Has the Predictability of the Yield Spread Changed?

Dong Heon Kim and Euihwan Park

This paper examines the stability of the predictive power of the yield spread for future GDP growth. We find that the ability of the spread to predict future GDP growth has weakened since 1984:Q1. Given the decomposition of the yield spread into the expectation component and the term premium component, we investigate the change in the predictability of both components and find that the term premium component appears to have lost the predictive power significantly while the predictive power of the expectation component has remained. We conjecture that since the 1984:Q1, the cyclical movement of the term premium seems to have been reduced due to the significant reduction in the volatility of US macroeconomy.

Keywords: Yield spread, Break, Predictability, Expectations effect, Term premium effect, Great Moderation

JEL Classification: E32, E43, C53

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I. Introduction

A large literature has shown that the yield spread between the long- and short-term interest rates is useful for forecasting future economic activity. Examples include Harvey (1988, 1989), Estrella and Hardouvelis (1991), Plosser and Rouwenhorst (1994), Haubrich and Dombrosky (1996), Dueker (1997), Estrella and Mishkin (1998), Hamilton and Kim (2002), Estrella *et al.* (2003), Estrella (2005), Ang *et al.* (2006), Cho and Lee (2014), and among others.

Recently, however, there were some evidence on the instability of the predictive relationships. Chauvet and Potter (2002, 2005), provide an evidence that there was a break in 1984 in the stability of recession forecasts of the yield spread and Bordo and Haubrich (2008) find that the predictive ability of the yield spread was less accurate during 1985–1997 than before. Nevertheless, Estrella *et al.* (2003) show that the predictive relationships between the yield curve and subsequent real activity are stable in both Germany and the United States. Thus, there is controversial on the stability of the predictive power of the yield spread for future real economic activity.

The purposes of this paper are twofold. First of all, we examine whether the predictive ability has weakened along with rigorous methodology for the structural break test. Secondly, if so, we try to explain why the predictability changed. We find using Bai and Perron (1998)'s multiple structural break test that the predictive power of the yield spread for future real GDP growth has declined since 1984:Q1 at all forecasting horizons. Following Hamilton and Kim (2002), we decompose the spread into the expectation component and the term premium and find that the term premium component appears to have lost the predictive power significantly while the predictive power of the expectation component has remained.

The structure of this paper is as follows. In section II, we briefly explain Bai and Perron (1998)'s methodology and present estimation results. In section III, we estimate the change in the predictive power of both the expectation component and the term premium over the subsample. The concluding remark is provided in Section IV.

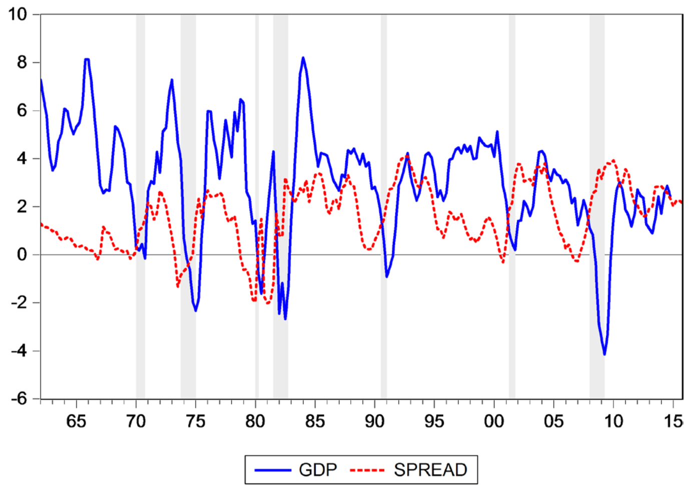

Figure 1
The 4-Quarter Growth Rate of Real GDP and the Yield Spread.
The Shaded are NBER Recession Dates

II. Structural Break Test

A. Predictive Regression

In this paper, we use 10-year Treasury bond rate, 3-month Treasury bill rate and real GDP growth rate from 1962:Q1 to 2017:Q4. Real GDP growth rate data are from FRB of St Louis FRED database and interest rates are 3-month Treasury bill rate from FRB of St Louis FRED and 10-year zero-coupon bond yield from Gurkaynak *et al.* (2007). Figure 1 displays the spread between 10-year bond yield and 3-month bill rate and 4-quarter real GDP growth rate. The Shaded areas indicate NBER recession dates. The figure indicates that the significant decrease in the yield spread appears to have preceded every recessions although the magnitude and the timing in the decrease of the yield spread seem to be different among different recessions.

However, we cannot clearly identify that the predictive power of the yield spread has changed only using this figure. In order to investigate the change in the statistical correlation between the yield spread and future real GDP growth rate, we consider the predictive regression as follows:

$$y_t^k = \alpha_0 + \alpha_1 spread_t + \varepsilon_t \tag{1}$$

$$y_t^k = \left(\frac{400}{k}\right) * \left(\ln Y_{t+k} - \ln Y_t\right) \tag{2}$$

$$spread_t = i_t^{40} - i_t^1 \tag{3}$$

where Y_{t+k} is real GDP in quarter $t + k$, y_t^k is the annualized real GDP growth over the next k quarters, i_t^{40}, i_t^1 are the ten-year Treasury bond rate and the three-month Treasury bill rate at time t. We use equation (1) to test whether the predictive power of the yield spread changed based on the coefficient α_1.

B. Break Test

Following Bai and Perron (1998), we consider the following equation:

$$y_t = x_t'\beta + z_t'\gamma_j + u_t, \quad t = T_{j-1} + 1, \dots, T_j, \quad j = 1, \dots, m, m+1 \tag{4}$$

where y_t is observed dependent variable at time t, x_t, and z_t are $(p \times 1)$ and $(q \times 1)$ independent vectors of covariates, β, γ_j are corresponding vector of coefficients. The equation (4) indicates that there are m unknown break points and our objective is to estimate unknown m break points $(T_1, T_2, \ldots, T_m)$ and coefficients γ_j. This case is called as partial structural break test and if x_t are zero vector, we call a pure structural break test.

The estimation strategy is based on least-squares. For each m-partitioned $(T_1, T_2, \ldots, T_m)$, the associated least-square estimates of β and γ_j are obtained by minimizing the sum of squared residuals as follows:

$$S_T\left(T_1, T_2, \dots, T_m\right) = \sum_{i=1}^{m+1} \sum_{t=T_{i-1}+1}^{T_i} \left(y_t - x_t'\beta - z_t'\gamma_i\right)^2 \tag{5}$$

We can estimate the break points minimizing the $S_T(T_1, T_2, \ldots, T_m)$, namely:

$$(\hat{T}_1, \hat{T}_2, \dots, \hat{T}_m) = \arg \min_{T_1,\dots,T_m} S_T\left(T_1, T_2, \dots, T_m\right) \tag{6}$$

Then, we apply this methodology to equation (1) in order to detect how many breaks are. Bai and perron (1998) propose a test that the null hypothesis of l breaks against the alternative (l + 1) breaks. The test is applied to each segment containing the observations ($\hat{T}_1$, $\hat{T}_2$, …, $\hat{T}_{l+1}$). We conclude for a rejection in favor of a model with (l + 1) breaks if the overall minimal value of the sum of residuals is sufficiently smaller than the sum of squared residuals from the l breaks model. More specifically, the test is defined as follows:

$$F_T(l+1\mid l) = \{S_T(\hat{T}_1, \hat{T}_2, \dots, \hat{T}_m) - \min_{1\le i\le l+1} \inf_{\tau\in\Lambda_{i,\eta}} S_T(\hat{T}_1, \dots, \hat{T}_{i-1}, \tau, \hat{T}_i, \dots, \hat{T}_l)\} / \hat{\sigma}^2 \quad (7)$$

where, $\Lambda_{i,\eta} = \{\tau; \hat{T}_{i-1} + (\hat{T}_i - \hat{T}_{i-1})\eta \le \tau \le \hat{T}_i - (\hat{T}_i - \hat{T}_{i-1})\eta\}$ and $\hat{\sigma}^2$ is consistent estimate of σ^2 under the null hypothesis.

C. Break Test Results

We estimate equation (1) for the full sample of 1962:Q1 - 2017:Q4 and investigate the stability of the coefficient on the spread, α_1 by using Bai and Perron (1998)'s multiple structural break test.[1]

We set the trimming value that means a minimal length of a segment as 0.2 and the maximum break point l as 3.[2]

The estimation results show that the estimated coefficient on the spread is statistically significant over 1 – 8 quarters forecasting horizons confirming the results of existing literature such as Estrella and Mishkin (1998) and Hamilton and Kim (2002). The predictive ability of the yield spread, however, appears to be weak as the values of R^2 are lower in most forecasting horizons than those of Hamilton and Kim (2002).

The fifth and sixth columns of the Table 1 show the estimated break date and the test statistics respectively. At 1-4 quarters forecasting horizon, the number of breaks is estimated one break and the break

[1] We perform the partial break test on the α_0 and α_1 in equation (1), which are specified as β and γ_j in equation (4) respectively.

[2] A distinct advantage of the Bai and Perron (1998) test is that they directly take into account potential serial correlation in the errors and heterogeneity across segments. If serial correlation and heterogeneity in the data or error are allowed in the estimated regression, Bai and Perron (2003a) recommend setting the trimming value as 0.2. The critical values for this test can be obtained from Bai and Perron (2003b).

TABLE 1
THE PREDICTIVE POWER OF THE YIELD SPREAD AND SEQUENTIAL TEST FOR MULTIPLE STRUCTURAL BREAKS
$$y_t^k = \alpha_0 + \alpha_1 spread_t + \varepsilon_t$$

k (forecasting horizons)	α_0	α_1	$\bar{R}^2$	estimated break dates	$SupF_T(l+1\|l)$ statistic
1	2.362*** (0.482)	0.381* (0.217)	0.020	1984:q1	$SupF_T(1\|0)$ = 24.923*** $SupF_T(2\|1)$ = 4.681
2	2.221*** (0.485)	0.477** (0.216)	0.052	1984:q1	$SupF_T(1\|0)$= 45.888*** $SupF_T(2\|1)$ = 2.712
3	2.213*** (0.464)	0.489** (0.202)	0.068	1984:q1	$SupF_T(1\|0)$= 39.264*** $SupF_T(2\|1)$ = 2.849
4	2.217*** (0.450)	0.496*** (0.191)	0.082	1984:q1	$SupF_T(1\|0)$= 42.556*** $SupF_T(2\|1)$ = 7.987
5	2.247*** (0.438)	0.486*** (0.181)	0.089	1983:q4 2005:q3	$SupF_T(1\|0)$= 56.863*** $SupF_T(2\|1)$ = 12.582**
6	2.305*** (0.424)	0.458*** (0.169)	0.088	1984:q1 2005:q2	$SupF_T(1\|0)$= 31.458*** $SupF_T(2\|1)$ = 25.406***
7	2.366*** (0.409)	0.425*** (0.158)	0.084	1984:q1 2005:q2	$SupF_T(1\|0)$ = 15.129*** $SupF_T(2\|1)$ = 16.772***
8	2.446*** (0.398)	0.382** (0.149)	0.074	1984:q1 2004:q4	$SupF_T(1\|0)$= 13.282*** $SupF_T(2\|1)$ = 10.860**

Note: a. In parentheses are Newey and West(1987) HAC standard errors. b. ***, ** and * denote statistically significant at the 1%, 5%, and 10% levels respectively. c. Row k is based on estimation for t = 1962:Q1 through 2017:Q4 – k.

date is the first quarter in 1984 (1984:Q1) at the 1% significant level. The number of breaks is two at 5-8 quarters forecasting horizon but the first break date is 1984:Q1 and significant at 1% level. The simple break test indicates that the predictive power of the yield spread has changed since 1984:Q1.

Based on the break test result, we divide the full sample into two subsamples according to the estimated break date and estimate equation (1) for two subsamples. The estimation results are reported in Table 2.

The estimated coefficients on the yield spread over 1-8 quarters forecasting horizons in the pre-break sample (1962:Q1–1984:Q1) are all statistically significant and $\bar{R}^2$'s are much higher than those of the full sample whereas those in the post-break sample (1984:Q2–2017:Q4) are

TABLE 2
THE PREDICTIVE POWER OF THE SPREAD PRE- AND POST-BREAK
$y_t^k = \alpha_0 + \alpha_1 spread_t + \varepsilon_t$

k (forecasting horizons)	pre-break(1962:q1:1984:q1)			post-break(1984:q2-2017:q4)		
	α_0	α_1	$\bar{R}^2$	α_0	α_1	$\bar{R}^2$
1	2.310*** (0.571)	1.425*** (0.314)	0.164	2.433*** (0.498)	0.095 (0.206)	-
2	2.154*** (0.315)	1.646*** (0.264)	0.342	2.309*** (0.513)	0.167 (0.196)	0.002
3	2.194*** (0.465)	1.620*** (0.225)	0.415	2.212*** (0.542)	0.227 (0.192)	0.012
4	2.249*** (0.432)	1.577*** (0.199)	0.471	2.131*** (0.572)	0.279 (0.192)	0.024
5	2.333*** (0.412)	1.489*** (0.177)	0.479	2.053*** (0.601)	0.332 (0.201)	0.039
6	2.461*** (0.396)	1.345*** (0.162)	0.442	1.960*** (0.628)	0.388* (0.212)	0.059
7	2.583*** (0.370)	1.217*** (0.140)	0.391	1.870*** (0.645)	0.438* (0.224)	0.083
8	2.690*** (0.354)	1.071*** (0.132)	0.342	1.923*** (0.652)	0.423* (0.225)	0.084

Note: a. In parentheses are Newey and West(1987) HAC standard errors. b. *** and * denote statistically significant at the 1% and 10% levels respectively.

statistically significant only over 6-8 quarters forecasting horizons and $\bar{R}^2$'s are substantially lower than the pre-break sample. For example, the highest value of $\bar{R}^2$ is above 47.9% in 5 quarters ahead forecasting horizon in the pre-break sample while that is 8.4% in 8 quarters ahead forecasting horizon in the post-break sample.

For the robustness check, we perform the structural break test using the quarterly industrial production from FRB of St Louis FRED database as output growth and present the estimation results in table 3. The estimation results are very similar with in table 1. There is only one break for the 1-6 forecasting horizons and all break dates are beginning of the 1980s.[3]

[3] As the anonymous referee suggests, we also investigate the predictive power

TABLE 3

THE PREDICTIVE POWER OF THE YIELD SPREAD AND SEQUENTIAL TEST FOR MULTIPLE STRUCTURAL BREAKS

$$IP_t^k = \alpha_0 + \alpha_1 spread_t + \varepsilon_t$$

k (forecasting horizons)	α_0	α_1	$\bar{R}^2$	estimated break dates	$SupF_T(l+1\|l)$ statistic
1	1.495 (0.946)	0.677 (0.434)	0.018	1980:3q	$SupF_T(1\|0)$ = 45.9202*** $SupF_T(2\|1)$ = 1.5626
2	1.227 (0.935)	0.855** (0.414)	0.041	1980:3q	$SupF_T(1\|0)$ = 29.9107*** $SupF_T(2\|1)$ = 1.3686
3	1.149 (0.902)	0.919** (0.380)	0.059	1983:q4	$SupF_T(1\|0)$ = 36.545*** $SupF_T(2\|1)$ = 2.2953
4	1.101 (0.871)	0.962*** (0.353)	0.075	1983:q3	$SupF_T(1\|0)$ = 40.3101*** $SupF_T(2\|1)$ = 0.5320
5	1.088 (0.846)	0.987*** (0.332)	0.089	1983:q3	$SupF_T(1\|0)$ = 54.8671*** $SupF_T(2\|1)$ = 0.8363
6	1.104 (0.820)	0.970*** (0.313)	0.101	1983:3q	$SupF_T(1\|0)$ = 42.6221*** $SupF_T(2\|1)$ = 1.1984
7	1.170 (0.788)	0.932*** (0.290)	0.104	-	$SupF_T(1\|0)$ = 3.7654
8	1.285* (0.756)	0.865*** (0.268)	0.100	-	$SupF_T(1\|0)$ = 3.3161

Note: a. In parentheses are Newey and West(1987) HAC standard errors. b. ***, ** and * denote statistically significant at the 1%, 5%, and 10% levels respectively. c. Row k is based on estimation for t = 1962:Q1 through 2017:Q4 – k.

Therefore, we interpret that the predictive power of the yield spread for future real economic activity has weakened significantly since 1984:Q1. Why has the predictability of the spread declined?

of the yield spread for the industrial production pre- and post-break date. The estimated coefficients on the yield spread over 1-6 quarters forecasting horizons in the pre-break are all statistically significant at the 1% significant level whereas those in the post-break sample are statistically significant only 2, 5 and 6 quarters forecasting horizons at the 10% significant level.

III. Decomposition of the Yield Spread

As in Hamilton and Kim (2002), consider the following definition of the time-varying term premium TP_t:

$$i_t^n = \frac{1}{n}\sum_{j=0}^{n-1} E_t i_{t+j}^1 + TP_t \tag{8}$$

where $E_t\, i^1_{t+j}$ denotes the market's expectation at time t of the value of i^1_{t+j}. Equation (8) can be written

$$i_t^n - i_t^1 = \left(\frac{1}{n}\sum_{j=0}^{n-1} E_t i_{t+j}^1 - i_t^1\right) + TP_t \tag{9}$$

Hamilton and Kim (2002) show that Equation (9) implies that the spread can be decomposed into two terms: the expectation component and the term premium component. Based on the equation (9), we derive the predictive regression as follows:

$$y_t^k = \beta_0 + \beta_1\left(\frac{1}{n}\sum_{j=0}^{n-1} E_t i_{t+j}^1 - i_t^1\right) + \beta_2\left(i_t^n - \frac{1}{n}\sum_{j=0}^{n-1} E_t i_{t+j}^1\right) + \varepsilon_t \tag{10}$$

$$\text{EP} = \frac{1}{n}\sum_{j=0}^{n-1} E_t i_{t+j}^1 - i_t^1 \tag{11}$$

$$TP = i_t^n - \frac{1}{n}\sum_{j=0}^{n-1} E_t i_{t+j}^1 \tag{12}$$

where EP in equation (11) is the difference between expected short-term interest rates over the next n periods and the current short rate and is called the expectations component, and TP in equation (12) is the time-varying term premium. From equations (1) and (10), if a fall in the spread predicts U.S. recessions, it could be either be because (1) a temporarily high short-term rate suggests a coming recession, or (2) a fall in the term premium on long-term bonds relative to short-term bonds suggests an economic recession. Hamilton and Kim (2002) interpret equation (10) as the question that given that the short rate rises relative to the long rate prior to a recession, to what extent this is

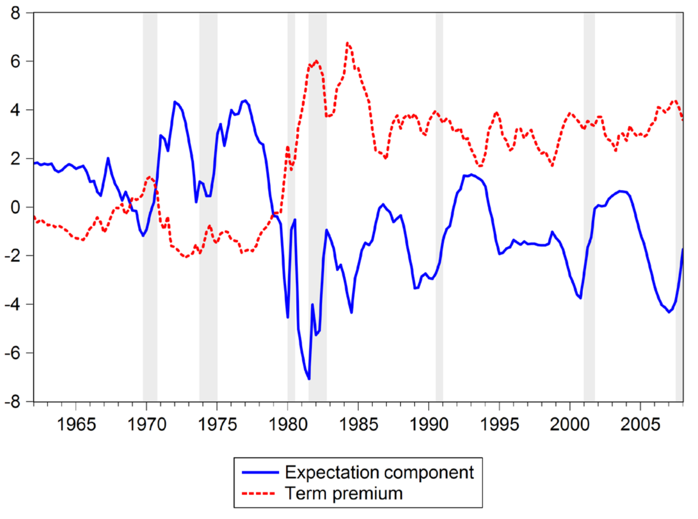

Figure 2
The Expectation Component and the Term Premium.
The Shaded are Nber Recession Dates

because future short rates are rationally expected to fall, and to what extent it is because the forecastable excess yield from holding long-term bonds has fallen.

In order to examine what happened to the expectation component and the term premium component in the yield spread since 1984, we plot the ex-post two components which are constructed by using the ex-post interest rate data in the equations (11) and (12) and they are shown in the Figure 2.

In the case of the expectation component, the average level of the expectation component appears to be higher before 1984 than after 1984 but we evidence that the expectation component tends to decline in advance before each recession over both periods, indicating that the expectation component helps to predict the recessions.

For the term premium component, while the average level of the term premium component is clearly higher before 1984 than after 1984, the cyclical movement of the term premium component appears not to show clear difference between two different periods.[4]

[4] Following the referee's suggestion, we perform the partial and pure structural

TABLE 4

THE PREDICTIVE POWER OF THE EXPECTATION COMPONENT AND THE TERM PREMIUM

$$y_t^k = \beta_0 + \beta_1\left(\frac{1}{n}\sum_{j=0}^{n-1}E_t i_{t+j}^1 - i_t^1\right) + \beta_2\left(i_t^n - \frac{1}{n}\sum_{j=0}^{n-1}E_t i_{t+j}^1\right) + \varepsilon_t$$

(USING AS INSTRUMENTS A CONSTANT, i_t^{40} AND i_t^1)

k (forecasting horizons)	pre-break(1962:1q-1984:1q)				post-break(1984:q2-2008:1q)			
	β_0	β_1	β_2	$\bar{R}^2$	β_0	β_1	β_2	$\bar{R}^2$
1	2.443***	1.400***	0.854	0.168	1.896	0.310*	0.468	-
	(0.627)	(0.433)	(0.569)		(1.445)	(0.177)	(0.432)	
2	2.281***	1.610***	1.082**	0.350	1.807	0.340*	0.502	-
	(0.553)	(0.373)	(0.500)		(1.452)	(0.186)	(0.432)	
3	2.313***	1.583***	1.103**	0.409	1.432	0.374*	0.618	-
	(0.508)	(0.332)	(0.458)		(1.614)	(0.196)	(0.473)	
4	2.364***	1.544***	1.102**	0.446	1.005	0.423**	0.755	-
	(0.473)	(0.296)	(0.424)		(1.793)	(0.204)	(0.526)	
5	2.440***	1.472***	1.065***	0.444	0.698	0.475**	0.862	-
	(0.449)	(0.266)	(0.397)		(1.877)	(0.211)	(0.558)	
6	2.452***	1.353***	0.999***	0.405	0.581	0.525**	0.910	-
	(0.428)	(0.239)	(0.373)		(1.877)	(0.216)	(0.566)	
7	2.659***	1.229***	0.926***	0.357	0.512	0.563**	0.943*	-
	(0.399)	(0.202)	(0.341)		(1.823)	(0.189)	(0.558)	
8	2.785***	1.089***	0.823**	0.292	0.418	0.586**	0.978*	-
	(0.379)	(0.184)	(0.318)		(1.777)	(0.227)	(0.551)	

Note: a. In parentheses are Newey and West (1987) HAC standard errors. b. ***, ** and * denote statistically significant at the 1%, 5%, and 10% levels respectively.

We estimate Equation (10) using instrumental variable estimation along with constant, i_t^{40} and i_t^1 as instruments and assuming the rational expectation. The estimation results are shown in the Table 4.

The estimation coefficients on the EP component (β_1) and the

break test for the equation (10) proposed by Perron and Yamamoto (2015) who deal with the test for structural change in models with endogenous regressors. The results of the pure structural break test show only one break around 1985 only in the 6-quarter ahead forecasting horizon while partial structural break test results show that the breaks were around early 1970s in both components. Since both tests show the mixed results, we consider the structural break test results suggested in the section II.

TP component (β_2) are statistically significant most of 1-8 quarters ahead forecasting horizons in the pre-break sample whereas only the estimated coefficients on the EP component are statistically significant over 1-8 quarters ahead forecasting horizon in the post-break sample. In particular, the coefficients on the TP components are only statistically significant over 7-8 quarters in the post-break sample. In Hamilton and Kim (2002), the TP component was helpful for forecasting future real economic activity over 1-8 quarters ahead.

These estimation results imply that the decrease in the predictive power of the yield spread for the future real economic activity mainly results from the significant reduction of the forecasting power of the TP component although the predictive power of EP component also appears to be weak. In other words, since 1984:Q1, the TP component has not shown cyclical movement before the business cycle as before.[5]

Why did the TP component lose its predictive power since 1984:Q1? In terms of existing literature, we may link the cyclical movement of the term premium with the Great Moderation. Kim and Nelson (1999) and McConell and Quiros (2000) find that the US GDP was more stable since 1984 which is called as "Great Moderation." We conjecture that the significant reduction in the uncertainty of US GDP may result in less cyclical movement in the term premium. In line with this conjecture, Rudebusch and Wu (2007) argue similar claim that the stability of overall macroeconomics conditions affect the term premium. Nevertheless, since the model estimated in this paper is not structural enough to draw any specific implications for the relationship between the Great Moderation and the change in the predictability of the yield spread, more structural model would be studied.

IV. Concluding Remarks

In this paper, we investigate the stability of the predictive power of the spread for future real economic activity by employing the rigorous structural break test and find that there is an evidence on the break in 1984:Q1. Furthermore, the predictive power of the spread was strong in the pre-break subsample whereas the predictability decreased

[5] Rudebusch and Wu (2007) and Dewatcher *et al.* (2014) show the similar results in the macro-finance model framework. In contrast, Favero *et al.* (2005) still emphasize the role of the term premium to predict future economic growth.

significantly in the post-break subsample.

Following the decomposition of the spread into the EP component and the TP component as in Hamilton and Kim (2002), we find that main reason of why the predictive ability of the yield spread decreased since 1984:Q1 results from the significant reduction in the predictive power of the TP component.

Why did the cyclical variation of the term premium component decrease since the mid-1980s? Hamilton and Kim (2002) mention that one factor that should matter for the term premium is the volatility of interest rates. From this point of view, the reduction in the uncertainty of output growth since the mid-1980s, may have contributed to the reduction in the uncertainty of the interest rate and thus to the reduction of the cyclical movement in the term premium. Giacomini and Rossi (2006) mention that regime change in monetary policy may be responsible for the change in the predictability of the yield spread. Nevertheless, since our model is not structural enough to provide structural interpretations, the issue needs to be investigated based on a more structural model in which the uncertainty in main macro variables such as the interest rates and output is associated with monetary policy. We leave this as a future research agenda.

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