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The Relations Among Canadian, Mexican, and United States Short-Term Money Markets: A Pre-NAFTA Analysis

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Abstract

This study uses cointegration analysis to examine linkages among the financial markets of Canada, Mexico, and the United States. The results are consistent with the hypotheses that long-run equilibrium relationships exist between Canadian and U.S. short-term interest rates and between Canadian and U.S. money supplies, respectively. The U.S. variables are found to have been a dominant force in the respective relationships with the Canadian variables. However, our findings do not lend support to the hypotheses of a cointegrating relationship between the short-term interest rate or the monetary base of Mexico and those of the United States or Canada. Nevertheless, we anticipate that the financial markets of Mexico will become more closely connected with those of the other two NAFTA countries in the future as NAFTA is implemented. (JEL Classification: F36)

I. Introduction

Changes in the tariff structure of trade in goods across the borders of Canada, Mexico, and the United States are the parts of the recently imple-

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mented North American Free Trade Agreement (NAFTA) that dominate the headlines ot the popular press.[1] However, the provisions of the Agreement with respect to financial services may have equal, or even greater, effects on the economy of each country. NAFTA provided for a gradual opening over a six-year period of the Mexican financial services industry and specified a set of regulations whereby wholly-owned subsidiaries of Canadian and U.S. banking, securities, and insurance firms can be established in Mexico. Although market share limits are placed on foreign-owned firms, these firms can engage in nearly all of the same activities as can domestic firms. NAFTA also provided for a reciprocal opening of the Canadian and U.S. financial systems to Mexican financial intermediaries.[2]

The purpose of this paper is to examine the nature of the short-term money market linkages among Canada, Mexico, and the United States before the NAFTA was implemented using the tool of cointegration analysis and corresponding error correction models. As explained later in the paper, the employed data are consistent with the hypothesis that both short-term interest rates and the monetary bases in Canada and the United States are cointegrated, but that a cointegration relationship involving these variables does not exist for Mexico and the United States or for Mexico and Canada. In the case of Canada and the United States, the data are consistent with the hypothesis that changes in the United States are the dominant variable in the short-term money market relationships. These results suggest the hypothesis that the financial sectors of Canada and the United States were closely connected during the period of study, but that those of Mexico and the United States or Canada were not. However, it seems reasonable to anticipate that this situation will change with respect to Mexico and the other two NAFTA countries as the provisions of NAFTA gradually bring about a much closer relationship among the financial markets in these countries. Thus, the long-term impact of the NAFTA will probably be far greater with respect to

1. See, for example, "Congress Passes North American Free Trade Pact," *OAS Trade News*, Vol. XVIII, No. 11 (November 1993), pp. 1-2.
2. Certain functions related to the monetary and exchange policies of each country and actions on behalf of such agencies as a country's development banks and social security system are excepted. See Banamex [1992, pp. 616-618]; Kehoe and Kehoe [1994, p. 21]; and the *North American Free Trade Agreement*, Chapter 14.

the financial services sector in Mexico than in the other two countries.

Using a variety of techniques, other authors have studied the relationships among interest rates, the monetary bases, and exchange rates for the countries of the European Monetary System (EMS) and, in some cases, for the United States and a subset of the European countries.[3] In a recent paper, Boothe [1991] used cointegration tests to investigate whether uncovered interest rate parity held between Canadian and U.S. interest rates. He used end-of-month constant maturity yields for Canadian and U.S. government three-month treasury bills, and two-, ten-, and twenty-year bonds. The results of his study were consistent with the hypothesis of uncovered interest rate parity in an open economy as an explanation of the term structure of interest rates in Canada.

There are also some studies on monetary growth linkages. The results, to some extent, differ with data under fixed exchange rate regimes compared with data covering flexible exchange rate systems. For example, Sheehan [1987] used quarterly data covering both a fixed exchange rate period (I/1960 – II/1971) as well as a floating exchange rate period (III/1973 to IV/1985). Using a cross-correlation statistical technique developed by Haugh [1976], Sheehan's results during the fixed exchange rate period were consistent with the hypothesis of monetary growth interdependence between the United States and Belgium, Canada, France, Germany, Japan, the Netherlands, and the United Kingdom. They did not support the hypothesis of interdependence between monetary growth in the United States and in Italy and Switzerland. However, during the flexible exchange

3. One issue that has been pervasive in a number of these studies is the hypothesis of German Dominance in the EMS. The studies by von Hagen and Fratianni [1990]; Fratianni and von Hagen [1990]; Cohen and Wyplosz [1989]; De Grauwe [1988]; Mastropasqua, Micossi, and Rinaldi [1988]; and von Hagen [1989] do not support the hypothesis of strict German dominance of the EMS. For a contrary view, see Fischer [1987], Giavazzi and Giovannini [1987], and Russo and Tullio [1988]. Kirchgassner and Wolters [1987] examined the interest rate linkage between short-term Euro-market and long-term bond market interest rates among Germany, Switzerland, and the United States over two subperiods between 1974 and 1984. Their results were consistent with the hypothesis that there was a strong linkage among the interest rates of these three countries during the second subperiod, but that the linkage was weak, if it existed at all, during the first subperiod.

rate period, Sheehan's study was consistent with the hypothesis of monetary growth interdependence only for the United States, Canada, and Japan.

Using data covering a flexible exchange rate period, Batten and Ott [1985] found evidence to support the hypothesis that United States monetary growth impacted monetary growth in Canada, Germany, Japan, and the Netherlands, but not in France, Italy, and Switzerland. Also using earlier data but covering both fixed and flexible exchange rate systems, Sheehan [1983] obtained results consistent with the hypothesis that monetary growth in the United States affected its counterparts in Australia and Germany, but had no bearing on monetary growth in Canada, Italy, Japan, and the United Kingdom.

This study differs from the previous research in the area of financial market integration in a number of ways. First, we employ the technique of cointegration analysis to investigate the nature of the short-term money market linkages among Canada, Mexico, and the United States. We do not know of any previous studies that have used cointegration analysis to examine the relationship between the financial markets in Mexico and the United States or among those of the three NAFTA countries. Second, both interest rate linkages and monetary base linkages are studied in the cointegration framework. We have not found any previous studies that have used that technique to examine the monetary supply relationships between any pair of these countries. Third, Johansen's [1988] maximum likelihood approach is employed to carry out the cointegration analysis. Unlike the other cointegration tests which have to be performed on the basis of an arbitrary choice of dependent variable in the regression equation and an assumption that there exists a unique cointegrating vector between the variables, Johansen's approach allows us to treat all the variables as endogenous and explicitly test for the number of cointegrating vectors.[4] Finally, in addition to the money market linkages, the relative exogeneity (endogeneity) of each variable in the revealed cointegration relationship is also examined through the error correction mechanism.

The next section presents a brief description of cointegration analysis,

4. Engle and Granger's [1987] residual based regression method of cointegration analysis is also applied to show the robustness of the results from the Johansen tests. Since the Engle-Granger tests are now well known, the descriptions of the tests are omitted.

and Section III develops the estimated interest rate relationships. Section IV summarizes the empirical findings of the study, and Section V discusses our conclusions.

II. A Brief Review of Cointegration Analysis

The technique of cointegration can be utilized to see if there is evidence consistent with the hypothesis of a long-run equilibrium relationship among a set of integrated variables. A time series X_t is said to be integrated of order d, $X_t \sim I(d)$, if it becomes stationary after d differencings. A set of integrated variables are said to be cointegrated if the variables do not drift "too far" apart, and there exists an equilibrium relationship among them.

If a vector β is a cointegrating vector of a vector of n time series variables, X, then $\beta'X = Z_t$, where Z_t has a lower order of integration than do the variables in X. An X vector of n variables may have more than one cointegrating vector, but a maximum of $n-1$.[5] Z_t can be considered as defining the long-run equilibrium relationship(s) among the variables where nonzero values represent short-run deviations from the long-run equilibrium relationship.

For a set of cointegrated variables, it must be possible to estimate an error correction model specifying the short-run dynamics of the variables as they move toward a set of values consistent with their long-run equilibrium relationship. Cointegration/error correction models have a number of advantages over other techniques using time series data to estimate relationships among variables whose values are not stationary over time. It has been shown that with an ordinary least squares approach a spurious relationship may be estimated between two variables that are nonstationary; in this case parameter estimates will also be inconsistent.[6] An alternative way of dealing with nonstationary time series variables has been to estimate their relationship using first differences of the variables. Unfortunately, although this technique may remove the nonstationarity problem, data regarding the long-run levels of the variables is lost in the process.[7] Cointegration/error correction models not only allow us to capture the long-run

5. These cointegrating vectors may be arranged in an $n \times r$ matrix of rank r, where r is the number of cointegrating vectors.

6. See Banerjee, Dolado, Hendry, and Smith [1986] for more detailed discussions.

information conveyed in the data of nonstationary variables through the estimation of the cointegrating vectors, $\beta'X$, but also enable us to examine the short-run dynamics of a cointegrated system through the analysis of error correction parameters.

The step-by-step procedures in cointegration analysis are summarized below.

1. Initial Stationarity Tests

Test each of the individual variables to be considered in the cointegration analysis to ascertain if they are stationary, or if not, their order of integration (d) [denoted by $I(d)$]. The augmented Dickey-Fuller (ADF) tests are applied for this purpose.[8] However, as pointed out by Perron [1989], Christiano [1992], Banerjee, Lumsdaine, and Stock [1992], and Zivot and Andrews [1992], the standard ADF tests are not appropriate for the variables with obvious structural breaks. We use the methods of Zivot and Andrews [1992] to investigate the nonstationarity of variables with an apparent structural break.

The sequential ADF tests developed by Zivot and Andrews [1992] not only allow us to examine the order of integration for the variables with a structural break but also to test for a possible break-point rather than assuming it exists. Basically, their tests are represented by the following augmented regression equations:

$$\textit{Model A}: \ \Delta x_t = \mu + \theta DU_t + \beta t + \phi x_{t-1} + \sum_{j=1}^{p} c_j \Delta x_{t-j} + e_t,$$

$$\textit{Model B}: \ \Delta v_t = \phi v_{t-1} + \sum_{j=1}^{p} c_j \Delta v_{t-j} + e_t,$$

$$\textit{Model C}: \ \Delta x_t = \mu + \theta DU_t + \beta t + \gamma DT_t + \phi x_{t-1} + \sum_{j=1}^{p} c_j \Delta x_{t-j} + e_t,$$

7. See Durlauf and Phillips [1988]; Nelson and Plosser [1982]; Nelson and Kang [1981]; Maddala [1992, pp. 584-590]; Ouliaris, Park, and Phillips [1989]; Perman [1991, pp. 11-12]; and Stock [1987] for a more detailed discussion of these issues.

8. Since the standard ADF tests are well known, the descriptions of the tests are omitted here. Interested readers are referred to Dickey and Fuller [1979 and 1981] and Said and Dickey [1984] for the details.

where T_B is the possible break-point, the level dummy variable $DU_t = 1$ if $t > T_B$ and zero otherwise, and the slope dummy variable $DT_t = t - T_B$ if $t > T_B$ and zero otherwise. Models A and C are estimated by one-step regressions. The estimation of Model B has a two-step procedure, where the series v_t is the residual series from a regression of x_t on a constant, a time trend, and a slope dummy variable DT_t.[9] Note that if DU_t is not included, Model A turns out to be the standard *ADF* test with a time trend, and if both DU_t and the time trend are excluded, Model A becomes the standard *ADF* test without a time trend.

The sequential *ADF* procedure estimates a regression equation for every possible T_B within the sample and calculates the t statistics for the estimated coefficients. The null of nonstationarity is rejected in favor of the alternative of stationarity if ϕ is significantly different from zero. The chosen break-point for each series is that T_B for which the t statistic for ϕ is minimized. Critical values for these tests, which are greater (in absolute value) than those for the standard *ADF* tests, are tabulated in Zivot and Andrews [1992]. Model specification (*i.e.*, which model, A, B, or C, is appropriate) is determined by first running each series on Model C, with the possibility of both a slope and a level break. Model C is chosen if both dummy variables are significant. If only the slope dummy variable is significant, Model B is estimated. If the level dummy variable is significant, Model A is estimated.

For either standard *ADF* tests or sequential *ADF* tests, the choice of lag length p may affect the test results. We follow the procedure suggested by Campbell and Perron [1991]. Start with an upper bound, p_{max}, on p. If the last included lag is significant, choose $p = p_{max}$. If not, reduce p by one until the last lag becomes significant. We set $p_{max} = 8$ for the quarterly data we use.

2. Cointegration Tests

Conditional on finding that the variables hypothesized to be cointegrated are nonstationary, tests for cointegration of the variables are

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9. Perron and Vogelsang [1991] give an explanation of why Model B is estimated differently than Models A and C.

conducted. The cointegration testing procedures utilized in this paper were developed by Johansen [1988 and 1991] and Johansen and Juselius [1990]. This method gives consistent maximum-likelihood estimates of the entire cointegrating matrix and yields a likelihood ratio statistic regarding the maximum number of cointegrating vectors.

The first step in the Johansen cointegration test procedures is to estimate a vector error correction model of the following form:

$$\Delta x_t = \sum_{i=1}^{k-1} \Gamma_i \Delta x_{t-i} + \Pi x_{t-k} + \mu + \eta D_t + \varepsilon_t,$$

where the ε_t is a vector of white noise and D_t represents some seasonal or other dummy variables. The hypotheses of interest involve Π; if the X_t are cointegrated with (at most) r cointegrating vectors, then the rank of Π is r, and Π can be decomposed into the two matrices α and β. That is, $\Pi = \alpha\beta'$, where β is the matrix of cointegrating vectors and α contains the error correction coefficients ⟨Johansen [1991], p. 1553⟩. Thus, the Johansen procedure combines the estimation of and testing for the presence of cointegrating vectors with the estimation of error correction processes within the framework of a single unified vector error correction model.[10]

The Johansen methodology utilizes two test statistics regarding the number of cointegrating vectors, r. The null hypothesis for the trace test is that the number of cointegrating vectors is less than or equal to r, where r = successively 0, 1, 2, . . . , $n - 1$. In the *maximum eigenvalue test*, the null hypothesis is that the number of cointegrating vectors is some explicit number, say r, with the alternative hypothesis that the number of cointegrating vectors is equal to $r + 1$.

Conditional on the outcomes of one or both of the above tests consistent with the hypothesis of at least one cointegrating vector, the esti-

10. Monte Carlo studies have produced evidence consistent with the hypothesis that the Johansen simultaneous estimation procedure performs more reliably than an alternative two-step procedure suggested by Engle and Granger [1987]. See Maddala [1992, pp. 262-264 and 591-592], Gonzalo [1994], and Perman [1991, pp. 17-18].

mated coefficients in the cointegrating relationship are tested to see if the data are consistent with the hypothesis that they are significantly different from zero. Other hypotheses regarding the estimated coefficients of the cointegrating relationship may also be tested, using the maximum likelihood methodology. Again conditional on evidence consistent with the hypothesis of the existence of a cointegrating vector, the coefficients of the error correction term are then tested to see whether the data are consistent with the hypothesis that they are significantly different from zero; that is, whether there is a process whereby disequilibrium values adjust back toward the long-run equilibrium relationship(s).[11]

III. International Parity Conditions and Money Market Connections

This paper investigates the relationship(s) among three-month treasury bill rates for Canada, Mexico, and the United States. Later in the paper the relationship(s) among the monetary bases of these countries is also examined.

In the initial part of the analysis, it is implicitly assumed that uncovered interest rate parity holds with respect to the interest rates for each country pair. The uncovered interest rate parity theory basically argues that the domestic interest rate must be equal to the foreign interest rate for the same time period, adjusted for any expected change in the exchange rate between the two countries. It implies the following relationship between the yields on two *n*-period pure discount bonds:

$$(1+R_{n_t}^{D})^n = \frac{S_{t+n}^{E}}{S_t}(1+R_{n_t}^{F})^n(1+L_t), \tag{1}$$

where

$R_{n_t}^{D}$ = the domestic interest rate on an n-period bond at time t,

S_{n+t}^{E} = the expected spot exchange rate (units of domestic currency

11. The *Microfit 3.0* econometric package was utilized in this paper for the cointegration tests except for the last step of testing the significance of the estimated error correction coefficients. The *RATS* econometric package was used for that procedure and also for the standard *ADF* tests as well as the sequential *ADF* tests.

per unit of foreign currency) at time $t+n$,

S_t = the spot exchange rate at time t,

R^F_{n+t} = the foreign interest rate on an n-period bond at time t, and

L_t = a risk premium.

The relationship in Equation (1) is usually approximated as follows:

$$n(R^D_{n_t} - R^F_{n_t}) = E_t(e_{t+n}) - e_t + L_t, \tag{2}$$

where

$E_t(e_{t+n})$ = the log of the expected spot exchange rate (units of domestic currency per unit of foreign currency) at time $t+n$, and

e_t = the log of the spot exchange rate at time t.

The risk premium could be positive, negative or zero, depending on how an investor viewed the riskiness of holding an asset denominated in foreign currency relative to a domestic currency asset.[12] To test whether the uncovered interest rate parity condition holds in any meaningful empirical sense (as opposed to an identity, where L_t merely accounts for any differential between the two yields, adjusted for exchange rate differentials), it must be assumed that the liquidity risk premium is stationary over the time period under consideration. It follows that if L_t is stationary, it will have no effect on the existence of a long-run equilibrium relationship between the other variables (the two yields and the expected change in (the log of) the exchange rate).

To estimate the above relationship one needs a proxy for the unobservable expected change in the spot exchange rate. For $n = 1$, $E_t(e_{t+n}) - e_t$ becomes the expected first difference of the log of the exchange rate. It is often found that the exchange rate is integrated of order one. That is, the first difference of the exchange rate is stationary. In that case, it is reasonable to assume $E_t(e_{t+1}) - e_t$ is also stationary. Thus, it follows that

$$R^D_t - R^F_t = \text{a stationary term.} \tag{3}$$

12. Strictly speaking, for uncovered interest rate parity to hold the risk premium would equal zero.

It follows that if the interest rates of two countries are related in accordance with the interest rate parity relationship, there must be a mechanism by which the interest rate parity relationship is maintained. The flow of international financial capital is recognized in the literature as one factor that is affected by interest rate differentials and these flows, in turn, act to move international interest rates toward an equilibrium relationship. Such international capital movements, if not offset by domestic monetary policies, will also affect the money supply of each country.

For example, suppose that as a result of a tightened monetary policy in the United States, the interest rate in the United States increases relative to those in foreign countries. *Ceteris paribus*, the higher U.S. domestic interest rates will attract new foreign capital, for example, from Canada. If not offset by Canadian monetary policy, the outflow of financial capital from Canada will decrease the Canadian money supply which will, in turn, result in higher Canadian interest rates. Simultaneously, the entry of foreign capital into the United States will to some extent offset the initial effects of the monetary policy, exerting a downward force on U.S. interest rates. With free capital movements, this process theoretically will continue until the foreign and U.S. interest rates are once again in a parity relationship.[13] It follows from this analysis, then, that if the interest rates of two countries have a long-term equilibrium relationship we should expect that their money supplies will have a similar long-term relationship as well. Thus, the presence of a linkage between the money supplies of two countries would be additional evidence consistent with the hypothesis that the money markets of two countries are connected with one another.

IV. Empirical Results

In the first part of this section, cointegration analysis is applied to the interest rate linkage expressions in equation (3), above, to see if the resulting evidence is consistent with the hypotheses that (a) this condition holds bilaterally in the form of long-run equilibrium relationships for Canada and the United States and for Mexico and the United States, and (b) a corre-

13. For a more in depth discussion of this issue see Frenkel and Mussa [1985].

sponding long-run relationship(s) exists among the interest rates of these three countries. Later, cointegration analysis also will be utilized to investigate the relationships between the monetary bases of each country pair and among those of the three countries.[14]

A. Interest Rate Linkages

Both standard and sequential *ADF* tests are employed to test the nonstationarity of the interest rate variables. The sequential *ADF* tests were utilized because of the existence of an apparent structural break in the Mexican interest rate. It has an upward trend in period 1978:1 – 1988:1 from about 10% to 134%. In the second quarter of 1988, the rate dropped, by more than 80%, to 52% and after that it shows a downward trend from 52% to 14% in period 1988:3 – 1993:3. There is no such dramatic structural change in the data of Canada and the U.S.

As a primary test, we apply Model *C* of the sequential *ADF* tests to all the relevant variables. There is no significant structural break to be found for either the Canadian data or the U.S. data.[15] Therefore, the standard *ADF* tests seem appropriate for the variables of these two countries. The results, presented in Table 1, indicate that the null hypothesis that the interest rate of Canada or the U.S. was nonstationary cannot be rejected at the 5 percent level of significance, but the same null hypothesis could be rejected for the first differences of the interest rates of these two countries. Therefore, we may conclude that the interest rates of Canada and the U.S. were I(1). Table 1 also shows that the null hypothesis of nonstationarity could be rejected for the change in the log of the Canada/United States exchange rate at the 5

14. As indicated earlier in the paper, three-month treasury bill rates, expressed on an annual basis in percentage terms for Canada, the United States, and Mexico were the interest rate data utilized in the study. The International Monetary Fund's *International Financial Statistics* was the source of the data. The sample period for the Canada-United States relationship was the first quarter of 1973 until the third quarter of 1993. The sample period for the Mexico-United States relationship was the first quarter of 1978 also through the third quarter of 1993. These time periods were chosen so that the data would come from periods characterized by at least somewhat flexible exchange rates in each respective case.

15. The results are not reported but are available from the authors upon request.

Table 1
Augmented Dickey-Fuller (ADF) Stationarity Tests for the Canadian and the U.S. Interest Rates and the Change in Their Exchange Rate

	t Test Statistic[a] (H_0: x is Nonstationary.)	
Variable (Lags)	Without Trend	With Trend
	A. Interest Rate Levels	
$R^C(3)$	–2.141	–2.021
$R^{US}(5)$	–2.216	–2.342
	B. First Differences	
$\Delta R^C(1)$	–6.728**	–6.900**
$\Delta R^{US}(6)$	–3.570**	–3.754*
$\Delta S^C(2)$	–3.727**	–3.703*

Notes: * Significant at the 5 percent level of significance.
** Significant at the 1 percent level of significance.
a: The 1-percent and 5-percent critical values for the t statistic of the *ADF* tests are –3.51 and –2.89 respectively for the model without a time trend, and are –4.04 and –3.45 respectively for the model with a time trend. They are taken from Fuller [1976].

percent level of significance.[16] Since this is consistent with the hypothesis that the variable was stationary over the time period covered by this study, it will be omitted from the estimated cointegrating relationship.

For Mexico, when we apply Model C of the sequential *ADF* tests to the

16. The critical values listed in Table 1 are taken from Fuller [1976] for a pure autoregressive process. Schwert [1987] argues that pure autoregressive processes have different sample distributions than when the true process is an autoregressive integrated moving average (ARIMA) process. He derives critical values for only two lag structures (4 and 12), corresponding to the 5 percent level of significance for an ARIMA(0,1,1) process for a range of moving average parameters from –0.8 to 0.8. However, a comparison of the test statistic values reported in Tables 1 indicates that a correction of the critical values for the presence of a moving average parameter with a range of values from –0.5 to 0.5 would not affect the decision to accept or reject H_0 for any of the cases in Table 1 or in Table 2.

Table 2
Stationarity Tests for the Interest Rate of Mexico and for the Change in the Mexico/U.S. Exchange Rate

A. Sequential ADF Tests of Zivot and Andrews			
Variable (Lags)	Date of Break	t Test Statistic[a] (H_0: x is Nonstationary.)	Model
$R^M(4)$	3/88	–1.014	A
$\Delta S^M(0)$	4/87	–7.594**	A

B. Standard ADF Tests		
	t Test Statistic[b] (H_0: x is Nonstationary.)	
Variable (Lags)	Without Trend	With Trend
$\Delta R^M(4)$	–3.600**	–3.851*

Notes: * Significant at the 5 percent level of significance.
** Significant at the 1 percent level of significance.
a: The 1-percent and 5-percent critical values for the t statistic of the sequential *ADF* tests are –5.34 and –4.80, respectively for Model A. They are taken from Zivot and Andrews [1992].
b: See note to Table 1.

Mexican interest rate, a dummy variable was added to reflect the outlier value mentioned above for the second quarter of 1988. The results suggest that the slope dummy variable is not significant while the level dummy variable is significant for both the interest rate of Mexico and the change in the log of the Mexico/United States exchange rate, but neither of these two dummy variables is significant for the first difference of the interest rate. Therefore, Model A is estimated for the former two variables while the standard *ADF* tests are performed for the latter. The results are reported in Table 2. The t statistics for ϕ from the estimation of Model A show that the interest rate of Mexico was nonstationary but the log difference of the Mexico/United States exchange rate was stationary. The break date for the Mexican interest rate is found to be the third quarter of 1988. This reflects the fact that the variable had a structural change which turned the upward sloping trend to a downward sloping trend after the third quarter of 1988.

The break date for the change in the Mexico/United States exchange rate, the fourth quarter of 1987, is consistent with an observable trend break in this variable occurring in the first quarter of 1988. The t statistics for ϕ from the standard *ADF* tests indicate that the first difference of the Mexican interest rate was stationary. Thus, it could be concluded that the interest rate of Mexico was $I(1)$ with a structural break. Possible reasons for these structural breaks are discussed later in the paper.

The results of the estimation procedures for the cointegrating vectors and the error correction process are presented in Table 3.[17] Whenever the estimation involves the Mexican interest rate, a dummy variable reflecting the outlier of 1988:2 and a level dummy variable based on the results of Table 2 are included in the model. The results of the trace test are consistent with the hypothesis that there is one cointegrating vector for the Canada-United States interest rate relationship. However, no cointegrating relationship is found between the interest rates of Mexico and the U.S. When the tests include all three interest rates, only one cointegrating relation is found. The maximum eigenvalue test yields virtually identical results. The one difference between the results of the maximum eigenvalue test and those of the trace test is that for the interest rate relationship involving Canada and the United States or the relationship involving all three countries the null hypothesis that there was no cointegrating vector could be rejected at the ten, rather than the five, percent level of significance.

For the two cases where the results suggest the existence of a cointegrating vector, the estimated coefficients (*i.e.*, β's) are normalized on the Canadian interest rate. In both cases, the estimated coefficients for the Canadian variable and for the U.S. variable were significantly different from zero at the one or the five percent level of significance in the estimated relation-

17. The Bayesian Information Criterion (BIC) of Schwartz [1978] is used to select the number of lagged terms to include in the model to be estimated. These lags numbered six in the Canada-United States interest rate parity relationship, five in the Mexico-United States interest rate parity relationship, and three in the relationship involving the three interest rates. The critical values for the interest rate relationship come from Table D.2 of Osterwald-Lenum [1990]; these are applicable to the case where there is no deterministic trend in the data generating process but the variables in X may have deterministic trends. Also see Osterwald-Lenum [1992].

Table 3
Cointegration Relationship Tests Interest Rate Linkages

	Trace (5% Critical Values)			λ_{Max} (5% Critical Values)		
H_0:	$r=0$	$r\leq 1$	$r\leq 2$	$r=0$	$r=1$	$r=2$
Canada & The U.S.	18.292* (17.953)	4.777 (8.176)	N.A.	13.516 (14.900)	4.777 (8.176)	N.A.
Mexico & The U.S.	14.348 (17.953)	0.893 (8.176)	N.A.	13.455 (14.900)	0.893 (8.176)	N.A.
Canada & Mexico & The U.S.	33.278* (31.525)	14.219 (17.953)	0.087 (8.176)	19.059 (21.074)	14.133 (14.900)	0.087 (8.176)

Estimated Cointegration Coefficients

	β_{R^C}	β_{R^M}	$\beta_{R^{US}}$	Restriction
Canada & The U.S.	–1.000	N.A.	1.017	H_0: $\beta_{R^C}+\beta_{R^{US}}=0$
$\chi^2(1)^a$	(7.921)**		(7.198)**	$\chi^2(1)^a=0.010$
Canada & Mexico & The U.S.	–1.000	0.017	1.349	
$\chi^2(1)^a$	(4.264)*	(1.323)	(4.912)*	

Estimated Error Correction Coefficients

	$\alpha_{\Delta R^C}$	$\alpha_{\Delta R^M}$	$\alpha_{\Delta R^{US}}$	
Canada & The U.S.	0.275	N.A.	0.036	
$\chi^2(1)^a$	(6.804)**	(0.172)		
Q Statistic[b]	6.875	10.469		
Canada & Mexico & The U.S.	0.359	-0.924	-0.107	
$\chi^2(1)^a$	(4.003)*	(1.209)	(0.384)	
Q Statistic[b]	7.626	13.108	16.322	

Notes: N.A. Not applicable.

* Significant at the 5% level of significance.

** Significant at the 1% level of significance.

a: The 1-percent and 5-percent critical values of the $\chi^2(1)$ statistic are 6.635 and 3.841, respectively, for H_0: β_j (or α_j) = 0 and for H_0: $\beta_{R^C}+\beta_{R^{US}}=0$ as well.

b: The 1-percent and 5-percent critical values of the Q statistic (12 lags) are $\chi^2(12)=26.217$ and 21.026, respectively, for H_0: ε_t is white noise.

ships, but the coefficient for the Mexican variable was not significant. Moreover, consistent with the interest rate parity hypothesis, the null hypothesis that the sum of the estimated coefficients in the bilateral cointegrating relationship was equal to zero could not be rejected for the Canada-U.S. interest rate relationship. These results are further evidence consistent with the hypothesis that the two variables had a long-run equilibrium relationship during this period of time.[18] The Q statistics for the error correction model are such that the null hypothesis that ε_t was a white noise process could not be rejected.[19] Moreover, the error adjustment term for the ΔR^C variable was significantly different from zero at the one or five percent level of significance while that of the ΔR^{US} variable was not. The outcomes of these significance tests are consistent with the hypothesis that in the linkage between the interest rates of Canada and the United States, the Canadian variable is relatively endogenous while the U.S. interest rate is weakly exogenous and thus seems to have been a driving force in the co-movements of the two variables.[20] On the other hand, the results do not support any close relationship between the interest rate of Mexico and that of the United States or Canada.

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18. To investigate the robustness of the test results, the residual based Engle-Granger tests were also conducted for the two bilateral relationships. For the residuals from the regression of the interest rate of Canada on that of the U.S., the null hypothesis of no cointegration was rejected at the five percent significance level because the *ADF* test statistic was found to be –3.54 while the five percent critical value was –3.17 ⟨See Engle and Granger [1987]⟩. The estimated coefficient of the U.S. interest rate in the cointegrating regression was 0.955. These results are very similar to those reported in Boothe [1991, p. 600]. In the regression of the interest rate of Mexico on that of the U.S., the dummy variables mentioned in the text were included. The corresponding *ADF* test statistic was found to be –1.23 and thus we fail to reject the null at any reasonable level of significance. The lag lengths of the *ADF* tests were again chosen on the basis of Campbell and Perron's method. They were set equal to four for both regression equations. These results of the Engle-Granger tests are basically consistent with those of the Johansen tests.
19. The Q statistic reported here is the modified Q statistic of Ljung and Box [1978], adjusted for finite sample sizes. There is some disagreement regarding the power of the Q statistic in testing for model adequacy in the case of autoregressive models. See Greene [1993, pp. 557-558] and Maddala [1992, pp. 540-542].
20. See Engle, Hendry, and Richard [1983]; and Johansen [1992] for a discussion of weak exogeneity.

B. Monetary Base Linkages

As explained earlier, if the interest rates of two countries are related, ceteris paribus, we would expect changes in their money supplies to be correlated as well. Therefore, if one is testing whether the interest rates of two countries are cointegrated and if such an interest rate linkage partly reflects the interdependence of two countries' monetary policies, it is also appropriate to investigate whether there exists a long-run relationship between the money supplies of the two countries. Evidence of a linkage between the money supplies of two countries would be further evidence consistent with a hypothesis of interrelated interest rates. As anticipated, the results obtained from this research are consistent with those regarding the existence of cointegrating relationships among the interest rates of the three countries.

In this study, a monetary base measure is employed to represent the money supply variable.[21] Cointegrating relationships between the monetary bases of Canada and the United States, between those of Mexico and the United States, and among those of the three countries are estimated to see if the outcomes of these procedures would be consistent with the results in the estimated interest rate relationships. Since the data were consistent with the existence of a long-run equilibrium relationship between the interest rates of Canada and the United States but not between those of Mexico and the United States, one might anticipate that a cointegrating relationship would be found in the monetary bases in the former pair of countries, but not in the latter pair.

Tables 4 and 5 present the results of the standard and sequential *ADF* tests for stationarity in the monetary base variables using exactly the same methodology and the same sample periods as those for the interest rate

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21. The monetary base measure used in this study is *reserve money* as defined in the *International Financial Statistics.* This figure represents the *liabilities of the monetary authority* including currency in circulation (held outside banks), reserves at banks (vault cash plus required and excess reserves deposited with the central bank), and demand deposits of the rest of the domestic economy, excluding the central government. See *International Financial Statistics: Supplement on Money*, Supplement Series No. 5, Washington, D.C.: International Monetary Fund, 1983, p. vi.

Table 4
ADF Stationarity Tests for the Canadian and the U.S. Monetary Bases

	t Test Statistic[a] (H_0: x is Nonstationary.)		
	Variable (Lags)	Without Trend	With Trend
A. Monetary Base	$ln\ M^C(2)$	–2.282	–3.081
	$ln\ M^{US}(3)$	–0.849	–2.886
B. Monetary Base First Differences	$\Delta ln\ M^C(1)$	–3.156*	–4.302**
	$\Delta ln\ M^{US}(2)$	–4.360**	–4.341**

Notes: * Significant at the 5 percent level of significance.
** Significant at the 1 percent level of significance.
a: See note to Table 1.

variables. The variables utilized for the monetary base levels were the natural logarithms of the respective monetary bases, using seasonally adjusted quarterly data.

As indicated in Table 4, the augmented Dickey-Fuller tests are consistent with the hypothesis of nonstationarity of the levels of the Canadian and the U.S. monetary base variables and reject the hypothesis of nonstationarity of their first differences.[22] In Table 5, it is reported that, by applying the sequential *ADF* tests to the Mexican monetary base variable, Model C (the model with a level and a slope dummy variable) is found to be appropriate for the level of the variable while Model A (the model with a level dummy variable only) is appropriate for the first difference of the variable. The break date is found to be the second quarter of 1988. The results fail to reject the null hypothesis of nonstationarity of the levels of the variable but are able to reject the hypothesis of nonstationarity of its first difference. Thus, it seems reasonable to conclude that the data are consistent with the hypothesis that the monetary base variables were $I(1)$ so that the cointegration analysis could proceed.

22. A comparison of the test statistic values reported in Tables 4 indicates that a correction of the critical values (Schwert [1987]) for the presence of a moving average parameter with a range of values from –0.5 to 0.5 would not affect the decision to accept or reject Ho in any of these cases.

Table 5
Sequential ADF Stationary Tests for the Monetary Base of Mexico

Variable (Lags)	Date of Break	t Test Statistic[a] (H_0: x is Nonstationary.)	Model
$ln\ M^{MX}(4)$	2/88	–3.8433	C
$\Delta ln\ M^{MX}(4)$	2/88	–10.5713**	A

Notes: ** Significant at the 1 percent level of significance.[a]

a: The 1-percent and 5-percent critical values for the t statistic are –5.34 and –4.80 respectively for Model A, and are –5.57 and 5.08 respectively for Model C. They are taken from Zivot and Andrews [1992].

The results of the cointegration tests are presented in Table 6.[23] Whenever the estimation involves the Mexican monetary base variable, a level and a slope dummy variable based on the results of Table 5 are included in the model. With respect to the Canada-United States relationship and the relation of the three countries' variables, the trace tests and the maximum eigenvalue tests are consistent with the hypothesis that one cointegrating relationship exists between the monetary base variables. The estimated coefficients of the cointegrating vectors were significantly different from zero at the one percent level of significance for the Canadian and the U.S. variables but not significant for the Mexican variable. With respect to the error adjustment model, the Q statistic is such that the null hypothesis that ε_t was white noise could not be rejected at the five percent level of significance. The estimated error correction coefficient for the Canadian monetary base was significantly different from zero at the one percent level of significance whereas the estimated error adjustment term for the U.S. monetary base was not, again consistent with the hypothesis that the Canadian monetary base is relatively endogenous while the U.S. monetary base is a

23. Based on the Schwartz' BIC results, four lags are included in the estimation of the monetary base relationship between Canada and the U.S. and three lags are included in the estimation of the relationships between Mexico and the U.S. and among the variables of the three countries. The critical values for the monetary base relationships come from Table D.1 in Osterwald-Lenum [1990]; these are applicable to the case where the variables in X as well as the data generating process have linear deterministic trends. Also see Osterwald-Lenum [1992].

Table 6
Cointegration Relationship Tests Monetary Bases

	Trace (5% Critical Values)			λ_{Max} (5% Critical Values)		
H_0:	$r=0$	$r\leq 1$	$r\leq 2$	$r=0$	$r=1$	$r=2$
Canada & The U.S.	20.678* (15.410)	0.053 (3.762)	N.A.	20.625* (14.069)	0.053 (3.762)	N.A.
Mexico & The U.S.	8.262 (15.410)	3.217 (3.762)	N.A.	5.045 (14.069)	3.217 (3.762)	N.A.
Canada & Mexico & The U.S.	52.467* (29.680)	13.386 (15.410)	3.530 (3.762)	39.081* (20.967)	9.856 (14.069)	3.530 (3.762)

Estimated Cointegration Coefficients

	β_{M^C}	β_{M^M}	$\beta_{M^{US}}$
Canada & The U.S.	−1.000	N.A.	0.643
$\chi^2(1)$[a]	(17.226)**		(11.985)**
Canada & Mexico & The U.S.	−1.000	−0.054	0.939
$\chi^2(1)$[a]	(16.105)**	(3.120)	(12.732)**

Estimated Error Correction Coefficients

	$\alpha_{\Delta M^C}$	$\alpha_{\Delta M^M}$	$\alpha_{\Delta M^{US}}$
Canada & The U.S.	0.039	N.A.	0.009
$\chi^2(1)$[a]	(18.303)**		(1.143)
Q Statistic[a]	8.859		6.231
Canada & Mexico & The U.S.	0.231	−0.290	−0.077
$\chi^2(1)$[a]	(16.049)**	(0.732)	(2.183)
Q Statistic[a]	12.797	16.282	12.329

Notes: a: See notes to Table 3 for the meanings and the critical values of the $\chi^2(1)$ statistic as well as of the Q statistic.

dominant variable in this relationship.

These results, consistent with the cointegration of the Canadian and U.S. money supplies as well as of the Canadian and U.S. interest rates, are also given credence by the findings of other research regarding the relationship between the Canadian and U.S. money markets. For example, Ahmad and Sarver [1994, p. 333] obtained evidence consistent with the conclusion that "innovations in the U.S. market account for about 52 percent of Canadian variance." Moreover, Crow [1988] and Booth, Mustafa, and Akgiray [1990] state that the Bank of Canada in the past has promoted exchange rate stability by intervening in the foreign exchange markets while allowing domestic interest rates to vary. Such a policy would certainly be consistent with the results obtained in this study.

As was expected, the trace and maximum eigenvalue tests are not consistent with the hypothesis of a cointegrating relationship between the Mexican and United States monetary bases. Such a result is consistent with the hypothesis that variables other than U.S. monetary policy and interest rates were the primary factors influencing the Mexican money supply and interest rates.[24]

Undoubtedly, the period under study in this paper was one during which some major changes occurred in the Mexican economy. From an annual rate of 11.2 percent in 1975, the rate of inflation rose to 98.8 percent in 1982 and as high as 159.2 percent in 1987.[25] The increasing rate of inflation in 1976 accompanied by a decline in the rate of growth of real GDP, and a serious decrease in the Banco de México's reserves led to speculation against the peso and capital flight, estimated at $4 billion during 1976 (see Ramírez,

24. The residual based Engle-Granger tests were again conducted for the two bilateral relationships of the monetary bases. For the pair of the Canadian-U.S. monetary bases, the null hypothesis of no cointegration was rejected at the five percent level of significance when the ADF test statistic was found to be –3.58. The estimated coefficient of the U.S. monetary base in the cointegrating equation was 0.718. The ADF test statistic for the residuals from the regression of the Mexican-U.S. monetary bases was -2.42, and thus we were unable to reject the null of no cointegration. The lag lengths of the ADF tests were chosen equal to three for both cases based on the method of Campbell and Perron. The results are again consistent with those of the Johansen tests.

25. Nacional Financiera, *La economia mexicana en cifras* [1990, p. 529].

p. 85, and Solís [1982], p. 344⟩. Consequently, President Echeverría was forced to devalue the peso on August 31, 1976.

In order to obtain a loan from the IMF, the López Portillo administration agreed to a three-year package of austerity measures, implemented in 1977 with corresponding negative effects on employment and real minimum wages. However, these policies were virtually abandoned in 1978. The discovery of large oil reserves in Tabasco and Chiapas had convinced the government that oil would be the solution to Mexico's economic problems. Subsequently, in 1979, the administration announced its Global Development Plan (Programa de Desarrollo Global), calling for continued growth in employment and income, combined with a more equitable distribution of the benefits of growth. Between 1978-1981, real output rose rapidly, and employment and real minimum wages also rose. However, the country became increasingly dependent on oil profits as well as the foreign exchange generated from the sales of petroleum. The public sector debt grew rapidly, and budget deficits were financed by the monetization of the debt as well as by borrowing from private and public foreign sources.

By 1982, the money supply (M1) was growing at an annual rate of 61.9 percent and the external public debt had grown to 61.4 percent of GDP. The 1981 fall in oil prices combined with a recession in the United States, capital flight, and rising interest rates (much of the foreign debt was financed at floating rates) resulted in another foreign exchange crisis. In August of that year the López Portillo administration once again devalued the peso. That move was followed by the imposition of exchange controls and the nationalization of the banking system in September.[26]

After taking office in December 1982, the de la Madrid Hurtado administration established a dual exchange rate system: a free rate applicable to relatively inconsequential transactions (for example, expenditures by tourists), and a *controlled rate*. The latter rate was applicable to export receipts, most payments by maquiladoras, payments for some imports, and certain other government-approved transactions. The amounts of foreign exchange in the free market sold by banks to individuals and business firms were also limited, and only exchange facilities in international airports or the border area

26. See Ramirez [1989, pp. 85-95] for a more detailed account of these events.

were allowed to sell foreign currency.[27]

President de la Madrid also authorized the sale of thirty-four percent of the banks to the public, abandoned López Portillo's Global Development Plan, and implemented another IMF stabilization plan embracing the principles of fiscal austerity, liberalization and restructuring of the domestic economy, export promotion, and promotion of foreign direct investment. Nevertheless, the public sector deficit began to increase once more in 1985, partly the result of expansionary policies undertaken for political reasons combined with another substantial fall in world oil prices. The rate of inflation fell from 98.8 percent in 1982 to 59.2 percent in 1984, but rose once again to 105.7 percent in 1986 and 159.2 percent in 1987.[28] In the midst of this turmoil, Mexico made a decision to join the General Agreement on Tariffs and Trade, and was admitted in July 1986. Also, the controlled peso exchange rate was devalued by 60 percent. Capital flight continued, and it was estimated that the $12.1 billion of Mexican deposits in the United States in 1985 exceeded those of Canada, the United Kingdom, Italy, and France, *combined* ⟨Ramírez [1989], pp. 105-107 and 111⟩.[29]

With such dramatic swings in the inflation rate and subsequent devaluations of the peso, it is not difficult to imagine that neither Mexican nor foreign investors were much influenced by relative movements in the United States and Mexican short-term interest rates. It appears that many potential investors viewed the riskiness associated with Mexican financial assets denominated in pesos as simply too great to be offset by any interest rate premiums in Mexico.

In December of 1987 the Pacto de Solidaridad Económica (Economic Solidarity Pact or PSE) was implemented. The purpose of this agreement (among representatives of the public, business, labor and farm sectors) was to eliminate inflationary expectations through price controls, to limit wage rate increases, and to control the rate of depreciation of the peso. Price

27. International Monetary Fund, Annual Report on Exchange Arrangements and Exchange Restrictions, 1983, p. 321.

28. See Ramirez [1989, pp. 97-101] and Nacional Financiera, La economia mexicana en cifras [1990, p. 529].

29. Other summaries of Mexican economic events during this period can be found in Philip [1988] and Looney [1985].

increases became subject to monthly reviews beginning in March 1988, which is consistent with the observed outlier in the Mexican interest rate series in the second quarter of 1988 and subsequent structural break in the third quarter of that year. The implementation of the Pacto is also consistent with an observable trend break in the Mexico/United States exchange rate in the first quarter of 1988 as well as a structural break in the Mexican monetary base in the second quarter of 1988.[30]

After taking office in 1988, the Salinas de Gortari administration succeeded in substantially reducing inflation, to an annual rate of 11.9 percent in December of 1992.[31] Domestic interest rates also fell dramatically after 1987, and in November of 1991 all foreign exchange controls were eliminated.[32] Nevertheless Dornbusch [1990, p. 155], speaking of the prospects for the return of capital during the Salinas administration, stated:

> Investors have an option to postpone the return of flight capital and they will wait until the frontloading of returns is sufficient to compensate for the risk of relinquishing the liquidity option of a wait-and-see position. This is the case even when interest rates are high and rewarding. Moreover, when capital does return it stays highly liquid, sitting so to speak in the parking lot (or on the tarmac), with the engine running.

In other words, it is not realistic to expect that investor confidence in the stability of the Mexican financial system on a comparable basis with what exists in Canada and the United States will be restored quickly, particularly given the events of late 1994 and early 1995.

V. Conclusions

30. The PSE was renamed the Pacto para la Estabilidad y el Crecimiento Economico (Pact for Stability and Growth or PECE) in late 1988. For more information on the PSE (PECE) see Ramirez [1993, p. 179], and Banco Nacional de México, *Review of the Economic Situation of Mexico* (March 1988, pp. 104-105; July 1988, pp. 273-275; September 1988, pp. 369-370; July 1989, pp. 281-282; and January 1992, pp. 2-3).
31. Nacional Financiera, *La economia mexicana en cifras* [1992, p. 150].
32. Ibid., pp. 122-123; and International Monetary Fund, *Exchange Arrangements and Exchange Restrictions*, 1992, p.319.

This study takes a comprehensive view of the linkages between money markets, including both interest rates and money supply, across three countries: Canada, Mexico, and the United States before the NAFTA was implemented. The results of this study with respect to the interdependence of the Canadian and United States financial sectors are consistent with the findings of Boothe [1991] regarding interest rate parity between the two countries. They are also consistent with the conclusions of Batten and Ott [1985] and the later Sheehan [1987] research regarding the impact of United States monetary growth on that of other countries, including Canada.[33] In addition, this study reveals that the U.S. variables have been a dominant force in the respective relationships with the Canadian variables.

As far as we know, no other studies have been undertaken using cointegration analysis to examine the nature of the relationship between the United States and Mexican financial markets or to investigate the integration of the financial markets of the three NAFTA countries. Our finding is consistent with the hypothesis that the financial markets of Mexico were not closely connected with those of the U.S. or Canada. Clearly, in the fifteen years prior to the implementation of NAFTA in 1994, the Mexican financial markets were not as open to international investors as were those of some other major industrialized nations. No futures market existed in Mexico, and, technically, the peso was not freely convertible. Furthermore, only Mexican banks could make loans in pesos, and borrowing in pesos to deposit in dollars was relatively expensive.[34] However, as Mexico's financial system, as well as its economy, becomes more open and more stable in the future accompanied by the implementation of the NAFTA, it seems reasonable to hypothesize that its financial markets will further develop and become more closely related to those of the United States and Canada.

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33. These results are also consistent with the results of a much earlier study by Caves and Reuber, *et. al.* (1971, especially Chapter 3), utilizing Canadian data from 1951-1962.
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