The Changing Relation Between the Canadian and U.S. Yield Curves

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Abstract

The term structures of Canada and of the United States, two countries with historically close economic ties, have been closely linked. We investigate the link between Canadian and U.S. yield curves and show previously strong correlations between yield curve components dissipate after Canadian monetary policy reforms in the early 1990s. First, the effect is particularly evident in the diminished cross-country correlations of the short term bond yields. Secondly, cross-country yields are cointegrated before the reforms, but not afterwards. Lastly, the results on the term structure are shown using a vector autoregression with an endogenously determined break date for Canadian and U.S. estimates of the three-factor Nelson-Siegel (1987) yield curve model.

Keywords: yield curve, Canada/US, monetary policy

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1. Introduction

Canada and the United States historically maintain close economic ties and have inter-related economic policies. This paper describes the link between the term structures of interest rates in the United States and Canada. We find strong correlation between U.S. and Canadian bond yields until a change in Canadian monetary policy induced a structural break in the determinants of the Canadian term structure. After this break date, the cross-border relationship breaks down as short term Canadian yields begin to follow a path independent of the U.S. However, we find continuing dependence at longer maturities.

The approach adopted in this paper is to estimate yield curves for both countries at each period by using the three-factor Nelson-Siegel (1987) model. We then examine the evolution of these time varying yield curve factors by estimating an unrestricted VAR(1) for each country. Testing for a structural break in yield curve parameters suggests January 1993 as the break point, a date following closely on the heels of Canadian monetary policy reform. We find the relation between U.S. and Canadian yields is strong before the structural break, but greatly diminishes afterwards for short term rates. In contrast, correlation remains throughout the sample for the long term rates. In addition, cointegration tests on yields for both countries find strong evidence of cointegration before the structural break and no evidence of cointegration after the structural break. The results of these tests imply the long run relationship between U.S. and Canadian bond yields breaks down after a change in Canadian monetary policy. Lastly, by accounting for a structural break in the determination of short term rates in Canada, the influence of U.S. determinants on the Canadian term structure becomes substantially weakened.

The paper continues as follows. In Section 2 we discuss the Nelson-Siegel (1987) model as well as relevant literature. We describe the characteristics of data used in constructing term structures in Section 3. Section 4 discusses the specification of the VAR(1) model of Nelson-Siegel factor loadings that we estimate. Our results are reported in Section 5 and lastly the conclusion is given in Section 6.

2. Literature Review/Nelson-Siegel Section

The Canadian and U.S. economies are closely tied through trade and capital markets, as well as policy interactions. Simply because the U.S. economy is (roughly) ten times the size of the Canadian economy in terms of GDP, one expects U.S. activity to drive Canadian activity more than Canadian activity drives U.S. activity. However, there is more room for independence in monetary policy than there is on the real side. For example, by estimating a joint U.S.-Canada unobserved component model Fung and Remolana (1998) suggest that monetary policy carries across borders. They show that inflation shocks affect each country independently, but real shocks affect both countries. They further show differences in inflation risk premium and expectations in each country explain yield and inflation spreads.

Nelson and Siegel (1987) develop a parsimonious model of the yield curve that is flexible enough to take on the various shapes of the yield curve: monotonic, humped, and inverted. The Nelson-Siegel curve is often written as in equation (1).

$$y_t(\tau) = \hat{\mathrm{L}}_t + \hat{\mathrm{S}}_t\left(\frac{1-e^{-\lambda\tau}}{\lambda\tau}\right) + \hat{\mathrm{C}}_t\left(\frac{1-e^{-\lambda\tau}}{\lambda\tau} - e^{-\lambda\tau}\right) \tag{1}$$

$y_t(\tau)$ is the time t yield on a τ-period bond. L, S, and C, estimated each period, are interpreted as the level factor ($y(\infty)$), the slope factor ($y(0) - y(\infty)$), and a curvature factor. Diebold and

Li (2006) and Diebold, Rudebusch, and Aruoba (2006) extract latent factors from estimates of Nelson Siegel yield curves and use VARs to study the driving forces of these latent factors. We use similar methodology to study the cross-border driving relations between Canadian and U.S. yield curves.

3. Data

Data and Term Structure Characteristics

The term structure in either country consists of a set of zero-coupon yields for bonds maturing τ periods away. Our choice of data is based on convenience of availability and comparability with other studies. We consider bonds with maturity lengths of 3, 6, 9, 12, 15, 18, 21, 24, 30, 36, 48, 60, 72, 84, 96, 108 and 120 months. The sampling period is 216 monthly observations from January 1986 through December 2003.

There exist several methods for decouponing bonds in order to obtain zero coupon yields. For the U.S. we use yields computed by the unsmoothed Fama and Bliss (1987) method.[1] (See also Bliss (1997)) The Fama-Bliss method uses forward prices to interpolate end-of-month bond prices with maturities between one and five years. Zero-coupon yields for Canada are obtained from the Bank of Canada and are generated using Merrill Lynch Exponential Splines, a technique outlined by Bolder and Gusba (2002) and extended by Bolder, Johnson and Metzler (2004). The latter introduced a comprehensive database of constant maturity zero-coupon yield curves for the Government of Canada bond market, which is kept current and publicly available on the Bank of Canada's website.

[1] We are grateful to Robert Bliss for providing the U.S. yield data.

We begin with descriptive statistics and then move to more formal modeling. Univariate descriptive statistics of data from both countries are in Table 1a and Table 1b. Observing the descriptive statistics, the mean yield curve is upward sloping and concave for both countries. However, the average Canadian yield is greater than American yields at all maturities. Furthermore, Canadian yields exhibit a higher variance and persistence at 12 and 30 month horizons.

Figure 1 plots Canadian and U.S. yields for selected maturities, illustrating two important points. First, the U.S. and Canadian zero-coupon yields are strongly correlated during the early years of the sample. During this earlier period, Canadian yields are above the corresponding U.S. yields. According to Clinton (1998), higher Canadian yields are a result of a less-liquid market, a perception of greater risks and a frequent expectation the Canadian dollar will decline in value. In addition, the Canadian yields appear to peak in the first half of the 1990s. Clinton (1998) attributes this fact to increasingly unsustainable growth in public debt, historical problems with inflation and structural problems in the public and private sectors. Second, there appears to be a structural break in Canadian data that occurs in the early 1990's. After this point, the Canadian yields for the various maturities fall below those of the U.S., where previously they were above the U.S. yields by a fairly consistent 2-3%. In addition to falling below the U.S. yields, the Canadian yields now follow their own path; independent of the U.S. This independent relationship weakens as the maturity date lengthens implying the structural break is in the determination of the Canadian short-run rates.

Table 2 presents the cross-country correlations between bond yields at each maturity. The right-most column shows correlations for the full sample. All correlations are high, and increasing with maturity from 0.81 to 0.95. The left two columns give correlations for our two sub-periods. The evident change is that the correlation of short maturities falls in the second sub-period. For example, the correlation in the 3-month maturity falls from 0.91 to 0.66. In contrast, there is little change at long maturities. We attribute this effect to a fundamental change in the determination of shorter term bond yields by the Bank of Canada.

Nonstationarity and Cointegration in the Term Structure

If yields are nonstationary, correlations computed on levels may be spurious. VAR estimation needs to account for nonstationarity if it exists. Table 3 presents unit root tests running augmented Dickey-Fuller tests on all available maturities with the null hypothesis that yields are I(1) without drift. The results reported in Table 3 provide no significant evidence against a unit root for either country. Results reported in Tables 4a and 4b are similarly unable to reject unit roots for our sample sub-periods.

Following a two-step procedure described by Engle-Granger (1987) and outlined by Boothe (1991), we test for a cointegrating relationship between the Canadian and US yields at the different maturities. The first step is to run two de-meaned cointegrating regressions: 1) regress the Canadian yields on a constant and the U.S. yields for each maturity and 2) regress the U.S. yields on a constant and the Canadian yields for each maturity. The second step is to perform a unit root test on the residuals of these OLS regressions, where the null hypothesis states the residuals are distributed I(1). Therefore, rejection of the null implies the yields are cointegrated. Table 5 reports results for the cointegration of yields of varying maturity between countries for the full sample period. In addition, Table 6 reports the results for the two sub-samples. In the full sample results, observe the Augmented Engle-Granger (AEG) test statistics from the full sample approach the Phillips-Ouliaris (PO) critical value as the length of maturity increases.[2] Therefore, there is stronger statistical evidence of cointegration between the two countries as the maturity period lengthens. Specifically, there is significant evidence of cointegration between the 108 month and 120 month zero-coupon yields under these two regressions.

In regards to the sub-sample results, there is strong evidence of cointegration between Canadian and U.S. yields of varying maturity before the structural break and no evidence of cointegration after the structural break. This implies the Canadian and U.S. yields are strongly correlated before the structural break, but no longer share a long term trend after the fact. This can be seen in Figure 1 above. Graphs of the Canadian and U.S. yields appear to follow a similar trend until after the early 1990's. This further strengthens the theory of a structural break in Canadian monetary policy. After Canada changed its monetary policy and financial markets, the correlation between U.S. and Canadian yields appears to diminish.[3]

3. Model

We estimate separate Nelson-Siegel term structure models for Canada and for the United States during each period in our sample. As in Diebold-Li (2006), λ is taken to be 0.0609, the value maximizing the medium term loading at 30 months. Equation (1) states the yield at time t

[2] Augmented Dickey Fuller unit root tests applied to estimated cointegrating residuals do not follow the usual Dickey-Fuller distribution under the null hypothesis. Instead Phillips and Ouliaris (199) find the unit root tests follow asymptotic distributions which are functions of Wiener processes. The Phillips-Ouliaris critical values are reported in Phillips and Ouliaris (1990).

[3] As a robustness check we also ran a Cointegrating Regression Durbin Watson (CRDW) test as outlined by Engle and Granger (1987). The results of this test also find increasing evidence of cointegration as the maturity period lengthens and evidence of cointegration in the first sub-sample, but none in the second sub-sample.

for a newly issued bond maturing τ periods in the future is a function of a long term level loading, short term slope loading, and medium term curvature loading.

Given the time series for three factors for each country, and because of the presence of unit roots in level and slope factor loadings for both countries, we estimate an unrestricted VAR(1) for the first-differences of the six estimated factors (see Diebold, Rudebusch, and Aruoba (2006)). Specifically,

$$\Delta\widehat{\mathrm{B}}_{\mathrm{t}} = \Gamma\,\Delta\widehat{\mathrm{B}}'_{\mathrm{t-1}} + \varepsilon_{\mathrm{t}} \quad (2)$$

$$\mathrm{B}_t = \begin{bmatrix} L_{US} \\ S_{US} \\ C_{US} \\ L_{CAN} \\ S_{CAN} \\ C_{CAN} \end{bmatrix}, \quad \Gamma = \begin{bmatrix} \gamma_{1,1}^{US} & \gamma_{1,2}^{US} & \gamma_{1,3}^{US} & \gamma_{1,1}^{C} & \gamma_{1,2}^{C} & \gamma_{1,3}^{C} \\ \gamma_{2,1}^{US} & \gamma_{2,2}^{US} & \gamma_{2,3}^{US} & \gamma_{2,1}^{C} & \gamma_{2,2}^{C} & \gamma_{2,3}^{C} \\ \gamma_{3,1}^{US} & \gamma_{3,2}^{US} & \gamma_{3,3}^{US} & \gamma_{3,1}^{C} & \gamma_{3,2}^{C} & \gamma_{3,3}^{C} \\ \gamma_{4,1}^{US} & \gamma_{4,2}^{US} & \gamma_{4,3}^{US} & \gamma_{4,1}^{C} & \gamma_{4,2}^{C} & \gamma_{4,3}^{C} \\ \gamma_{5,1}^{US} & \gamma_{5,2}^{US} & \gamma_{5,3}^{US} & \gamma_{5,1}^{C} & \gamma_{5,2}^{C} & \gamma_{5,3}^{C} \\ \gamma_{6,1}^{US} & \gamma_{6,2}^{US} & \gamma_{6,3}^{US} & \gamma_{6,1}^{C} & \gamma_{6,2}^{C} & \gamma_{6,3}^{C} \end{bmatrix}$$

where $\widehat{\mathrm{B}}_{\mathrm{t}}$ is the 6x1 vector estimated factors from the U.S. and Canada; Γ is a 6x6 matrix of factor loading coefficients; and $E(\varepsilon_t \varepsilon_t') = \Omega$ allowing for correlation of errors across equations.

4. Results

Structural Breaks and Sub-Samples

The historical events pertaining to Canadian monetary policy reform provide a basis for investing the existence of a structural break in bond yield determination. To determine whether a structural break exists in the Canadian slope factor loading, the Andrews (1993) test is employed. The Andrews test allows for any criterion test statistic, of which we construct a Wald statistic based on the null hypothesis that coefficients in the Canadian slope equation are equal before and after the structural break. While historical events suggest a range of possible break dates, the exact date is not known with certainty. In fact, there exists the possibility that no break occurred, hence we let the data speak for itself. The break point search range is from January 1990 to December 1993. The maximized test statistic value of 18.565 corresponding to the period January 1993 allows for rejection of the null hypothesis of no structural change at the 5% level. The timing is consistent with the expectation that the structural break coincide with the change in monetary policy in Canada during the period 1989 – 1993, but is likely to occur at the latter part of this period at the completion of all monetary reforms.

Historical Connection

Canada implemented a major change in its conduct of monetary policy roughly a third of the way into our sample. Specifically, in 1991 Canada adopted and implemented inflation targets that resulted in a deep contraction of the Canadian economy shown by falling economic growth and increasing unemployment through 1999 as outlined in Curtis (2005). The Bank of Canada implemented inflation targets by setting the short term interest rate. Bond yields had a lagged response to this policy tool. Also, this type of monetary policy had a reduced effect on long term interest rates according to Clinton (1998) and Bhuiyan and Lucas (2007), implying the structural break is in the determination of Canadian short-run rates. Canada's policy was in sharp contrast to U.S. monetary policy at the time, which emphasized a balanced approach to controlling inflation, real output growth and unemployment, resulting in a boom in the U.S. economy during the same period. In addition, Canada looked to foster smooth adjustments in short term interest rates. The goal was to prevent uncertainty about future monetary policy. Bolder, Johnson and Metzler (2004) deem the Canadian bond market a less risky place during the mid-to-late '90s as the level of volatility in the various Canadian yield curve measures fall significantly during this time. In fact, in 1991 we observe a dramatic increase in the volume of open market operations conducted by the Bank of Canada. Racette and Raynauld (1994) analyze the Canadian monetary policy during 1989-1993 and find monetary policy in Canada primarily emphasized severe reductions in inflation in the early 1990s. The anecdotal evidence suggests a structural break in policy occurred sometime during this four year period and is most likely to be evidenced in the dynamics of short rates.

VAR Estimation Results

Consider the estimation of model (2) using three samples. First, we estimate the joint U.S.-Canada VAR model using the entire sample. We proceed by dividing observations into two distinct sub-samples representing pre and post monetary policy shift. Because the relationship between the short term and long term aspects of the term structure are of interest, and the curvature component of the Nelson-Siegel model is not well identified, interpretation of results focuses on the slope and level components. Distinguishing effects between two distinct periods of Canadian monetary policy supports the observation that the correlation between U.S. and Canadian short term bond yields once existed, but ceased after our estimated structural change date. Differing effects resulting from dichotomizing the full sample exist only for the Canadian short rate. In each of the three samples, all estimated coefficients in the Canadian level equation are insignificant. Hence, the dynamics governing Canadian long rate remain the same even after the transition to a different monetary policy regime. We also find the U.S. slope factor is a major determinant of the U.S. term structure throughout the sample.

Table 7 presents coefficient estimates from the full sample from February 1986 to December 2003. The most important results involve the effects of the U.S. level and slope factors. Both terms are strongly significant in the equations for short term factors in both countries with the effect more pronounced in Canada. For example, the coefficient for lagged change in the U.S. slope factor is 0.761 in the Canadian slope equation and 0.368 in the U.S. slope equation. The result in itself is notable—albeit perhaps unsurprising—when evaluating the behavior of short term rates in Canada, suggesting a larger influence from U.S. than Canadian monetary policy. The remaining two significant coefficient estimates involve the lagged U.S. slope factor in the U.S. level equation and the lagged U.S. level factor in the U.S. curvature equation with both marginally significant at the 5% level.

Accounting for a structural break in term structure dynamics yields far different results. Table 8a summarizes results for the unrestricted VAR(1) model using the first sub-sample from January 1986 to January 1993. The primary observation from the first sub-sample is the significant effect of term structure components from both countries on the Canadian slope component. In the Canadian slope equation, the estimated coefficient for the Canadian level factor is -0.584 and the coefficient for the Canadian slope factor is -0.258. The full sample fails to capture the significant effects of lagged Canadian level and lagged Canadian slope in the first sub-sample. In addition, coefficients for U.S. level and slope in the Canadian slope equation remain significant, and have a larger effect relative to the full sample. The U.S. level coefficient is 1.441 compared to 1.186 for the full sample and the U.S. slope coefficient is 0.946 compared to 0.761 for the full sample. The implication is U.S. determinants more strongly influence Canadian short rates before monetary policy reforms. The importance of the U.S. slope component also extends to the U.S. term structure. The U.S. slope component drives the term structure as lagged changes in U.S. slope positively affect changes in U.S. slope with a coefficient estimate of 0.341 and negatively affects changes in U.S. level with a coefficient estimate of -0.271. In contrast, Canadian factors have little effect on U.S. rates.

Table 8b presents results of the unrestricted VAR(1) model for the second sample period January 1993 to December 2003. The results further emphasize the changing dynamics of the Canadian short rate after monetary reform. First, observe changes in U.S. term structure coefficients in the Canadian slope equation. Specifically, the decrease in the magnitude from 1.441 to 0.801 in the level component and 0.946 to 0.600 in the slope component between sub-samples show an attenuation of U.S. monetary policy on the Canadian short rate. Second, the significance of the lagged U.S. slope factor continues to strongly affect contemporaneous U.S. slope. The estimate of 0.547 for $\gamma_{2,1}^{US}$ suggests an increasing importance of the level factor in determining U.S. short rates after 1993.

Canadian factors generally have a small effect on the term structure in both countries, regardless of the sample period. Specifically, U.S. factors remain unaffected by lagged Canadian factors throughout the sample. The notable exception is the negative effect on the Canadian short rate exhibited in the first sub-sample. After the estimated break date, this negative effect ceases. The shift in signs for $\gamma_{5,1}^{C}$ and $\gamma_{5,2}^{C}$ coefficient estimates suggest new Canadian policies for determining short rate. Coefficient estimates in the Canadian slope equation are not significant at the 5% level; however $\widehat{\gamma_{5,2}^{C}}$ is significant at the 10% level. Finally, separating the data set into two samples introduces significance in determinants of the medium-term factor for the U.S. However, shifts in significance before and after the break date fail to follow any meaningful pattern.

Table 9 summarizes results of F-tests for hypotheses of joint significance in U.S. and Canadian VAR coefficients respectively in the Canadian slope equation. This test identifies collective country effects on the Canadian short rate. Test statistics with and without the curvature factor are included. The results coincide with VAR results. Before the estimated break date, both the U.S. and Canadian effect are significant at the 1% level. After the structural break, the U.S. effect remains significant, however, the effect weakens. For example, the F-statistic including U.S. level and slope is now only significant at the 5% level. Furthermore, the Canadian effect is no longer significant.

5. Conclusion

This paper shows there is a correlated relationship between U.S and Canadian term structures. Due to increased risk perceptions, historically high inflation and weak confidence in the Canadian dollar, the Canadian yields for all maturities lie above those in the U.S. until a shift in monetary policy caused a structural break in the Canadian yields. The result of this structural break was a termination in the correlated relationship between the term structures in the two countries.

The changing relationship between the U.S. and Canadian term structure is displayed in our results. The VAR results show U.S. determinants strongly influence Canadian short rates before the structural break. Specifically the Canadian short term rate is largely determined by the U.S. level and slope factors. After the structural break the correlation between U.S. and Canadian bond yields breaks down as new Canadian monetary policy takes effect. In addition, tests for cointegration between the yields of the two countries display strong evidence of cointegration before the structural break and no evidence of cointegration after the structural break. This implies the long-run relationship between U.S. and Canadian bond yields degrades after the shift in Canadian monetary policy.

According to the above results changes in monetary policy can have strong effects on the integration of neighboring countries. Canada, a country historically influenced by the U.S. economy, asserts its independence in the bond market by changing monetary policy to place a hard target on inflation rates. Further research on additional shifts in monetary policy, both in Canada and the U.S. is needed to strengthen the argument on the correlated term structure of these two countries. In addition, it would be interesting to see if the degeneration of a correlated term structure in response to changing monetary policy exists for other integrated countries.

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Table 1a: Descriptive Statistics, Canadian Yield Curves: 1986-2003

Maturity	Mean	Std. Dev.	Minimum	Maximum	$\hat{\rho}(1)$	$\hat{\rho}(12)$	$\hat{\rho}(30)$
3	6.436	2.979	1.979	13.484	0.978	0.730	0.373
6	6.454	2.853	1.939	13.073	0.978	0.736	0.376
9	6.504	2.747	2.035	12.979	0.977	0.738	0.377
12	6.561	2.653	2.223	12.858	0.976	0.738	0.379
15	6.617	2.569	2.429	12.786	0.975	0.737	0.382
18	6.672	2.494	2.574	12.715	0.974	0.736	0.387
21	6.723	2.429	2.730	12.628	0.973	0.735	0.393
24	6.771	2.371	2.887	12.534	0.973	0.734	0.399
30	6.858	2.277	3.100	12.352	0.972	0.736	0.414
36	6.935	2.204	3.249	12.199	0.972	0.738	0.427
48	7.070	2.098	3.532	11.992	0.972	0.746	0.453
60	7.187	2.022	3.813	11.861	0.972	0.755	0.477
72	7.287	1.964	4.043	11.729	0.973	0.763	0.499
84	7.370	1.919	4.223	11.563	0.973	0.770	0.519
96	7.439	1.886	4.358	11.376	0.974	0.777	0.534
108	7.496	1.864	4.459	11.205	0.975	0.782	0.545
120	7.549	1.851	4.542	11.087	0.976	0.787	0.552

Table 1b: Descriptive Statistics, U.S. Yield Curves: 1986-2003

Maturity	Mean	Std. Dev.	Minimum	Maximum	$\hat{\rho}(1)$	$\hat{\rho}(12)$	$\hat{\rho}(30)$
3	4.911	1.916	0.876	9.131	0.978	0.599	-0.029
6	5.039	1.932	0.958	9.324	0.978	0.603	-0.007
9	5.145	1.943	0.979	9.343	0.977	0.610	0.018
12	5.290	1.963	1.040	9.641	0.975	0.613	0.027
15	5.431	1.966	1.066	9.698	0.975	0.616	0.052
18	5.516	1.936	1.144	9.660	0.974	0.614	0.074
21	5.588	1.902	1.219	9.544	0.973	0.613	0.095
24	5.626	1.858	1.299	9.524	0.972	0.610	0.113
30	5.785	1.808	1.447	9.510	0.971	0.610	0.155
36	5.899	1.747	1.618	9.461	0.969	0.611	0.188
48	6.111	1.660	1.999	9.348	0.966	0.620	0.262
60	6.234	1.595	2.351	9.293	0.965	0.627	0.313
72	6.390	1.549	2.663	9.286	0.965	0.646	0.366
84	6.487	1.495	3.003	9.395	0.966	0.656	0.399
96	6.590	1.471	3.221	9.521	0.967	0.672	0.435
108	6.641	1.459	3.389	9.594	0.966	0.681	0.461
120	6.638	1.437	3.483	9.527	0.965	0.688	0.485

Table 2: Correlation Coefficients between U.S. and Canadian Bond Yields

Maturity	Correlation 1986-1993	Correlation 1993-2003	Correlation 1986-2003
3	0.90786	0.65940	0.813
6	0.90744	0.70085	0.824
9	0.89944	0.72601	0.834
12	0.89419	0.73613	0.839
15	0.89596	0.74267	0.848
18	0.89706	0.75196	0.857
21	0.89586	0.76046	0.864
24	0.89432	0.76743	0.869
30	0.89081	0.77732	0.883
36	0.88860	0.78804	0.893
48	0.88427	0.80384	0.912
60	0.88700	0.81150	0.923
72	0.88267	0.8102	0.931
84	0.87408	0.81635	0.935
96	0.86478	0.83310	0.942
108	0.85860	0.85295	0.947
120	0.84677	0.87742	0.953

Table 3: Augmented Dickey-Fuller (ADF) Tests: 1986-2003 [4]

Maturity (months)	ADF (U.S.)	P-value	ADF (Canada)	P-value
3	-0.631	0.860	-1.310	0.625
6	-0.601	0.866	-1.216	0.668
9	-0.680	0.848	-1.162	0.691
12	-0.701	0.843	-1.149	0.697
15	-0.752	0.830	-1.155	0.694
18	-0.820	0.811	-1.170	0.688
21	-0.875	0.795	-1.188	0.680
24	-1.004	0.752	-1.207	0.672
30	-1.092	0.719	-1.238	0.658
36	-1.145	0.698	-1.260	0.648
48	-1.244	0.655	-1.282	0.638
60	-1.353	0.605	-1.293	0.633
72	-1.426	0.569	-1.298	0.631
84	-1.420	0.572	-1.287	0.636
96	-1.410	0.577	-1.258	0.649
108	-1.535	0.514	-1.216	0.668
120	-1.673	0.443	-1.167	0.689

[3] The ADF critical values are calculated from MacKinnon (1996). The ADF lags are defined according to the Schwarz Info Criterion.

Table 4a: Augmented Dickey-Fuller (ADF) Tests: January 1986-January 1993

Maturity (months)	**ADF** (U.S.)	**P-value**	**ADF** (Canada)	**P-value**
3	0.210	0.972	-1.069	0.725
6	0.178	0.970	-1.076	0.722
9	0.106	0.964	-1.134	0.699
12	-0.140	0.941	-1.210	0.667
15	-0.133	0.942	-1.284	0.634
18	-0.179	0.936	-1.350	0.603
21	-0.257	0.926	-1.409	0.574
24	-0.434	0.898	-1.460	0.549
30	-0.568	0.871	-1.544	0.507
36	-0.746	0.829	-1.607	0.475
48	-1.081	0.720	-1.707	0.424
60	-1.328	0.613	-1.815	0.371
72	-1.580	0.489	-1.926	0.319
84	-1.595	0.481	-2.025	0.276
96	-1.705	0.425	-2.111	0.241
108	-2.017	0.279	-2.187	0.213
120	-2.249	0.191	-2.239	0.195

Table 4b: Augmented Dickey-Fuller (ADF) Tests: February 1993-December 2003

Maturity (months)	**ADF** (U.S.)	**P-value**	**ADF** (Canada)	**P-value**
3	-0.330	0.916	-2.031	0.274
6	-0.894	0.788	-1.924	0.321
9	-0.430	0.900	-1.859	0.351
12	-0.387	0.907	-1.825	0.367
15	-0.460	0.894	-1.808	0.375
18	-0.497	0.887	-1.796	0.381
21	-0.573	0.872	-1.784	0.387
24	-0.693	0.844	-1.768	0.395
30	-0.804	0.815	-1.723	0.418
36	-0.873	0.794	-1.368	0.596
48	-0.733	0.834	-1.363	0.598
60	-0.919	0.780	-1.352	0.604
72	-1.086	0.720	-1.353	0.603
84	-1.188	0.678	-1.372	0.594
96	-1.308	0.625	-1.399	0.581
108	-1.360	0.600	-1.426	0.568
120	-1.515	0.523	-1.448	0.557

Table 5: Residual-based No-Cointegration Tests: Full Sample[5]

Maturity (Months)	AEG Regression 1	AEG Regression 2
3	-1.7965	-1.0687
6	-1.6211	-0.9877
9	-1.6296	-1.1440
12	-1.6512	-1.2302
15	-1.7772	-1.6863
18	-1.8586	-1.8055
21	-1.9194	-1.8883
24	-2.2083	-1.9265
30	-2.3542	-2.1336
36	-2.4319	-2.2627
48	-2.6633	-2.5987
60	-2.7392	-2.7537
72	-2.8632	-2.9314
84	-2.8424	-2.9105
96	-2.9422	-3.0164
108	-3.0764*	-3.2294*
120	-3.3399*	-3.5688**

[5] Henceforth, significance at the 10% level, the 5% level and the 1% level are represented by one asterisk, two asterisks and three asterisks respectively. The augmented Engle Granger (AEG) test statistics are calculated using test regressions that include a constant. The lags are selected according to the Schwarz Info Criterion. The critical values are calculated from Phillips and Ouliaris (1990), where the 1% critical value is -3.9618, the 5% critical value is -3.3654, and the 10% critical value is -3.0657.

Table 6: Residual-based No-Cointegration Tests: Sub-Samples

Maturity	AEG: 1986-1993 Regression 1	AEG: 1986-1993 Regression 2	AEG: 1993-2003 Regression 1	AEG: 1993-2003 Regression 2
3	-2.1499	-1.6651	-2.0603	-1.2043
6	-1.9149	-1.4538	-2.0715	-1.8673
9	-1.9165	-1.4583	-2.6300	-2.0266
12	-1.9521	-1.5050	-2.6686	-2.1151
15	-2.1327	-1.6582	-2.6297	-2.0806
18	-2.2482	-1.7619	-2.6748	-2.0958
21	-2.3346	-1.8193	-2.6312	-2.1263
24	-3.1276*	-1.8520	-2.6065	-2.1410
30	-3.2485*	-1.9883	-2.5166	-2.0734
36	-3.3101*	-2.0233	-2.3986	-1.9780
48	-3.4085**	-3.0792*	-2.1282	-1.7919
60	-3.3596*	-3.0846*	-1.9617	-1.7544
72	-3.4360**	-3.2404*	-1.7402	-1.6550
84	-3.1127*	-2.8248	-1.7147	-1.7238
96	-3.1274*	-2.8472	-1.7188	-1.8053
108	-3.1221*	-1.9466	-1.7715	-1.8891
120	-3.1728*	-3.2008*	-2.0011	-2.1605

Table 7: Vector Autoregression Estimates: Full Sample

	DB1_U.S.	DB2_U.S.	DB3_U.S.	DB1_CAN	DB2_CAN	DB3_CAN
DB1_U.S.(-1)	-0.2421* (0.1248)	0.4531*** (0.1510)	0.7225** (0.3490)	0.0424 (0.1475)	1.1856*** (0.2227)	-0.3560 (0.5336)
DB2_U.S.(-1)	-0.1714** (0.0801)	0.3683*** (0.0969)	0.2913 (0.2240)	-0.0806 (0.0947)	0.7614*** (0.1430)	-0.1738 (0.3426)
DB3_U.S.(-1)	0.0302 (0.0286)	0.0062 (0.0346)	-0.0176 (0.0799)	0.0340 (0.0338)	0.0107 (0.0510)	0.0824 (0.1222)
DB1_CAN(-1)	0.0746 (0.0930)	-0.0016 (0.1125)	-0.1253 (0.2599)	-0.1672 (0.1099)	-0.3089* (0.1659)	0.3984 (0.3975)
DB2_CAN(-1)	-0.0487 (0.0410)	0.0951* (0.0496)	-0.0222 (0.2599)	0.0105 (0.0485)	-0.1128 (0.0732)	-0.2205 (0.1753)
DB3_CAN(-1)	-0.0122 (0.0209)	0.0142 (0.0253)	0.0624 (0.0585)	-0.0036 (0.0247)	-0.0152 (0.0374)	-0.0142 (0.0895)

Table 8a: Vector Autoregression Estimates: 1986-1993

	DB1_U.S.	DB2_U.S.	DB3_U.S.	DB1_CAN	DB2_CAN	DB3_CAN
DB1_U.S.(-1)	-0.3008	0.3038	0.5198	0.1139	1.4412***	-0.9929
	(0.1876)	(0.1969)	(0.4971)	(0.2455)	(0.3444)	(0.9216)
DB2_U.S.(-1)	-0.2714**	0.3406***	0.1157	-0.1710	0.9460***	-0.6474
	(0.1211)	(0.1271)	(0.3209)	(0.1585)	(0.2223)	(0.5949)
DB3_U.S.(-1)	0.0148	0.0018	0.0462	0.0282	-0.0481	0.1132
	(0.0458)	(0.0481)	(0.1213)	(0.0599)	(0.0841)	(0.2249)
DB1_CAN(-1)	0.0913	0.0602	-0.1586	-0.1777	-0.5839**	0.3946
	(0.1390)	(0.1459)	(0.3682)	(0.1818)	(0.2551)	(0.6826)
DB2_CAN(-1)	-0.0267	0.0528	0.1332	0.0040	-0.2583**	-0.1939
	(0.0566)	(0.0594)	(0.1499)	(0.0741)	(0.1039)	(0.2780)
DB3_CAN(-1)	0.0020	-0.0028	0.1431*	0.0179	-0.0241	-0.0593
	(0.0292)	(0.0306)	(0.0773)	(0.0382)	(0.0535)	(0.1433)

Table 8b: Vector Autoregression Estimates: 1993-2003

	DB1_U.S.	DB2_U.S.	DB3_U.S.	DB1_CAN	DB2_CAN	DB3_CAN
DB1_U.S.(-1)	-0.1132	0.5415**	1.3702***	-0.0273	0.7950***	0.5786
	(0.1826)	(0.2386)	(0.5144)	(0.1951)	(0.3017)	(0.6771)
DB2_U.S.(-1)	-0.0399	0.3960**	0.7226**	-0.0487	0.5960***	0.5719
	(0.1211)	(0.1583)	(0.3412)	(0.1294)	(0.2001)	(0.4491)
DB3_U.S.(-1)	0.0584	-0.0067	0.0206	0.0559	0.0271	0.0800
	(0.0388)	(0.0507)	(0.1094)	(0.0415)	(0.0641)	(0.1439)
DB1_CAN(-1)	0.0531	0.0261	-0.3227	-0.1488	0.2644	0.6103
	(0.1406)	(0.1836)	(0.3959)	(0.1501)	(0.2322)	(0.5211)
DB2_CAN(-1)	-0.0829	0.1403	-0.2534	0.0347	0.1847*	-0.2764
	(0.0659)	(0.0860)	(0.1855)	(0.0703)	(0.1088)	(0.2441)
DB3_CAN(-1)	-0.0377	0.0443	-0.0762	-0.0378	0.0167	0.0320
	(0.0318)	(0.0416)	(0.0897)	(0.0340)	(0.0526)	(0.1180)

Table 9: F-Test Results[6]

	Full Sample	Pre-1993	Post-1993
U.S.	12.3486***	8.6049***	3.2904**
	(0.000)	(0.000)	(0.023)
Canada	1.7686	4.1872***	1.0581
	(0.154)	(0.008)	(0.370)
U.S Without Curvature	17.7570***	11.8728***	4.7460***
	(0.000)	(0.000)	(0.010)
Canada without Curvature	2.4281*	5.7965***	1.5413
	(0.091)	(0.005)	(0.218)

[6] P-values are in parentheses.

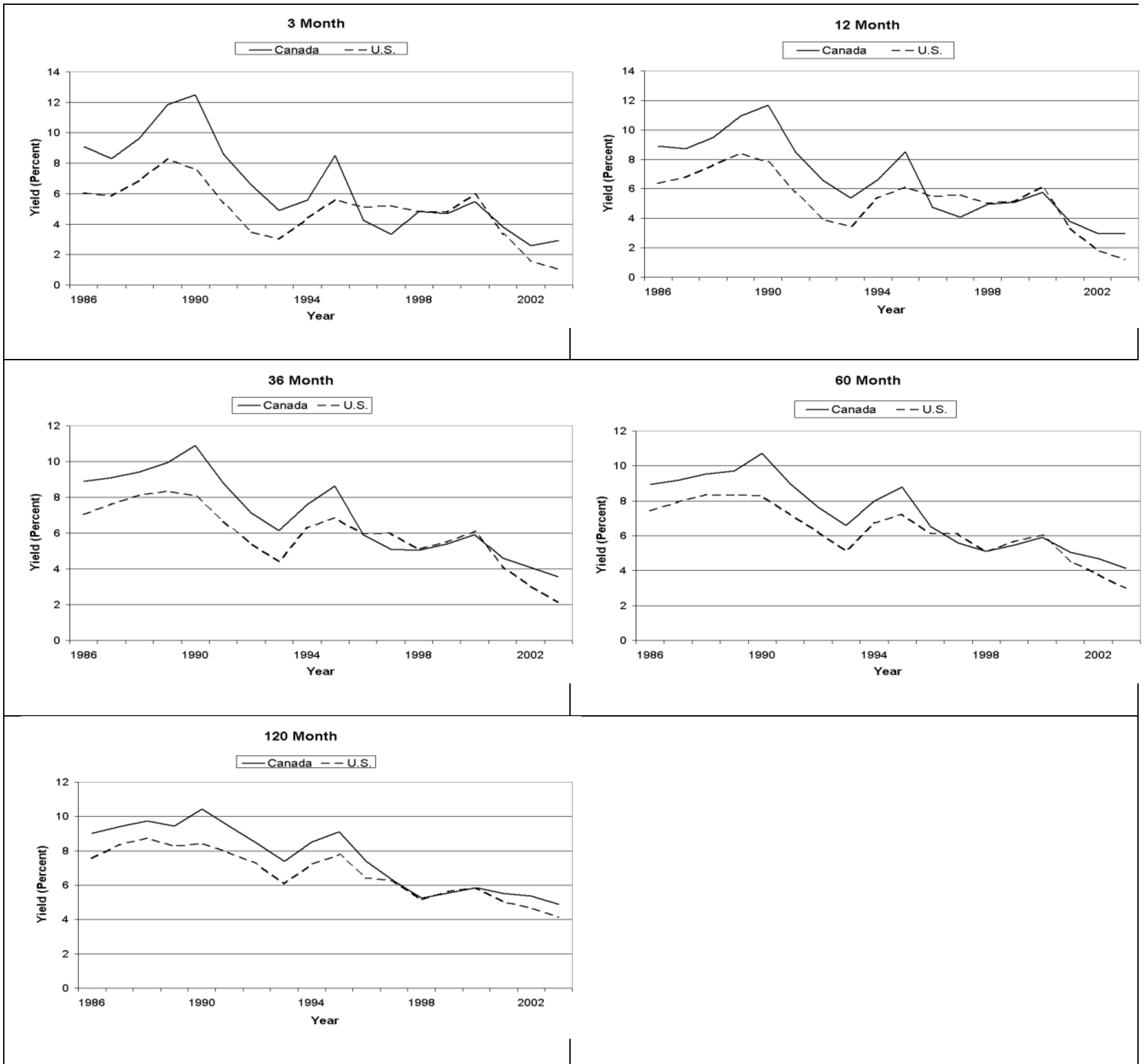

Figure 1. Yearly average zero-coupon Yields for both Canada and the U.S., 1986-2003.