

The turning black tide: Energy prices and the Canadian dollar

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Abstract

This paper revisits the relationship between energy prices and the Canadian dollar in the Amano-van Norden (1995) equation which showed a negative relationship such that higher real energy prices led to a depreciation of the Canadian dollar. Based on structural break tests, we 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.

Résumé

L'effet des prix des produits énergétiques sur le dollar canadien identifié par Aamno et van Norden (1995) est ré-examiné dans ce document de travail. Ces derniers avaient trouvé une relation négative entre les deux variables, telle qu'une hausse des cours des prix énergétiques conduit à une dépréciation du dollar canadien. En testant pour des bris structurels dans l'équation, nous découvrons que la relation a changé de signe dans le temps. Initialement négative, la relation tourne positive au début des années 1990. Le changement de la relation reflète, selon nous, l'évolution du commerce frontalier des produits énergétiques et l'assouplissement de la réglementation dans le secteur au fil du temps.

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1. Introduction

This paper examines the relationship between the Canadian-US 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 as 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 for the determination 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. 2000), reduced-form structural VAR models (Gauthier et 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. Renihart and Rogoff (2003) 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 counter-intuitive 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 if 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).

The outline of this paper is 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 the 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

While the implementation of the Canada-US 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 is still largely comprised of 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% of Canada's energy exports, with natural gas becoming relatively more important since the mid-1990s (Chart 4).²

These trends reflect changes in world energy prices and the growth of US net external demand for oil and natural gas. They also reflect a change in Canada's energy policies, initially shaped in response to the 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 1985.³ 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 subsequently amended in July 1975 and again in April 1982.⁴ 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, the Canadian government adopted in October 1980 the *National Energy Policy* (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 on March 1985 between a new 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

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2. Most of Canada's energy-related exports go to the United States, accounting for about 25% of US energy imports in recent years.
 3. The notes in this section are 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.
 4. 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.

deregulated domestic oil prices and it lifted controls on short-term oil exports.⁵ The 1985 Agreement also led to a progressive deregulation of the natural gas industry that continued into the early 1990s.⁶ The Canada-US Free Trade Agreement (in 1989), and the broader agreement that included Mexico, i.e., the North-American Free Trade Agreement (NAFTA) that came into being in 1994, 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 supply disruption (Mexico requested to be excluded from this provision).⁷

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 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 of a net exporter of energy is expected to remain firm given 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 last 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 to a number of papers trying to use world commodity price shocks to explain the exchange rate movements of commodity

5. In December 1985, world prices collapsed following Saudi Arabia's attempts to recapture market share. The price collapse (a decline of over 50%) 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.

6. 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 March 25, 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% of the upstream oil and gas industry. On March 19, 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.

7. For more details, see: *North American Free Trade Agreement Overview and Description* (1992). Ottawa. Government of Canada.

producers, such as Australia (Blunndell-Wignall and Gregory 1989; Gruen and Wilkinson 1994), Canada (Amano and van Norden 1995) and Norway (Akram 2004).⁸

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.⁹ 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.

Our approach builds on Amano and van Norden (1995). Over their sample period (1973-1993), 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 their findings to see if 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 do diagnostic tests to detect the possibility of a structural break. In 3.2, we test for parameter stability treating the timing of the break as unknown. Based on the results, we re-test 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.

8. 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 appear 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, only adopting inflation targeting in 2001.

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

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-US exchange rate (CAD/USD), com is a real commodity price index excluding energy, ene is a real energy price index, and int is the Canada-US 90-day commercial paper rate differential. Except for int , all variables are expressed in logarithms. Data sources are provided in Appendix 1. Estimates of Eq. 1 for various sample periods suggest that the relationship between real energy prices and the CAD 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.

One way to look at the evolving relationship between real energy prices and the CAD is through recursive estimates of Eq. 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 HAC standard errors.¹⁰ It 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 Eq. 1 reveal that β_e is significantly negative for the most recent sample period (Chart 6).¹¹

The same exercise is repeated in Chart 7, except 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:

$$rfx_t = \beta_t x_t + \varepsilon_t, \quad (2)$$

10. 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.

11. In this case, the regressions start with the most recent sample period and is extended backwards in time.

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 (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.¹²

The *SupF* corresponds to the largest value of a series of Chow-type tests based on recursive estimates of Eq. 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 $\tau \subset T$.¹³ 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.¹⁴

The finite sample performance of the three statistics was evaluated by Kuo (1998), who also extended 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 null of parameter stability at the 5 per cent level.^{15,16} Five cases of the null hypothesis are considered: (1) $\mu_t = \mu$, (2) $\beta_{e,t} = \beta_e$, (3) $\beta_{c,t} = \beta_c$, (4) $\beta_{c,t} = \beta_c$ and $\beta_{e,t} = \beta_e$, and (5) $\mu_t = \mu$ and $\beta_{c,t} = \beta_c$ and $\beta_{e,t} = \beta_e$. The results clearly suggest the presence of

12. 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 as 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), and 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.

13. Andrews (1993) suggested 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.

14. 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.

15. We thank Biing-Shen Kuo for making his GAUSS program available to us, which was used to generate the test statistics in Table 2.

16. As 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 bandwidth 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 bandwidth parameter take on values from 0 to 9. We also tested with the plug-in bandwidth estimator recommended by Andrews (1991). Both procedures generally yielded similar timing for the break date.

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.

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 $\tau=1993Q3$.

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 analysed the impact of structural breaks on the performance of unit root tests. He showed that standard unit root tests, like the augmented Dickey-Fuller (ADF) test, have dramatically reduced power when the underlying process undergoes a structural break.¹⁷ As a solution, Perron developed 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), henceforth PN, this assumption is relaxed as 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 PN considers the same unit root null of the ADF tests, he considers three different alternative hypothesis. These are that the underlying process is trend stationary with a one-time break in (1) mean, (2) trend and (3) mean and trend. PN referred 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 PN tests along with the modified ADF test due to Elliot et al. (1996), henceforth ERS, to determine whether the variables in the model are non-stationary.¹⁸ Estimation details for 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

17. Intuitively, a unit root process is defined with 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 argued 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’.

18. 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.

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, as 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.¹⁹

We approached the issue of cointegration in two different ways. First, we followed the approach of Quintos (1995) and applied 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 second subsample, which may be due to a smaller sample size in the second period.²⁰

Our second approach was 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 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 error correction (ECM) based tests do not.²¹ 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 affect 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

19. For the Canada-US 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 suggest that they are mean stationary below the 10 per cent level. Based on this result and the fact that non-stationary implies a counter-intuitive conclusion that the interest rate differential is explosive, we assume *int* to be stationary.

20. 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 affect on the CAD.

21. 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 re-arranging 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 ignores 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 et al. (1996) for further details.

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.

ILM has the following specification:

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

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

Chart 8 plots dynamic simulations from 1973 onwards of AvN and ILM and compares them to 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 last 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 9, 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.²² 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.

22. 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 test whether the value is significantly different than zero.

4. Conclusion

Based on a variety of tests, we find that the ability of Amano and van Norden equation to tract the evolution of the CAD 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 suggest, 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 found 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, our findings suggest that traditional relationships which can lose their value in analysing 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.

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Appendix 1

1. *rfx*: Real Canadian per US dollar exchange rate.

Quarterly average of the daily spot rate recorded by the Bank of Canada at 12:00 p.m. E.S.T., multiplied by the ratio of US to Canada implicit GDP deflator. US and Canadian deflators are indexed to 1997=1.0 and are obtained from the US Bureau of Economic Analysis and Statistics Canada series v1997756, respectively.

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

Quarterly average of the daily non-energy commodity price index (1982-1990=100) calculated by the Bank of Canada. Deflated by the US implicit GDP deflator from 1.

3. *ene*: Real energy price index.

Quarterly average of the daily energy commodity price index (1982-1990=100) calculated by the Bank of Canada. Deflated by the US implicit GDP deflator from 1.

4. *int*: Canada - US nominal interest rate differential.

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

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
Mean F	2.79	4.34	3.16	8.09	9.23
Sup F	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**). The test statistics are obtained from FM-OLS estimates using a Bartlett kernel with bandwidth 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 case 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 - <i>C</i>	τ_C	PN - <i>G</i>	τ_G	PN - <i>B</i>	τ_B
<i>ene</i>	-1.229 [4]	-3.730 [3]	1985Q3	-3.484 [3]	2000Q2	-3.189 [3]	2001Q1
<i>com</i>	-2.027 [0]	-4.995 [4]	1993Q1	-5.009 [4]	2000Q3	-5.492 [4]	2001Q1
<i>rfx</i>	-2.532 [3]	-2.743 [3]	1992Q4	-2.907 [3]	2000Q4	-3.300 [3]	2000Q2
<i>int</i>	-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 Dicky-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 et al. (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 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 was 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. Specification for the estimated models are as follows:

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

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

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

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 Michelies (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 less than 10 per cent. Test type assume 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, SAAR

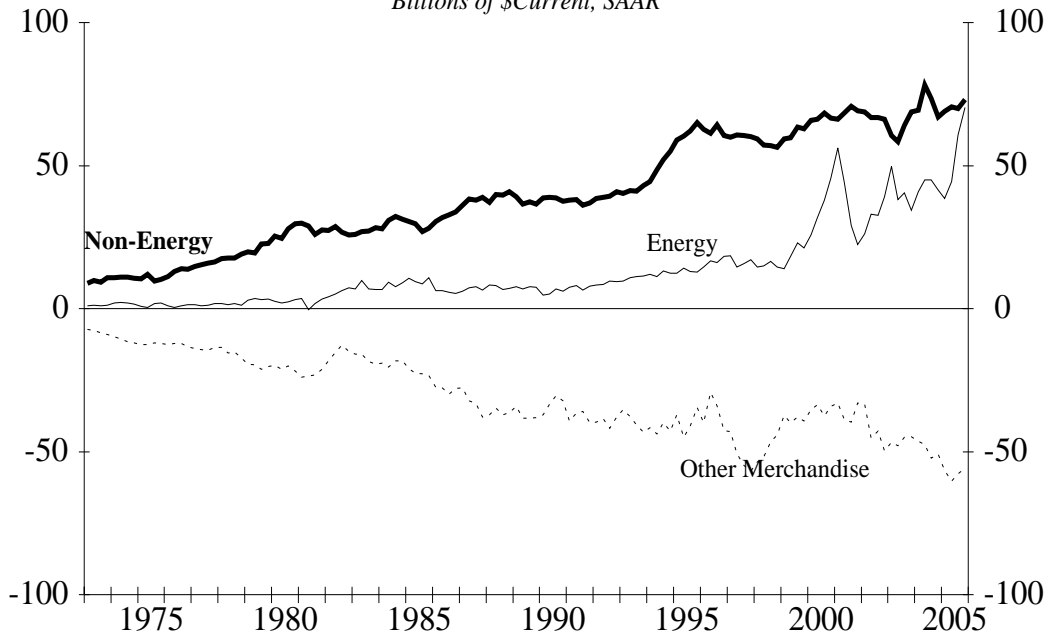


Chart 2
Canada's Real Net Exports
Billions of Chained \$1997, SAAR

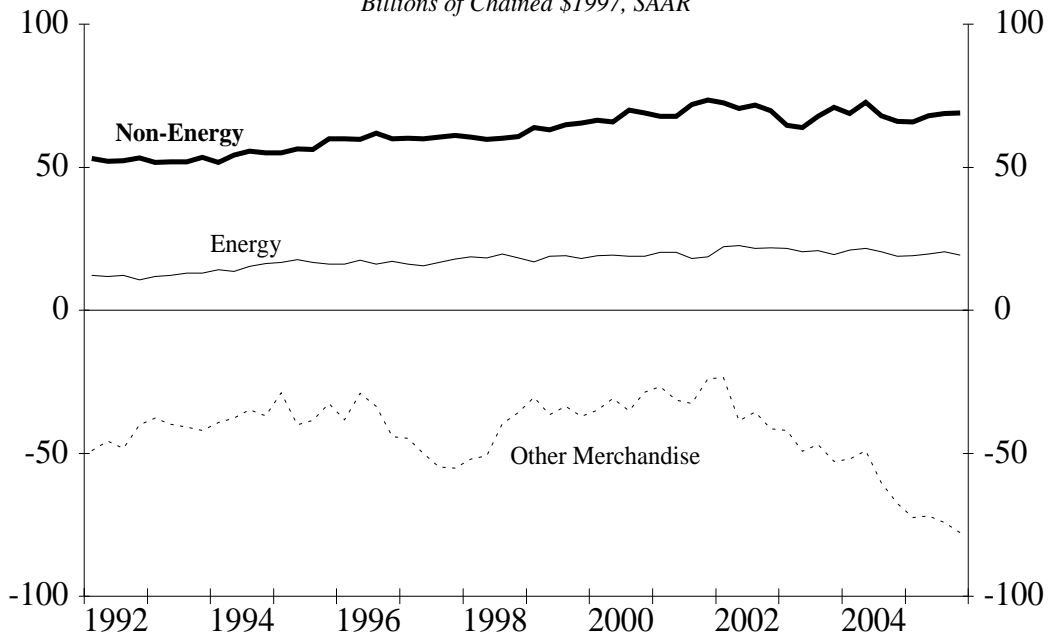


Chart 3
 Canada's Crude Oil Net Exports
Billions of \$Current, SAAR

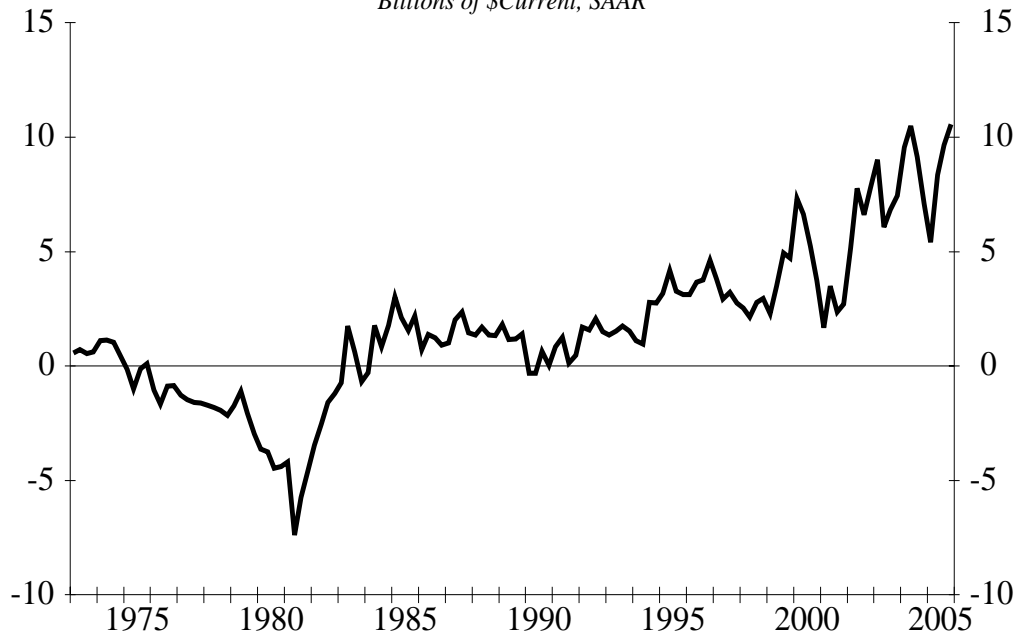


Chart 4
 Canada's Decomposition of Energy Exports
% of Total Energy Exports

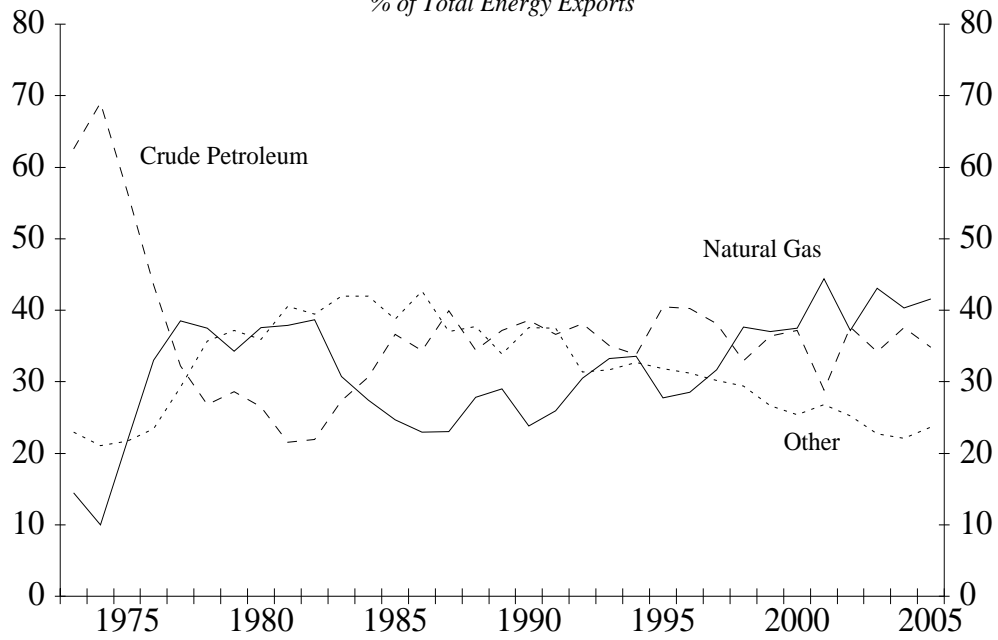


Chart 5
Energy Parameter Value: Forward Recursive Regression
Beginning 1973Q1 with P-Values > 0.1 Dotted

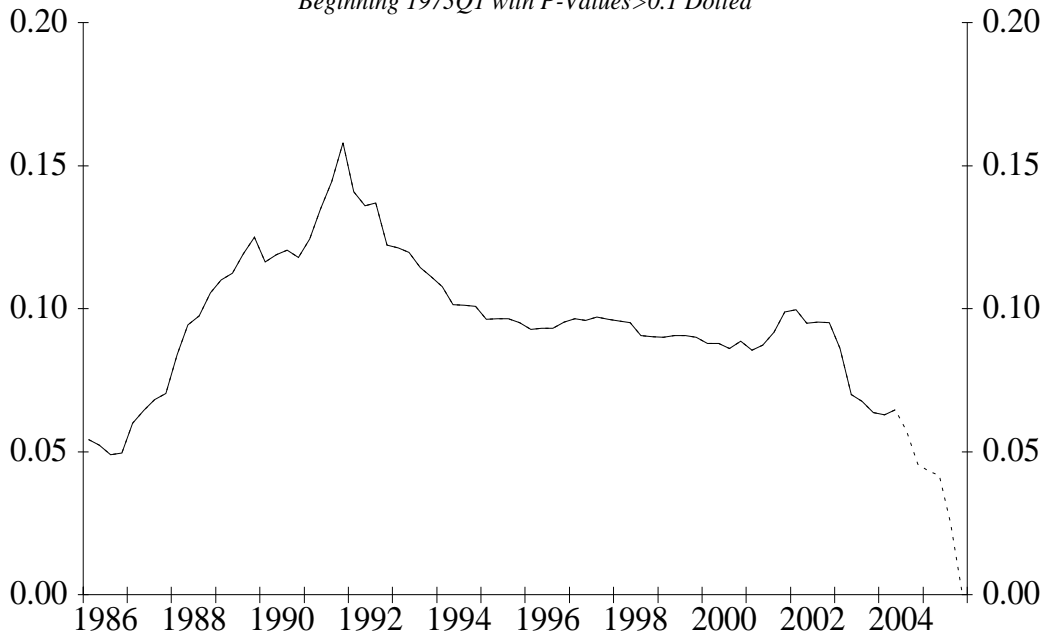
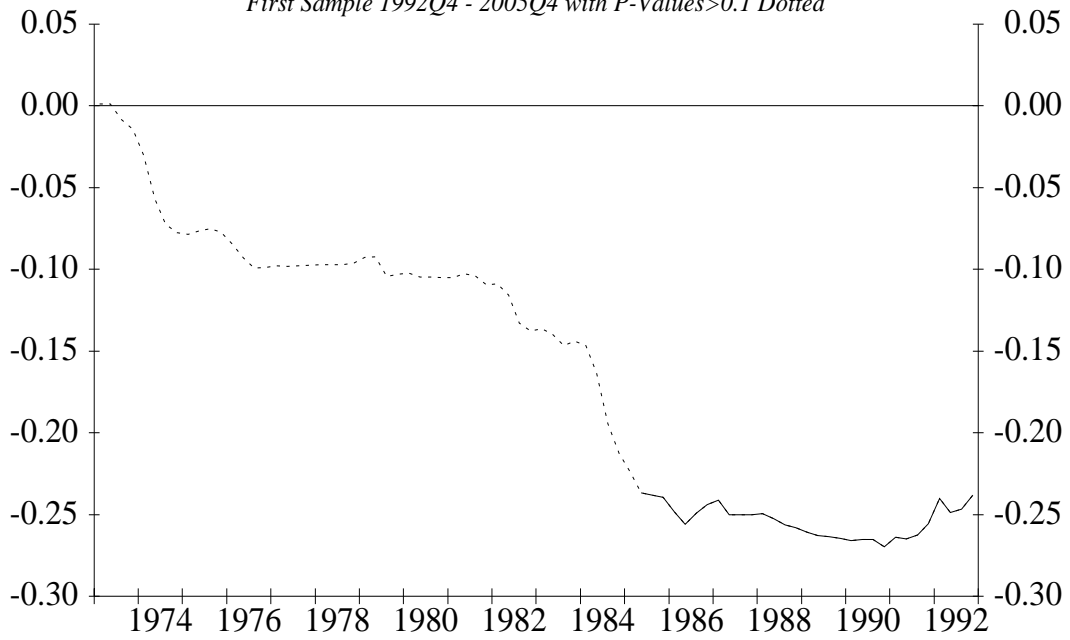


Chart 6
Energy Parameter Value: Backward Recursive Regression
First Sample 1992Q4 - 2005Q4 with P-Values > 0.1 Dotted



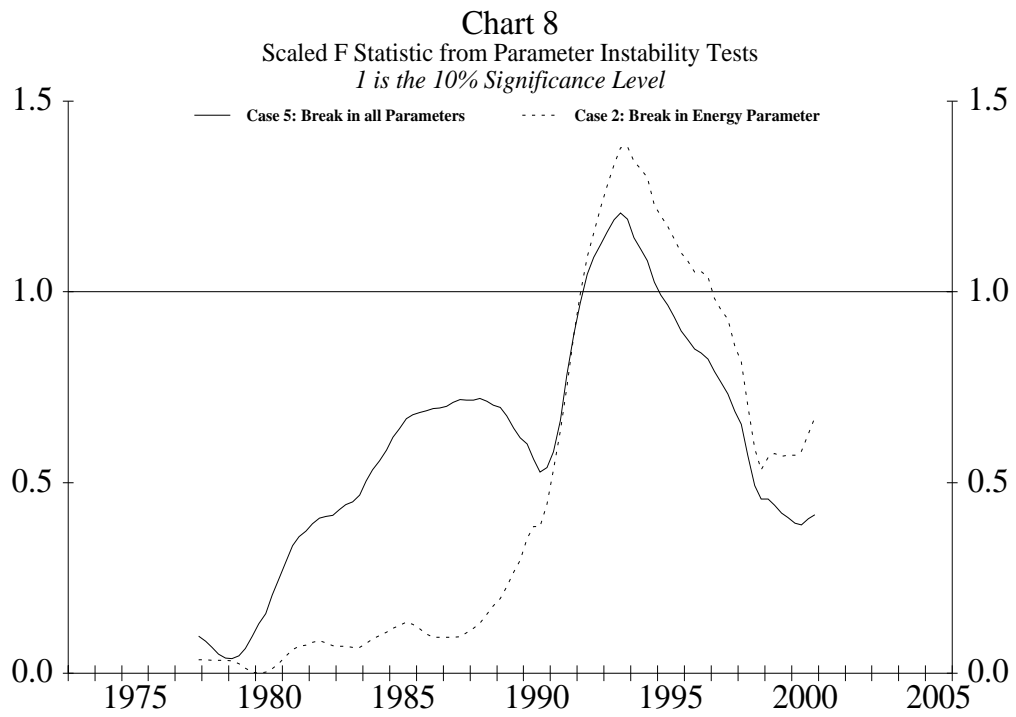
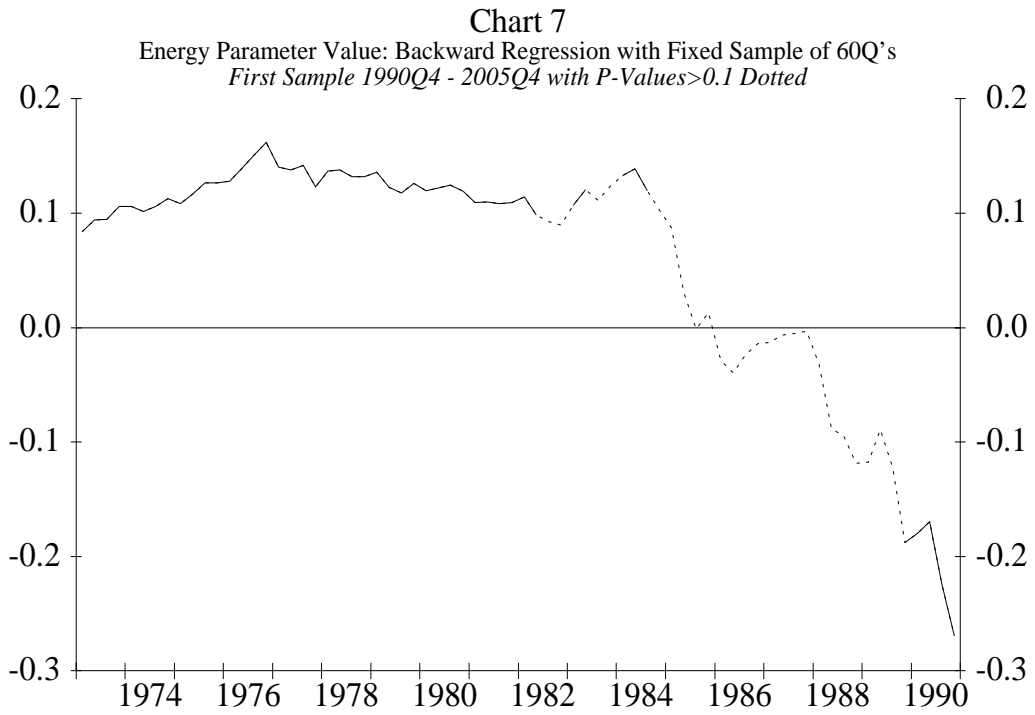


Chart 9
 Actual vs. Dynamic Simulation
 Estimated 1973Q1-2005Q4, Simulation Begins 1973Q1, US Cents

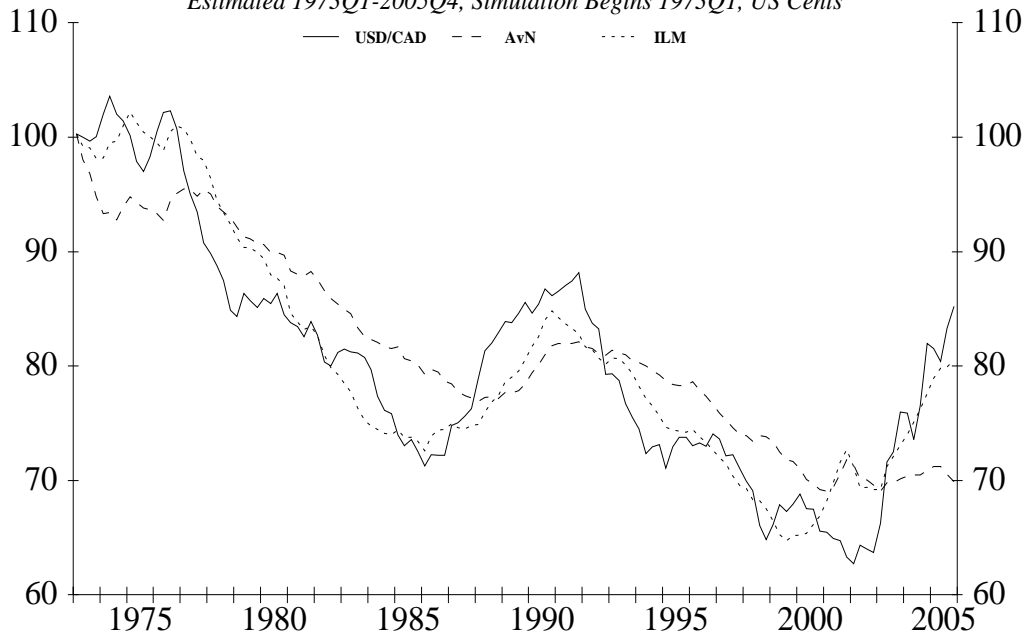


Chart 10
 Parameter Values with Regime Change: Forward Moving Dummy
 First Dummy at 1983Q1 with P-Values > 0.1 Dotted, Estimated 1973Q1-2005Q4

