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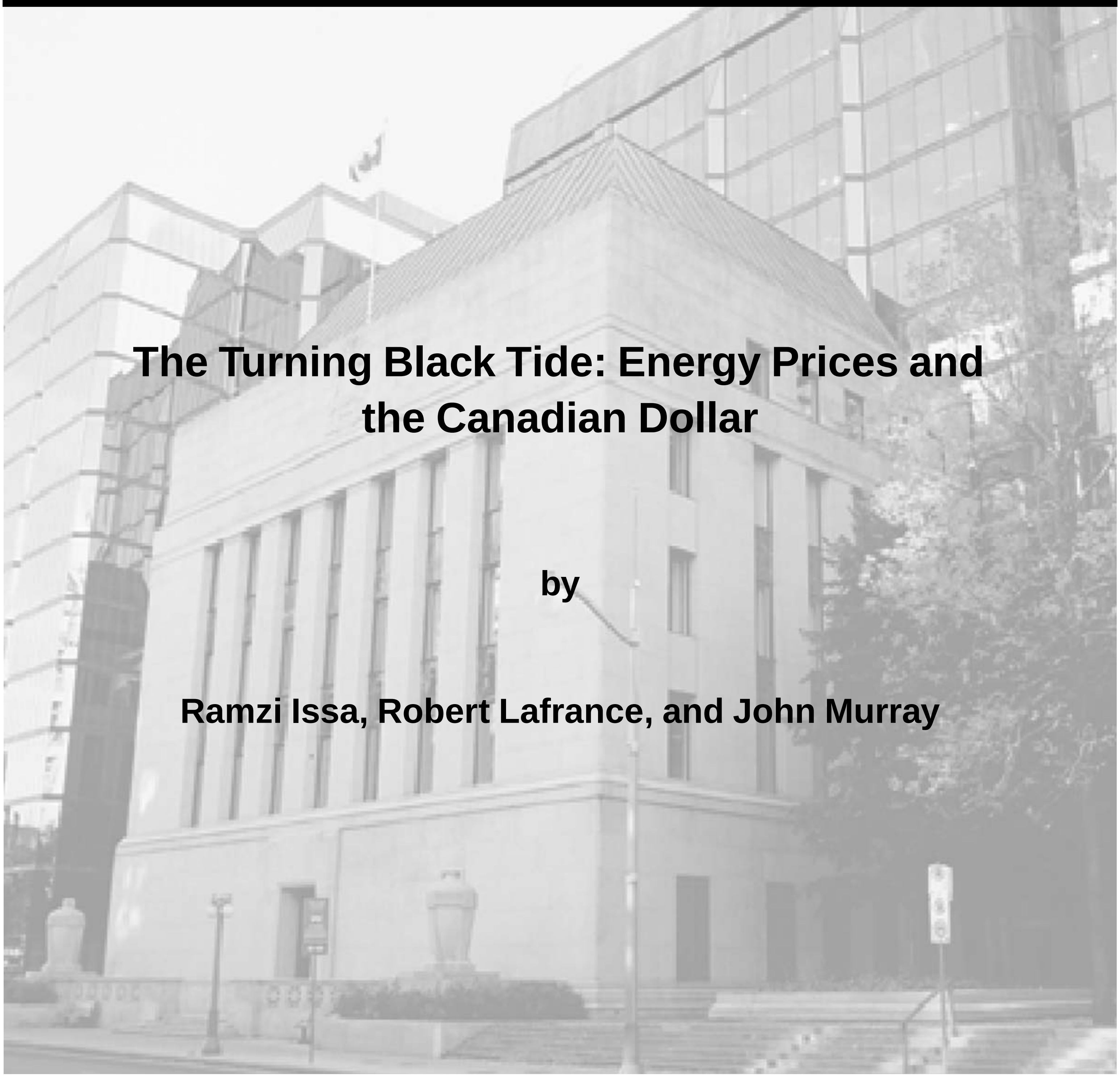

The Turning Black Tide: Energy Prices and the Canadian Dollar

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Ramzi Issa, Robert Lafrance, and John Murray

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Ramzi Issa, Robert Lafrance, and John Murray

International Department
Bank of Canada
Ottawa, Ontario, Canada K1A 0G9
rlafrance@bankofcanada.ca

Contents

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Abstract

The authors revisit the relationship between energy prices and the Canadian dollar in the Amano and van Norden (1995) equation, which shows a negative relationship such that higher real energy prices lead to a depreciation of the Canadian dollar. Based on structural break tests, the authors find a break point in the sign of this relationship, which changes from negative to positive in the early 1990s. The break in the effect between energy prices and the Canadian dollar is consistent with major changes in energy-related cross-border trade and in Canada's energy policies.

JEL classification: F31
Bank classification: Exchange rates; Econometric and statistical methods

Résumé

Les auteurs réexaminent la relation entre les prix réels de l'énergie et le dollar canadien établie par Amano et van Norden (1995), à savoir qu'une hausse de ces prix entraîne une dépréciation du dollar. À l'issue de tests de rupture structurelle, ils constatent que la relation entre les deux variables a changé de signe : initialement négative, elle est devenue positive au début des années 1990. Ce changement cadre avec l'évolution du commerce frontalier des produits énergétiques et l'assouplissement de la réglementation dans le secteur au fil du temps.

Classification JEL : F31
Classification de la Banque : Taux de change; Méthodes économétriques et statistiques

1. Introduction

This paper examines the relationship between the Canadian-U.S. dollar real exchange rate and real energy prices. Our inquiry is motivated by the surge in energy prices, mainly for natural gas and oil, and the recent growing importance of Canadian energy exports. Many analysts see a close link between the appreciation of the Canadian dollar since 2003 and the rise in energy prices. This is not surprising, since the Canadian dollar, like those of Australia and New Zealand, is viewed by markets as a commodity currency; that is, one whose external value mainly reflects the global movements of commodity prices.

The importance of commodity prices in determining of the evolution of the Canadian bilateral exchange has long been recognized at the Bank of Canada. The linkages have been examined in structural models (Macklem 1993; Macklem et al. 2001), reduced-form structural vector autoregression (VAR) models (Gauthier and Tessier 2002), and single-equation models (Amano and van Norden 1995; Helliwell et al. 2005).

Doubts have been raised, however, about Canada's membership in the commodity currency club. Laidler and Aba (2001) find that the relevance of commodity prices for exchange rate movements seems to have declined substantially between the 1970s and the 1990s. Reinhart and Rogoff (2002) find limited correspondence between commodity price and exchange rate movements for Canada. Chen and Rogoff (2003) conclude from their results that, while there may be a long-run relationship between commodity prices and the real exchange rate in Canada's case, this may be due to deterministic rather than stochastic trends, and that there is relatively weak co-movement in the short run. Many have also found it counterintuitive that in the seminal Bank of Canada exchange rate equation developed by Amano and van Norden (1995), henceforth AvN, the sign of energy prices is such that higher energy prices lead to a depreciation of the Canadian dollar.

This paper proposes to address these doubts and solve the mystery of the counter-intuitive results in AvN by showing that the effect of energy prices on the Canadian dollar has shifted over time, from negative to positive, as the importance of Canada's energy exports has grown. A number of tests diagnose this empirical fact and suggest that the shift occurred in the early 1990s. This corresponds to a period when Canada began to become an important net energy exporter; most of the deregulation of the Canadian energy sector had taken place and trade agreements had encouraged a North American view on energy production and consumption. These new results can help to decipher whether movements of the Canadian dollar respond to underlying demand and supply fundamentals, rather than temporary market aberrations, and thus help to inform the monetary authorities on whether special consideration should be given to exchange rate movements in setting the bank rate (Dodge 2005).

This paper is organized as follows. Section 2 reviews the major developments of the Canadian energy market, focusing on policy initiatives and the growing importance of energy exports for Canada's trade balance. Motivated by these ongoing changes, the possibility of a structural break in the AvN equation is explored in section 3. A variety of tests are used to determine the most probable break point. The results suggest a significant break date that reflects ongoing changes in Canada's energy policies and trade, such that the earlier negative cointegration relationship between energy prices and the real exchange rate is reversed after the break point. Section 4 concludes.

2. A Look at the Data: Canadian Energy Trade Trends and Policies

Although the implementation of the Canada-U.S. Free Trade Agreement in 1989 and of the North American Free Trade Agreement (NAFTA) in 1994 led to a growing share of manufactured goods exports, Canada's international trade in *net* terms still largely comprises commodities (Chart 1). Moreover, net exports of energy-related commodities (mainly oil, natural gas, and electricity) have grown in importance since the late 1990s, much more so in value than in volume terms (Charts 1 and 2). Canada was a net importer until the early 1980s, but in recent years, export sales have increased sharply with higher oil prices (Chart 3). Crude petroleum and natural gas currently represent about 80 per cent of Canada's energy exports, with natural gas becoming relatively more important since the mid-1990s (Chart 4).[1]

These trends reflect changes in world energy prices and the growth of U.S. net external demand for oil and natural gas. They also reflect a change in Canada's energy policies, initially shaped in response to the Organization of Petroleum Exporting Countries (OPEC) oil-price shocks of the 1970s, towards a freer market-based system that promoted exploration and investment.

In response to the OPEC oil-price shocks, Canada tried to achieve energy security through self-sufficiency. Canada's response to the tripling of oil prices in the fall of 1973 was a policy of supply management that lasted in various forms until March 1985.[2] In December 1973, the Government of Canada announced an energy policy to protect Canadians from the volatility of the world oil market and provide producers with sufficient incentives to develop new energy sources. January 1974 saw the introduction of the Oil Import Compensation Program, which was

1. Most of Canada's energy-related exports go to the United States, accounting for about 25 per cent of U.S. energy imports in recent years.
2. This section is based on Helliwell et al. (1989) and sources from the Natural Resources Canada website at <http://www2.nrcan.gc.ca/es/es/EnergyChronology/index_e.cfm>.

subsequently amended in July 1975 and again in April 1982.[3] Security of supply was a key issue and the main reason why the National Energy Board assumed an active role in allocating Canadian production among domestic and export markets. In 1979–80, the second OPEC shock resulted in a doubling of world oil prices. In response, in October 1980 the Canadian government adopted the National Energy Program (NEP), which introduced several new energy taxes and a broad range of policy initiatives.

The NEP was not well received, however, by the industry and the governments of the producing provinces, and was subsequently amended a number of times. The Western Accord, which was signed in March 1985 between a new federal government in Ottawa and the governments of the producing provinces, marked the effective end of these policies. As far as Canadian oil markets were concerned, the Accord accomplished two main objectives: it deregulated domestic oil prices and it lifted controls on short-term oil exports.[4] The 1985 Accord also led to a progressive deregulation of the natural gas industry that continued into the early 1990s.[5] The Canada-U.S. Free Trade Agreement (in 1989), and the broader NAFTA agreement (in 1994) that included Mexico, gave additional momentum to the creation of an integrated North American energy market. Provisions in the NAFTA were conditioned, however, by national differences in public versus private ownership of the energy sector among the signatory countries. The treaty required non-discriminatory treatment among the signatories with respect to the imposition of taxes or duties on energy or basic petrochemical goods. It also guaranteed access, under certain provisions, for all the countries in times of supply disruption (Mexico requested that it be excluded from this provision).[6]

Since the early 1990s, the Canadian energy sector has experienced solid growth. Supported by investment, which has nearly doubled over this period, production output has significantly

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3. The program had three main components; (1) direct regulation of domestic crude oil prices through federal-provincial agreements, (2) subsidies for imported oil so that consumers in Eastern Canada would also benefit from lower oil prices, and (3) controls over the prices and quantities of crude oil and products in the export market.
4. In December 1985, world prices collapsed following Saudi Arabia's attempts to recapture market share. The price collapse (a decline of over 50 per cent) spawned a number of proposals that called for a return to government involvement in the industry: import tariffs, minimum prices, and various income stabilization plans. The federal government resisted these suggestions, and aid by both the federal and provincial governments was confined to various tax relief and incentive programs.
5. Measures included the relaxation and reform of export regulations, the removal of wellhead price controls, the reform of pipeline regulation, and the freeing up of pipeline capacity. On 25 March 1992, the Canadian government announced new rules for foreign investment in the oil and gas industry. In particular, there was no longer a requirement to ensure Canadian ownership of 50 per cent of the upstream oil and gas industry. On 19 March 1993, the governments of Canada and Alberta introduced new initiatives designed to support the development of Alberta's oil sands and the heavy oil resources of Alberta and Saskatchewan through research and development partnerships.
6. For more details, see Government of Canada (1992).

increased. Total oil production has increased by nearly 33 per cent to around 3.1 million barrels per day (bbl/d), while natural gas production has increased by over 20 per cent. As a result, Canada has become an important player in world energy markets and is currently the seventh and third largest producer of oil and natural gas, respectively. Canada's position as a net exporter of energy is expected to remain firm given its plans to continue developing the vast oil sand deposits. The inclusion of these deposits would give Canada the world's second largest proven reserves. While the costs of extracting and processing the oil sands are larger than for conventional oil, technological improvements over the past decade have resulted in a doubling of production to more than 1 million bbl/d.

3. Empirical Results: A Reduced-Form Equation

The oil shocks of the 1970s spurred numerous research papers on how economies might respond to commodity price shocks and how commodity price shocks would affect real exchange rates in small open economies, including the implications of the Dutch disease for commodity producers (e.g., Neary and Purvis 1982; Bruno and Sachs 1982; Corden 1984). This line of enquiry also led a number of researchers to try to use world commodity price shocks to explain the exchange rate movements of commodity producers, such as Australia (Blunndell-Wignall and Gregory 1989; Gruen and Wilkinson 1994), Canada (Amano and van Norden 1995), and Norway (Akram 2004).[7]

In multi-country studies, the evidence of a link between commodity prices and the real exchange rate is mixed. In particular, Cashin, Céspedes, and Sahay (2003) test for cointegration of the real exchange rate and commodity export prices for 58 countries, finding evidence of cointegration for 19 countries—though not for Canada.[8] Of these, ten countries show a shift in the cointegration relationship based on the Gregory-Hansen (1996) test. The authors also show that this relationship can explain why purchasing power parity (PPP) has limited explanatory power for the long-run real exchange rates of commodity exporting countries. The long-run real exchange rate of commodity currencies is not constant (as PPP would imply) but time-varying, being dependent on movements in real commodity prices.

7. In the case of Norway, however, the empirical evidence does not support its membership in the commodity currency club. Akram (2004) fails to find a long-run relationship between oil prices and the Norwegian real exchange rate. His analysis shows that the behaviour of the Norwegian real and nominal exchange rates appears to be consistent with the purchasing power parity theory. He offers three possible explanations. First, Norway is more open than other industrialized countries and thus more subject to PPP arbitrage pressures. Second, this is reinforced by the fact that Norway mainly exports commodities whose prices are determined in world markets. Third, and more plausibly, Norway managed its exchange rate for most of the post–Bretton Woods period, adopting inflation targeting only in 2001.

8. Of the five industrialized countries in their sample, Australia and Iceland are found to be the only commodity currencies. Canada, New Zealand, and Norway are not.

Our approach builds on Amano and van Norden (1995). Over their sample period (1973–93), they find that, in the long run, oil-price movements lead to a real depreciation of the Canadian dollar, even though the United States is more dependent on energy imports than Canada. They conjecture that the negative energy price effect might be reflecting the effects of Canadian domestic energy policies that were put in place. Using a longer and richer data set, it is now possible to review Amano and van Norden's findings to see whether this relationship still holds, given the growth of Canada's energy exports.

Thus, picking up where Amano and van Norden left off, this paper tests whether the effect of energy prices on the Canadian dollar has changed over time. Based on a variety of diagnostic tests and tests for structural breaks with an unknown break date, we find that the effect of energy prices on the Canadian dollar, which was negative as Amano and van Norden had reported in the 1970s and 1980s, turned positive in the early 1990s.

In section 3.1, we perform diagnostic tests to detect the possibility of a structural break. In section 3.2, we test for parameter stability, treating the timing of the break as unknown. Based on the results, we retest the basic statistical assumptions of the model, such as non-stationarity and cointegration, in section 3.3. We conclude with estimation results and sensitivity analysis in section 3.4.

3.1 Diagnostic tests

The standard Amano-van Norden (AvN) equation is specified as an error-correction model (ECM):

$$\Delta rfx_t = \alpha(rfx_{t-1} - \mu - \beta_c com_{t-1} - \beta_e ene_{t-1}) + \phi int_{t-1} + \varepsilon_t, \quad (1)$$

where *rfx* is the real Canada-U.S. exchange rate (Can$/US$), *com* is a real commodity price index excluding energy, *ene* is a real energy price index, and *int* is the Canada-U.S. 90-day commercial paper rate differential. Except for *int*, all variables are expressed in logarithms. Data sources are provided in Appendix A. Estimates of equation (1) for various sample periods suggest that the relationship between real energy prices and the Canadian dollar has broken down over time, as the energy price coefficient is no longer significant and the R^2 of the equation over the 1973–2005 period is only 0.026 (see Table 1).

One way to look at the evolving relationship between real energy prices and the Canadian dollar is through recursive estimates of equation (1). Chart 5 plots the parameter value of real energy prices, β_e, from forward recursive estimates where the solid (dotted) line refers to a significant (insignificant) value at the 10 per cent level using robust Newey-West heteroscedasticity- and autocorelation-consistent (HAC) standard errors.[9] The chart shows that the positive value of β_e has steadily decreased from its peak in the early 1990s, and that β_e became statistically insignificant in 2004. In contrast, backward recursive estimates of equation (1) reveal that β_e is significantly negative for the most recent sample period (Chart 6).[10]

The same exercise is repeated in Chart 7, except that we use a fixed window of 60 observations for each estimate. Here, the transition of β_e from a positive to negative value is dramatic. The insignificant values in the mid- to late 1980s is consistent with the estimation bias due to mixing the post- and pre-break data samples. Overall, these initial diagnostic tests suggest that a break in the AvN equation is highly probable. The next step is to do formal tests of parameter stability.

3.2 Parameter stability tests and estimates of the break date

Tests for parameter stability in the case of non-stationary variables have been developed by Hansen (1992). Effectively, these are direct tests of the cointegrating vector and hence, for clarity, we rewrite the cointegrating relationship of the AvN equation as:

$$rfx_t = \beta_t x_t + \varepsilon_t, \tag{2}$$

where $\beta_t = (\mu_t, \beta_{ct}, \beta_{et})$ and $x_t = (1, com_t, ene_t)'$.

Hansen proposes three test statistics, the *SupF*, *MeanF*, and L_c, derived as Lagrange multiplier tests based on the fully modified ordinary least squares (FM-OLS) estimator of Phillips and Hansen (1990). All the tests have the same null of parameter constancy in the cointegrating vector; that is, $\beta_t = \beta$ for all *t*. However, each test addresses a different type of structural change with unknown timing.[11]

9. Regressions start with a given sample, which is increased by one quarter going forward in time. Hence, the first observation in Chart 5 corresponds to a regression over the 1973Q1–1986Q1 period, the second to the 1973Q1–1987Q2 period, and so forth.
10. In this case, the regressions start with the most recent sample period and are extended backwards in time.
11. The treatment of the break date as known or unknown is important. Hansen (2001) provides an elegant discussion illustrating the difference. The classical version of the Chow test requires that the break date be known a priori, which could be selected arbitrarily or based on some information on the data. In the first case, the Chow test would be uninformative, since the true break date could be missed. In the second case, the selection of the break date would be correlated with the data and, hence, potentially biased in finding a break when in fact none occurred. Moreover, selecting the ‘true’ break date can be a complicated or arbitrary decision when the null of stability is rejected for several dates. One solution would be to apply the Chow test systematically, which would require using appropriate critical values. These have been calculated by Andrews (1993) and Andrews and Ploberger (1994); Hansen (1997) provides a method to calculate *p*-values. The date that provides the most evidence against the null is then taken as the break date. This approach is used by many stability tests, including those in this paper.

The *SupF* corresponds to the largest value of a series of Chow-type tests based on recursive estimates of equation (2). It is directed against the alternative that the cointegrating relationship undergoes an abrupt one-time structural change at τ, or more formally that $\beta_{t_1} \neq \beta_{t_2}$ where $t_1 \leq \tau$ and $t_2 > \tau$ for $\tau \subset T$.[12] The *MeanF* is calculated as the average of the Chow-type tests used to compute the *SupF*, and thus uses the same trimmed sample period. Its alternative is similar to that of the *SupF*, although it captures the notion that structural change occurs more gradually. The L_c test is calculated using the full sample and is directed against the alternative of random-walk coefficients, which is equivalent to a test of cointegration.[13]

The finite sample performance of the three statistics is evaluated by Kuo (1998), who also extends Hansen's tests to the case of stability in a subset of the cointegrating parameters. In his Monte Carlo experiments, Kuo finds that both the full and partial stability tests perform relatively well. He also confirms that his tests have improved power in detecting instability when the structural change occurs only in a subset of the cointegrating vector.

Results of the parameter stability tests are reported in Table 2, with values in bold denoting rejection of the null hypothesis of parameter stability at the 5 per cent level.[14,15] Five cases of the null hypothesis are considered: (1) $\mu_t = \mu$, (2) $\beta_{e,t}$=β_e, (3) $\beta_{c,t}$=β_c, (4) $\beta_{c,t}$=β_c and $\beta_{e,t}$= β_e, and (5) $\mu_t = \mu$, $\beta_{c,t}$=β_c, and $\beta_{e,t}$= β_e. The results clearly suggest the presence of parameter instability. The *MeanF* and L_c statistics are significant at the 5 per cent level in all cases. In comparison, the *SupF* statistic is significant only in the cases where β_e is involved. This suggests that an abrupt break, as captured by the *SupF*, likely occurred between the real exchange rate and real energy prices.

12. Andrews (1993) suggests restricting the range of candidate break dates to ensure that the test statistics converge. This trimmed sample drops the first and last 15 per cent of observations. The retained sample is 1977Q4–200Q3.
13. The error term from a system of variables that are not cointegrated can be decomposed into a random-walk and stationary component. But this is equivalent to the intercept following a random walk.
14. We thank Biing-Shen Kuo for making his GAUSS program available to us, which was used to generate the test statistics in Table 2.
15. Since both forms of the tests use the FM-OLS estimator of Phillips and Hansen (1990), we need to select a bandwidth parameter to estimate the covariance matrices. Kuo finds in his Monte Carlo experiments that the results of the stability tests may be sensitive to the value of the bandwith parameter. In particular, he finds that, while choosing a large value such as 12 (which is considered to be an extreme choice) results in a power loss, it does not affect size. To protect against falsely rejecting the null, we repeated our tests by letting the bandwith parameter take on values from 0 to 9. We also tested with the plug-in bandwith estimator recommended by Andrews (1991). Both procedures generally yielded similar timing for the break date.

The results of the *SupF* tests also provide an estimate of the break date. Chart 8 plots the recursive Chow-type test statistic for cases 2 (dashed line) and 5 (solid line) scaled to their respective 10 per cent critical value such that a value larger than one indicates instability. In both cases, the test statistics follow the same pattern and reach their maximum value (i.e., the *SupF*) for τ=1993Q3.

3.3 Non-stationarity and cointegration

The previous results indicate that the cointegrating relationship in the standard AvN equation is no longer supported. Thus, it is necessary to test two basic assumptions before respecifying the model with the structural break. The first is that the three long-run variables are non-stationary, which we check using unit root tests robust to structural breaks. The second is that these three variables are cointegrated when the break is incorporated.

Perron (1989) first analyzed the impact of structural breaks on the performance of unit root tests. He shows that standard unit root tests, like the augmented Dickey-Fuller (ADF) test, have dramatically reduced power when the underlying process undergoes a structural break.[16] As a solution, Perron develops ADF-type tests that directly incorporate the structural break. One drawback of these tests, however, is that the timing of the structural break is assumed to be known. In Perron (1997), this assumption is relaxed, since the break date is selected using a data-dependent procedure similar to the Hansen (1992) *SupF* test described above. That is, the break date corresponds to the largest test statistic, in absolute terms, in a series of ADF-type tests. While Perron considers the same unit root null of the ADF tests, he considers three different alternative hypotheses. These are that the underlying process is trend stationary with a one-time break in (1) mean, (2) trend, and (3) mean and trend. Perron refers to these as the *crash*, *changing growth*, and *broken trend* models, respectively.

Given the advantage of not having to prespecify the break date, we use the Perron (1997), or PN, tests along with the modified ADF test due to Elliott, Rothenberg, and Stock (1996), henceforth ERS, to determine whether the variables in the model are non-stationary.[17] Estimation details for

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16. Intuitively, a unit root process is defined as having a continuously changing trend function. Hence, a rejection of the unit root null in the standard tests implies that the trend function never changes. Perron argues that these standard tests are biased towards the conclusion that the trend function ‘always’ changes versus the alternative that it ‘never’ changes, when it actually changes ‘sometimes.’
17. Ng and Perron (2001) show that the ERS tests with lag selection based on a modified information criteria, which is the approach we use, have vastly improved finite sample properties. Under the ERS approach, the variable is first detrended and then tested for the presence of a unit root in the usual ADF manner.

the PN models and test results are reported in Table 3, where bold values denote rejection of the unit root null at the 5 per cent level. The ERS tests suggest that the real energy prices (*ene*), real commodity price index excluding energy (*com*), and real exchange rate (*rfx*) are, as expected, non-stationary. This holds when allowing for one-time breaks, since all but one PN test statistic is not rejected. Overall, the unit root test results tend to confirm our priors and hence support our treatment of each variable in the model.[18]

We approach the issue of cointegration in two different ways. First, we follow the approach of Quintos (1995) and apply standard Johansen (1992) tests over the full sample and our two identified subsamples based on the estimated break date of 1993Q3. Results are reported in Table 4, with values in bold denoting rejection of the null of no cointegration at the 10 per cent level. Consistent with results from the structural break tests, we find no evidence of cointegration in the full period. In contrast, the subsample tests provide evidence in favour of two cointegrating regimes prior to, and after, 1993Q3. The evidence, however, is much stronger in the first than in the second subsample, which may be due to a smaller sample size in the second period.[19]

Our second approach is to use the residual-based tests developed by Gregory and Hansen (1996), henceforth GH. These tests are extensions of the Engle and Granger (1987) test where a unit root test is applied to the residual error from an OLS regression of a cointegrating equation that directly incorporates the structural break. Thus, GH consider the usual null of no cointegration. Like the PN tests discussed earlier, the GH tests treat the timing of the break as unknown and estimate it in a similar manner. Three structural break models are considered by GH, which are a shift in (1) constant, (2) constant and trend, and (3) constant and slope in the cointegrating vector.

The results are reported in Table 5. They, unlike the Johansen test results, are not supportive of a cointegrating relationship. This could reflect, however, problems associated with residual-based tests. Specifically, these tests impose a common factor between the regressand and regressors, which other cointegration tests like the Johansen (1992) tests reported above, or the ECM-based

18. For the Canada-U.S. interest rate differential (*int*), the ERS test suggests that it contains a unit root. However, applying the ERS test to the individual interest rate series suggests that they are mean stationary below the 10 per cent level. Based on this result, and the fact that non-stationary implies a counterintuitive conclusion that the interest rate differential is explosive, we assume *int* to be stationary.

19. The conclusion remains the same when we extend the second-period sample to the full period, in which β_e is estimated to have a significant and estimated negative effect on the Canadian dollar.

tests, do not.[20] Kremers, Ericsson, and Dolado (1992) show that if this restriction is rejected, which it typically is, then the residual-based tests will suffer a loss of power. This is also shown by Campos, Ericsson, and Hendry (1996), who study the effect of structural breaks in the marginal process of a variable on cointegration tests. In particular, they show that the residual-based tests have substantially lower power relative to the ECM-based tests in the presence or absence of breaks if the common factor restriction is violated. Based on these considerations and the Johansen results, we interpret the evidence mildly in favour of two cointegrating regimes over the periods 1973Q1–1993Q3 and 1993Q4–2005Q4. We recognize, however, that the evidence is far from conclusive, but leave this issue open to future research.

3.4 Parameter estimates and sensitivity analysis

In Table 6, we report estimation results from various specifications of the AvN equation that incorporates the structural break. We begin with allowing all cointegrating parameters to take on new values after the break (6.1). In 6.3 (6.2), we restrict the change to the real energy (and constant) parameter(s) only. Consistent with the prior *SupF* results, the real commodity price index excluding energy is not significantly different in the post- and pre-break subsamples (6.1). However, we find that both the constant and real energy price parameters are significant in both subsamples. Overall, this general-to-specific approach suggests that specification 6.2 has the best statistical properties. This is our preferred equation, which we designate as ILM (for Issa, Lafrance, and Murray).

ILM has the following specification:

$$\Delta rfx_t = \alpha(rfx_{t-1} - \mu - \beta_c com_{t-1} - \beta_e ene_{t-1} - \delta_e I(t>\tau)ene_{t-1} - \delta_\mu I(t>\tau)) + \phi int_{t-1} + \varepsilon_t ,$$

where $I(\bullet)$ is an indicator function (or a multiplicative dummy variable) that takes the value of unity when $t > \tau$ for τ=1993Q3 and 0 otherwise.

20. To demonstrate the implications of the common factor restriction, we provide a simple example taken from Kremers, Ericsson, and Dolado (1992). Consider a bivariate system (y_t, z_t) with a cointegrating vector of $(1, -1)$. The ECM-based approach tests the null of b=0 in $\Delta y_t = a\Delta z_t + bw_{t-1} + \varepsilon_t$, where $w_t = y_t - z_t$ is the cointegrating error. Subtracting Δz_t from this equation yields the formulation under the residual-based approach, which, after rearranging, is $\Delta(y-z)_t = b(y-z)_{t-1} + [(a-1)\Delta z_t + \varepsilon_t]$, or in more familiar form, $\Delta w_t = bw_{t-1} + e_t$, where $e_t = (a-1)\Delta z_t + \varepsilon_t$. While the null remains the same as before, the errors from these two equations will be the same only if $\Delta z_t = 0$ or $a = 1$. Hence, the residual-based tests impose that the short- and long-run elasticities are equal. If this assumption is violated, then the residual-based tests ignore this information, resulting in some loss of power. This loss of information is potentially greater when the cointegrating relationship is non-unity, or $(1, -\lambda)$. Refer to Campos, Ericsson, and Hendry (1996) for further details.

Chart 9 plots dynamic simulations from 1973 onwards of AvN and ILM and compares them with the actual values of the exchange rate. ILM is clearly an improvement over AvN and does a good job tracking the historical data, notably in the past few years.

Finally, to determine whether the estimation results are sensitive, we estimate ILM over our full sample period of 1973Q1–2005Q4, but vary the break date, τ, between 1983Q1 and 1995Q4. The values of β_c, β_e, and $\beta_e+\delta_e$ are plotted in Chart 10, where the X-axis indicates the value of τ and a solid (dotted) line refers to a significant (insignificant) estimate at the 10 per cent level using robust standard errors.[21] The chart shows that β_c remains significantly negative and stable, and that the real energy price term consistently takes on a positive sign in the first period (β_e) but a negative sign in the second period ($\beta_e+\delta_e$). The results also suggest that the estimates for the real energy price term in both periods do not vary as a function of the break date.

4. Conclusion

Based on a variety of tests, we find that the ability of the Amano and van Norden equation to track the evolution of the Canadian dollar since the most recent floating period can be significantly improved if one allows for a structural break in the effect of real energy prices on the real exchange rate. The parameter stability tests estimate the break date to be in 1993Q3. The parameter sensitivity analysis suggests, however, that the break date could have happened earlier.

The ILM equation can account for most of the Canadian dollar's appreciation since 2003. We recognize, however, that the equation could be improved. The transition between the negative and positive contribution of energy prices could be made smoother. Other factors could be considered to address some of the evident serial correlation in the residuals. Perhaps additional data over time will allow a more conclusive statement on the cointegration relationship between commodity prices and the exchange rate in the context of a structural break. We leave these refinements to future research.

In a broader context, our results are another illustration that parameter instability can be a major problem in economic time-series modelling, as noted by Stock and Watson (1996) and Rossi (2005), who find that the inability of fundamental exchange rate models to beat a random walk in forecasting exercises could be attributed to parameter instability. On a more positive note,

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21. Note that δ_e is an additive variable by definition. Therefore, to obtain the coefficient value of *ene* in the second period, we construct $\beta_e' = \delta_e + \beta_e$ and perform a Wald test on the same relationship to determine whether the value is significantly different than zero.

our findings suggest that traditional relationships, which can lose their value in analyzing economic data, such as the AvN equation, may be redeemed by re-examining the data and allowing for structural breaks that are motivated by institutional changes and evolving trade relationships.

References

Akram, Q.F. 2004. “Oil Prices and Exchange Rates: Norwegian Experience.” *Econometrics Journal* 7(2): 476–504.

Amano, R. and S. van Norden. 1995. “Terms of Trade and Real Exchange Rates: The Canadian Evidence.” *Journal of International Money and Finance* 14(1): 83–104.

Andrews, D.W.K. 1991. “Heteroscedasticity and Autocorrelation Consistent Covariance Matrix Estimations.” *Econometrica* 59(3): 817–58.

———. 1993. “Tests for Parameter Instability and Structural Change with Unknown Change Point.” *Econometrica* 61(4): 821–56.

Andrews, D.W.K. and W. Ploberger. 1994. “Optimal Tests when a Nuisance Parameter is Present Only Under the Altenative.” *Econometrica* 62(6): 1383–1414.

Blunndell-Wignall, A. and R.G. Gregory. 1989. “Exchange Rate Policy in Advanced Commodity-Exporting Countries: The Case of Australia and New Zealand.” Paper presented at the IMF Conference, *Exchange Rate Policy in Selected Industrial Countries.*

Bruno, M. and J. Sachs. 1982. “Energy and Resource Allocation: A Dynamic Model of the Dutch Disease.” *Review of Economic Studies* 49(5): 845–59.

Campos J., N.R. Ericsson, and D.F. Hendry. 1996. “Cointegration Tests in the Presence of Structural Breaks.” *Journal of Econometrics* 70(1): 187–220.

Cashin, P., L.F. Céspedes, and R. Sahay. 2003. “Commodity Currencies and the Real Exchange Rate.” Central Bank of Chile Working Paper No. 236.

Chen, Y. and K. Rogoff. 2003. “Commodity Currencies.” *Journal of International Economics* 60 (1): 133–60.

Corden, W.M. 1984. “Booming Sector and Dutch Disease Economics: Survey and Consolidation.” *Oxford Economic Papers* 36(3): 359–80.

Dodge, D. 2005. “Monetary Policy and Exchange Rate Movements.” Remarks to the Vancouver Board of Trade, Vancouver, British Columbia. Ottawa: Bank of Canada.

Elliott, G., T.J. Rothenberg, and J.H. Stock. 1996. “Efficient Tests for an Autogressive Unit Root.” *Econometrica* 64(4): 813–36.

Engle, R.F. and C.W.J. Granger. 1987. “Cointegration and Error Correction: Representation, Estimation, and Testing.” *Econometrica* 55(2): 251–76.

Gauthier, C. and D. Tessier. 2002. “Supply Shocks and Real Exchange Rate Dynamics: Canadian Evidence.” Bank of Canada Working Paper No. 2002–31.

Government of Canada. 1992. *North American Free Trade Agreement Overview and Description*. Ottawa.

Gregory, A.W. and B.E. Hansen. 1996. “Residual Based Tests for Cointegration in Models with Regime Shifts.” *Journal of Econometrics* 70(1): 99–126.

Gruen, D. and J. Wilkinson. 1994. "Australia's Real Exchange Rate - Is it Explained by the Terms of Trade or by Real Interest Rate Differentials?" *Economic Record* 70: 204–19.

Hansen, B.E. 1992. "Tests for Parameter Instability in Regressions with I(1) Processes." *Journal of Business and Economic Statistics* 10(3): 321–35.

———. 1997. "Approximate Asymptotic P Values for Structural-Change Tests." *Journal of Business and Economic Statistics* 15(1): 60–67.

———. 2001. "The New Econometrics of Structural Change: Dating Breaks in U.S. Labour Productivity." *The Journal of Economic Perspectives* 15(4): 117–28.

Helliwell, J., R. Issa, R. Lafrance, and Q. Zhang. 2005. "NEMO: A Canadian-U.S. Dollar Exchange Rate Equation." In *Canada and the Global Economy.* Proceedings of a conference held at the Bank of Canada, November 2004.

Helliwell, J.F., M.E. MacGregor, R.N. McRae, and A. Plourde. 1989. "Oil and Gas in Canada: The Effects of Domestic Policies and World Events." Canadian Tax Paper No. 83, Canadian Tax Foundation, Toronto.

Johansen, S. 1992. "Cointegration in Partial Systems and the Efficiency of Single-Equation Analysis." *Journal of Econometrics* 52(3): 389–402.

Kremers, J.J.M., N.R. Ericsson, and J.J. Dolado. 1992. "The Power of Cointegration Tests." *Oxford Bulletin of Economics and Statistics* 54(3): 325–48.

Kuo, B.S. 1998. "Test for Partial Parameter Instability in Regressions with I(1) Processes." *Journal of Econometrics* 86: 337–68.

Laidler, D. and S. Aba. 2001. "The Canadian Dollar - Still a Commodity Currency." C.D. Howe Institute Backgrounder. No. 47 (January 11).

MacKinnon, J.G., A.A. Haug, and L. Michelis. 1999. "Numerical Distribution Functions of Likelihood Ratio Tests For Cointegration." *Journal of Applied Econometrics* 14(5): 563–77.

Macklem, T. 1993. "Terms-of-Trade Shocks, Real Exchange Rate Adjustment, and Sectoral and Aggregate Dynamics." In *The Exchange Rate and the Economy*, 1–60. Ottawa: Bank of Canada.

Macklem, T., P. Osakwe, H. Pioro, and L. Schembri. 2001. "The Economic Consequences of Alternative Exchange Rate Regimes in Canada." In *Revisiting the Case for Flexible Exchange Rates.* Proceedings of a conference held at the Bank of Canada, November 2000. Ottawa: Bank of Canada.

Neary, J. and D.D. Purvis. 1982. "Sectoral Shocks in a Dependent Economy: Long-Run Adjustment and Short-Run Accommodation." *Scandanavian Journal of Economics* 84(2): 229–53.

Ng, S. and P. Perron. 2001. "Lag Length and the Construction of Unit Root Tests with Good Size and Power." *Econometrica* 69(6): 1519–54.

Perron, P. 1989. “The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis.” *Econometrica* 57(6): 1361–1401.

———. 1997. “Further Evidence on Breaking Trend Functions in Macroeconomic Variables.” *Journal of Econometrics* 80(2): 355–85.

Phillips, P.C.B. and B.E. Hansen. 1990. “Statistical Inference in Instrumental Variables Regression with I(1) Processes.” *Review of Economic Studies* 57: 99–125.

Quintos, C.E. 1995. “Sustainability of the Deficit Process with Structural Shifts.” *Journal of Business and Economic Statistics* 13(4): 409–17.

Reinhart, C. and K. Rogoff. 2002. “The Modern History of Exchange Rate Arrangements: A Reinterpretation.” NBER Working Paper No. 8963.

Rossi, B. 2005. “Are Exchange Rates Really Random Walks? Some Evidence Robust to Parameter Instability.” *Macroeconomic Dynamic*s 10(1): 20–38.

Stock, J. and M. Watson. 1996. “Evidence on Structural Instability in Macroeconomic Time Series Relations.”*Journal of Business and Economic Statistics* 14(1): 11–30.

Table 1: Estimates for the Standard AvN Specification of the Real Exchange Rate [a]

	1973Q1–1992Q1	**1973Q1–1999Q4**	**1973Q1–2005Q4**
α	-0.165 (0.043)	-0.151 (0.035)	-0.086 (0.048)
μ	0.356 (0.046)	0.442 (0.035)	0.363 (0.066)
β_c	-0.298 (0.087)	-0.447 (0.078)	-0.307 (0.154)
β_e	0.141 (0.036)	0.090 (0.037)	**0.001 (0.091)**
ϕ	-0.525 (0.161)	-0.639 (0.119)	-0.340 (0.213)
Statistics			
$\bar{R}^2$	0.268	0.228	0.026
DW	1.136	1.232	1.159

a. Newey-West HAC standard errors in parentheses, with *p*-values larger than 0.10 denoted in bold. The dependent variable is the real exchange rate expressed in Canadian terms, so that a positive coefficient implies a real depreciation.

Table 2: Tests for Parameter Stability[a]

Tests	(1) $\mu_t = \mu$	(2) $\beta_{e,t}=\beta_e$	(3) $\beta_{c,t}=\beta_c$	(4) $\beta_{c,t}=\beta_c$ and $\beta_{e,t}= \beta_e$	(5) $\mu_t = \mu$ and $\beta_{c,t}=\beta_c$ and $\beta_{e,t}= \beta_e$
L_c	**0.21**	**0.64**	**0.20**	**0.90**	**1.09**
MeanF	**2.79**	**4.34**	**3.16**	**8.09**	**9.23**
SupF	6.41	**12.75**	7.18	**14.86**	**18.47**

a. Sample period is 1973Q1–2005Q4, although the effective sample drops the first and last 15 per cent of the sample (see footnote 15 in main text). The test statistics are obtained from FM-OLS estimates using a Bartlett kernel with bandwith set at 4. Values in bold denote rejection of the null hypothesis of parameter stability at the 5 per cent level. Critical values are from Table 1 in Kuo (1998) [indexed by (m, p, m_2, $p(\pi_2)$) = (2,1,0, 0) for case 1, (2,1,1,1) for cases 2 and 3, and (2,1,2,1) for case 4], and Tables 1 to 3 in Hansen (1992) [indexed by (m_2, p) = (2,1) for case 5].

Table 3: Unit Root Tests[a]

	ERS	PN - *C*	τ_C	PN - *G*	τ_G	PN - *B*	τ_B
ene	-1.229 [4]	-3.730 [3]	1985Q3	-3.484 [3]	2000Q2	-3.189 [3]	2001Q1
com	-2.027 [0]	-4.995 [4]	1993Q1	**-5.009** [4]	2000Q3	-5.492 [4]	2001Q1
rfx	-2.532 [3]	-2.743 [3]	1992Q4	-2.907 [3]	2000Q4	-3.300 [3]	2000Q2
int	-1.324 [8]	--	--	--	--	--	--
Net Energy Exports	-0.822 [4]	**-5.262** [1]	1999Q4	**-6.366** [1]	1997Q2	**-6.405** [1]	1999Q3

a. Sample period is 1973Q1–2005Q4 for all series. Bold values for the GLS and PN tests indicate rejection of unit root null at the 5 per cent level, with number of lags in [].
ERS: Elliott, Rothenberg, and Stock (1996) modified augmented Dickey-Fuller test with constant and linear time trend assumed for all series except *int*, which assumes a constant only. The corresponding 5 per cent critical values are -3.00 and -1.94, taken from Table 1 in Elliott, Rothenberg, and Stock (1996). Lag selection based on modified Akaike information criterion.
PN: Perron (1997) test with unknown break date. Lag selection based on general-to-specific *t*-test procedure such that the last included lag has a marginal significance of less than 10 per cent. PN - *i* and the estimated break date, τ_i, for $i = C, G, B$ correspond to a structural change model with a one-time break in mean (crash), trend (changing growth), and mean and trend (broken trend). A trim of 15 per cent is used (see footnote 15 in text). Finite sample based critical values are taken from Table 1 in Perron (1997). These are -5.10, -4.65, and -5.55 for model *C*, *G*, and *B*, respectively. Specifications for the estimated models are as follows:

Crash: $y_t = \mu^C + \delta_\mu^C I(t > \tau_C) + \lambda I(t = \tau_C + 1) + \theta^C t + \alpha^C y_{t-1} + \sum_{i=1}^{k} \gamma_i^C \Delta y_{t-i} + \varepsilon_t$.

Growth: $\tilde{y}_t = \alpha^G \tilde{y}_{t-1} + \sum_{i=1}^{k} \gamma_i^G \Delta \tilde{y}_{t-i} + \varepsilon_t$ where $\tilde{y}_t = y_t - \mu^G - \theta^G t - \delta^G (t - \tau_G) I(t > \tau_G)$.

Broken: $y_t = \mu^B + \delta_\mu^B I(t > \tau_B) + \lambda I(t = \tau_B + 1) + \theta^B t + \delta_\theta^B \, t I(t > \tau_B) + \alpha^B y_{t-1} + \sum_{i=1}^{k} \gamma_i^B \Delta y_{t-i} + \varepsilon_t$.

Table 4: Standard Johansen Cointegration Tests[a]

	1973Q1–2005Q4	1973Q1–1993Q3	1993Q4–2005Q4	1985Q1–2005Q4
No. of CV(s)		Trace Statistics		
Less than 1	17.52	**47.87**	22.99	25.68
Less than 2	8.52	9.41	3.73	5.09
Less than 3	2.91	0.81	0.36	1.42
		λ^{max} Statistics		
Less than 1	9.00	**38.45**	**19.27**	**20.59**
Less than 2	5.62	8.60	3.36	3.67
Less than 3	2.91	0.81	0.36	1.42

a. Bold values denote rejection of the null of no cointegration at the 10 per cent significance level based on critical values calculated by MacKinnon, Haug, and Michelis (1999). Reported test statistics assume a constant in cointegrating vector and linear trend in the data. Lag selections based on sequential modified likelihood ratio test statistic.

Table 5: Residual-Based Cointegration Tests with Structural Change[a]

Type	Break Date	ADF Test Statistic	5% Critical Value	Lags
Regime	1995Q4	-4.54	-5.50	8
Trend	1994Q4	-4.34	-5.29	7
Level	1985Q2	-3.73	-4.92	7

a. Sample period is 1973Q1–2005Q4 with trim of 15 per cent (see footnote 15 in text). Lag length based on general-to-specific *t*-test procedure such that the last included lag has a marginal significance of less than 10 per cent. Test type assumes three different types of structural breaks. Regime: shift in intercept and slope of cointegrating vector, Trend: shift in intercept only with time trend, Level: shift in intercept. Critical values are from Table 1 in Gregory and Hansen (1996).

Table 6: Real Exchange Rate Equations: 1973Q1–2005Q4[a]

	6.1	**6.2 [ILM]**	**6.3**
α	-0.150 (0.050)	-0.156 (0.049)	-0.161 (0.050)
μ	0.410 (0.053)	0.414 (0.050)	0.322 (0.038)
β_c	-0.390 (0.098)	-0.400 (0.091)	-0.252 (0.072)
β_e	0.118 (0.040)	0.120 (0.041)	0.113 (0.042)
δ_c	**-0.041 (0.203)**	--	--
δ_e	-0.369 (0.083)	-0.371 (0.088)	-0.303 (0.099)
δ_μ	-0.123 (0.059)	-0.129 (0.052)	--
ϕ	-0.623 (0.140)	-0.614 (0.148)	-0.399 (0.161)
Statistics			
$\bar{R}^2$	0.200	0.204	0.150
DW	1.334	1.337	1.246

a. Newey-West HAC standard errors in parentheses, with *p*-values larger than 0.10 denoted in bold. Equations 6.1 to 6.3 are based on:

$$\Delta rfx_t = \alpha(rfx_{t-1} - \mu - \beta_c com_{t-1} - \beta_e ene_{t-1} - \delta_e I(t > \tau)ene_{t-1} - \delta_c I(t > \tau)com_{t-1} - \delta_\mu I(t > \tau)) + \phi int_{t-1} + \varepsilon_t ,$$

where $I(\bullet)$ is the indicator function that takes the value of unity when $t > \tau$ and 0 otherwise. The lag of the dependent variable is added to this equation for 6.4. The value (standard error) of *ene* in the second period, $\delta_e + \beta_e$, is -0.251 (0.072), -0.252 (0.074), -0.190 (0.083), and -0.197 (0.078) for 6.1, 6.2, 6.3, and 6.4, respectively.

Chart 1
Canada’s Nominal Net Exports
Billions of Current Dollars,
seasonally adjusted annual rate

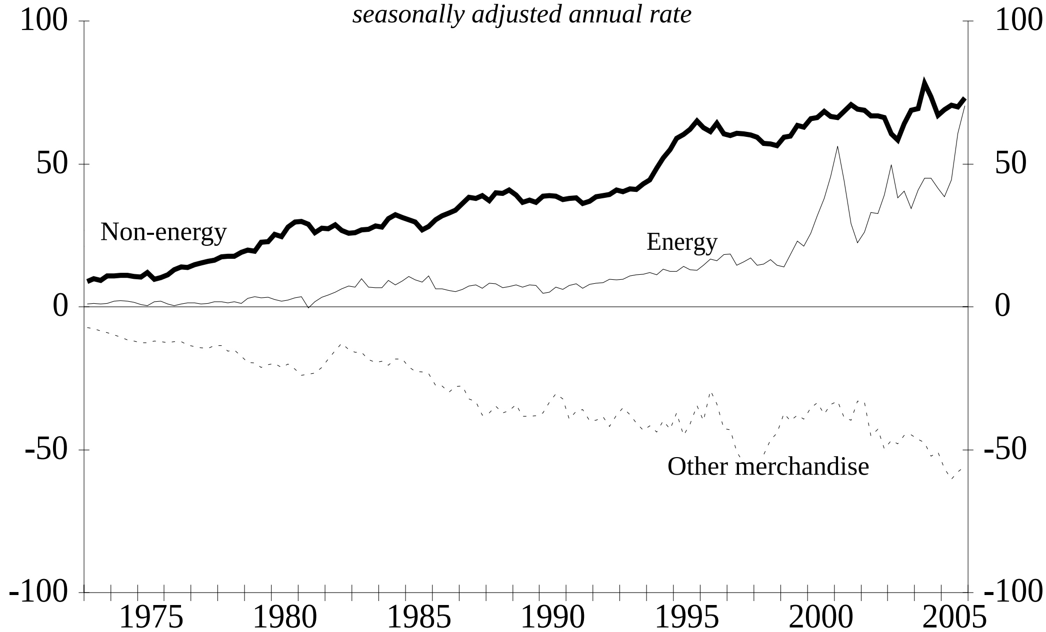


Chart 2
Canada’s Real Net Exports
Billions of Chained $1997,
seasonally adjusted annual rate

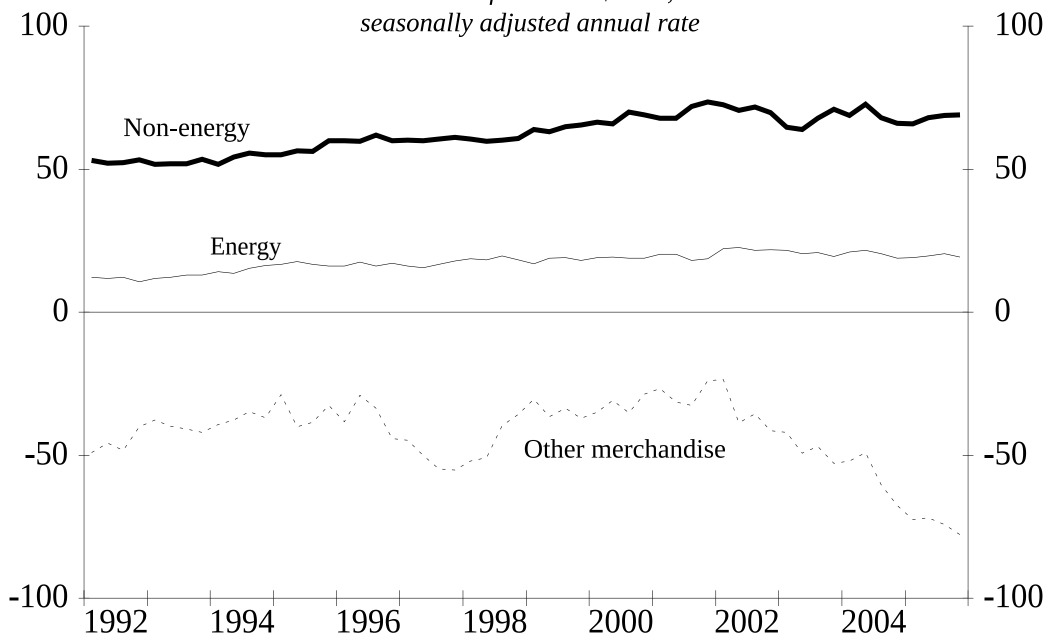


Chart 3
Canada's Crude Oil Net Exports
Billions of Current Dollars, seasonally adjusted annual rate

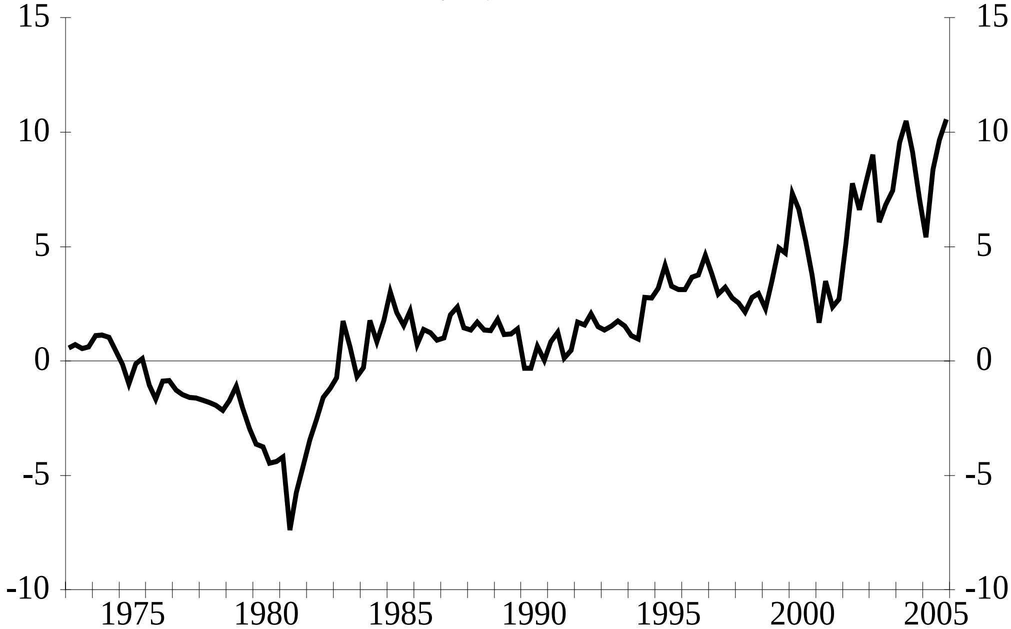


Chart 4
Canada's Decomposition of Energy Exports
% of Total Energy Exports

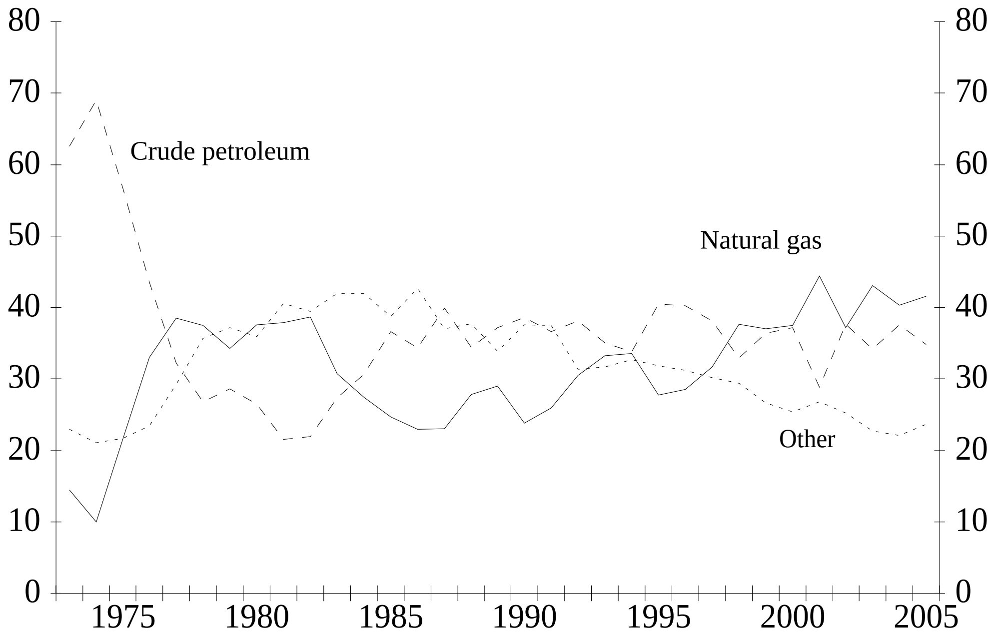


Chart 5
Energy Parameter Value: Forward Recursive Regression
Beginning 1973Q1 with P-Values>0.1 Dotted

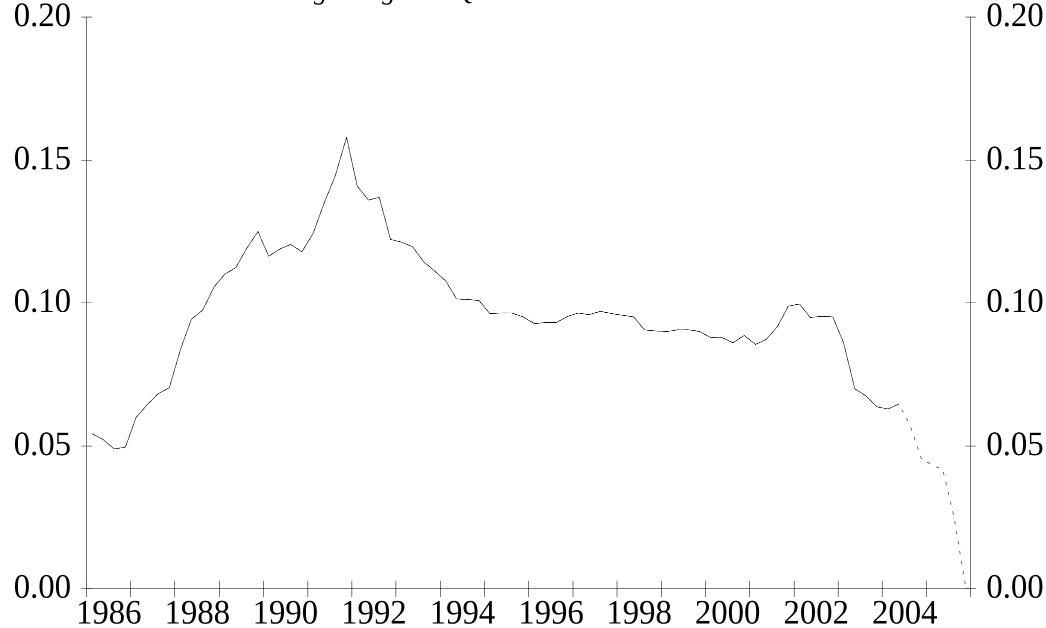


Chart 6
Energy Parameter Value: Backward Recursive Regression
First Sample 1992Q4–2005Q4 with P-Values>0.1 Dotted

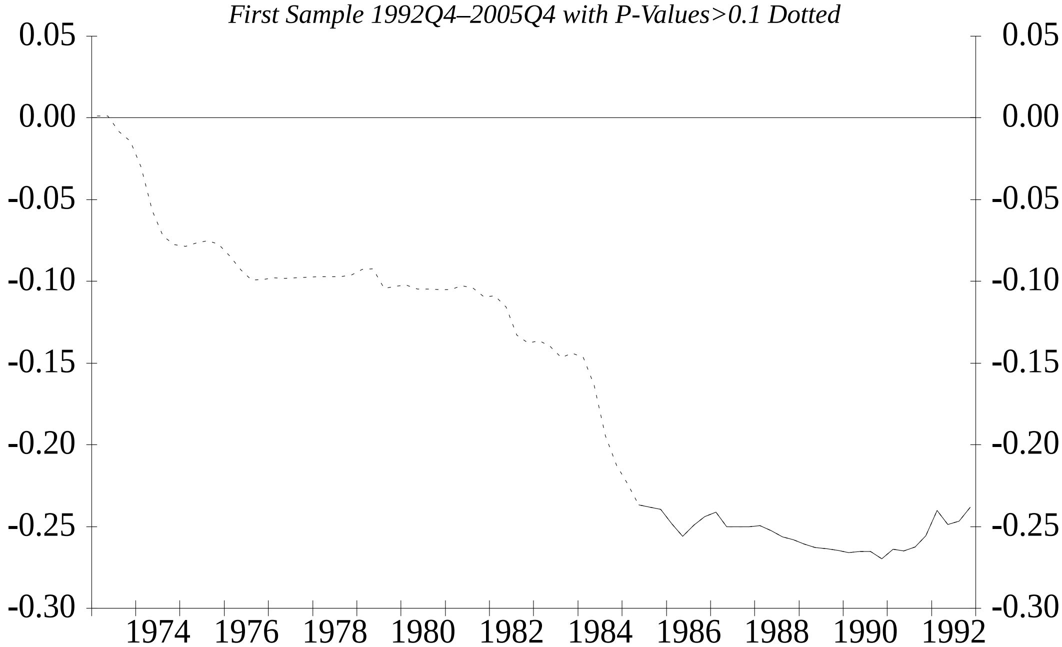


Chart 7

Energy Parameter Value: Backward Regression with Fixed Sample of 60Q's

First Sample 1990Q4–2005Q4 with P-Values>0.1 Dotted

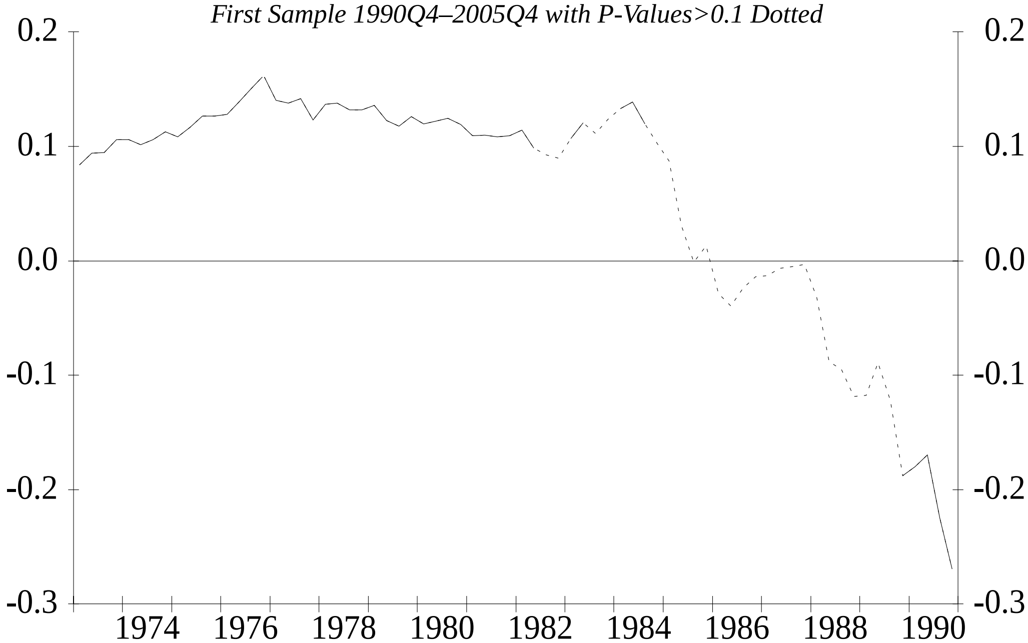


Chart 8

Scaled F Statistic from Parameter Instability Tests

1 is the 10% Significance Level

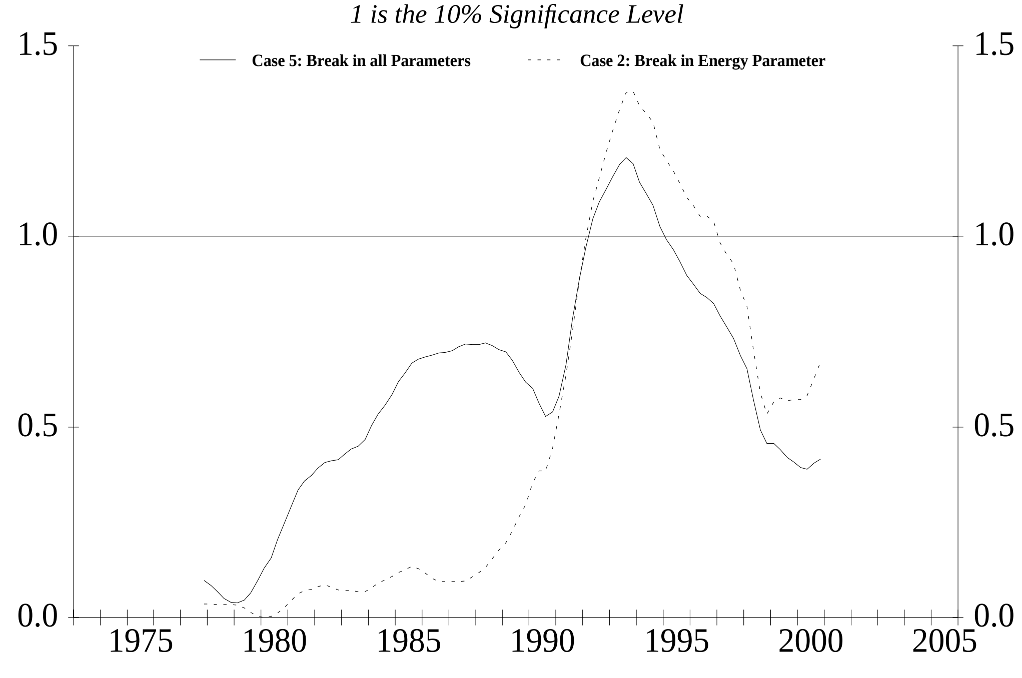


Chart 9
Actual vs. Dynamic Simulation
Estimated 1973Q1–2005Q4, Simulation Begins 1973Q1, U.S. Cents

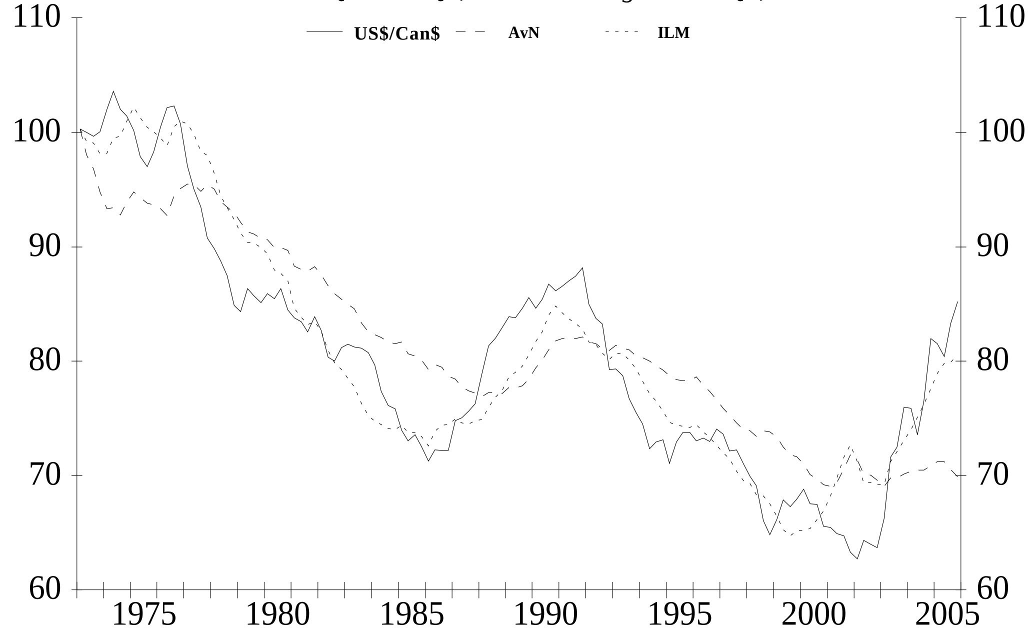
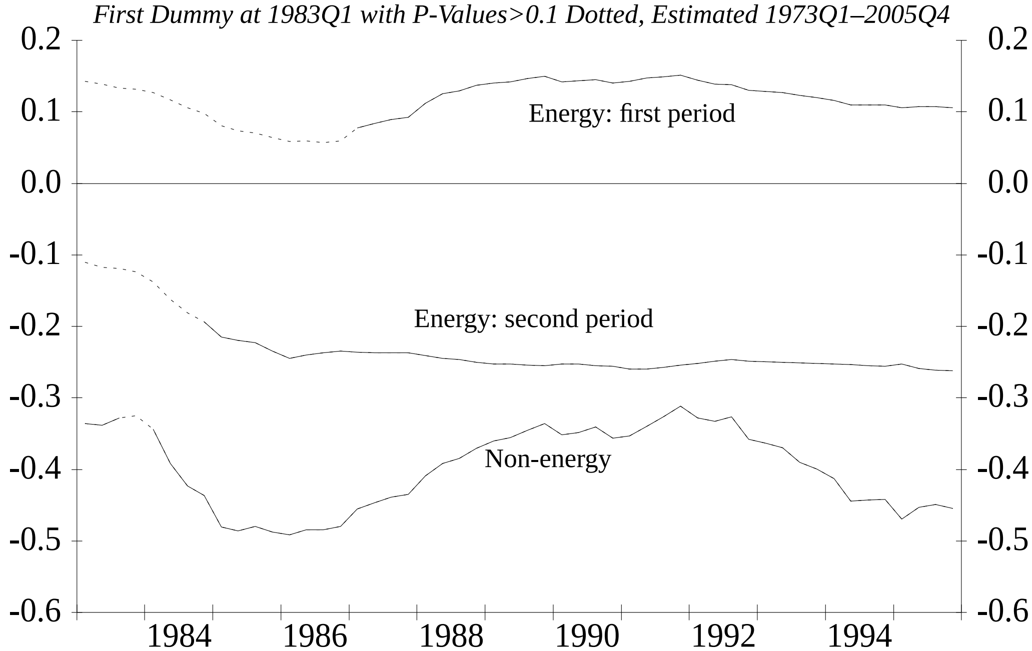


Chart 10
Parameter Values with Regime Change: Forward Moving Dummy
First Dummy at 1983Q1 with P-Values>0.1 Dotted, Estimated 1973Q1–2005Q4



Appendix A

1. *rfx*: Real Canadian per U.S. dollar exchange rate.

Quarterly average of the daily spot rate recorded by the Bank of Canada at 12 p.m. EST, multiplied by the ratio of U.S. to Canada implicit GDP deflator. U.S. and Canadian deflators are indexed to 1997=1.0 and are obtained from the U.S. Bureau of Economic Analysis and the Statistics Canada series v1997756, respectively.

2. *com*: Real commodity price index excluding energy.

Quarterly average of the daily non-energy commodity price index (1982–90=100) calculated by the Bank of Canada. Deflated by the U.S. implicit GDP deflator from 1.

3. *ene*: Real energy price index.

Quarterly average of the daily energy commodity price index (1982–90=100) calculated by the Bank of Canada. Deflated by the U.S. implicit GDP deflator from 1.

4. *int*: Canada–U.S. nominal interest rate differential.

Quarterly average of daily Canadian three-month prime corporate paper rate collected by the Bank of Canada and U.S. 90-day AA non-financial commercial paper closing rate (expressed in equivalent Canadian yield basis) from the Federal Reserve Board.

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