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INSTITUTE FOR ECONOMIC RESEARCH

QUEEN'S UNIVERSITY

Kingston, Ontario, Canada K7L 3N6

NATIONAL SAVINGS AND DOMESTIC INVESTMENT IN THE LONG RUN: Some Time Series Evidence for the U.S. and Canada

Glenn Otto and Tony Wirjanto*

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*Department of Economics, Queen's University, Kingston, Ontario, Canada K7L 3N6. We would like to thank, without implicating, Michael Devereux, Allan Gregory, James MacKinnon, Arman Mansoorian, Doug Purvis and Gregor Smith for helpful comments.

Abstract

An empirical puzzle in international finance is the finding of a high correlation between national savings and domestic investment. This result is widely interpreted as evidence of low international capital mobility. In this paper we examine the long run behaviour of national savings and domestic investment for the United States and Canada in the time series context, by testing for evidence of cointegration between the two series. Overall we find little evidence of a cointegrating relationship between the levels of national savings and domestic investment, except for Canada during the Bretton Woods era.

1. Introduction

Since the original paper by Feldstein and Horioka (1980) there has been a large amount of effort, both theoretical and empirical, directed towards investigating potential relationships between national savings and domestic investment. Interest in the area was initially generated by Feldstein and Horioka's empirical finding that, for OECD countries, national savings (ie. private and government savings) appear to be highly correlated with domestic investment. This result is typically interpreted as evidence of low international capital mobility. Feldstein and Horioka's findings have been confirmed by numerous other studies,[1] using both pooled cross-section and time series data and pure time series data.

Since international capital is widely believed to be highly mobile, the close relationship between aggregate measures of national savings and investment is viewed as a puzzle. One response to the puzzle has been to develop models that give rise to such a correlation, despite perfect capital mobility, see Obstfeld (1986), Murphy (1986) and Cardia (1988). The purpose of this paper is to study the long run relationship of time series data on savings and investment. Essentially we seek to take advantage of the fact that they are likely to be non-stationary, due to the presence of a unit root. If this is the case, it is of interest to test for evidence of cointegration between national savings and investment. Finding cointegration

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[1] Cross-section studies include Feldstein and Horioka (1980), Feldstein (1983), Fieleke (1982), Penati and Dooley (1984), Murphy (1984), Caprio and Howard (1984), Summers (1985) and Dooley, Frankel and Mathieson (1987). Time series studies are less common, but include Sachs (1981), Frankel (1985) and Obstfeld (1986).

between the two time series suggests that there is a long run co-movement between the two variables. We would view this result as being favourable to the hypothesis of capital immobility as formulated by Feldstein and Horioka (1980).

In section 2 of this paper we discuss some of the econometric problems that can arise in the standard approach to testing the relationship between savings and investment. We argue that testing for cointegration is robust to virtually all of these difficulties. The technique is applied to quarterly time-series data for the United States and Canada, and the empirical results are presented in section 3. Briefly our results suggest that for the U.S. there is little support for the existence of a long run co-movement between national savings and domestic investment. For Canada there is evidence of cointegration between the two variables during the Bretton Woods regime, but this relationship appears to breakdown with the advent of floating exchange rates. In section 4 we conclude by indicating some potential refinements and extensions to this work.

2. Econometric Issues

Following Feldstein and Horioka (1980) the standard approach to testing for the degree of international capital mobility is to use a model in terms of saving and investment rates:

$$\left(\frac{I}{Y}\right) = \alpha + \beta \left(\frac{S}{Y}\right) + e \qquad (1)$$

In (1) I is a measure of private domestic investment, S is national savings and Y is a measure of national income, typically gross national product (GNP).[2] Under perfect capital mobility there should be no systematic relation between national savings and domestic investment. National savings should respond to international investment opportunities, while investment for an economy can be financed from the pool of international capital. Under perfect capital mobility and the small country assumption, the value of β should be close to zero. Alternatively if capital is not highly mobile then national savings should be highly correlated with domestic investment, and this will be reflected in a value of β that is greater than zero and possibly close to unity.

Although model (1) is a relatively simple specification, obtaining consistent estimates of α and β and their respective standard errors has been the focus of much of the empirical literature on this issue. Under standard econometric assumptions about the covariance stationarity of the variables in (1), there are a number of possible pitfalls that are likely to arise in any attempt to obtain consistent estimates of the parameters in (1). As has been recognized in most studies it is unrealistic to assume that $(\frac{S}{Y})$ is orthogonal (even asymptotically) to the error term in (1). Factors such as measurement errors in national accounts data and the endogeneity of national savings - due either to the procyclical nature of savings and investment or to the endogeneous response by the government to current account imbalances - all suggest that the regressor and error term

[2] Typically the series for S and I are scaled by GNP as a means of deflating the data and ensuring common units of measurement in cross-sectional studies. In time series studies dividing by GNP is a relatively simple way of attempting to control for the effects of business cycle fluctuations on the estimates of β.

in (1) will be correlated. In this case ordinary least squares (OLS) estimates will be biased and inconsistent.

Recent papers by Dooley, *etc.* (1987) and Frankel (1989) conclude that there is little reason to believe, either for time-series or cross-sectional data, that savings or investment can be treated as exogenous variables. The importance of this point is that even under conditions of perfect capital mobility the estimate of β obtained by OLS may differ from zero. One way of minimizing the potential biases that may arise due to the lack of orthogonality between the regressor and the error term in (1) is to estimate the model by an instrumental variable (IV) procedure. However when such an approach is employed, for example by Feldstein and Horioka and also by Dooley, Frankel and Matheison, it has not tended to alter the general finding of low capital mobility.

A second problem that has received less attention in empirical studies, is that the error term in (1) is unlikely to be well-behaved. If (1) is specified for cross-sectional data then the error term is likely to exhibit heteroscedasticity - which requires some adjustment to the usual formulas for OLS and IV standard errors, if valid inferences are to be made, see White (1980). From a time series perspective, (1) is a purely static regression model and the omission of any dynamics will typically result in serially correlated errors. This point is recognized by Obstfeld (1986) who does not estimate (1), but rather directly estimates the correlation coefficient between $\Delta(\frac{S}{Y})$ and $\Delta(\frac{I}{Y})$, where Δ is a quarterly first difference operator.

All of the aforementioned complications in estimating and testing equation (1) arise whenever it is assumed that $(\frac{S}{Y})$ and $(\frac{I}{Y})$ are jointly covariance stationary. Even if this is a reasonable assumption to make, Fieleke (1982) argues that the use of saving and investment rates may be inappropriate. He recommends the use of the levels of savings and investment instead, but recognizes the inferential problems they pose. In the time series context, non-stationarities of the variables will require the use of non-standard econometric methods. Following the work of Nelson and Plosser (1982) there is a body of empirical evidence that suggests many macroeconomic time series may be non-stationary due to the presence of a unit root. A time series that contains a single unit root is said to be integrated of order one, ie. I(1), and needs to be differenced once to induce stationarity. Recent work on time series regression with non-stationary variables, notably by Phillips (1986), Phillips and Durlauf (1986) and Engle and Granger (1987), points to an alternative means of testing for the presence of any long-run relationship between national savings and investment, based upon the concept of cointegration. The advantage of such an approach is that it is robust to many of the econometric problems that have confronted earlier studies.

Generally any linear combination of two I(1) series will itself be I(1). However following Engle and Granger (1987) two series are said to be cointegrated, if they are individually I(1), but there exists a linear combination of the series that is I(0). This implies that while the two series may drift apart in the short-run there exists a long-run or equilibrium relationship between them. It seems to us that testing for evidence of a long-run relationship between national savings and domestic investment is essentially what most empirical studies have been concerned with. As evidence of this consider the following quotation from Feldstein and

Horioka (1980, p323, our emphasis).

" ....., we view the investment-savings function of equation (1) as a long-run relation in which intercountry differences in savings rates reflect basic structural differences among countries. In this context, the estimate of β can represent the effect on investment of sustained changes in savings rates. "

To make the above discussion more explicit consider the following saving-investment regression in levels,

$$I_t = \delta_o + \delta_1 S_t + u_t \qquad (2)$$

where I_t and S_t are both I(1) processes. If the error term in (2) is also I(1) then the model is an example of a spurious regression, discussed in Phillips (1986) and Durlauf and Phillips (1988). In this case, the estimated parameter of δ_1 is inconsistent and the standard significance tests are biased towards rejection of no relationship and non-rejection of a spurious relationship. Moreover an unusually high R^2 and low Durbin Watson statistic will characterise a spurious regression such as (2).

We can write equation (2) in terms of the magnitude order in probability of the sample moments of the variables as: $O_p(T) = O_p(1) + O_p(T)$. Note that the equality of the order of magnitude between I_t and S_t gives rise to the possibility of cointegration between the two series. This will occur when the error term in (2) is I(0). In this case δ_1 is the cointegrating coefficient which is a measure of the long-run or low frequency relationship between

I_t and S_t. From Stock (1987) it is known that OLS applied to (2) will yield a consistent estimate of the cointegrating coefficient. In fact OLS is super-consistent since ($\hat{\delta}_1 - \delta_1$) is of $O_p(T^{-1})$. However Monte Carlo evidence reported by Banerjee *et al.* (1986) suggests that the bias term can be substantial in some cases. Neverthelesss the significance of this convergence result is that in (2) the OLS estimate of δ_1 is robust to: correlation between S_t and u_t (ie. due to the endogeneity of S_t or measurement errors), omission of relevant short-run dynamics, such as lags of ΔS_t and ΔI_t and serial dependence and heterogeneity in u_t.[3]

One way of avoiding the spurious regression problem is to take differences of non-stationary variables. However by focussing the analysis on differenced variables, information about the long run relationships is inadvertently lost. A less drastic measure is to use the cointegration methodology as a statistical tool for retaining variables in levels. Testing for cointegration can be seen as a pre-test designed to avoid the spurious regression problem discussed above. More specifically we are interested in testing the null of no cointegration between national savings and domestic investment. A non-rejection of the null is taken as evidence against long run co-movements between the two variables, casting doubt on the hypothesis of long run capital immobility as formulated by Feldstein and Horioka (1980).

3. Empirical Results for United States and Canada

All of the series used in the following empirical work are constructed

[3] For further technical details the reader should consult Phillips and Durlauf (1986).

from United States and Canadian national accounts. The data are quarterly, seasonally adjusted and cover the period 1956:1 to 1987:4. Both nominal and constant price data are used. For further details see the data appendix.

We begin the empirical analysis by testing whether the time series measures of national saving and gross domestic investment for the U.S. and Canada are non-stationary due to the presence of a unit root. To test for a unit root the augmented Dickey-Fuller (ADF) test[4] is used. Initially we allow for one and four lags of the first-difference of the dependent variable to ensure serially uncorrelated residuals in the test regression; where the presence of serial correlation is tested using a Lagrange Multiplier test. If four lags are found to be insufficient the lag length is increased until uncorrelated residuals are obtained. Statistics are computed allowing for both an estimated drift and time-trend and for drift alone. The results are presented in Table 1 for with representations of each series: in real terms and in nominal terms.

The results in Table 1 indicate that for the nominal series it is not possible to reject the null of a unit root for either U.S. or Canadian data. In fact excluding the time-trend from the test regression produces large positive test statistics (not reported), which would lead to rejection of the null of a unit root in favour of the explosive alternative.

When the constant price measures of the series are examined, the results

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[4] The augmented Dickey-Fuller tests in this paper are based on the t-value of the coefficient on y_{t-1} in the regression,

$$\Delta y_t = \alpha + \beta t + (1-\rho) y_{t-1} + \sum_{i=1}^{q} \gamma_i \Delta y_{t-i} + e_t$$

where q is chosen sufficiently large to whiten the error term e_t.

are less clear-cut. For the U.S., the test statistics from a test regression including both drift and trend, suggest that both RI_d and RS_n are stationary. However when the trend is excluded it is not possible to reject the null of a unit root for either series. For Canada there is less evidence against the unit root hypothesis when trend is included, and when the time trend is excluded the unit root hypothesis is rejected - but in favour of the explosive alternative.

While the results of the unit root tests are somewhat mixed, we consider there to be sufficient evidence that the series for national savings and domestic investment are non-stationary, to proceed with the cointegration tests. The approach used is that developed by Engle and Granger (1987), which involves estimating the static or cointegrating regression by OLS and testing if the residuals are I(1) or I(0), using the ADF test. The null hypothesis is that there is no cointegration, which should be interpreted to mean there is no long-run relationship between national savings and domestic investment.

Table 2 presents the static regression, which is estimated by OLS and the results from an ADF test of whether the residuals from the static regression contain a unit root. On the basis of the four ADF statistics reported in Table 2, it is not possible to reject the null hypothesis of no cointegration between national savings and gross domestic investment in nominal or real terms, for either the U.S. or Canada. It is also evident that the estimated saving-investment regressions reported in Table 2 feature high R^2 statistics and unusually low D-W statistics. While these are not valid tests they are characteristic of spurious regressions.

Since the estimated parameters of the static regression model have a non-degenerate limiting distribution, under the null hypothesis of a spurious relationship, there is no appropriate way of interpreting the magnitude of the estimated coefficients. However in the absence of cointegration between the level of national savings and domestic investment it is clear that over the long run the two series can drift apart without bound. We would therefore interpret this result as an evidence against the capital immobility hypothesis as formulated by Feldstein and Horioka (1980).

One possible criticism of the above results is that we make no allowance for the effects on the degree of capital mobility of the change in exchange rate regime that occurred with the breakdown of the Bretton Woods agreement in the early 1970's. The degree of capital mobility seems likely to be lower under the earlier period of fixed exchange rates, when many industrialized countries had capital controls and other restrictions on capital mobility. In fact some previous studies have suggested that exactly the opposite is the case. Dooley, Frankel and Mathieson (1987) present evidence that national savings and domestic investment are more highly correlated during the period since 1973, than before. We investigate this additional puzzle in Table (3).

For the U.S. the sample is split at 1971:2, which is when the U.S. abandoned fixed exchange rates. The results of testing for unit roots and cointegration over the two sample periods are presented in Table 3. For neither sample period is it possible to reject the null hypothesis of no cointegration between national savings and domestic investment, in either real or nominal terms. There is no evidence of a closer relationship between U.S. national savings and domestic investment over the floating period than over the fixed period. In fact the tests for cointegration yield considerably smaller test statistics (ie. a larger negative value) over the fixed exchange rate period compared to the floating period, however not small enough to be significant at the 5 % level of significance.

The results for Canada are different to those obtained for the United States. The sample for Canadian data is split in 1970:2 to reflect the earlier move to floating exchange rates.[5] The evidence from both the real and the nominal data suggest that national savings and domestic investment were cointegrated during the Bretton Woods period. However this cointegrating relationship appears to have disappeared during the period of generalized floating exchange rates. This finding is the opposite to those reported by Dooley et. al., but is consistent with prior expectations that increasing capital mobility will tend to weaken the link between national saving and domestic investment.

Before concluding, it is instructive to discuss Obstfeld's (1986) results in light of our findings. Using the full-sample of 1954:1-1984:2 Obstfeld finds that the estimated correlations between saving and investment rates for the U.S. and Canada are 0.908 and 0.550 respectively. Due to the regime switch over the full sample Obstfeld splits the sample at 1972:4 and finds that the correlations drop dramatically during the post Bretton Woods era. More specifically the correlations between saving and investment rates for the U.S. are estimated at 0.902 and 0.870 over the two sub-samples which correspond to the fixed and flexible exchange rate regimes. For Canada, the

[5] For the purpose of this study the earlier period of a floating exchange rate for Canada, from October 1950 to May 1962 is ignored.

estimated correlations for the two sub-samples are 0.716 and 0.399 respectively. Despite these similarities the results obtained by Obstfeld are not entirely comparable with the results reported in this paper. In particular it is important to recall that Obsfeld scales the variables by the GNP and then takes the first-difference of the resulting ratios. Thus the procedure used by Obstfeld deals with the short run behaviour of national savings and domestic investment. In contrast our results pertain to the long run behaviour of the two series in levels. As argued earlier we believe this to be more consistent with the hypothesis of capital immobility as formulated by Feldstein and Horioka (1980).

4. Conclusion

In this paper we presented evidence from time series data for the United States and Canada that suggests there is no long-run relationship between national savings and gross domestic investment, during the period of post-Bretton Woods floating exchange rates. In addition for the United States we cannot reject the hypothesis of no cointegration for the earlier period of fixed exchange rates. For Canada the data suggest evidence of cointegration between national savings and domestic investment during the Bretton Woods period.

While the above results are preliminary, we believe they are of interest since this are the first empirical study to find considerable evidence against what Dooley, Frankel and Matheison (1987) call a " robust empirical regularity ". In contrast to previous empirical findings, most of our empirical results are not inconsistent with the widely accepted theoretical assumption in the international finance literature, that there is a highly integrated international capital market. Although our methodology is different from that used in earlier studies it is in some sense more robust to many of the econometric problems (eg. measurement error, endogeneity, omitted stationary variables) that can arise in testing the capital mobility hypothesis, within the Feldstein-Horioka framework.

There are a number of ways of extending the above results. Most obvious is to apply the methodology to other countries. Secondly we have made no attempt to adjust the measures of national saving for the effects of inflation. While this is not a trivial exercise with quarterly data, it seems to be important given our focus on long-run relationships. Finally it is entirely possible that our finding of no cointegration between national savings and domestic investment may be due to the omission of relevant I(1) variables from the cointegrating regressions. For example Cardia (1988) has suggested a role for productivity shocks in explaining the correlation between national savings and investment. However since our objective has been to focus on the "puzzle" in its original form, we have not pursued this issue in the paper.

References

Banerjee, A., J. J. Dolado, D.F. Hendry and G. W. Smith, 1986, Exploring Equilibrium Relationships in Econometrics Through Static Models: Some Monte Carlo Evidence, Oxford Bulletin of Economics and Statistics 48, 253-277.

Caprio G., and D.H. Howard, 1984, Domestic Saving, Current Accounts and International Capital Mobility, International Finance Discussion Paper 244.

Cardia E., 1988, The Dynamics of Savings and Investment in Response to Monetary, Fiscal and Productivity Shocks, University of Montreal, CRDE Working Paper, 8821.

Dooley M., J. Frankel, and D.J. Mathieson, 1987, International Capital Mobility. What do Saving-Investment Correlations Tell Us?, IMF Staff Papers 34, 503-530.

Durlauf, S. N. and P. C. B. Phillips, 1988, Trends Versus Random Walks in Time Series Analysis, Econometrica 56, 1333-1354.

Engle R.F. and C.W.J. Granger, 1987, Cointegration and Error Correction: Representation, Estimation and Testing, Econometrica 55, 251-176.

Feldstein M. and C. Horioka, 1980, Domestic Saving and International Capital Flows, Economic Journal 90, 314-329.

Feldstein M., 1983, Domestic Saving and International Capital Movements in the Long Run and the Short Run, European Economic Review 21, 129-151.

Fieleke N., 1982, National Saving and International Investment, in Saving and Government Policy, Conference Series No. 25, Federal Bank of Boston.

Frankel J.A., 1985, International Capital Mobility and Crowding Out in the

U.S. Economy: Imperfect Integration of Financial Markets or of Goods Markets? NBER Working Paper 1773.

Frankel J.A., 1989, Quantifying International Capital Mobility in the 1980s, NBER Working Paper 2856.

Fuller, W.A.,1976, Introduction to Statistical Time Series (Wiley and Sons: New York).

Murphy R., 1984, Capital Mobility and the Relationship between Savings and Investment in the OECD Countries, Journal of International Money and Finance 3, 327-342.

Nelson C.R. and C.I. Plosser, 1982, Trends and Random Walks in Macro-Economic Time Series: Some Evidence and Implications, Journal of Monetary Economics 10, 139-162.

Obstfeld M., 1986, Capital Mobility in the World Economy: Theory and Measurement, Carnegie-Rochester Conference Series on Public Policy 24, 105-14.

Penati A. and M. Dooley, 1984, Current Account Imbalances and Capital Formation in Industrial Countries, IMF Staff Papers 31, 1-24.

Phillips P.C.B., 1986, Understanding Spurious Regressions in Economics, Journal of Econometrics 33, 311-340.

Phillips P.C.B. and S.N. Durlauf, 1986, Multiple Time Series Regression with Integrated Processes, Review of Economic Studies 53, 473-495.

Sachs J.D., 1981, The Current Account and Macroeconomic Adjustment in the 1970's, Brookings Papers on Economic Activity 1, 201-68.

Stock J.H., 1987, Asymptotic Properties of Least Squares Estimators of Cointegrating Vectors, Econometrica 55, 1035-1056.

Summers L., 1985, Tax Policy and International Competitiveness, in:

Table 1

Unit Root Tests for United States and Canadian Series

Series	United States ADF(1)	United States ADF(4)	Canada ADF(1)	Canada ADF(4)
	(1) Test with Drift and Trend			
\$ S_n	-2.58 (14.00)	-1.83 (8) (1.91)	-1.34 (0.62)	-1.24 (6.94)
\$ I_d	-1.58 (16.66)	-1.15 (6.88)	-0.95 (20.75)	-0.48 (7) (1.81)
RS_n	-3.45* (4.53)	-3.83* (0.88)	-3.64* (4.07)	-3.31 (8.40)
RI_d	-3.73* (5.60)	-3.98* (2.71)	-3.25 (10.88)	-2.88 (7.52)
	(2) Test with Drift			
RS_n	-1.24 (4.37)	-1.19 (4.50)	0.41* (3.78)	0.67** (8.02)
RI_d	-0.69 (5.78)	-0.52 (5.83)	0.38* (13.93)	0.77** (8.89)

Critical values for the ADF test statistics are taken from Fuller (1976). A * and ** imply significance at the 5 % and 1 % levels respectively. The numbers in brackets are statistics from a Lagrange Multiplier test for serial correlation (up to lag four) in the residuals of the ADF test regression. They should be compared to a Chi-Squared four, (eg. 9.488 is the 5% critical value). A \$ indicates a nominal value. Note that for \$ S_n and \$ I_d eight and seven lags in the ADF test were required to obtain uncorrelated residuals.

J. Frankel, ed., Aspects of Fiscal Policies (University of Chicago Press: Chicago).

White H., 1980, A Heteroscedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroscedasticity, Econometrica 48, 817-838.

Table 2

Tests for Cointegration Between National Savings and Investment

- United States

(1a) Nominal ($) — Sample 1956:1 to 1987:4

$$I_d = -28.005 + 1.1625\, S_n$$
$$\quad (4.98) \qquad (64.19)$$

$\bar{R}^2 = 0.970$ DW = 0.056 T = 128

ADF(4) = - 1.00

LM	1	2	3	4
	1.38	1.87	1.10	1.48

(1b) Real — Sample 1956:1 to 1987:4

$$RI_d = -18.670 + 1.1424\, RS_n$$
$$\quad (1.23) \qquad (30.77)$$

$\bar{R}^2 = 0.882$ DW = 0.064 T = 128

ADF(4) = - 2.55

LM	1	2	3	4
	1.40	1.40	0.97	1.05

- Canada

(1a) Nominal ($) — Sample 1956:1 to 1987:4

$$I_d = 94.939 + 1.0669\, S_n$$
$$\quad (0.21) \qquad (100.8)$$

$\bar{R}^2 = 0.988$ DW = 0.345 T = 128

ADF(4) = - 1.40

LM	1	2	3	4
	2.09	0.59	2.51	0.95

(1b) Real — Sample 1956:1 to 1987:4

$$RI_d = -1921.0 + 0.9954\, RS_n$$
$$\quad (2.76) \qquad (62.87)$$

$\bar{R}^2 = 0.952$ DW = 0.266 T = 128

ADF(4) = - 2.31

LM	1	2	3	4
	0.89	0.23	0.83	0.22

Critical values for the cointegration tests are from Engle and Granger (1987), (eg. the ADF(4) statistic should be compared to (-3.17) at the 5% level). LM refers to the results of a Lagrange Multiplier test for serial correlation (in lags one to four) of the ADF test regression.

Table 3

Split-Sample Tests for Unit Roots and Cointegration

	United States		Canada	
	56:1 - 71:2	71:3 - 87:4	56:1 - 70:2	70:3 - 87:4
(1) Unit Root Tests				
	ADF(q)	ADF(q)	ADF(q)	ADF(q)
$ S_n	-3.34 (1)	-3.46 (3)	-2.50 (1)	-3.19 (1)
S_n	-2.54 (1)	-3.25 (3)	-3.10 (1)	-2.61 (1)
$ I_d	-2.88 (1)	-3.09 (4)	-2.48 (1)	-3.47 (2)
I_d	-3.54* (2)	-3.11 (3)	-2.85 (1)	-3.60* (2)
(2) Cointegration Tests				
Nominal ($)				
Coeff	0.9952	1.2879	0.9380	1.0797
$\bar{R}^2$	0.995	0.930	0.982	0.971
D-W	0.314	0.086	0.650	0.550
ADF(q)	-2.35 (3)	-1.25 (4)	-4.00* (2)	-1.02 (4)
Real				
Coeff	1.1449	0.9993	0.6930	1.1069
$\bar{R}^2$	0.997	0.494	0.931	0.869
D-W	0.236	0.052	0.448	0.339
ADF(q)	-2.24 (3)	-1.52 (4)	-3.24* (2)	-2.79 (2)

Critical values for the unit root tests are from Fuller (1976) and from Engle and Granger (1987) for the cointegration tests. The unit root tests are based upon a regression which includes a constant and a linear time trend.

Data Appendix

Series for the United States are taken from the national accounts data available in *Citibase* and the *Survey of Current Business.* The series for I_d is gross private domestic investment, which is available in both current and constant prices (1982=100). The series for national savings S_n is obtained by adding gross private domestic investment and net foreign investment. There is no published constant price data for S_n. To construct a constant price series for net foreign investment: the series for net transfer payments to foreigners and interest paid by the government to foreigners are deflated by the GNP price deflator, and then deducted from net exports of goods and services , in constant prices. Constant price national savings is obtained by adding gross private domestic investment and net foreign investment, both at constant prices.

For Canada series are obtained from the national accounts data in the *Cansim 87* database. The series I_d is gross business investment plus inventories. The Canadian national accounts distinguish between current government expenditure and government investment expenditure. To be consistent with the United States data we treat all government expenditure as being for current consumption purposes. Including the government investment figures with those for gross business investment does not affect the above results. To obtain series for S_n in current and constant prices we follow the same approach as for the United States, except that there is no published deflator for GNP, so where necessary we use the GDP deflator.