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The Expectations Hypothesis of the Term Structure of LIBOR US Dollar Interest Rates[†]

A b s t r a c t. Using the monthly sampled data on LIBOR US dollar interest rates and maturities ranging from 1 to 12 months from 1995 to 2009 we provide with a number of tests of the expectations hypothesis based on a 3-variable VAR allowing for a time-varying term premium. We find some evidence against the expectations hypothesis. The term premia appear to vary in time and the yield spread has a good predictive power, however the long rates under-react to current information about future short rates. Unexpected changes in holding period returns to large extent depend upon revisions to forecasts about future short rates and to small extent upon revisions to future term premia.

K e y w o r d s: term structure of interest rates, expectations hypothesis, term premium, LIBOR, VAR.

J E L Classification: E43.

Introduction

The expectations hypothesis (EH) of the term structure of interest rates credited to Fisher (1886, 1930) and Lutz (1940) states that the expected one-period holding period return on a bond that has n periods to maturity (long bond) equals to the return on one-period (short) bond increased by the term premium. If valid it has two important implications: the yield on a long bond (long rate) equals to the average of expected yields on the short bond (short rates) over the life of the long bond plus the rolling-over term premium, and the

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actual yield spread between the long and the short rate is an optimal predictor of the next period's change in the long rate as well as future changes in the short rate.

The early tests of the EH invented by Campbell and Shiller (1991) examine the ability of the yield spread to predict future changes in the short and the long rates. Embedded in either a single equation or VAR setting they provide with a very limited support for the EH when performed on the US and the other data. The long rates appear to move in the opposite direction to that predicted by theory. The short rates move in the correct direction, however the yield spread is their poor predictor at the shorter end of maturity spectrum (see Campbell, Shiller, 1991; Hardouvelis, 1994; Gerlach, Smets, 1997, among many others).

The empirical failure of the EH is explained in a number of ways. It is usually accounted for the existence of a time-varying term premium which is assumed constant in traditional tests. The other explanations include a small sample bias of the EH tests remaining severe in large samples, the over-reaction of long rates to current short rates as well as the asset pricing anomaly disappearing once it is widely recognized to the public (Tzavalis, Wickens, 1997; Bekaert et al., 1997; Garganas, Hall, 2011; Bulkley et al., 2011). It is also stressed that the predictive power of the yield spread depends upon monetary policies implemented by the central bank being much stronger at the times of monetary targeting than interest rates smoothing (see Mankiw, Miron, 1986; McCallum, 2005, among many others).

In this paper we report on that whether the LIBOR US dollar interest rates behave according to the EH. We assume that the term premium vary over time and nest the analysis within a 3-variable VAR of Tzavalis and Wickens (1997). We estimate it on the monthly sampled data from 1995 to 2009. In doing so we use maturities ranging from 1 to 12 months. The data come from Thomson Reuters¹. To test for the time-varying term premium, the ability of the yield spread to predict future changes in the short rate and the link between the current yield spread and that predicted from the VAR we set restrictions on its parameters and statistics. We provide with some evidence against the EH. The results reported in the paper complement those of Hurn et al. (1995) and Miłobędzki (2010) who analyzed the LIBOR interest rates in sterling and using the VAR methodology found much support for the EH at the whole maturity spectrum.

The remainder of the paper proceeds as follows. Section 1 introduces the EH of the term structure of interest rates and shows its implications. Section 2 reviews the VAR based tests of the EH allowing for the time-varying term premium. Section 3 discusses our empirical findings. The last section briefly concludes.

¹ The data are supplied under the agreement between Thomson Reuters Poland and the University of Gdańsk.

1. EH of the Term Structure of Interest Rates and Its Implications

The EH of the term structure of interest rates may be formally stated as:

$$E_t h_{t+1}^{(n)} = E_t \left[\ln P_{t+1}^{(n-1)} - \ln P_t^{(n)} \right] = R_t^{(1)} + \theta_t^{(n)}, \quad (1)$$

where $P_t^{(n)}$ is the price at time t of pure discount bond with face value of \$1 and n periods to maturity, $R_t^{(1)}$ is the certain (riskless) one-period interest rate, and $\theta_t^{(n)}$ is a term premium which compensates for the risk of investing in long bonds. The term premium is admitted to vary in time but presumed to be a stationary random variable. Variants of the EH include the pure ($\theta_t^{(n)} = 0$, PEH), constant ($\theta_t^{(n)} = const$, CEH) and liquidity preference versions ($\theta_t^{(n)} > \theta_t^{(n-1)} > \dots > \theta_t^{(2)}$, LPEH) (for all t and n). The term premium, $\theta_t^{(n)}$, according to Eq. (1), is reflected by the expected excess one-period holding period return, $E_t h_{t+1}^{(n)} - R_t^{(1)}$.

To demonstrate implications of the EH for the interest rates the following is usually undertaken (Campbell, Shiller, 1991; Cuthbertson, 1996; Tzavalis, Wickens, 1997; Cuthbertson, Nitzsche, 2003). Firstly, a continuous compounding is assumed, i.e. $\ln P_t^{(n)} = -nR_t^{(n)}$, where $R_t^{(n)}$ is the spot yield on the long bond. Then some manipulations of Eq. (1) result in:

$$R_t^{(n)} = (1/n) \sum_{i=0}^{n-1} E_t R_{t+i}^{(1)} + \Theta_t^{(n)}, \quad (2)$$

where $\Theta_t^{(n)} = (1/n) \sum_{i=0}^{n-1} E_t \theta_{t+i}^{(n-i)}$. Subtracting $R_t^{(1)}$ from both sides of Eq. (2) and rearranging terms yield:

$$S_t^{(n)} = E_t \sum_{i=1}^{n-1} (1-i/n) \Delta R_{t+i}^{(1)} + \Theta_t^{(n)}. \quad (3)$$

Eq. (3) shows that the observed yield spread should equal to the sum of the optimal forecast of future changes in the short rate, $E_t \sum_{i=1}^{n-1} (1-i/n) \Delta R_{t+i}^{(1)}$, and the average of term premia expectations, $\Theta_t^{(n)}$ (rolling-over term premium). Thus it can be concluded that at time t no other information apart from that contained in both variables should help predict future changes in the short rate.

The immediate consequence of the latter is twofold: $S_t^{(n)}$ should Granger cause $\Delta R_{t+i}^{(1)}$, and in the case term premium $\theta_t^{(n)}$ is not time-varying the expected excess one-period holding period return is constant and should not depend upon its past values as well as past values of the actual spread and changes in the future short rate.

Last but not least, substituting Eq. (2) into the unanticipated change (‘surprise’) in the one-period holding period return, $eh_{t+1}^{(n)} = h_{t+1}^{(n)} - E_t h_{t+1}^{(n)}$, gives (Tzavalis, Wickens, 1997):

$$eh_{t+1}^{(n)} = -(E_{t+1} - E_t) \sum_{i=1}^{n-1} R_{t+i}^{(1)} - (E_{t+1} - E_t) \sum_{i=1}^{n-1} \theta_{t+i}^{(n-i)} = -[eR_{t+1}^{(1)} + e\theta_{t+1}^{(n)}], \quad (4)$$

where $eR_{t+1}^{(1)} = (E_{t+1} - E_t) \sum_{i=1}^{n-1} R_{t+i}^{(1)}$ exhibits the ‘news’ about future short rates and $e\theta_{t+1}^{(n)} = (E_{t+1} - E_t) \sum_{i=1}^{n-1} \theta_{t+i}^{(n-i)}$ exhibits the ‘news’ about future term premia. Hence, unanticipated change in the one-period holding period return must be due to either a revision to expectations about future short rates or a revision to expectations about future term premia.

2. Three-variable VAR Based Tests of the EH

The VAR based tests of the EH solving for the time-varying term premium hinge on the extended 2-variable VAR of Campbell and Shiller (1991) in which the yield spread, $S_t^{(n)}$, and the change in the short rate, $\Delta R_t^{(1)}$, are supplemented by the excess one-period holding period return, $h_t^{(n)} - R_{t-1}^{(1)}$. Such a VAR of order p with vector $Z_t^* = [S_t^{(n)} \Delta R_t^{(1)} h_t^{(n)} - R_{t-1}^{(1)}]'$ containing stationary variables is stacked into companion form as a first order VAR (see Tzavalis, Wickens, 1997; Cuthbertson, Bredin, 2001; Cuthbertson, Nitzsche, 2003; Blangiewicz, Miłobędzki, 2008):

$$Z_t = AZ_{t-1} + u_t, \quad (5)$$

where A is a square $(3p \times 3p)$ matrix of coefficients, Z_t is a $(3p \times 1)$ vector of regressors like $Z_t = [S_t^{(n)} \dots S_{t-p+1}^{(n)} \Delta R_t^{(1)} \dots \Delta R_{t-p+1}^{(1)} h_t^{(n)} - R_{t-1}^{(1)} \dots h_{t-p+1}^{(n)} - R_{t-p}^{(1)}]'$, and u_t is a $(3p \times 1)$ vector of errors. Variables included in the VAR can be picked up from the system using $(3p \times 1)$ selection vectors $e1'$, $e2'$ and $e3'$ with unity in the first, second and third row, respectively, and zeros elsewhere so that $S_t^{(n)} = e1'Z_t$, $\Delta R_t^{(1)} = e2'Z_t$ and $h_t^{(n)} - R_{t-1}^{(1)} = e3'Z_t$. Their predictions from the VAR can be computed throughout the chain rule of forecasting as:

$$E(Z_{t+k} | Z_t) = A^k Z_t. \quad (6)$$

The tests of interest verify whether the excess one-period holding period return is not time-varying, what the sources of ‘surprise’ in its performance are (if there are any), as well as whether long rates properly react to current information about future short rates. Construction of the appropriate test statistics is based on the assumption that predictions from the VAR system are adequate.

The prediction of the expected excess one-period holding period return from the VAR is $E_t h_{t+1}^{(n)} - R_t^{(1)} = e3'Z_{t+1} = e3'AZ_t$ which in the case of time-invariant term premium should equal to some constant. In terms of the VAR with demeaned variables it requires a set consisted of $3p$ linear restrictions be such that $e3'A = 0$. This is tested with the use of a Wald test. Under the null the relevant test statistics is distributed as χ^2 variable with $3p$ degrees of freedom.

The prediction of the yield spread from the VAR ('theoretical spread') is (Cuthbertson et al., 2000):

$$S_t^{*(n)} = E_t \sum_{i=1}^{n-1} (1-i/n) \Delta R_{t+i}^{(1)} = e2' \Lambda Z_t, \tag{7}$$

where $\Lambda = A \left[I - (1/n)(I - A^n)(I - A)^{-1} \right] (I - A)^{-1}$. It should track the actual spread, $S_t^{(n)} = e1'Z_t$, provided expectations about the future term premia, $E_t \theta_{t+i}^{(n-i)}$, are constant over time. In such circumstances it is expected that $S_t^{*(n)} = S_t^{(n)}$, which implies the following set of VAR metrics:

$$e1' - e2' \Lambda = 0, \tag{8}$$

$$VR = \sigma^2 \left[S_t^{*(n)} \right] / \sigma^2 \left[S_t^{(n)} \right] = 1. \tag{9}$$

$$\rho = corr \left[S_t^{*(n)}, S_t^{(n)} \right] = 1, \tag{10}$$

where VR and ρ stand for a variance ratio and a correlation coefficient, respectively.

The set of nonlinear cross-equation restrictions from Eq. (8) can be tested for with the use of a Wald test. The relevant test statistics is:

$$W = f(a)' \times \left[\frac{\partial f(a)}{\partial a'} \hat{\Sigma}_{aa} \frac{\partial f(a)}{\partial a} \right]^{-1} \times f(a), \tag{11}$$

where $f(a) = e1' - e2' \Lambda = 0$ and $\hat{\Sigma}_{aa}$ is either the standard or the Eicker-White heteroscedasticity consistent variance-covariance matrix of VAR parameters estimator. Under the null (and standard properties of error term u_t) it is distributed as the χ^2 variable with $3p$ degrees of freedom.

To proceed with the metrics contained in Eq. (9) and (10) it is worth noting that under the PEH the series of theoretical and actual spread should move together. A high degree of co-movement indicates that variation in the spread is mainly due to rationally expected changes in future short rates with no or only minor variation in the premia (Engsted, 1996). The validity of PEH can be informally deduced plotting $S_t^{*(n)}$ versus $S_t^{(n)}$, while it can be more formally evaluated using the above two metrics. Since $\beta = \rho / \sqrt{VR}$ is the OLS estimator

of the slope in the regression of actual spread onto theoretical spread, which should also be unity, both the numerator and denominator should be close to unity or one of them must be approximately the inverse of the other. Thus the rejection of $\beta=1$ is to be accounted for either the over-reaction (under-reaction) hypothesis or the presence of the time-varying term premium. If $VR < (>) 1$ and $\rho \approx 1$, then the slope would be more (less) than unity and the actual spread is more (less) volatile than the theoretical spread, the optimal predictor of future short rates. Hence, although there is a strong relationship between $S_t^{*(n)}$ and $S_t^{(n)}$, the long rate is over-reacting (under-reacting) to current information about future short rates. In the case neither are close to unity, the actual spread behaves differently from the theoretical spread and the over-reaction (under-reaction) could be the consequence of a time-varying term premium (Campbell, Shiller, 1991; Hardouvelis, 1994).

Bekaert et al. (1997), Bekeart and Hodrick (2001) and Garganas and Hall (2001) show that a 2-variable VAR based tests of the EH with the exception of $\text{corr}[S_t^{*(n)}, S_t^{(n)}]$ are biased in small samples in the case the short rate is persistent. The Wald test tends to over-reject the null, while the variance bound ratio favours it too often. The bias in these tests increases with the degree of short rate persistence.

The rejection of the EH may be also due to noise traders. Their excessive activity relative to that of smart money traders increases interest rates volatility which results in a downward bias of all VAR metrics (Cuthbertson et al., 1996).

We are now to assess what a portion of ‘surprise’ in the one-period holding period return, $eh_{t+1}^{(n)} = h_{t+1}^{(n)} - E_t h_{t+1}^{(n)}$, is due to the ‘news’ about future short rates, $eR_{t+1}^{(1)}$, and the ‘news’ about future term premia, $e\theta_{t+1}^{(n)}$. Such a decomposition is based on residuals from the VAR system. To see this note that the error from the second VAR equation, $u_{2,t+1} = e2'u_{t+1}$, represents the ‘surprise’ in the future change of the short rate, while the error from its third equation, $u_{3,t+1} = e3'u_t -$ the ‘surprise’ in the excess one-period holding period return. Since (see Tzavalis, Wickens, 1997):

$$eR_{t+1}^{(1)} = (E_{t+1} - E_t) \sum_{i=1}^{n-1} R_{t+1}^{(i)} = (E_{t+1} - E_t) \left[(n-1)R_t^{(1)} + \sum_{i=1}^{n-1} \sum_{j=1}^i \Delta R_{t+j}^{(1)} \right] = \\ (E_{t+1} - E_t) \sum_{i=1}^{n-1} \sum_{j=1}^i \Delta R_{t+j}^{(1)}, \quad (12)$$

the ‘surprise’ in the term premia can be calculated from:

$$e\theta_{t+1}^{(n)} = -eR_{t+1}^{(1)} - eh_{t+1}^{(n)} = \quad (13)$$

$$e2' \left[(n-1)I + (n-2)A + (n-3)A^2 + \dots + (n-(n-1))A^{n-2} \right] u_{t+1} - u_{3,t+1},$$

using the appropriate VAR residuals. The first term on the right hand side

of Eq. (13) stands for the weighted sum of the ‘surprises’ in future short rates so that matrices A^s exhibit the degree of persistence in the ‘news’ about future short rates ($s = 1, 2, \dots, n - 2$).

Suppose further that a revision to expectations about future term premia are negligible ($e\theta_{t+1}^{(n)} \approx 0$). This yields $eh_{t+1}^{(n)} \approx -eR_{t+1}^{(1)}$, and the following metrics also apply:

$$\sigma^2[eR_{t+1}^{(1)}] / \sigma^2[eh_{t+1}^{(n)}] \approx 1, \tag{14}$$

$$\text{corr}[eR_{t+1}^{(1)}, eh_{t+1}^{(n)}] \approx -1. \tag{15}$$

In addition, from Eq. (1) and (4) we obtain:

$$h_{t+1}^{(n)} - R_t^{(1)} = \theta_t^{(n)} - eR_{t+1}^{(1)} - e\theta_{t+1}^{(n)}. \tag{16}$$

Hence we can conclude that $(1 - R^2)$ of the one-period holding period return equation in the VAR system indicates a proportion of the excess one period holding period return that is due to variation in the ‘news’ about future short rates.

3. Empirical Results

Since interest rates are believed to be integrated of order one variables the use of the VAR based tests in the applied work is limited to cases in which all variables in the VAR system (actual yield spread, change in the short rate, ex-post excess one-period holding period return) are stationary². This is to be empirically confirmed, however. Hence the analysis sets off with testing for (non) stationarity of the individual US dollar LIBORs and the variables entering the VAR. For testing purposes we employ the DF-GLS and KPSS tests (see Elliot et al., 1996; Kwiatkowski et al., 1992). Their results (available to readers upon a request) prove that the variables in question are integrated of order zero.

The results from the VAR models are stacked in Table 1 (see Appendix). VAR order p for each maturity is set with the use of Schwarz information criterion but occasionally increased to remove autocorrelation in residuals³. The

² In such circumstances a 3-variable VAR of Tzavalis and Wickens (1997) implies a vector error correction model with the yield spreads and excess one-period holding period return being the co-integrating vectors; see Appendix C in King and Kurmann (2002) for details regarding a 2-variable VAR of Campbell and Shiller (1991).

³ There is some unclear picture of autocorrelation for the yield spread and the change in short rate equations (for $n = 4$ and $n = 12$, respectively). The estimates of the Breusch-Godfrey test statistics used to test for no-autocorrelation of up to the 12-th order are just equal to the critical value of the F variable with 12 and $T - (3p + 1) - 12$ degrees of freedom, while the estimates of the Ljung-Box test statistics for these maturities are far away from the critical value of the χ^2 variable with 12 degrees of freedom at the conventional 5 per cent significance level. We are not able

first two equations in the system have from a relatively moderate to large explanatory power as reflected by their coefficient of determination R^2 estimates. Nevertheless quite a lot of unexplained variation in the ex-post excess one-period holding period return equation is left to be attributed to a revision to the expectations about future short rates and future term premia (estimates of $1 - R^2$ in the third equation range from 0.58 to 0.81). In addition, the estimates of Granger non-causality test statistics prove the ability of the yield spread to predict future changes in the one-month US dollar LIBOR.

Table 2 (see Appendix) reports the results of testing for the EH using the restrictions set on the VAR parameters and other metrics. Restriction $e3'A = 0$ is rejected for all maturities so that we suspect the term premia are time-varying.

Turning now to the VAR metrics, a graph of the actual and theoretical spread for both $n = 4$ and 12 show their rather poor correspondence over time with some visual evidence of under-reaction of the actual spread relative to the expected changes in future short rates (see Fig. 1-2, right panels, Appendix). The same somewhat poor correspondence is apparent when the first spread is scattered versus the latter (see Fig. 1-2, left panels, Appendix). Empirical points on these panels are much dispersed along the straight 45-degree line indicating that for both maturities correlation between $S_t^{(n)}$ and $S_t^{*(n)}$ may substantially differ from one. The null stating that $S_t^{*(n)} = S_t^{(n)}$ is also rejected at 5 per cent significance level for all interest rates but not for the 12-month US dollar LIBOR (in this case it is rejected at 10 per cent significance level) which assures that the term premia are time-varying.

A more formal measures of the relationship between the actual and theoretical spread to much extent support the under-reaction hypothesis. While for all maturities the VR estimates does not depart from unity by more than its 2 standard deviations (the relevant 95 per cent confidence interval obtained from the bootstrap covers unity in all cases apart from those of $n = 8, 10$ and 11), the correlation coefficient estimates are less than unity by more than its two standard deviations for all maturities except $n = 12$.

Given the result that a lot of unexplained variation in the ex-post excess one-period holding period return equation is due to a revision to the expectations about future short rates and future term premia (see estimates of $1 - R^2$ from the third equation of the VAR in Table 1, Appendix) we are to evaluate the size of their contribution to the overall effect. Formally, the estimates of $\rho[eR_{t+1}^{(1)}, eh_{t+1}^{(n)}] \approx -1$ and for all maturities they do not differ from

to remove autocorrelation without over-parametrizing the system. Hence, for these two maturities some caution should be retained when further predictions about the theoretical spread as well as predictions based upon all VAR metrics employing the change in short rate are made.

minus unity by more than its 2 standard deviations, and those of $\sigma^2 [eR_{t+1}^{(1)}] / \sigma^2 [eh_{t+1}^{(n)}]$ are close to but above unity and slightly differ from that by more than its 2 standard deviations (its 95 per cent confidence interval from the bootstrap does not cover unity for all maturities except the 11-month US dollar LIBOR). This indicates that a portion of ‘surprise’ in the one-period holding period return due to ‘news’ about future term premia for all n is not negligible, however small.

Taking into account the above findings we can conclude that the evidence we have gathered against the EH in the London interbank market is strong enough to reject it due to under-reaction of long rates to current information about the future short rate. On the other hand, as Campbell and Shiller (1991) have argued, rejection of the cross-equation parameter restrictions is not a final argument against the EH on economic grounds as long as the theoretical spread closely tracks the actual spread. Out of all VAR metrics we primarily trust correlation coefficient ρ which properties are not much distorted by the short rate persistence and for that its estimates substantially depart from unity for all maturities. This is supported by the variance ratio metrics for some rates at the longer end of maturity spectrum.

The rejection of the EH due to under-reaction could be totally erroneous, however. If agents use the VAR methodology for forecasting purposes, when forming expectations about the future short rates are expected to utilize information on a more frequent basis (minute-to-minute, hourly, daily; see Cuthbertson et al., 1996). In addition, large banks can meet their longer-term needs for monies at a lower cost outside London. Information regarding interest rates of both origins is not exhibited in our data set so that the predictions we have obtained from the VAR system might be heavily biased. In particular, using the theoretical spread we could substantially underestimate agents’ expectations about the future short rates. Our predictions could also poorly track the true expectations.

Conclusions

In this study using the monthly sampled data on LIBOR US dollar interest rates from 1995 to 2009 and a wide range of maturities we find a rather conclusive evidence against the EH. However the term premia appear to vary in time and the yield spread has a good predictive power, the long rates under-react to current information about the future short rates. Unexpected changes in the holding period returns to a large extent depend upon revisions to forecasts about the future short rates and to a small extent upon revisions to the future term premia.

The results reported in this paper are in some respects in contrast to those of Hurn et. al. (1995), Engsted (1996), Cuthbertson (1996), Cuthbertson et al. (1996), Blangiewicz and Miłobędzki (2009, 2010) and Miłobędzki (2010) who

analyzed the term structure of interest rates at the Danish, Polish and the UK money markets with the use of either a 2 or 3-variable VAR and thus provide with the nearest comparison to our work. The main difference revealed in their work is that of a time-invariant term premium which is consistent with the PEH (Hurn et al., 1995; Cuthbertson, 1996; Miłobędzki, 2010 – for pound sterling, Engsted, 1996 – for Danish kroner; Blangiewicz, Miłobędzki, 2010 – for Polish zloty, for all or some maturities), and the main similarity – a strong predictive power of the yield spread (all authors except from Engsted, 1996, for Danish kroner during the period of interest rates smoothing).

Appendix

Table 1. Summary statistics for VAR $Z_t^* = [S_t^{(n)} \quad \Delta R_t^{(1)} \quad h_t - R_{t-1}^{(1)}]'$

n	p	Autocorrelation						R ²			Granger noncausality
		LM(12) ^{a)}			Ljung-Box(12) ^{b)}			S _t ⁽ⁿ⁾	ΔR _t ⁽¹⁾	h _t ⁽ⁿ⁾ - R _{t-1} ⁽¹⁾	
		S _t ⁽ⁿ⁾	ΔR _t ⁽¹⁾	h _t ⁽ⁿ⁾ - R _{t-1} ⁽¹⁾	S _t ⁽ⁿ⁾	ΔR _t ⁽¹⁾	h _t ⁽ⁿ⁾ - R _{t-1} ⁽¹⁾				
3	9	1.43 (0.15)	0.96 (0.49)	1.07 (0.38)	7.00 (0.87)	6.63 (0.88)	7.39 (0.83)	0.69	0.45	0.28	135.69 (0.00)
4	20	1.79 (0.05)	1.56 (0.11)	1.42 (0.16)	3.92 (0.99)	3.68 (0.99)	3.78 (0.99)	0.79	0.53	0.42	96.89 (0.00)
5	7	1.13 (0.34)	0.98 (0.47)	0.94 (0.51)	5.73 (0.93)	9.69 (0.64)	7.35 (0.83)	0.71	0.36	0.20	101.96 (0.00)
6	22	1.76 (0.06)	1.29 (0.23)	1.57 (0.10)	2.58 (0.99)	3.73 (0.99)	3.93 (0.99)	0.8	0.53	0.23	88.09 (0.00)
7	12	1.72 (0.07)	1.18 (0.30)	0.99 (0.46)	3.64 (0.99)	3.98 (0.98)	3.57 (0.99)	0.76	0.41	0.22	70.66 (0.00)
8	18	0.99 (0.46)	1.36 (0.187)	1.33 (0.21)	1.63 (1.000)	3.92 (0.99)	3.86 (0.99)	0.80	0.49	0.38	76.60 (0.00)
9	16	0.85 (0.60)	1.76 (0.06)	1.70 (0.07)	1.97 (0.99)	4.97 (0.96)	3.47 (0.99)	0.80	0.48	0.27	54.61 (0.00)
10	18	0.86 (0.59)	1.08 (0.38)	1.15 (0.33)	2.24 (0.99)	3.41 (0.99)	3.39 (0.99)	0.81	0.46	0.35	52.90 (0.00)
11	17	1.07 (0.39)	1.13 (0.34)	1.02 (0.43)	1.84 (1.00)	3.56 (0.99)	2.32 (0.99)	0.79	0.42	0.30	48.80 (0.00)
12	12	0.68 (0.78)	1.78 (0.05)	1.57 (0.10)	1.226 (1.00)	4.35 (0.98)	3.21 (0.99)	0.78	0.37	0.19	54.18 (0.00)

Note: ^{a)} Estimates of the Breusch-Godfrey [Ljung-Box] test statistics for autocorrelation of order 12 under the null of no-autocorrelation distributed as $F[12, T - (3p + 1) - 12]$ [$\chi^2(12)$], T – number of observations; relevant p -values in brackets under the estimates; ^{b)} Estimates of the Wald test statistics for Granger noncausality from $S_t^{(n)}$ to $\Delta R_t^{(1)}$ under the null distributed as $\chi^2(p)$ variable; relevant p -values in brackets under the estimates.

Table 2. VAR restrictions and other metrics, variance decomposition

n	Excess one period holding period return not time varying $e3'A=0^{a)}$	Actual $S_t^{(n)}$ and theoretical $S_t^{*(n)}$ spread				News about future short rates and one period returns			
		$S_t^{*(n)}=S_t^{(n)}$	$\sigma^2[S_t^{(n)}]$		$\rho[S_t^{*(n)}, S_t^{(n)}]$ b)	$\sigma^2[eR_{t+1}^{(1)}]$		$\rho[eR_{t+1}^{(1)}, eh_{t+1}^{(n)}]$ b)	
			VR	CI		VR	CI		
3	W(27)=90.37 (0.00)	W(27)=93.57 (0.00)	0.91 (0.26)	0.61 1.62	0.71 (0.10)	1.28 (0.08)	1.12 1.45	-0.99 (0.00)	
4	W(60)=138.15 (0.00)	W(60)=182.80 (0.00)	1.22 (0.41)	0.69 2.28	0.60 (0.10)	1.31 (0.14)	1.04 1.60	-0.98 (0.01)	
5	W(21)=60.33 (0.00)	W(21)=50.64 (0.00)	0.90 (0.31)	0.51 1.73	0.72 (0.12)	1.33 (0.15)	1.03 1.63	-0.98 (0.01)	
6	W(66)=278.42 (0.00)	W(66)=161.53 (0.00)	1.28 (0.47)	0.76 2.60	0.58 (0.12)	1.64 (0.23)	1.20 2.10	-0.99 (0.01)	
7	W(36)=114.86 (0.00)	W(36)=73.69 (0.00)	1.14 (0.47)	0.55 2.36	0.70 (0.13)	1.57 (0.24)	1.10 2.03	-0.98 (0.02)	
8	W(54)=177.26 (0.00)	W(54)=135.30 (0.00)	1.82 (0.75)	1.09 3.96	0.52 (0.09)	1.50 (0.29)	1.02 2.00	-0.96 (0.03)	
9	W(48)=161.83 (0.00)	W(48)=103.48 (0.00)	1.42 (0.61)	0.67 3.03	0.64 (0.13)	1.58 (0.27)	1.04 2.10	-0.99 (0.02)	
10	W(54)=139.17 (0.00)	W(54)=96.06 (0.00)	1.99 (0.90)	1.48 4.80	0.51 (0.08)	1.53 (0.26)	1.04 2.08	-0.99 (0.02)	
11	W(51)=89.64 (0.01)	W(51)=123.33 (0.00)	2.14 (1.46)	1.36 6.83	0.50 (0.11)	1.41 (0.24)	0.96 1.92	-0.99 (0.01)	
12	W(36)=52.23 (0.04)	W(36)=49.85 (0.06)	1.41 (0.69)	0.60 3.24	0.74 (0.15)	1.70 (0.35)	1.03 2.36	-0.99 (0.01)	

Note: ^{a)} Relevant p -values in brackets under the Wald test statistics estimates; ^{b)} ρ – linear correlation coefficient. Relevant standard errors from the bootstrap under the variance ratio VR and correlation coefficient ρ estimates. CI – 95 per cent confidence interval from the bootstrap.

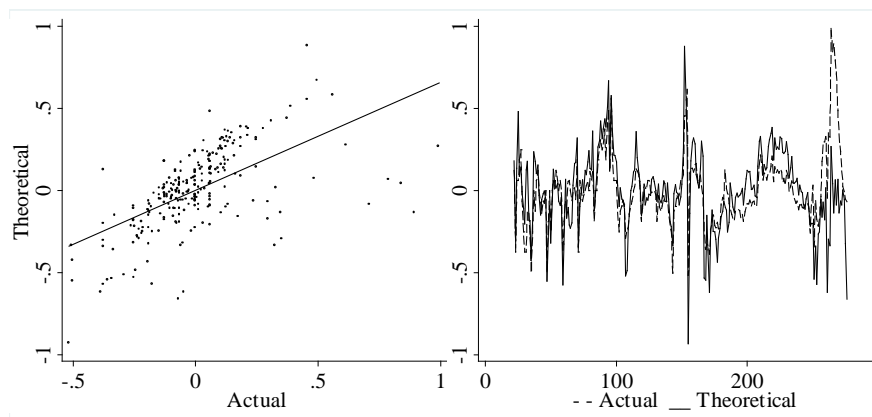


Figure 1. Actual and theoretical spread (4 vs. 1-month US dollar LIBOR)

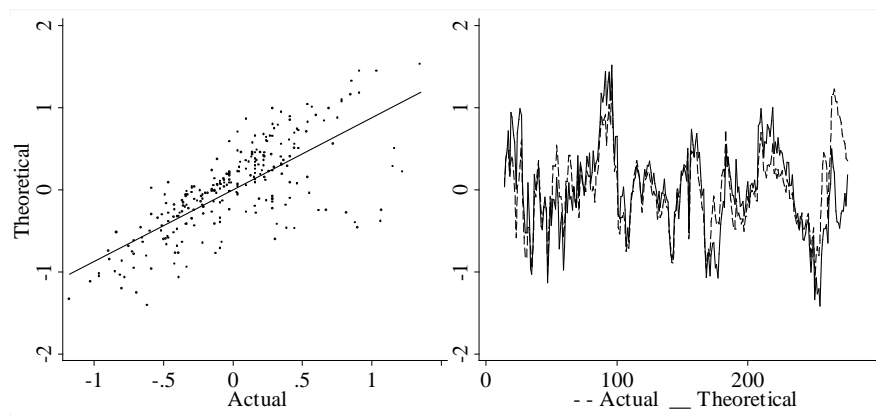


Figure 2. Actual and theoretical spread (12 vs. 1-month US dollar LIBOR)

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Hipoteza oczekiwań struktury terminowej stóp procentowych LIBOR dla dolara USA

Zarys treści. W artykule przedstawia się wyniki testów hipotezy oczekiwań struktury terminowej stóp LIBOR dla dolara USA opartych na 3-wymiarowym modelu VAR. Model ten oszacowano na podstawie miesięcznych szeregów czasowych stóp procentowych z lat 1995-2009 i zapadalności od 1 do 12 miesięcy. Zależono kilka przesłanek świadczących przeciwko tej hipotezie. Choć premie płynności okazały się być zmiennymi w czasie, a spready stóp procentowych – mieć silne własności prognostyczne, niemniej stopy długie w niedostateczny sposób reagowały na bieżące informacje odnośnie do przyszłych stóp krótkich. Niespodziewane zmiany w okresowych stopach zatrzymania były w dużej mierze spowodowane rewizjami prognoz przyszłych stóp krótkich, a tylko w skromnej mierze rewizjami prognoz przyszłych premii płynności.

Słowa kluczowe: struktura terminowa stóp procentowych, hipoteza oczekiwań, premia płynności, LIBOR, VAR.

