premia at different maturities (basis points)

Months Ahead	1 month	3 months	6 months	9 months
ESTIMATIONS:				
Monthly data (German LIBOR /EURIBOR):				
1989.12-1998.12 (Germany)	5	10	15	27
1989.12-2001.8: (Germany and euro area, pre terrorist attacks)	5	n.a[1)]	n.a[1)]	n.a[1)]
1989.12-2002.2: (Germany and euro area)	5	n.a[1)]	n.a[1)]	n.a[1)]
Daily data (EONIA SWAP rates):				
Jan. 1999-Sept. 2001:	0	2	6	10
Jan. 1999-June 2002:	1	2	4	13
OTHER STUDIES:				
Cassola and Luis (Germany, 1972-1998)[2)]	0-5	5-10	20-25	25-35
Brooke et. al (UK, 1993-2000)[3]			23	
Gravelle et.al. (Canada, 1988-1998)	6	20	58	100

1) Not available because "expectations hypothesis" is formally rejected

2) Approximate values, see ECB working paper No. 46, page 51.

3) Risk premia embedded in interbank implied forward rates compared with the two-week official monetary policy repo rate. This study finds that the risk premia are 45 and 109 basis points for the one-year and the two-year maturities respectively

[4] The credit risk spread stems from the fact that they use interbank forward rates, which contain credit risks, and compare with the outturns of the official repo rate of the Bank of England, which is given on loans with collateral, thus almost free of credit risks.

2. Data

Two different data sets are used in our estimations: A daily data set of EONIA swap rates from beginning 1999 to 5 June 2002 and a data set of monthly averages of EURIBOR rates extended backwards with German LIBOR rates, covering the period between December 1989 and February 2002. EONIA swap rates are superior to the EURIBOR in providing information about market expectations. However, before the introduction of the euro money market the national swap markets were not liquid enough to provide as reliable information as could be derived from the LIBOR markets.

2.1 Daily data - EONIA swap rates

In an EONIA swap, two parties agree to exchange the difference between the interest accrued at an agreed fixed interest rate for a fixed period on an agreed notional amount and interest accrued on the same amount by compounding EONIA daily over the term of the swap.[5] The 'fixed leg' of this agreement is referred to as the EONIA swap rate. Hence, the EONIA swap rates reflect the expected average EONIA rate over the maturity of the swap contract. EONIA swap rates are traded for maturities from one to three weeks and one to twelve monthsplus maturities of 15, 18, 21 and 24 months.

Chart 1: ECB interest rates and the EONIA

(annual percentages, daily data)

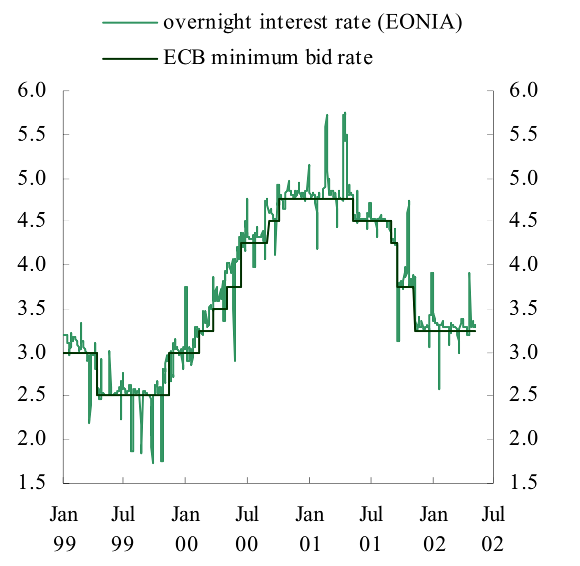
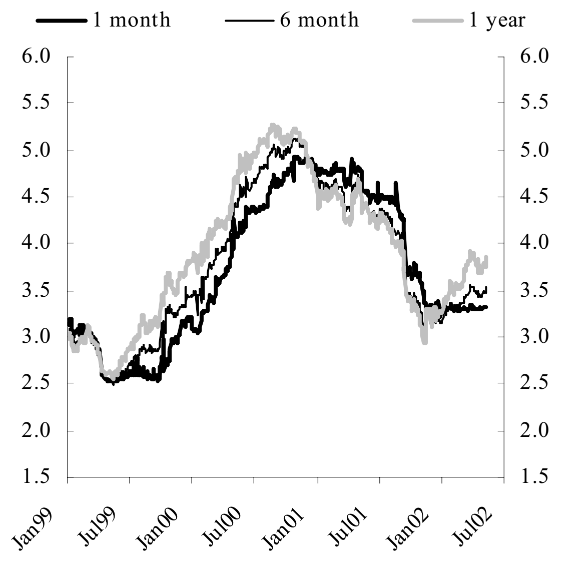


Chart 2: EONIA swap rates

(annual percentages, daily data)



Sources: ECB and Reuters

[5] It is only possible to roll over investments on workdays. Thus, on working days the EONIA swap rates are based on daily compounding, while it is treated as a simple rate over weekends.

The EONIA is computed, by the ECB, as a weighted average of all overnight unsecured lending transactions made in the euro area interbank market by a panel of primary banks. Normally the EONIA is traded close to the ECB minimum bid rate in the main refinancing operations, see Chart 1.

The EONIA sometimes differs markedly from the minimum bid rate. Interest rate expectations[6] or the liquidity conditions can usually explain these discrepancies. Calendar effects (at the last business day of the month, quarter or year) from balance sheet considerations of financial institutions can also affect the liquidity conditions.

In Chart 2 the EONIA swap rates with maturities of one, six and twelve months show the same pattern as the EONIA in Chart 1. When the minimum bid rate is expected to increase, as in most of 2000, the twelve-month swap rates are generally higher than the shorter maturities of one- to six-months due to higher expected averages of the future EONIA.

The nature of the swap arrangement also limits the credit risk since no principle amounts are exchanged.[7] The swap rates therefore provide a more exact indication of the term premium the market adds due to increasing uncertainty for future interest rates.

Since early 1999 the swaps linked to the EONIA have replaced swaps linked to the EURIBOR as the main reference swap rate in the euro money market.[8] Thus, in the regression analysis for the sample period from 1 January 1999 onwards we use implied forward rates computed from the term structure of the EONIA swaps.[9]

Charts 3 and 4 display the implied one-month forward rate in one and six months and the one-month spot rate with a one-month and six-month lead respectively. The predictions by the forward rates are quite accurate on the one-month horizon whereas the difference between actual and predicted rates is somewhat

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[6] If market participants expect changes in the ECB rates within the reserve maintenance period the liquidity demand is affected in the direction of the expectations. For example, expectations of an increase in interest rates in the reserve maintenance period can lead to higher liquidity demand and drive up the EONIA as market participants expects the cost of lending to be higher in the following period.

[7] The credit risk embedded in EONIA swap rates reflect therefore almost uniquely the credit risk premium on overnight transactions as included in the EONIA.

[8] For further information see Santillan et.al. (2000).

[9] Implied forward rates can be computed using the following equation:

$$f_t^{\mathrm{i,j}} = \left[\left(\frac{1 + r_{i+j,t} * \frac{m_{i+j}}{1200}}{1 + r_{i,t} * \frac{m_i}{1200}} \right) - 1 \right] * \frac{1200}{m_{i+j} - m_i}$$

Where $f_t^{\mathrm{i,j}}$ represents the implied forward rate at time t for the i-month interest rate in j months. The rates $r_{i+j,t}$ and $r_{i,t}$ represents the spot rates with maturity i and i+j at time t, respectively. m refers to the maturity of the EONIA swap.

larger on the six-month horizon, as actual out turns normally are harder to predict the further ahead one looks.

Chart 3: Implied one-month forward rate in one month and one-month EONIA swap rates (with a one-month lead)

(annual percentages, daily)

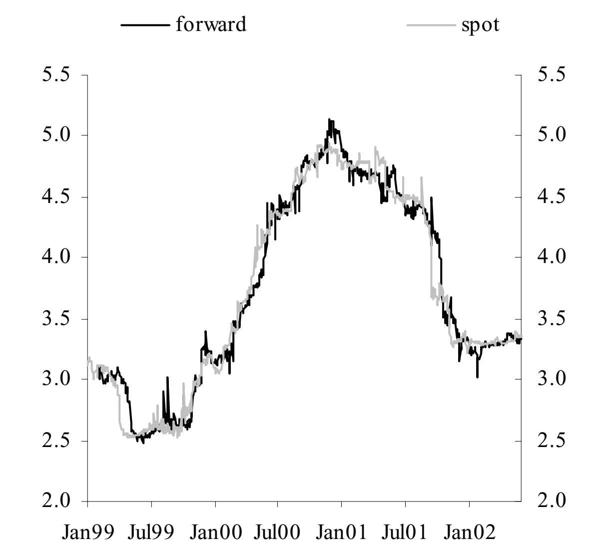
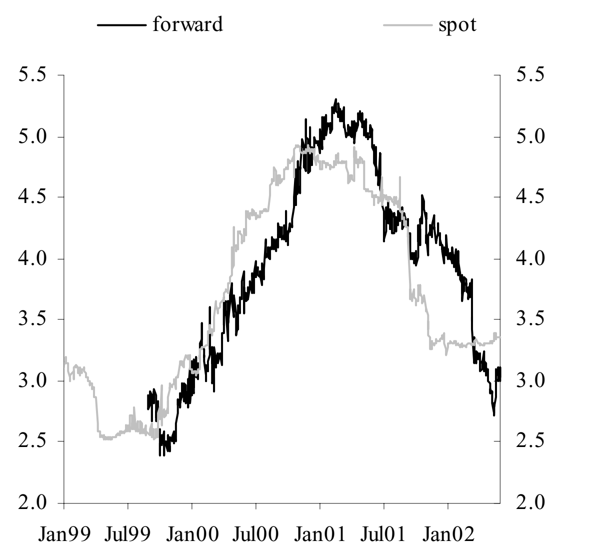


Chart 4: Implied one-month forward rate in six months and one-month EONIA swap rates (with a six-month lead)

(annual percentages, daily)



Sources: Reuters and own calculations.

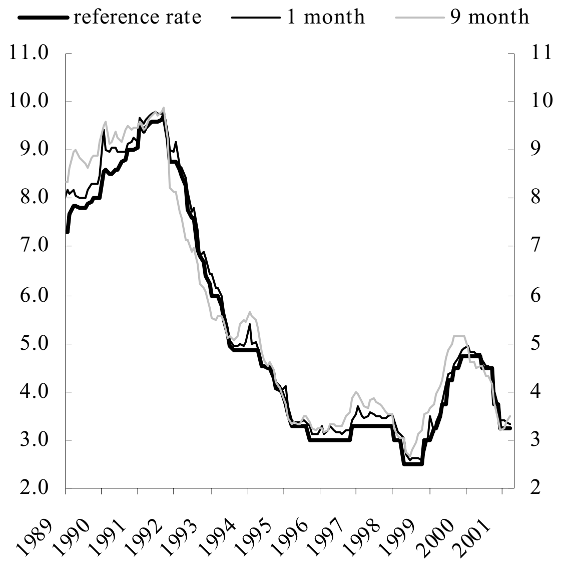
2.2 Monthly data - GERMAN LIBOR and EURIBOR

The LIBOR and the EURIBOR rates are weighted average interbank offered rates and both are regarded to be highly representative benchmark rates respectively for Germany until December 1998 and the euro area thereafter. LIBOR and EURIBOR are interest rates on exchange of uncollateralised short-term liquidity quoted by primary banks. The unsecured nature of these loans implies that some premium for credit risk is likely to be embedded in the rates.

We used monthly averages for German LIBOR rates from December 1989 until December 1998 and EURIBOR rates from 1999 until February 2002. Chart 5 below illustrates the developments of these monthly data for the one- and nine-month maturities together with the Bundesbank's official tender rate spliced with the ECB fixed rate/minimum bid rate from 1999. The LIBOR/EURIBOR rates follow the refinancing rate set by the central bank closely.

Chart 5: Central bank reference rate and interbank offered rate

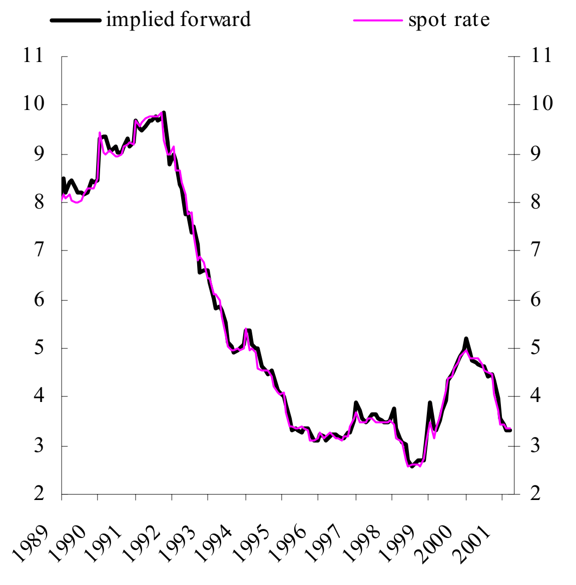
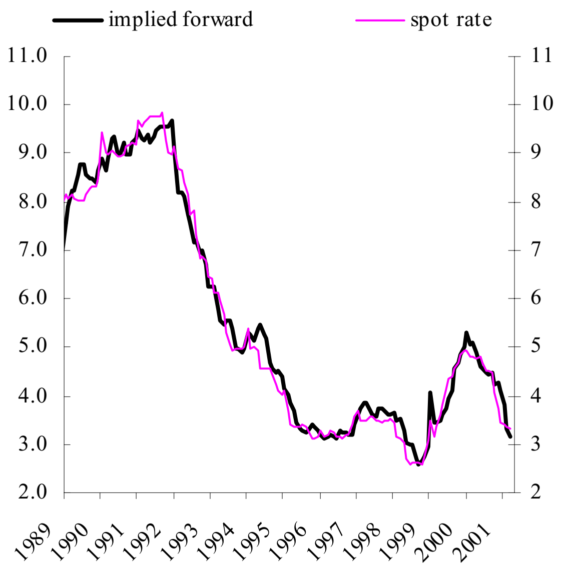
(annual percentages, monthly averages)



Sources: Bundesbank, ECB, Reuters and Bloomberg
Note: From December 1989 to December 1998 Bundesbank tender rate and German LIBOR rates are used. Since 1999 the ECB minimum bid rate and EURIBOR rates are used.

As for the EONIA swap rates, market expectations are extracted using implied forward rates based on the LIBOR/EURIBOR money market yield curve. The one-month implied forward rate in one and three months are presented in Charts 6 and 7 together with the actual one-month spot rate with one and three months lead. "Predicted" and actual outturns of the interest rates are relatively close for most of the period. The main exceptions are at the beginning of the sample.[10]

[10] The average historical difference between the one-, three-, six- and nine-month German LIBOR/EURIBOR rates and the corresponding implied forward rates (i.e. with a lead of one, three, six and nine months) were 4, 7, 5 and 12 basis points respectively in the period between December 1989 to February 2002.

Chart 6: Implied one-month forward rate in one-month and one-month LIBOR/EURIBOR rates (with a one-month lead) *(annual percentages, monthly)*	**Chart 7: Implied one-month forward rate in three months and one-month LIBOR / EURIBOR rates (with a three-month lead)** *(annual percentages, monthly)*
	

Sources: Bloomberg, Reuters and own calculations.

3. Empirical evidence from other studies

In general, most studies test the EH theory by estimating the relationship between actual outturns of interest rates and the implied forward rates derived from the yield curve:[11]

$$(1) \quad R(n)_t = \alpha + \beta f^k(n)_{t-k} + e_t ,$$

where $f^k(n)_{t-k}$ is the forward rate at time t-k of an n period instrument beginning in k periods, $R(n)_t$ is the spot rate at time t of an n period instrument, and e_t is the residual. Here, postulating the expectations hypothesis is the same as assuming that β=1. Thus, if β is significantly different from unity, the EH theory is rejected. The constant term α can be interpreted as a constant risk premium if β is strictly equal to one.[12] Moreover, the 'strong' version of the EH theory holds true only if β=1 and at the same time α=0. The 'weak' version of the EH theory holds true if the estimated β=1 and at the same time $\alpha \neq 0$.[13]

The empirical literature on the term structure of interest rates contains disparate pieces of evidence about the predictive power of the yield curve in the money and the capital market. Most studies seem to agree on the following "stylized facts": First, whereas the strong version of the EH (the spread between maturities or the implied forward rates are unbiased predictors of the future spot rate) is rejected in most cases, the studies, however, suggest that there is an important element of "truth" to the expectations theory of the term structure. Most studies find that the forward rates have a high predictive power of actual outturns of spot interest rates, i.e., when regressing implied forward rates on the actual outturns, the estimated coefficient (the estimated β-value) is typically found to be significantly positive. Secondly, the EH seems to receive more support in the money market (with maturities up to 12 months) than for longer maturities. Thirdly, at the very short end of the yield curve (up to 3 months), the weak version of the EH (with a time-invariant but maturity dependent risk premium) seems to gain support in most studies. Finally, the predictive content of the yield curve appears to be higher in European countries than in the United States.

Brooke et.al. (2000) estimate the size of term premia in UK data by comparing implied two-week interbank forward rates derived from money market instruments with actual outturns of the two-week

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[11] Some studies normally avoid the possibility of spurious regression by running the regression in Equation (1) in first differences but the possibility of integrated data gives the opportunity of constructing information tests based on the existence of a cointegrating relationship between the 'forecast' and outturn.

[12] See for instance Cuthbertson et al. (2000) and Bhundia et al. (1998).

[13] The expectations hypothesis is explained in Shiller (1973), Shiller et al. (1983), Engle (1987), Mishkin (1988), Gravelle et al. (1998) and Jondeau (2001).

repo rate. If term premia are broadly stable and if the sample period is sufficiently long, expectational error should average out to zero. Any remaining bias should then represent the average term premium, though this method also will pick up differences between the money market instruments used and the repo rate that are related to liquidity and credit quality. They find that term premia embodied in interbank forward rates are significant beyond a six-month horizon. They tentatively indicate that the average risk premia over the period from January 1993 to September 2001 for six-month, one-year and two-year maturities were 23, 45 and 109 basis points respectively. However, credit risk considerations may account for 20-25 basis points, on average for these maturities.

Cassola et al. (2001), find that a "two-factor constant volatility"-model on German data quite well describes the yield curve from 1972 to 1998. Their results indicate that term premia are negligible at the maturity of one month but increasing to around 40-50 basis points at the twelve-month maturity.

Using a single-equation model and "imposing" that the EH theory holds true, Gravelle et al. (1998) find on Canadian data that the constant term premium increases with the maturity of the forward rate, going from 6 to 20, and from 58 to 100 basis points when the maturity increases from 1 to 3, and from 6 to 9 months respectively. In addition, they find that the risk premium is 47 basis points at the maturity of nine months in the United States. This study is based on FRA interest rates (forward interest rate agreements) which may include some additional risk premia compared with EONIA swaps and EURIBOR rates, i.e. higher liquidity and credit risk premia.

Studies using Cointegrating regressions

The empirical literature on the term structure of interest rates received a new impulse from the pioneer works of Campbell and Shiller (1987, 1991). They were the first to test the EH by testing for a cointegration relationship between interest rates in a VAR framework, using US money market and capital market data. Their idea was that, if the EH holds strictly true, then the slope of the yield curve does only depend on the future fluctuations of the short rates. A necessary (but not sufficient) condition for the EH to hold true is that, if both the short-term and the long-term interest rates are integrated of the same order, there must exist a cointegrating relationship between these rates. As a second condition, the sum of the coefficients inside the cointegrating vector must be equal to zero, i.e. there must be a long term one-to-one relationship between the interest rates at different maturities (homogeneity of degree 1). The studies by Campbell and Shiller were followed by several studies using this framework to test the validity of the EH theory. Some of them are reported in Table 2. In particular, these studies show that nominal interest rates are characterised as being integrated of order one and that they tend to move together enabling them to be cointegrated.

Campbell and Shiller (1991) did not find the spread between interest rates of maturities up to 12 months significant in explaining the developments of the spot rates, except for at the one-month maturity. And even for the one-month maturity the restriction imposed by the EH theory, i.e. a long-term unit relationship between yield spreads and future changes in interest rates, was rejected. However, using different samples and specifications, Jondeau and Ricart (1999) and Jondeau (2001), both on US data, find that the coefficient for the spread is significantly different from zero but also significantly different from unity, thus rejecting the EH.

The study from Gravelle et al. primarily tests the EH theory on Canadian data in a VAR model. Using such model they reject the EH theory. Their evidence suggests that the formal rejection of the EH theory is due to the existence of a time-varying term premium. The existence of a time-varying risk premium also gains support in some other studies.[14]

For studies on European countries, the picture is mixed. While most studies cannot reject the cointegration hypothesis, Bundhia and Chadha (1998) cannot find evidence that $\beta=1$ for the sterling futures market, i.e the EH theory is rejected. By contrast, the study of Cuthbertson, Hayes and Nitzsche (2000) cannot reject the EH theory for the German money market. Jondeau (2001) cannot reject the expectations theory on French and British data. On the other hand, the study suggests that the EH theory is in most cases rejected on German data.

[14] See for example Campbell and Shiller (1987), Engle and Ng (1993), Iyer (1997) and Jondeau (2001).

Table 2: Results of testing the EH in the literature using cointegrating regressions

Maturity of one-period rate	$\hat{\beta}$	$s.e.(\hat{\beta})$	Sample period	Authors
1 month	0.50*	0.12	1952-87	Campbell&Shiller (1991) on US data
	1.02*	0.01	1984-96	Bhundia&Chadha (1998) on UK data
	0.99*	0.01	1976-93	Cuthbertson, Hayes&Nitzsche (2000) on German data
2 months	0.20	0.28	1952-87	Campbell&Shiller (1991) on US data
	1.73*	0.50	1976-86	Engsted&Tanggaard (1995) on Danish data
	1.04*	0.02	1984-96	Bhundia&Chadha (1998) on UK data
	0.99*	0.01	1976-93	Cuthbertson, Hayes&Nitzsche (2000) on German data
3 months	-0.15	0.20	1952-87	Campbell&Shiller (1991) on US data
	1.59*	0.41	1976-86	Engsted&Tanggaard (1995) on Danish data
	1.05*	0.01	1984-96	Bhundia&Chadha (1998) on UK data
	0.45*	0.16	1975-97	Jondeau&Ricart (1999) on US data
	0.71*	0.15	1975-97	Jondeau&Ricart (1999) on UK data
	0.56*	0.08	1975-97	Jondeau&Ricart (1999) on German data
	0.71*	0.23	1975-97	Jondeau&Ricart (1999) on French data
	1.03*	0.01	1976-93	Cuthbertson, Hayes&Nitzsche (2000) on German data
	0.67*	0.17	1982-97	Jondeau (2001) on US data
	0.72*	0.23	1982-97	Jondeau (2001) on UK data
	0.60*	0.11	1982-97	Jondeau (2001) on German data
	0.74*	0.14	1982-97	Jondeau (2001) on French data
6 months	0.04	0.33	1952-87	Campbell&Shiller (1991) on US data
	1.13*	0.26	1976-86	Engsted&Tanggaard (1995) on Danish data
	1.13*	0.08	1984-96	Bhundia&Chadha (1998) on UK data
	0.12	0.28	1975-97	Jondeau&Ricart (1999) on US data
	0.71*	0.18	1975-97	Jondeau&Ricart (1999) on UK data
	0.63*	0.13	1975-97	Jondeau&Ricart (1999) on German data
	0.82*	0.17	1975-97	Jondeau&Ricart (1999) on French data
	1.03*	0.01	1976-93	Cuthbertson, Hayes&Nitzsche (2000) on German data
	0.40	0.24	1982-97	Jondeau (2001) on US data
	0.64*	0.21	1982-97	Jondeau (2001) on UK data
	0.64*	0.11	1982-97	Jondeau (2001) on German data
	0.69*	0.20	1982-97	Jondeau (2001) on French data
12 months	-0.02	0.37	1952-87	Campbell&Shiller (1991) on US data
	0.84*	0.35	1976-86	Engsted&Tanggaard (1995) on Danish data
	1.22	3.75	1984-96	Bhundia&Chadha (1998) on UK data
	0.55*	0.24	1975-97	Jondeau&Ricart (1999) on US data
	0.77*	0.20	1975-97	Jondeau&Ricart (1999) on UK data
	0.91*	0.21	1975-97	Jondeau&Ricart (1999) on German data
	0.72*	0.12	1975-97	Jondeau&Ricart (1999) on French data
	1.07*	0.03	1976-93	Cuthbertson, Hayes&Nitzsche (2000) on German data
	0.21	0.30	1982-97	Jondeau (2001) on US data
	0.68*	0.22	1982-97	Jondeau (2001) on UK data
	0.68*	0.15	1982-97	Jondeau (2001) on German data
	0.73*	0.21	1982-97	Jondeau (2001) on French data

* denotes the estimates of β's are significantly different from zero. While these studies employ different methodologies for testing the EH, all of them are based on the assumption that the money market interest rates at different maturities are cointegrated.

4. Results from estimating risk premia in a cointegrating vector autoregressive framework

We first test the EH in a framework using cointegration analysis estimating the long run relationship between the spot interest rates and the forward rates using the Phillips-Hansen methodology[15] (in a single-equation regression). Second, we will estimate the dynamic relationship between the spot and forward rates using a cointegrating vector autoregressive model (CIVAR).

While the Phillips-Hansen methodology allows to test explicitly the EH using the long-run equilibrium, the CIVAR methodology takes into account also the short term dynamics and is useful in determining forecasts of the expected spot rate in the future. The former methodology corrects for non-normality (and autocorrelation) in the residuals and is therefore useful as a cross check of the results of the VAR estimations. Moreover, the VAR methodology will be particularly useful for testing the null hypothesis of a constant risk premium. It will also give an estimation of the size of the risk premia given that the EH is not rejected. Both methods take into account the non-stationary feature of the interest rates in levels which may be useful in order to avoid some bias associated with standard single-equation estimations of the expectation hypothesis as suggested by Thornton (2001).[16]

4.1 German LIBOR and EURIBOR rates

The first step in the econometric analysis is to check the existence of cointegration between the spot rates and the implied forward rates. We computed two tests of cointegration, the Engle-Granger (1987) test and the Johansen and Juselius (1990) test.[17] Table 3 summarizes the results from the Johansen tests. The results suggest that for all horizons the null hypothesis of cointegration between spot and forward rates is not rejected. These results are confirmed by the Engle & Granger cointegration tests[18] and we therefore conclude that the one-month spot rate and the one-month implied forward rate with horizons of one to nine months are cointegrated.[19] Subsequently, we tested the EH theory taking into account the non-stationarity of the money market interest rates as suggested by the cointegration tests.

[15] See Phillips et al. (1990).

[16] We present a demonstration of his argument in Appendix 1.

[17] See Appendix 2 for more details about the methods for testing cointegration.

[18] See Appendix 3.

[19] These results are in line with the results for other countries on monthly data (see for example Engsted, 1999, 2000).

Table 3. Johansen test for cointegration (stochastic matrix test). German LIBOR

Horizon	H_0	trace-test	p-value	cointegrating vector restriction test[1] (p-value)
One month	r <= 0	13.349	0.033 *	0.69
$(r^{1m}_{t+1m} - f_t^{1m,1m})$	r <= 1	0.54702	0.526	
Three-month	r <= 0	12.348	0.049 *	0.06
$(r^{1m}_{t+3m} - f_t^{1m,3m})$	r <= 1	0.4831	0.556	
six-month	r <= 0	9.2972	0.153	0.00 **
$(r^{1m}_{t+6m} - f_t^{1m,6m})$	r <= 1	0.3891	0.604	
Nine-month	r <= 0	6.5206	0.377	0.00 **
$(r^{1m}_{t+9m} - f_t^{1m,9m})$	r <= 1	1.4415	0.269	

*Note: A number of 12 lags have been used. The test statistics for H(rank≤p) are listed with p-values based on Doornik and Hendry (2001); ** and * mark significance at 95% and 99%. Testing commences at H(rank=0), and stops at the first insignificant statistics. Sample period is 1989.12-1998.12.*

*[1)] The statistic shown for this test is the p-value of the null hypothesis test that the cointegrating vector β coefficient is equal to 1. ** indicates the rejection of the null hypothesis at a significance level of 1%.*

4.1.1 Estimations using the Phillips-Hansen methodology

By using the Phillips-Hansen methodology, we estimated the cointegrating relationship between forward and spot rates. Three different sample periods were used. Sample A cover the period from December 1989 to December 1998 (i.e. only German data). Sample B covers the period from December 1989 up to the terrorist attacks in September 2001 and sample C covers the period from December 1989 to February 2002. In sample B and C, monthly observations from the EURIBOR market (from January 1999 to February 2002) were spliced with data from the German LIBOR market. This allows us to check whether the observations since the introduction of the euro in January 1999 affect the results. Sample B excludes the period after the terrorist attacks on 11 September 2001 and the following turmoil in financial markets. Market participants could not possibly predict the following substantial decrease in money market interest rates. We therefore suspect that the results for the sample including the observations after these events could be distorted by significant expectational errors.

The Phillips-Hansen estimator (also called the Fully Modified Ordinary Least Squares (FM-OLS) estimators method) is appropriate for estimation and inference when there exists a single cointegrating relationship between a set of *I(1)* variables.[20] This econometric methodology does not follow the standard

[20] The methodology is briefly outlined in Appendix 4.

statistical inference. Broadly speaking, Phillips and Hansen (1990) proposed a method that corrects for both endogeneity in the data and asymptotic bias in the coefficient estimates. As mentioned, this methodology takes into account problems with autocorrelation and non-normality of the residuals.

Table 4: Phillips-Hansen cointegrating regressions: $R(n)_t = \alpha + \beta f^k(n)_{t-k}$

Explanatory Variable	Sample period	Lags of window	α	β	Wald statistic
1 month	A	13	-.0126 (.0603)	.9992 (.0093)	.0057 [.939]
	B	13	-.0326 (.0489)	1.0019 (.0081)	.0574 [.811]
	C	13	-.0383 (.0468)	1.0025 (.0078)	.1050 [.746]
3 month	A	12	-.3090 (.1569)	1.0473 (.0251)	3.546 [.060]
	B	12	-.2861 (.1285)	1.0449 (.0636)	4.2626 [.039]*
	C	12	-.2974 (.1239)	1.0462 (.0211)	4.7471 [.029]*
6 month	A	13	-.5690 (.4020)	1.1144 (.0509)	3.2310 [.072]
	B	13	-.5459 (.3103)	1.1124 (.0523)	4.6017 [.032]*
	C	13	-.5904 (.2996)	1.1181 (.0510)	5.3562 [.021]*
9 month	A	11	-1.1065 (.6094)	1.1901 (.0988)	3.6977 [.054]
	B	11	-.9555 (.4958)	1.1692 (.0851)	3.9545 [.047]*
	C	11	-1.0039 (.4842)	1.1752 (.0837)	4.3767 [.036]*

*Note: Column 3 reports the number of lags used for truncation. Columns 4 and 5 report the value of each coefficient with their standard errors in parentheses while the last column reports the Wald test statistics for the null hypothesis β=1 with the p-value between brackets. * denotes rejection at the 5% significance level.*

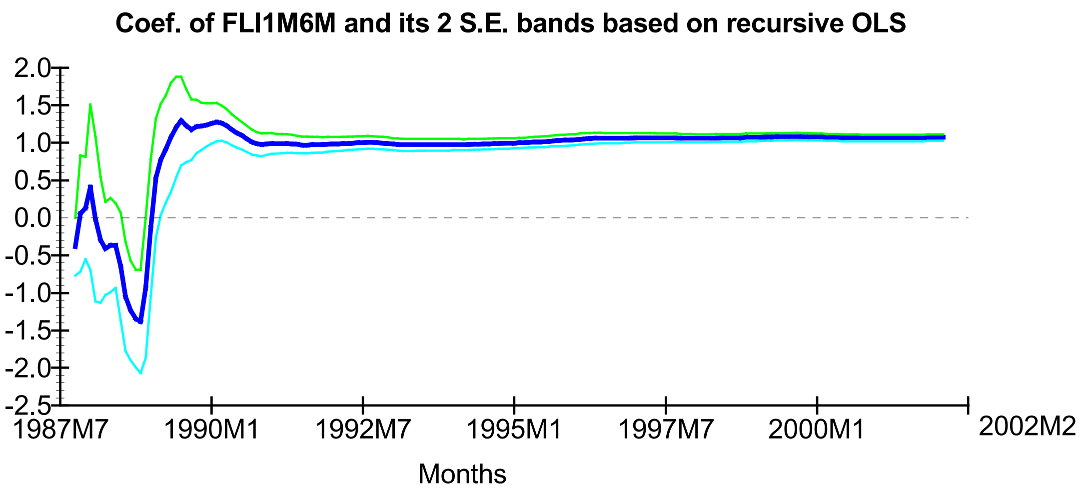
The Phillips-Hansen (1990) fully modified estimators of the cointegrating parameters are shown in Table 4 [21]. The Phillips-Hansen modified estimator allows a valid test (Wald test) that the cointegrating vector is (1,-1). This restriction, on a one-month horizon, is not rejected on any sample investigated. Moreover, on sample A the restriction imposed on the cointegrating vector (1,-1) is not rejected at any maturity. The results on the German sample are in line with the results in Cuthbertson et al. (2000) who find that German money market interest rates "appear to conform reasonably closely to the expectations hypothesis of the term structure with a constant term premium."

However, on sample B and C the restriction tests (Wald statistic) are formally rejected at a 5% significance level for the three-, six- and nine-month horizons. Strictly interpreted, this means that the expectations hypothesis is formally rejected when adding EURIBOR data to the German LIBOR dataset at maturities beyond one-month. Second, the value of the β- coefficient is increasing with maturity. Moreover, the p-value of the Wald test is smaller on sample B and C than on sample A, which may be interpreted as a change in the long-run relationship.

These results suggest that, at the very short end of the yield curve in the money market (i.e. at the one-month horizon), there is not much difference between the samples as the expectations hypothesis seems to hold true on all of them. From the three-month horizon and onwards, the hypothesis is rejected on the extended samples B and C, although only slightly. These results may suggest that "something" changes after January 1999 (see below). A possible explanation could be a different structure of risk premia before and after January 1999. Finally, the rejection of the test does not seem to be only explained by the disturbances from the "unexpected" decreases of interest rates following the terrorist attacks, as the EH is rejected both on sample B (which ends in August 2001) and C (which ends in February 2002).

The recursively estimated coefficient for the forward rates (β-coefficient) for the six-month maturity is shown in Chart 8. In Appendix 5 similar charts are displayed also for the other maturities investigated. The conclusions are the same for all maturities: The coefficients are broadly stable.

Chart 8: Recursive estimation of β using the Phillips-Hansen methodology



[21] Note that they use for that the results from Newey and West (1987) which is also the case with the autocorrelation- and

4.1.2 The Cointegrating Vector Autoregressive Estimations (CIVAR)

The VAR framework allows evaluating the potential impact of short-term dynamics on the long-term relationship. Given the inherent problem of overlapping observations, the chosen number of lags in the CIVAR was primarily selected to eliminate serial correlation and secondly, the lag structure was selected to minimise the AIC and HQ information criteria[22]. The VAR approach also allows to test the empirical content in the EH theory. A further description of the VAR methodology is presented in Appendix 6.

We estimated the two-equation CIVAR for three sample periods: First, we estimated only for the German period (i.e. from December 1989 to December 1998). Second, we re-estimated the same model for a sample period until September 2001 and a period until February 2002. Regarding the sample period ending in December 1998, we first estimated a cointegration relationship between the one-month money market interest rate and the one-month implied forward rate at different horizons with an unrestricted intercept as suggested by Johansen (1990). Inside this framework, we tested the null hypothesis of $\beta=1$ which was rejected at the six and nine-month horizon at a 5% significance level, but not rejected for the one- and three-month horizons (see Table 3).

However, the estimated cointegrating coefficient for the implied forward rates seems to be relatively close to unity[23] at all maturities investigated. The rejection of the second condition of the strict version of the EH theory (i.e. a cointegrating vector of (1,-1) and without an intercept term) suggests that the forward rate cannot fully explain the future levels of the money market rates. The interpretation of this result may be twofold. Either, we can postulate that only a weak version of the EH is supported by the data, i.e. there exists a 'constant' term premium,[24] or it may mean that there exists a time-varying risk premium. The latter interpretation has also received support in the literature.

In order to check the existence and size of a constant risk premium, we re-estimated the CIVAR model imposing the presence of the intercept in the cointegrating space. We then tested the joint hypothesis of $\beta=1$ and $\alpha\neq0$. For all maturities investigated except the three-month, the tests were not rejected, thus supporting a weak version of the EH theory. Although the restriction test was rejected at the three-month maturity, we chose to estimate with a constant term also at this maturity. This was because we find it a bit of a puzzle that there should exist risk premia at the horizons of one, six and nine months, but not for the three-month horizon. Finally, this results seems to be sensitive to the choice of sample period and by

heteroscedastic-consistent standard errors methods in Hendry and Doornik (2001).

[22] The problem of serial correlation is often provided by overlapping observations and can however explain the difficulty to obtain appropriate and reasonable estimation results in a single-equation framework (not reported here).

[23] See Appendix 7.

[24] See Gerlach and Smets (1997) and Hardouvelis (1988).

starting the estimation period later, the restriction test was not rejected, thus suggesting a constant term premium also for the three-month rate.

Under these circumstances (i.e. β=1) the estimated coefficient for α can be interpreted as a constant term premium.[25] Our findings indicate a constant risk premium that varies from 5 basis points at the one-month horizon, 10 basis points at the three-month horizon, 15 basis points at the six-month horizon and reaching 27 basis points at the nine-month horizon. However, it should be noted that the estimated standard deviations of the constant terms were found to be relatively high at the six- and nine-month maturities (17 and 32 basis points respectively), indicating high uncertainty about the "true" levels of risk premia. At the one- and three-month maturities the standard deviations were smaller: 1 and 6 basis points respectively. The results are reported in Table 5. The last row of Table 5 reports the LR test of restrictions which follows a χ^2 (1) distribution with the p-value in brackets (see Hendry and Doornik, 2001).

Broadly speaking, the results of the CIVAR estimations broadly confirm the results of the Phillips-Hansen estimations. The fact that the results from the Phillips-Hansen estimations and the CIVAR are in line suggests that non-normality in the residuals is not critical for the results from the cointegrating VAR[26].

Regarding the samples that end in September 2001 and February 2002, the results were mixed (see Appendix 8). This was not surprising given that the Phillips-Hansen estimations indicated that the restriction β=1 was rejected at maturities beyond one-month. This was broadly confirmed from the VAR estimations. However, the LR-restriction test in the VAR regressions did not reject the EH theory for the six-month maturity, where the estimated risk premia were 6 and 9 basis points for the two samples. Nevertheless, the tests were close to reject the hypothesis at a 5% significance level also at this maturity. Given that the unit relationship restriction tests were rejected using the Phillips-Hansen fully modified estimator we chose to not emphasise these particular estimates.

[25] Even if normally, as proposed by Johansen (1992), there is no need to have an intercept inside the cointegration relationship, its inclusion makes sense since the implied forward rates cannot explain alone all the fluctuations of the future spot rates. In other words, any disparities between interest rates that are not explained by the strict relationship is found in the constant term which therefore is defined as an excess return (for a detailed definition of excess return, see Chapter 10 of Campbell, Lo and MacKinlay, 1997) of the forward rates relative to the spot rates for similar maturity (i.e. a risk premium). Another argument for running cointegration analysis with a restricted constant is that we want to eliminate the linear trend from the interest rate level.

[26] The non-normality seems to stem from excess kurtosis, which is less problematic. See Godfrey and Orme (1991).

Table 5: Results and mis-specification tests for the CIVAR(ρ) for the one-, three-, six- and nine-months horizon. German Libor market December 1989- December 1998

Horizon	One-month	Three-month	Six-month	Nine-month
	$(f_t^{1m,1m} - r_{1m,t}^{t+1m} - \alpha)$	$(f_t^{1m,3m} - r_{1m,t}^{t+3m} - \alpha)$	$(f_t^{1m,6m} - r_{1m,t}^{t+6m} - \alpha)$	$(f_t^{1m,9m} - r_{1m,t}^{t+9m} - \alpha)$
Lags	13	12	13	11
Cointegrating vector	1 - 1 - 0.053	1 - 1 - 0.101	1 - 1 - 0.146	1 - 1 - 0.272
S.E. of α	(0.014)	(0.057)	(0.167)	(0.322)
Portmanteau test (26)				
spot rate	27.29	29.38	30.28	36.78
forward rate	34.00	22.61	17.55	15.98
ARCH 1-7				
spot rate	0.58	0.41	0.24	0.40
forward rate	0.98	0.42	0.78	1.60
Hetero-test				
spot rate	0.41	0.60	0.47	0.47
forward rate	0.59	0.64	0.72	0.73
Normality -test				
spot rate	21.81 **	31.73 **	36.41 **	39.09 **
forward rate	28.92 **	9.70 **	7.24 **	0.40
LR test of restrictions	0.02	7.72 **	1.49	2.82
p-value	(0.89)	(0.01)	(0.22)	(0.09)

*Note: * and ** denote the rejection of the null hypothesis at 5% and 1% significance levels respectively. The Portmanteau test follows a χ^2 distribution. The critical value of the test is 38.9 and 45.6 respectively at 5% and 1% significance level. See Hamilton (1994) for the critical values. The selection of the number of lags used in the CIVAR is based on two elements: (i) having a sufficient number of lags to remove autocorrelation and heteroscedasticity (ii) the AIC and HQ information criteria.*

4.1.3 Testing a level shift in the cointegrated vector autoregressive model

The results of both methods (Phillips and Hansen and the CIVAR) for the larger sample (including the EURIBOR observations) could be explained either by a time-varying risk premium, or by a change in the structure of the constant risk premium since the launch of the European single currency. Given the results of both the cointegrated estimations and the recursive estimations (see Appendix 5), only the latter of the two explanations are tested here[27].

In order to detect the potential shift in the level of the intercept in the long-run relationship and its timing as well, Durré (2003) follow the methodology proposed in Gregory and Hansen (1996) and Gregory, Nason and Watt (1996). This methodology proposes a residual-based test where α in Equation (1) is not exactly considered as time-invariant, but is allowed to shift to a new long-run relationship (i.e. with another constant risk premium). Moreover, by considering the timing of this shift as unknown, it would

[27] See Durré (2003) for further details.

be particularly interesting to check the evolution of the residuals for each equation around the launch of the European single currency. [28]

In order to take into account a potential shift in the intercept, Equation (1) is modified as follows:

$$(2) \quad R(n)_t = \alpha_1 + \alpha_2 \varphi_{t\tau} + \beta f^k(n)_{t-k} + \zeta_t$$

where α_1 is the intercept before the shift and α_2 represents the change in the intercept at the time of the shift τ. As previously, $f^k(n)_{t-k}$ is the forward rate at time $t-k$ of an n period instrument beginning in k periods, $R(n)_t$ is the spot rate at time t of an n period instrument, and ζ_t is the residual which displays a I(0) process. We then estimate this cointegrating relationship by ordinary least squares (OLS) and an augmented Dickey-Fuller unit root test is applied to the regression errors. Thereafter the cointegration test statistic for each possible regime shift $\tau \in T$ (with T as a subsample of all available observations) is computed. For each τ we estimate equation 2 by OLS, yielding the residual ζ_{tT} (where the subscript τ denotes the dependance of the residual sequence on the choice of change point τ). As suggested by Gregory and Hansen (1996), the smallest value (or the largest negative value) of the test statistic is a signal of a regime shift.[29] The results displayed in Figure 1 in Appendix 9 clearly show a structural change far and near the emergence of the European single currency.

According to these results, we reestimated the VAR systems for the sample covering the period from December 1989 to August 2001 for the LIBOR/EURIBOR market, by imposing the following specification for the cointegrating vector for all maturities : $X=(1,-1,-\alpha_i,-d_{emu})'$, in which the first two variables are the imposed values for respectively the implied forward rate and the spot rate at different horizons, α_i (i=1,3,6 and 9) is the restricted intercept for which we impose the value equal to the one prevailing during the German period (cfr Table 5) and d_{emu} is a dummy variable with value one for the euro period (i.e. from January 1999 onwards) and zero otherwise. The results are reported in Table 6.

[28] The results (not reported here) of the residuals sum of squares of the spot rate and the forward rate equations of the CIVAR showed indeed a jump around 1999 while the Chow test clearly signals a break in 1999 (see Durré, 2003).

[29] Gregory and Hansen (1996) propose to choose the subset T=(0.15,0.85).

Table 6: Results and mis-specification tests for the CIVAR(ρ) for the one-, three-, six- and nine-month horizon. German Libor and Euribor markets December 1989-August 2001

Horizon	One-month ($f_t^{1m,1m}$,$-r_{1m,t}$,$-\alpha$,$-d_{emu}$)	Three-month ($f_t^{1m,3m}$,$-r_{1m,t}$,$-\alpha$,$-d_{emu}$)	Six-month ($f_t^{1m,3m}$,$-r_{1m,t}$,$-\alpha$,$-d_{emu}$)	Nine-month ($f_t^{1m,3m}$,$-r_{1m,t}$,$-\alpha$,$-d_{emu}$)
Lags	13	12	13	11
Cointegrating vector	1, -1, -0.053, 0.0169	1, -1, -0.101, 0.052	1, -1, -0.146, 0.088	1, -1, -0.272, 0.1360
Portmanteau test				
- spot rate	4.9764	2.9743	3.3816	4.5517
- forward rate	8.7334	2.8748	2.2145	2.9509
ARCH 1-7				
- spot rate	0.7288	0.7180	0.3051	0.8486
- forward rate	2.8191	1.0778	1.7966	1.7878
Hetero-test				
- spot rate	0.8730	0.9809	0.939	0.9760
- forward rate	1.4482	0.6903	1.4840	1.0334
Normality-test				
- spot rate	22.92**	38.54**	36.94	44.12**
- forward rate	71.91**	102.2**	11.94	2.5415**
LR test of restriction	0.07	6.17	3.72	5.89
p-value	(0.96)	(0.05)	(0.15)	(0.05)

Note : * and ** denote the rejection of the null hypothesis at 5% and 1% significance levels respectively. The Portmanteau test follows a χ^2 distribution. The critical value of the test is 38.9 and 45.6 respectively at 5% and 1% significance level. See Hamilton (1994) for the critical values. The selection of the number of lags used in the CIVAR is based on two elements: (i) having a sufficient number of lags to remove autocorrelation and heteroscedasticity (ii) the AIC and HQ information criteria.

The results presented in Table 6 thus confirm the preliminary intuition from the residual-based test from Gregory and Hansen (1990). Indeed, at most horizons, the restriction tests do not reject the specification with a level shift from January 1999 onwards. In particular, the estimations of the CIVAR systems seem to indicate that the introduction of the European single currency has entailed a decrease of the constant risk premium of an amount 1.7, 5.2, 9 and 13.6 basis points respectively at the 1-, 3-, 6- and 9-month horizons compared to the estimated risk premiums using the "German" sample period.

4.2 Estimating a CIVAR for the EONIA swap market

The CIVAR was estimated on daily EONIA swap data using two different samples. The first sample spans the period from 1999[30] to 10 September 2001, while the second sample period includes data until 5 June 2002. The reason for using the sample period ending on 10 September was the significant impact on money market rates from the terrorist attacks, which of course could not be foreseen. Thus one could suspect that the results from including the observations following these events would be distorted by large expectational errors. This is obvious when looking at chart 4, which shows that expected rates three months ahead were much higher than the actual outturns For these reasons we chose to cross check the results by using these two sample periods. However, the results from the two different samples were in fact broadly the same. We note from the outset that an unavoidable problem in studying economic relationships in the euro area is the fact that the period since the introduction of the euro on January 1 1999 is relatively short. This is particularly relevant considering the use of cointegrating techniques to evaluate long-term relationships, which normally requires long sample periods. However, one could argue that the "long-term relationships" are established faster for financial markets data than for other more typical macroeconomic variables, like money and prices, consumption and income, et cetera. In any case, such estimations on a limited sample period might give some tentative indications, which need to be reestimated as more data become available.

[30] The exact start date depends on the maturity. This is because the first observations of the spot rates are used to compute implied forward rates, which are then regressed on the actual outturns of the spot rates. For the one-month regression, the first estimation observation is 2 February. For the three-, six- and nine-month horizons the regressions start at 7 June, 2 September and 30 November 1999 respectively.

Table 7. Johansen test for cointegration (stochastic matrix test). EONIA swap rates

Horizon	H_0	trace-test	p-value	cointegrating vector restriction test[1] (p-value)
One month	r <= 0	78.133	0.000 **	
$(r^{1m}_{t+1m} - f_t^{1m,1m})$	r <= 1	0.31117	0.649	0.311
Three-month	r <= 0	14.21	0.023 *	
$(r^{1m}_{t+3m} - f_t^{1m,3m})$	r <= 1	0.81921	0.423	0.000 **
six-month	r <= 0	14.819	0.018 *	
$(r^{1m}_{t+6m} - f_t^{1m,6m})$	r <= 1	0.35328	0.624	0.000 **
Nine-month	r <= 0	13.151	0.035 *	
$(r^{1m}_{t+9m} - f_t^{1m,9m})$	r <= 1	0.30171	0.655	0.000 **

*Note: A number of 21 lags have been used. The test statistics for H(rank≤p) are listed with p-values based on Doornik and Hendry (2001); ** and * mark significance at 95% and 99%. Testing commences at H(rank=0), and stops at the first insignificant statistics. Sample period is 1999 - 10.9 2001.*
*1) The statistic shown for this test is the p-value of the null hypothesis test that the cointegrating vector β coefficient is equal to 1. ** indicates the rejection of the null hypothesis at a significance level of 1%.*

We used the same methodology as for the monthly dataset outlined above. I.e., we first estimated a cointegration relationship between the EONIA swap one-month rate and the one-month implied forward rate at different horizons with an unrestricted intercept. We tested the null hypothesis of β=1 which was rejected, except at the one-month horizon (see Table 7). However, the cointegrating coefficient of the implied forward rates seems to be relatively close to unity at all maturities investigated (see Table 1 in Appendix 10). Furthermore, we tested the joint hypothesis of β=1 and $\alpha \neq 0$ by imposing the presence of the intercept in the cointegrating space. The results are reported in Table 8. The last row of Table 8 reports the LR-test of restrictions which follows a χ^2 (1) distribution with the p-value in brackets.

The results of the mis-specifications tests report basically no autocorrelation or heteroscedasticity (White, 1980, and Kalirajan, 1989) in the residuals. However, the normality test is strongly rejected in all cases. Nevertheless, as indicated by the charts in Appendix 11, this result does not seem to stem from 'excess skewness' but from 'excess kurtosis', thus, the distribution function of the residuals is symmetric, even if there is non-normality.

More interestingly, the joint restriction of β=1 and $\alpha \neq 0$ is not rejected for the three, six and nine month maturities. This suggests existence of a constant risk premium at these maturities. The risk premia are estimated to be 2, 6 and 10 basis points for the three-, six- and nine-month horizons respectively (see Table 8). For the one-month horizon the estimated risk premium was zero as the restricted presence of a constant term in the cointegrating space was rejected and therebyconfirms the results from Table 6.

Also in the case of the EONIA swap rates the standard deviations of the estimated constants were relatively high. At the three-, six- and nine-month maturities they were 5, 7 and 15 basis points respectively, about the same size as the point-estimates of the risk premia. This indicates high uncertainty about the size of risk premia. Consequently, estimations with different sub-samples may result in different risk premia (see Appendix 12).

Table 8: Results and mis-specification tests for the CIVAR(ρ) for the one- three-, six- and nine-months horizon. Sample period: 1999 - 10 September 2001 (see also footnote 28). EONIA swap rates

Horizon	One-month	Three-month	Six-month	Nine-month
	$(f_t^{1m,1m} - r_{1m,t}^{t+21} - \alpha)$	$(f_t^{1m,3m} - r_{1m,t}^{t+63} - \alpha)$	$(f_t^{1m,6m} - r_{1m,t}^{t+126} - \alpha)$	$(f_t^{1m,9m} - r_{1m,t}^{t+189} - \alpha)$
Lags	22	9	14	5
Cointegrating vector	1 - 1 - 0.0166	1 - 1 - 0.0159	1 - 1 - 0.058	1 - 1 - 0.096
S.E. of α	(0.013)	(0.051)	(0.068)	(0.149)
Portmanteau test (21)				
spot rate	0.53	14.59	8.82	16.26
forward rate	19.69	16.61	10.81	22.50
ARCH 1-1				
spot rate	33.08 **	20.63 **	19.18 **	15.14 **
forward rate	53.44 **	16.53 **	2.00	3.58
Hetero-test				
spot rate	0.76	1.43	1.49	1.31
forward rate	1.32 *	1.41	0.85	1.48
Normality -test				
spot rate	637.66 **	347.92 **	363.07 **	380.87 **
forward rate	187.69 **	119.47 **	39.46 **	29.38 **
LR test on restrictions	4.21	0.19	0.18	0.05
p-value	(0.04) *	(0.66)	(0.68)	(0.83)

Note: * and ** denote the rejection of the null hypothesis at 5% and 1% significance levels respectively. The Portmanteau test follows a χ^2 distribution with m=21 degrees of freedom. The critical value of the test is 32.7 and 38.9 respectively at 5% and 1% significance level. See Hamilton (1994) for the critical values. The selection of the number of lags used in the CIVAR is based on two elements: (i) having a sufficient number of lags to remove autocorrelation and heteroskedasticity (ii) the AIC and HQ information criteria.

Extending the sample

As mentioned, we also estimated the model on a longer sample, ending in early June 2002. The estimated risk premia on this sample are not much different from the results using the shorter sample. Table 9 presents the results from the Johansen cointegration tests and Table 10 presents the results from the estimated CIVAR. On this sample, the estimated risk premia were 1, 2, 4 and 13 basis points for the maturities of one-, three-, six- and nine months respectively. The standard deviations of the risk premia were 1, 5, 9 and 15 basis points respectively. Thus, risk premia seem to be very small at maturities up to three months, but more significant beyond that horizon. There is high uncertainty about the exact size of the risk premia at the six- and nine-month maturities.

Table 9: Johansen test (stochastic matrix) for cointegration

Horizon	H_0	trace-test	p-value	cointegrating vector restriction test[1] (p-value)
One month	r <= 0	107.93	0.000 **	
$(r^{1m}_{t+1m} - f_t^{1m,1m})$	r <= 1	0.013265	0.946	0.3616
Three-month	r <= 0	23.299	0.000 **	
$(r^{1m}_{t+3m} - f_t^{1m,3m})$	r <= 1	0.022347	0.926	0.000 **
six-month	r <= 0	13.912	0.026 *	
$(r^{1m}_{t+6m} - f_t^{1m,6m})$	r <= 1	0.017919	0.935	0.000 **
Nine-month	r <= 0	11.852	0.059	
$(r^{1m}_{t+9m} - f_t^{1m,9m})$	r <= 1	0.4258	0.584	0.000 **

Note: the test statistics for H(rank≤p) are listed with p-values based on Doornik and Hendry (2001); ** and * mark significance at 95% and 99%. Testing commences at H(rank=0), and stops at the first insignificant statistics.
1) The statistic shown for this test is the p-value of the null hypothesis test that the cointegrating vector β coefficient is equal to 1. ** indicates the rejection of the null hypothesis at a critical level of 1%. Sample period is 1.1 1999 - 5.6 2002.

Table 10: Results and mis-specification tests for the CIVAR(ρ) for the three-, six- and nine-months horizon with constant. Sample period: 1999- 5 June 2002 (see also footnote 28). EONIA swap rates

Horizon	One-month $(f_t^{1m,1m} - r_{1m,t}^{t+21} - \alpha)$	Three-month $(f_t^{1m,3m} - r_{1m,t}^{t+63} - \alpha)$	Six-month $(f_t^{1m,6m} - r_{1m,t}^{t+126} - \alpha)$	Nine-month $(f_t^{1m,9m} - r_{1m,t}^{t+189} - \alpha)$
Lags	22	9	14	5
Cointegrating vector	1 - 1 - 0.0132	1 - 1 - 0.0226	1 - 1 - 0.041	1 - 1 - 0.128
S.E. of α	(0.0103)	(0.0518)	(0.0948)	(0.1485)
Portmanteau test (21)				
spot rate	0.55	16.12	8.94	20.10
forward rate	19.19	22.80	6.88	21.22
ARCH 1-1				
spot rate	32.36 **	22.96 **	19.64 **	13.26 **
forward rate	61.35 **	23.08 **	1.73	8.44 **
Hetero-test				
spot rate	1.08	0.96	1.10	1.40
forward rate	1.84 **	1.37	0.88	0.93
Normality -test				
spot rate	1454.60 **	1233.60 **	1334.30 **	1459.50 **
forward rate	298.46 **	191.05 **	64.67 **	31.89 **
LR test on restrictions	2.47	0.06	0.01	0.55
p-value	(0.12)	(0.81)	(0.91)	(0.46)

Note: * and ** denote the rejection of the null hypothesis at 5% and 1% significance levels respectively. The Portmanteau test follows a χ^2 distribution with m=21 degrees of freedom. The critical value of the test is 32.7 and 38.9 respectively at 5% and 1% significance level. See Hamilton (1994) for the critical values. The selection of the number of lags used in the CIVAR is based on two elements: (i) having a sufficient number of lags to remove autocorrelation and heteroskedasticity (ii) the AIC and HQ information criteria.

5. Conclusions

The expectations hypothesis (EH) was tested on German data for the period December 1989- December 1998. The tests were conducted using the single-equation framework of Phillips-Hansen and in a cointegrated VAR framework. The results from both methodologies indicated that the weak version of the EH (i.e. with a constant risk premium) holds true for maturities of up to nine months.

The VAR-estimations on monthly German LIBOR data suggest risk premia of around 5 basis points at the one-month maturity. The premium increases to around 10 basis points at the three-month maturity, and reaches 15 and 27 basis points at the six- and nine-month maturities respectively.

When expanding the data by incorporating the period from 1999 until late 2001 using EURIBOR rates, the results of the estimations suggest that the specification of the cointegrating vector for the German period is slightly different after the start of EMU, i.e. the null hypothesis that the cointegrating vector between forward rates and spot rates is (1,-1) was rejected by close margin. These results may have two potential explanations. On the one hand, the rejection of the EH may come from the existence of time-varying risk premia following the launch of the European single currency. On the other hand, the results could be explained by a shift in the level of risk premia at different horizons, in the sense of Gregory and Hansen (1996).

This paper proposes to test the latter potential explanation mainly for two reasons. First, according to the convergence of the different national monetary policies towards the German approach, it would be surprising that time-varying risk premia should appear with the launch of the European single currency, while the results suggest that time-independent risk premia existed in German money markets before this. Second, according to the Phillips and Hansen results and the OLS recursive estimations, the rejection of the EH seems to be due to a change of a limited size. Indeed, the OLS recursive estimations of the relationship between the spot rate and the implied forward rate appears quite stable over the whole period, i.e. from December 1989 to August 2001. Therefore, we explicitly tested the existence of a structural change in the level of the constant risk premium and found that this change probably appeared in January 1999. By reestimating the preliminary models for the whole sample, we found that this last event has entailed a decrease of risk premia of an amount of 1.7, 5.2, 9 and 13.6 basis points respectively at the 1-, 3- ,6- and 9-month horizons compared to the estimated risk premiums of the German period.

Although the limited number of years for an cointegration analysis, we also estimated a cointegrating VAR on EONIA swap rates data with daily frequency starting in 1999, using two different samples (the first sample period ends on 10 September 2001 and the second sample period ends on 5 June 2002). The results broadly confirmed the level shift in the previous estimations. Indeed, they indicate somewhat lower term premia than those resulting from the estimations using monthly data on the German Libor market. More precisely, they tentatively suggest risk premia in ranges of 0-1, 2, 4-6 and 10-13 for the maturities of one, three, six and nine months respectively. Given the limited sample period, these results should be interpreted with caution.

The results from the EONIA swap rates and the LIBOR/EURIBOR rates are not 100% comparable. This is because, as explained in Section 2, credit risk in EONIA swap rates is more limited than for the LIBOR and EURIBOR rates. As credit risks also could be expected to increase with the maturity, the difference between risk premia embedded in EONIA swap rates and EURIBOR rates could also be expected to increase with the maturity. Thus, some of the discrepancy between the estimated term premia from the two different datasets may be due to different credit risk.

Finally, the results seem to be broadly in line with other comparable empirical evidence. For instance, Cuthbertson et. al. (2000) find that the expectations hypothesis is compatible with German money market data for the period from 1976 to 1993. Moreover, the results of Cassola and Luis (2001), using German data up to 1998, suggest that the risk premium increases with the maturity of the implied forward rates.

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Appendix 1

Potential bias involved by the conventional test of the Expectations Hypothesis

In his recent study, Thornton (2001) underlines problems that may arise from the conventional test of the expectations hypothesis, i.e. in which the short-term interest rate appears symmetrically on both sides of the estimation. By doing this, we test if the spread between interest rates with different maturities is a good predictor of the future change of the spot rate. The basic point of the argument may be summarized as follows. Traditionally, when testing the EH theory, we estimate the following relationship:

$$\frac{1}{k}\sum_{i=0}^{k-1} R^m_{t+mi} - R^m_t = \alpha + \beta(R^n_t - R^m_t) + \omega_t$$

where R^m_t is the (m-period) short-term interest rate, R^n_t is the longer-term interest rate, (k-1) is the period in the future and ω_t is an iid white noise error. If the EH theory holds true, the null hypothesis $\beta=1$ is not rejected. In his paper, Thornton argues that this relationship is correctly specified if and only if the EH holds true. Otherwise, this specification may entail bias in favour of the support the EH theory. Then, before subtracting the short-term interest on both sides of the equation, he parameterises the relationship between the long-term and the short-term interest rates according to the expectations theory:

$$\frac{1}{k}\sum_{i=0}^{k-1} R^m_{t+mi} = \theta R^n_t - \frac{1}{k}\sum_{i=0}^{k-1} V_{t+mi} - \pi$$

where the average weighted short-term interest rate in the future is a function of the actual long-term interest rate, minus the sum of the white noise errors and a constant term (defined here as a constant term premium). Now, when subtracting the actual short-term interest rate, R^m_t, on both sides of the previous equation, we obtain:

$$\frac{1}{k}\sum_{i=0}^{k-1} R^m_{t+mi} - R^m_t = \theta R^n_t - R^m_t - \frac{1}{k}\sum_{i=0}^{k-1} V_{t+mi} - \pi$$

Note that this equation is similar to the first one if we add the expression $\theta R^m_t - \theta R^m_t$ on the left-hand side of the previous equation. Doing so and using the same notation for the constant and the error terms, we obtain:

$$\frac{1}{k}\sum_{i=0}^{k-1} R^m_{t+mi} - R^m_t = \alpha + \theta(R^n_t - R^m_t) + (\theta - 1)R^m_t + \omega_t$$

which is similar to the first formulation of the EH theory only if $\theta=1$. If not, the third term of the previous equation does not disappear and then θ cannot be considered as the best estimator of β. Indeed, taking the expected value of the least squares estimator of β from the first equation, we obtain in this case:

$$E\hat{\beta} = \theta + (\theta - 1)E\frac{\sum(\overline{R^n_t} - \overline{R^m_t})\overline{R^m_t}}{\sum(\overline{R^n_t} - \overline{R^m_t})^2} - E\frac{\sum(\overline{R^n_t} - \overline{R^m_t})\omega_t}{\sum(\overline{R^n_t} - \overline{R^m_t})^2}$$

where the variables in the fractional expressions denote the adjustment for the mean. Then, if $\theta \neq 1$, the limit of $\hat{\beta}$ when $N \rightarrow \infty$ converges to:

$$P \lim_{N\rightarrow\infty} \hat{\beta} = \theta + (\theta - 1)\left[\frac{\sigma_{Rn,Rm} - \sigma^2_{Rm}}{\sigma^2_{Rn} - 2\sigma_{Rn,Rm} + \sigma^2_{Rm}}\right]$$

So, if $\theta \neq 1$, then $\hat{\beta}$ will also be influenced by the second term which is affected itself by the covariance and the relative variances of each interest rate. In the case where short-term and long-term interest rates are uncorrelated, i.e. $\sigma_{Rn,Rm} = 0$, the previous equation converges to $\theta + (1-\theta)\left[1 + \frac{\sigma^2_{Rm}}{\sigma^2_{Rn}}\right]$ and therefore, when the variance of the (short-term) m-period rate increases and is slightly higher than the variance of the (long-term) n-period interest rate, it is possible that $\hat{\beta}$ is biased toward one. On the other hand, in the case where $\sigma^2_{Rn} > \sigma_{Rn,Rm}$ the terms in brackets become less than one which increases the probability of a higher value for $\hat{\beta}$. That could explain why in some of our single-equation estimations we obtained estimated values for $\hat{\beta}$ larger than one. Even if long-term and short-terms interest rates are generated independently (i.e. a random walk process) or by an autoregressive process, the test of the EH in terms of spreads finds in most cases results in favour of the theory. Indeed, for the latter case, which seems to be valuable for instance for interest rates in the EONIA market, each interest rate is generated by the following process:

$$R^m_t = \mu_m + \varphi R^m_{t-1} + \varepsilon_t$$

$$R^n_t = \mu_n + \psi R^n_{t-1} + \gamma_t$$

This is indeed the process the EONIA swap rates seems to display using the standard methodology of Box and Jenkins. We then obtain in this case:

$$P \lim_{N\rightarrow\infty} \hat{\beta} = \frac{\frac{1}{k\sigma^2_\gamma}\left[\frac{\sigma_{Rn,Rm} * \sigma^2_{Rm}/\sigma^2_{Rn}}{1-\varphi\psi} - \frac{1}{1-\psi^2}\right]\left[\frac{\psi - \psi^k}{1-\psi} - (k-1)\right]}{\frac{\sigma^2_\varepsilon}{1-\psi^2} + \frac{\sigma^2_\gamma}{1-\psi^2} - 2\frac{\sigma_{\varepsilon,\gamma}}{1-\psi\varphi}}$$

where we see that the estimate of β will be positive in the particular case where $\sigma^2_{Rm}/\sigma^2_{Rn} < 1$ and $\varphi < \psi$. Moreover, the larger the variance of the short-term interest rate is compared to the variance of the long-term interest rate, the higher the estimate of $\hat{\beta}$ will be. If we look for instance at the EONIA market, it must be quoted that the short-term interest rates display on average higher volatilities than the long-term interest rates. This is especially the case during September 2001 where the variance of the EONIA spot rate was 39% higher than the variance of the EONIA one-month swap rate while the latter was only 10% higher on average than the variance of the implied forward rates of maturity from one month up to one year. Some explanation of our preliminary results inside the basic single equation (Section 3.2) could come from these features. Note for example that the estimations on the sample including the observations after 11 September 2001 display systematically higher values of $\hat{\beta}$ compared to the results on the sample without those observations.

Appendix 2

Methods for cointegration testing

The first step in the econometric analysis is to check the existence of cointegration between the spot rates and the implied forward rates. We computed two tests of cointegration, the Engle-Granger (1987) test and the Johansen and Juselius (1990) test.

The Engle-Granger methodology is to check the presence of unit-roots in the residuals of the long-run relationship between the interest rates. If both interest rates are integrated of order 1, there may exist a linear combination of these variables that is stationary. Accordingly, the first step is to estimate the following relationship:

$$(3)\ r_{i,t}^{\ t+k} = \alpha + f_t^{\ i,j} + e_t$$

where $r_{i,t}^{\ t+k}$ is the ex-post realised i-month spot rate in t+k days at time t (i.e. in our case the LIBOR/EURIBOR rate or the EONIA swap rate), $f_t^{\ i,j}$ is the implied forward rate at time t for the i-month interest rate in j months while α and e_t are the intercept and the residuals of the regression respectively. The second step is to compute the first-difference of the estimated residuals of this long-run relationship, $\hat{e}_t$:

$$(4)\ \Delta\hat{e}_t = a_1\hat{e}_{t-1} + \varepsilon_t$$

If the residuals, $\hat{e}_t$, are found to be stationary while the interest rates variables display an I(1) process, we can conclude that $r_{i,t}^{\ t+k}$ and $f_t^{\ i,j}$ are cointegrated of order (1,1).[31] In other words, if we reject the null hypothesis $|a_1| \neq 0$, we can reject the hypothesis that the variables are cointegrated.[32] We also perform an augmented Dickey-Fuller test in order to test the presence of a unit root in the residuals of equation (3). We test this hypothesis for the one-month spot rate and the implied forward rate with maturities of one

[31] Basically equation (4) is equivalent to perform a Dickey-Fuller test on these residuals to determine their order of integration. Note that since the sequence { $\hat{e}_t$ } is a residual from a regression equation, there is no need to include an intercept term.

[32] See Enders (1995).

month, three months, six months and nine months with a number of 21 lags on the daily data and 12 lags on the monthly data. The results for both data sets are reported in Tables 2 and 3 in Appendix 3.

The DF- and ADF-tests rejected the null hypothesis of non-stationary residuals. We can then conclude that the one-month spot interest rates and the corresponding implied forward rates are cointegrated of order (1,1) for both markets/data sets.

We also performed the Johansen test for cointegration. This four-step procedure was performed with an unrestricted vector autoregressive model using the levels of spot and forward rates, for each relationship between the spot rate and the implied forward rates. A maximum number of 12 lags for the Libor/Euribor market and 21 lags for the EONIA market were chosen. The difference in the number of lags is explained by the frequency used in each case (i.e. monthly and daily frequency respectively). The results are reported in Table 3 for the LIBOR/EURIBOR data and in Table 6 for the EONIA swaps in Section 4 of the paper.

Appendix 3

1. Descriptive statistics. LIBOR/EURIBOR market.

Table 1.1. Normality test for the 1-month spot rate

```
Sample size 173: 10 to 182
Mean                   5.442775
Std.Devn.              2.264577  (using T-1:        2.271151)
Skewness               0.616111
Excess Kurtosis       -1.092024
Minimum                2.570000
Maximum                9.840000
Normality Chi^2(2)=     72.882 [0.0000] **
(asymptotic form of normality test: 19.541)
```

Table 1.2. Normality test for the 1-month implied forward rate

```
Sample size 173: 10 to 182
Mean                   5.475708
Std.Devn.              2.255033  (using T-1:        2.261579)
Skewness               0.627543
Excess Kurtosis       -1.073164
Minimum                2.564508
Maximum                9.839318
Normality Chi^2(2)=      73.82 [0.0000] **
(asymptotic form of normality test: 19.657)
```

Table 1.3. Normality test for the 3-month implied forward rate

```
Sample size 173: 10 to 182
Mean                   5.482594
Std.Devn.              2.184902  (using T-1:        2.191244)
Skewness               0.634632
Excess Kurtosis       -1.035792
Minimum                2.563466
Maximum                9.660893
Normality Chi^2(2)=     71.484 [0.0000] **
(asymptotic form of normality test: 19.346)
```

Table 1.4. Normality test for the 6-month implied forward rate

```
Sample size 173: 10 to 182
Mean                   5.411143
Std.Devn.              2.041883  (using T-1:        2.047810)
Skewness               0.687032
Excess Kurtosis       -0.920810
Minimum                2.635735
Maximum                9.351318
Normality Chi^2(2)=     73.462 [0.0000] **
(asymptotic form of normality test: 19.722)
```

Table 1.5. Normality test for the 9-month implied forward rate

```
Sample size 173: 10 to 182
Mean                   5.441886
Std.Devn.              1.956032  (using T-1:        1.961710)
Skewness               0.691068
Excess Kurtosis       -0.879147
Minimum                2.694427
Maximum                9.207066
Normality Chi^2(2)=      70.27 [0.0000] **
(asymptotic form of normality test: 19.341)
```

2. LIBOR/EURIBOR market : Unit-root tests 23 to 182 with a maximum of 12 lags[33]

Table 2. Test for unit-roots in the residuals of the simple OLS regressions with a maximum of 12 lags

	Level	Δ
Residuals of 1 month equation	-3.4816**	-5.4873**
Residuals of 3 months equation	-3.5970**	-5.6761**
Residuals of 6 months equation	-3.4054*	-5.1205**
Residuals of 9 months equation	-2.9199*	-4.2049**

Source: European Central Bank. Δ denotes the first difference operator. The data sample covers the period from the January 1987 to February 2002. The unit root test used is the univariate augmented Dickey-Fuller for which the critical values are –2.18 and –3.473 at 5% and 1% level of significance, respectively which are marked by * and** (MacKinnon, 1991).

3. EONIA swap market : Unit-root tests

Table 3. Unit-root tests based on ADF test with 21 lags

	Level	Δ
Residuals of 1 month equation	-4.836**	-7.036**
Residuals of 3 months equation	-2.801*	-5.866**
Residuals of 6 months equation	-2.206*	-4.874**
Residuals of 9 months equation	1.374	-4.568**

Source: European Central Bank. Δ denotes the first difference operator. The data sample covers the period from the 4th January to the 10 September2001. The unit root test used is the univariate augmented Dickey-Fuller for which the critical values are –2.18 and –3.473 at 5% and 1% level of significance, respectively which are marked by * and** (MacKinnon, 1991).

[33] Critical values: 5%=-2.18 1%=-3.472; Constant included. ** and * indicates respectively the rejection of the null hypothesis at 5% and 1% significance level.

Appendix 4

The Phillips-Hansen methodology: A short Explanation

Consider for instance the following linear regression model:

$$y = \beta_0 + \beta_1 x_t + u_t, t = 1,2,\ldots,n$$

where the k*1 vector of I(1) regressors are not themselves cointegrated. Therefore, x_t has a first-difference stationary process given by

$$\Delta x_t = \mu + v_t, t = 1,2,\ldots,n$$

in which μ is a k*1 vector of drift parameters and v_t is a k*1 vector of *I(0)*, or stationary variables. It is also assumed that $\xi_t = (u_t, v_t)'$ is strictly stationary with zero mean and a finite positive-definite covariance matrix, Σ.

The computation of the FM-OLS estimator of β is carried out in two stages. In the first stage, yt is corrected for the long-run inter-dependence of u_t and v_t. For this purpose, let $\hat{u}_t$ be the residual vector in the first equation, and write

$$\xi t = (\hat{u}_t, v_t)', t = 2,3,\ldots,n$$

where $v_t = \Delta x_t - \hat{\mu}$, for t=2,3,…,n

and

$$\mu = (n-1)^{-1} \sum_{t=2}^{n} \Delta x_t .$$

A consistent estimator of the long-run variance of ξt is given by

$$\Omega = \Sigma + \Lambda + \Lambda' = \begin{bmatrix} \underset{1*1}{\Omega_{11}} & \underset{1*k}{\Omega_{12}} \\ \underset{k*1}{\Omega_{21}} & \underset{k*k}{\Omega_{22}} \end{bmatrix}$$

where

$$\Sigma = \frac{1}{n-1} \sum_{t=2}^{n} \xi t \xi t'$$

and

$$\Lambda = \sum \omega(s,m) \Gamma s$$

$$\Gamma s = n^{-1}\sum_{t=1}^{n-s}\xi_t\xi_{t+s}'$$

and ω(s,m) is the lag-window with horizon (or truncation) m.

Now let

$$\Delta = \Sigma + \Lambda = \begin{bmatrix}\Delta_{11} & \Delta_{12}\\ \Delta_{21} & \Delta_{12}\end{bmatrix}$$

$$Z = \Delta_{21} - \Delta_{22}\Omega_{22}^{-1}\Omega_{21}$$

$$\hat{y}_t^* = y_t - \Omega_{12}\Omega_{22}^{-1}\hat{v}_t$$

$$\underset{(k+1)*k}{D} = \begin{bmatrix}\underset{1*k}{0}\\ \underset{k*k}{I_k}\end{bmatrix}$$

In the second stage the FM-OLS estimator of β is given by

$$\hat{\beta}^* = (W'W)^{-1}(W'\hat{y}^* - nDZ)$$

where

$$\hat{y}^* = (\hat{y}_1^*, \hat{y}_2^*, \hat{y}_3^*, \ldots, \hat{y}_4^*)', W = (\tau_n, X), \tau_n = (1,1,1,\ldots,1)'$$

In general, we use the lag window of Tukey that is defined as

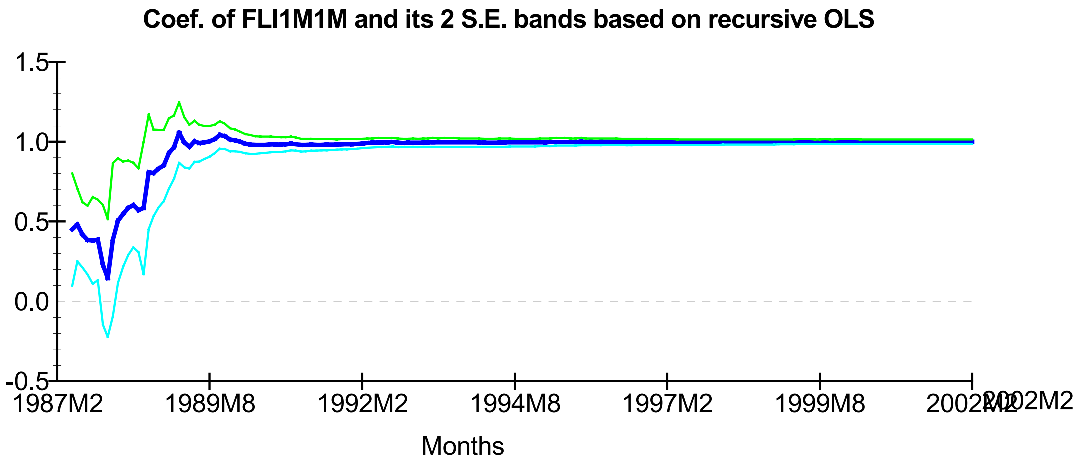
$$\omega(s,m) = 1/2\{1+\cos(\pi s/m)\} \text{ and } 0 \leq s \leq m/2$$

in which the value of m is arbitrary chosen. In the following of the analysis here, we choose a maximum value of m that not exceed 40, as recommended by Pesaran and Pesaran (1997). Moreover, to be consistent with the forthcoming analysis, we determine the exact number of lags for each model that is in line with the selection of the order in the VAR framework. That means a number of lags for truncation equal to 4, 5, 2 and 11 for the 1, 3, 6 and 9-month horizons respectively.

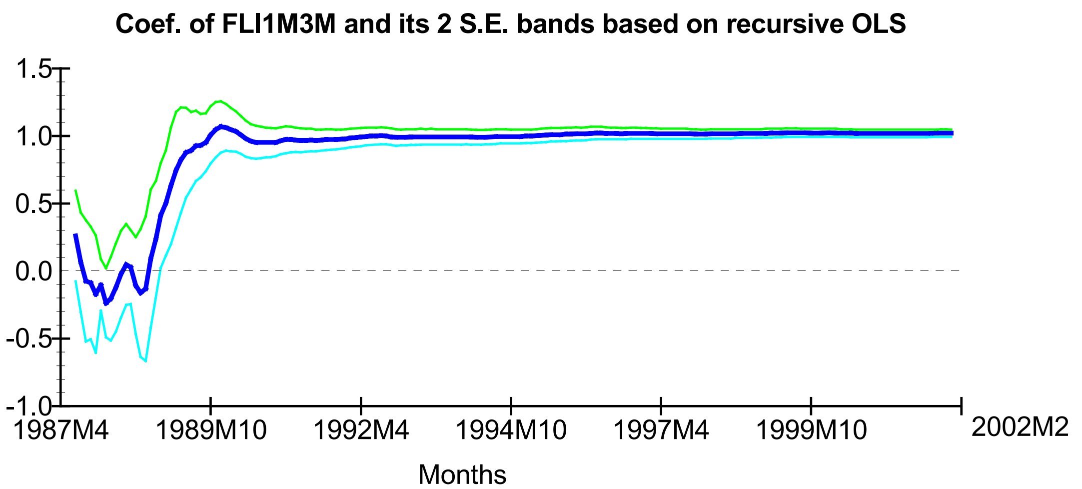
Appendix 5

Estimated β-coefficients and their two standard errors bands based on the recursive ordinary least squares (Phillips-Hansen methodology)

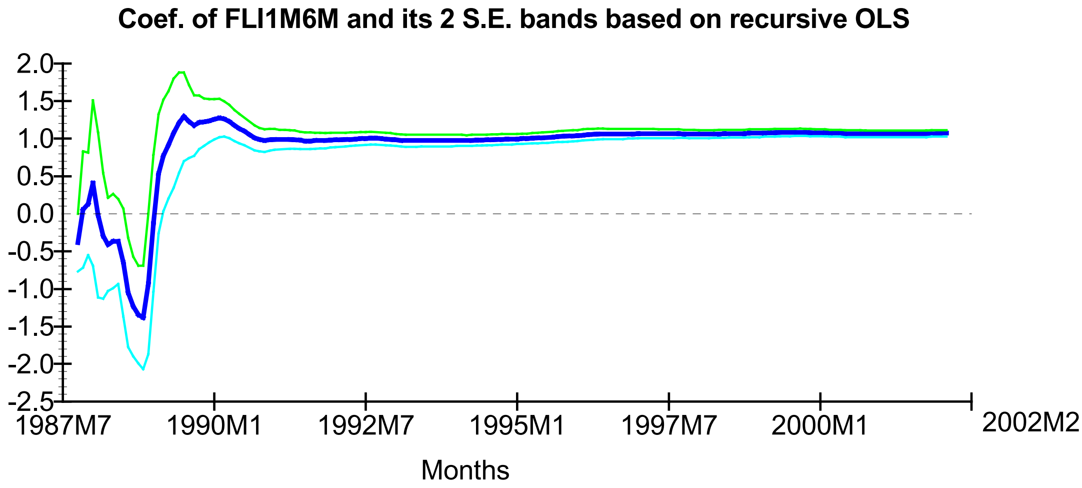
For the 1-month horizon



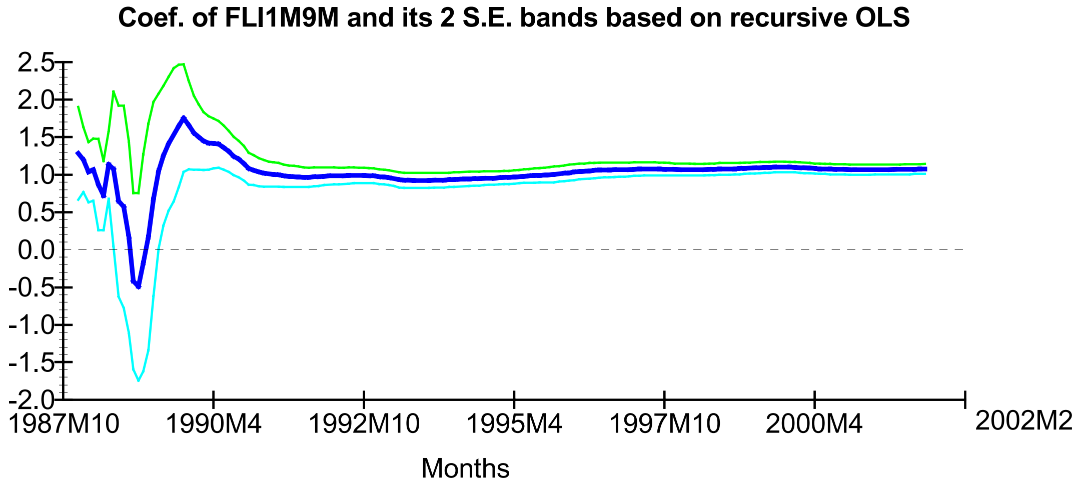
For the 3-month horizon



For the 6-month horizon



For the 9-month horizon



Appendix 6

The VAR methodology

If X_t is an n-sized vector of I(1) variables, the VAR description of the process of X_t can be written :

$$X_t = \mu + \sum_{l=1}^{p} \Phi_l X_{t-l} + u_t$$

where μ is a n-components vector, Φ_l are n x n matrices and u_t a vector of white noise error terms. This description can be re-parameterised as follows:

$$X_t - X_{t-1} = \mu - \Phi(1) X_{t-1} + \sum_{l=1}^{p-1} A_l (X_{t-l} - X_{t-l-1}) + u_t$$

where

$$\Phi(1) = I_n - \sum_{l=1}^{p} \Phi_l$$

and

$$A_l = -\sum_{j=l+1}^{p} \Phi_j$$

Unit roots imply that $\Phi(1)$ is singular. When the VAR is cointegrated, which means that there exist r cointegrating vectors, $\Phi(1)$ can be written as $-\alpha\beta'$ where α and β are n x r matrices of rank r. β is a matrix of cointegrating vectors of coefficients and $\beta' X_{t-1}$ is a stationary r-components vector of cointegrating relations, the fluctuations of which describe short-term departures from the long term relations. α is a matrix of adjustment coefficients, describing the relation of each I(1) variable to previous short term departures from each long term relation. As suggested previously by the cointegration tests, r is equal to 1 between the spot rate and the implied forward rates at different maturities. Imposing this restriction, the cointegrating vector is normalised by constraining the coefficient of the spot rate to 1.[34] Clements and Hendry (1995) show that the estimation of the VAR with this constraint, when true, strongly improves the forecasts. However, the interpretation of the VAR results on a limited number of years must be interpreted with caution as suggested by Campbell and Perron (1991).

[34] For more details on the methodology, see Dor and Durré (1997) and Jondeau (2001).

Appendix 7

Table 1: Results and mis-specification tests for the CIVAR(ρ) for the one-, three-, six- and nine-months horizon without a constant. German LIBOR market.

	One-month	Three-month	Six-month	Nine-month
	$(f_t^{1m,1m} - r_{1m,t}^{t+1m})$	$(f_t^{1m,3m} - r_{1m,t}^{t+3m})$	$(f_t^{1m,6m} - r_{1m,t}^{t+6m})$	$(f_t^{1m,9m} - r_{1m,t}^{t+9m})$
lags	13	12	13	11
Cointegrating vector	1-0.991	1-0.990	1-0.981	1-0.964
portmanteau test (26)				
spot rate	32.75	29.03	29.70	36.53
forward rate	32.63	22.22	17.60	16.26
ARCH 1-7				
spot rate	0.57	0.41	0.23	0.40
forward rate	1.02	0.42	0.77	1.61
Hetero-test				
spot rate	0.41	0.62	0.49	0.48
forward rate	0.62	0.66	0.74	0.73
Normality -test				
spot rate	24.245 **	32.35 **	36.54 **	39.16 **
forward rate	24.37 **	9.66 **	7.29 *	0.37

Note: * and ** denote the rejection of the null hypothesis at 5% and 1% significance levels respectively. The Portmanteau test follows a χ^2 distribution with m=21 degrees of freedom. The critical value of the test is 32.7 and 38.9 respectively at 5% and 1% significance level. These results were compared to the critical values reported in Hamilton (1994).

Appendix 8

Table 1: Results and mis-specification tests for the CIVAR(ρ) for the one, three-, six- and nine-months horizon without a constant. German LIBOR/EURIBOR.. Sample period 1989.12 - 2001.8.

Horizon	One-month	Three-month	Six-month	Nine-month
	$(f_t^{1m,1m} - r_{1m,t}^{t+1m} - \alpha)$	$(f_t^{1m,3m} - r_{1m,t}^{t+3m} - \alpha)$	$(f_t^{1m,6m} - r_{1m,t}^{t+6m} - \alpha)$	$(f_t^{1m,9m} - r_{1m,t}^{t+9m} - \alpha)$
Lags	13	12	13	11
Cointegrating vector	1 - 1 - 0.047	1 - 1 - 0.080	1 - 1 - 0.058	1 - 1 - 0.174
S.E. of α	(0.013)	(0.048)	(0.127)	(0.237)
Portmanteau test (26)				
spot rate	32.77	43.61	25.83	46.42
forward rate	30.57	22.10	17.77	15.18
ARCH 1-7				
spot rate	0.93	1.00	0.34	0.76
forward rate	3.91 **	2.02	1.47	1.83
Hetero-test				
spot rate	0.98	1.13	1.00	0.84
forward rate	1.71 *	0.86	1.38	0.98
Normality -test				
spot rate	17.51 **	35.79 **	35.37 **	48.58 **
forward rate	70.85 **	143.20 **	11.96 **	1.78
LR test of restrictions	0.01	4.15 *	2.10	5.07 *
p-value	(0.93)	(0.04)	(0.15)	(0.02)

Note: * and ** denote the rejection of the null hypothesis at 5% and 1% significance levels respectively. The Portmanteau test follows a χ^2 distribution with m=21 degrees of freedom. The critical value of the test is 32.7 and 38.9 respectively at 5% and 1% significance level. These results were compared to the critical values reported in Hamilton (1994).

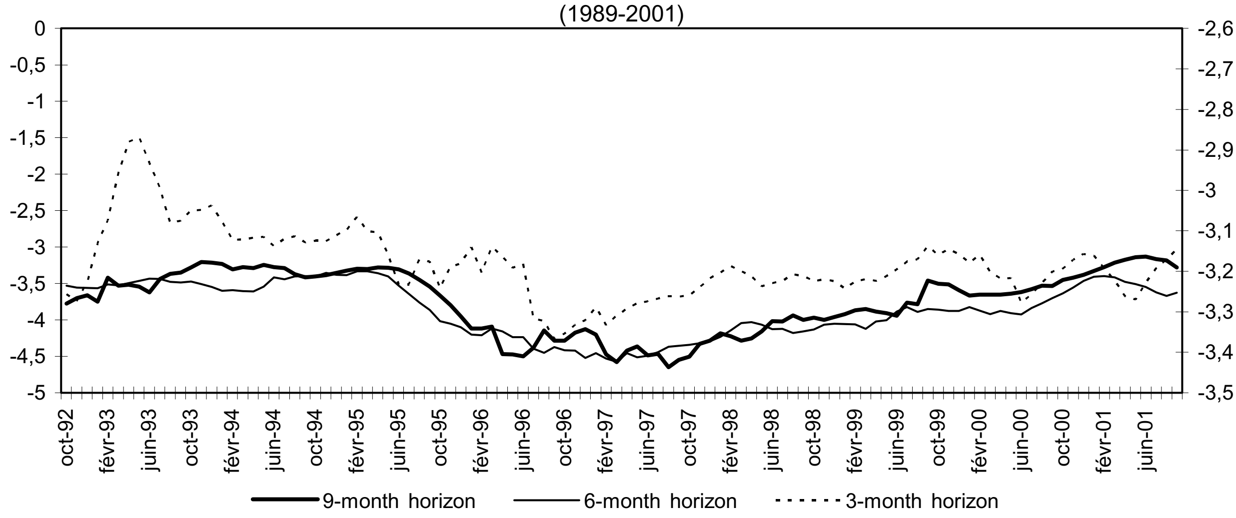
Table 2: Results and mis-specification tests for the CIVAR(ρ) for the one, three-, six- and nine-months horizon without a constant. German LIBOR/EURIBOR.. Sample period 1989.12 - 2002.2

Horizon	One-month	Three-month	Six-month	Nine-month
	$(f_t^{1m,1m} - r_{1m,t}^{t+1m} - \alpha)$	$(f_t^{1m,3m} - r_{1m,t}^{t+3m} - \alpha)$	$(f_t^{1m,6m} - r_{1m,t}^{t+6m} - \alpha)$	$(f_t^{1m,9m} - r_{1m,t}^{t+9m} - \alpha)$
Lags	13	12	13	11
Cointegrating vector	1 - 1 - 0.049	1 - 1 - 0.083	1 - 1 - 0.092	1 - 1 - 0.214
S.E. of α	(0.013)	(0.048)	(0.129)	(0.235)
Portmanteau test (26)				
spot rate	35.54	41.05	28.35	48.72
forward rate	29.16	24.24	16.41	12.80
ARCH 1-7				
spot rate	0.98	1.07	0.36	0.82
forward rate	4.23 **	2.15 *	1.65	1.81
Hetero-test				
spot rate	1.08	1.20	1.14	0.92
forward rate	1.89 **	0.91	1.45	1.01
Normality -test				
spot rate	18.40 **	36.28 **	32.77 **	48.75 **
forward rate	76.05 **	151.95 **	12.82 **	1.77
LR test of restrictions	0.09	4.56 *	3.54	4.83 *
p-value	(0.77)	(0.03)	(0.06)	(0.03)

Note: * and ** denote the rejection of the null hypothesis at 5% and 1% significance levels respectively. The Portmanteau test follows a χ^2 distribution with m=21 degrees of freedom. The critical value of the test is 32.7 and 38.9 respectively at 5% and 1% significance level. These results were compared to the critical values reported in Hamilton (1994).

Appendix 9

Figure 1 : Test Statistic for Structural Breaks in the Forward-Spot Rates Relationship (1989-2001)



Appendix 10

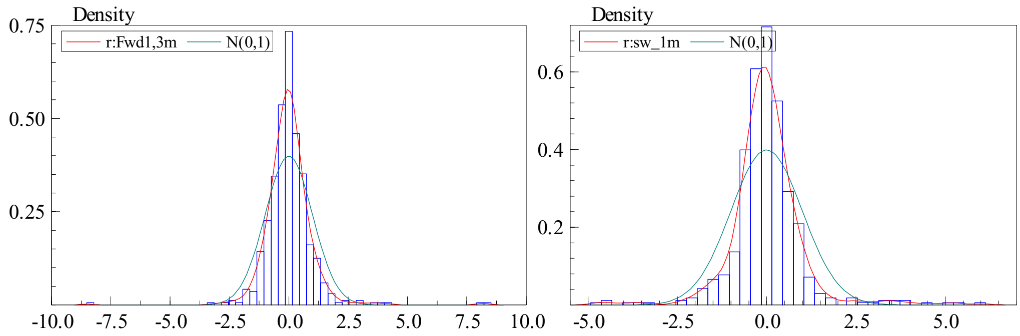
Table 1: Results and mis-specification tests for the CIVAR(ρ) for the three-, six- and nine-months horizon without a constant. EONIA swap market.

Horizon	One-month	Three-month	Six-month	Nine-month
	$r_{1m,t}^{t+21} - f_t^{1m,1m}$	$r_{1m,t}^{t+63} - f_t^{1m,3m}$	$r_{1m,t}^{t+126} - f_t^{1m,6m}$	$r_{1m,t}^{t+189} - f_t^{1m,9m}$
lags	22	9	14	5
Cointegrating vector	1-0.997	1-0.998	1-0.988	1-0.980
portmanteau test (21)				
spot rate	0.53	14.73	8.76	16.1746
forward rate	19.75	16.62	10.88	22.5066
ARCH 1-1				
spot rate	33.25 **	20.55 **	19.17 **	15.12
forward rate	53.65 **	16.42 **	1.97	3.57
Hetero-test				
spot rate	0.76	1.43	1.49 *	1.31
forward rate	1.31 *	1.42	0.85	1.48
Normality -test				
spot rate	637.81 **	347.30 **	363.46 **	381.50 **
forward rate	189.84 **	119.49 **	39.41 **	29.30 **

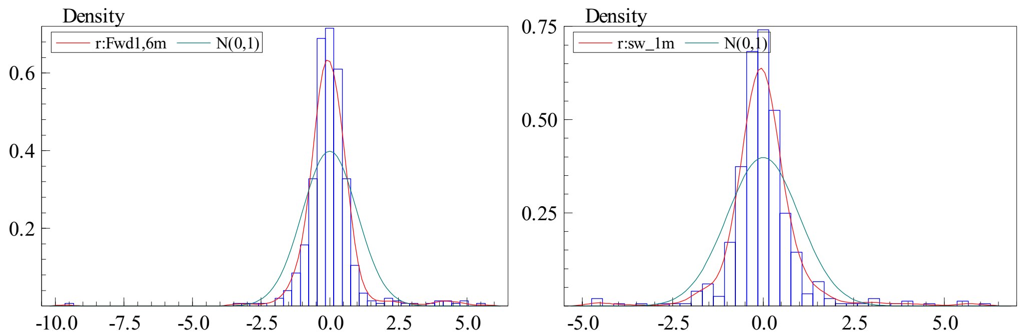
Note: * and ** denote the rejection of the null hypothesis at 5% and 1% significance levels respectively. The Portmanteau test follows a χ^2 distribution with m=21 degrees of freedom. The critical value of the test is 32.7 and 38.9 respectively at 5% and 1% significance level. These results were compared to the critical values reported in Hamilton (1994).

Appendix 11

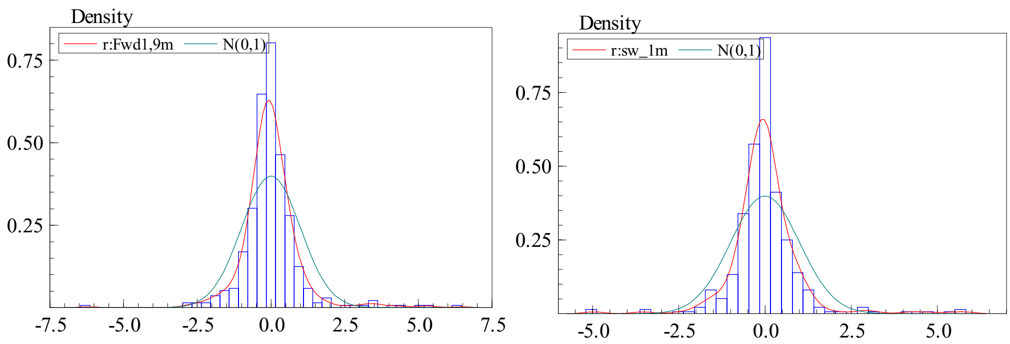
Distribution function of the residuals using a CIVAR(ρ) model – Three-month Horizon



Distribution function of the residuals using a CIVAR(ρ) model – Six-month Horizon



Distribution function of the residuals using a CIVAR(ρ) model – Nine-month Horizon



Appendix 12

Table 1: Estimated constant terms at sub samples, standard errors in parenthesis. EONIA swap market.

*Indicated risk premium measured in basis points = Constant term * 100*

Observation/Date	**3-month maturity**	**6-month maturity**	**9-month maturity**
650/ 16 Jul. 2001	0.068 (0.058)	0.111 (0.070)	0.170 (0.179)
689/ 10 Sep. 2001	**0.016** **(0.051)**	**0.058** **(0.068)**	**0.096** **(0.149)**
800/ 15 Feb. 2002	0.026 (0.056)	0.109 (0.099)	0.101 (0.163)
874/ 5 Jun. 2002	0.022 (0.052)	0.041 (0.095)	0.128 (0.148)

Appendix 13

Resulting tables when using the sample from 4 January 1999 to 5 June 2002. EONIA swap market.

Table 1: Results and mis-specification tests for the CIVAR(ρ) for the three-, six- and nine-months horizon without a constant

Horizon	One-month	Three-month	Six-month	Nine-month
	$r_{1m,t}^{t+21} - f_t^{1m,1m}$	$r_{1m,t}^{t+63} - f_t^{1m,3m}$	$r_{1m,t}^{t+126} - f_t^{1m,6m}$	$r_{1m,t}^{t+189} - f_t^{1m,9m}$
lags	22	9	14	5
Cointegrating vector	1-0.997	1-0.995	1-0.990	1-0.973
portmanteau test (21)				
spot rate	0.55	16.18	8.92	5.48
forward rate	19.29	22.80	6.86	11.43
ARCH 1-1				
spot rate	32.45 **	22.93 **	19.63 **	13.23 **
forward rate	61.36 **	22.95 **	1.74	8.53 **
Hetero-test				
spot rate	1.08	0.96	1.11	1.40
forward rate	1.84 *	1.37	0.88	0.93
Normality -test				
spot rate	1454.8 **	1233.60 **	1333.70 **	1458.60 **
forward rate	299.71 **	191.32 **	64.54 **	31.75 **

Note: * and ** denote the rejection of the null hypothesis at 5% and 1% significance levels respectively. The Portmanteau test follows a χ^2 distribution with m=21 degrees of freedom. The critical value of the test is 32.7 and 38.9 respectively at 5% and 1% significance level. These results were compared to the critical values reported in Hamilton (1994).

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