Quantifying Non-tariff trading barriers: What difference did the U.S. security precautions following 9/11 make to Canadian cross border shopping?¹

by

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July 2009

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¹ I would like to thank Jing Jing for research assistance on this project.
1. Introduction

With the success that the WTO and GATT agreements have had in lowering tariffs and subsidies as impediments to the free flow of trade between countries, attention has turned to the role of non-tariff trade barriers. Here the question is the extent to which non-price impediments—the use of quotas, differential quality standards, special labeling, origin requirements, etc.—lower rather than raise national income and welfare. In answering this question, policy makers accept that the mere existence of a trade related problem is not sufficient to justify the restriction of trade and now try to assess whether the welfare losses from such trade restrictions more than offset the benefit to domestic protection.\(^2\) One difficulty in coming to grips with this question is the problem of quantifying the impact on trade of non-tariff restrictions. This is often because the restriction predates the concern with lowering trading obstacles such that evidence of what trade flows would have been in their absence is unavailable. When a policy is changed, however, pre-trade patterns can be compared with those that followed the adoption of a new (or removal of an old) impediment. Hence revisiting the data allows the construction of a counterfactual that can provide information that was missing when that policy was first adopted.

In this paper I illustrate how one dimension of trade loss that follows the adoption of a new non-tariff trade barrier can be quantified. In this case it is the degree to which cross border shopping has been affected by the additional security measures adopted following the attack on the World Trade Center on September 11 2001. While the advantage of having a more secure border is widely accepted, it is also recognized that tighter security arrangements have raised the cost of cross border trade. A U.S. Department of Transportation study (2003) reports that the

\(^2\) Note that the welfare effect from a reduced volume of trade need not be negative. A reduction in cross border shopping would increase welfare if there was a negative externality associated with the volume of same day cross border trade.
cost to the Canadian and U.S. economies of the present border management system and trade policies is between $7.52 and $13.2 billion (U.S.) dollars annually and this includes only the direct costs of uncertainty and delay to the carriers and manufacturers involved (as well the administrative costs of greater border surveillance). However, by raising costs, the volume of trade should also have been affected. On this possibility recent research offers conflicting answers. Burt (2007) finds that with the exception of border crossings at Port Erie, there is little evidence that tighter border security has actually reduced export volumes, whereas Globerman and Storer (2008) conclude the exact opposite. In the following pages I use the results of a cross border shopping model to demonstrate that the restrictions following 9/11 did significantly reduce the amount of same-day cross border travel by automobile by Canadians. For at least this dimension of cross border trade, the concern with higher security has not only raised the cost but reduced the quantity of cross border trade.

The paper proceeds as follows. Section 2 presents a simplified version of a cross border shopping model (Ferris, 2000) that uses changes in relative prices, travel costs and institutional constraints to explain the evolution of Canadian cross border shopping into the U.S. and adapts that model for testing. In Section 3 estimates of the implied cointegration model are provided first for the original 1972:01-1997:12 time period and then for the longer time period (1972:01 – 2001:08) leading into 9/11. The results suggest that the key determinants of cross border shopping have remained constant in their effect over the somewhat longer time interval. Using the latter outcome as our empirical model, the pre-9/11 model is projected into the post 9/11 time period to become our counterfactual. Hence the difference between the forecast and the actual outcome becomes the measured effect of tighter border security on same-day cross border travel. An alternative approach is found by re-estimating the model over the entire 1972:01 to 2008:08
time period and using a 1/0 dummy variable to capture the discrete change made by the post 9/11 security precautions. The two results reinforce each other and the implication that the new security arrangements have had real economic impact and significance. In Section 4 the error correction models implied by the cointegration equations are presented and discussed for their insights into the short run adjustment process about the longer run equilibrium. Section 5 summarizes our conclusions.

2. A Simple Cross Border Shopping Model

To develop the counterfactual model of cross border shopping, I use a simplified version of a model developed by Ferris (2000). In that paper Canadian consumers are assumed to value two distinct types of consumer goods (goods that can and goods that cannot be smuggled) in addition to leisure time and government services (provided through commodity taxes). Hence depending on the cost of domestic goods that can be smuggled relative to their foreign alternative and on the cost of traveling to the U.S. to acquire these goods, Canadian consumers choose whether or not to cross the border to shop. In addition to the money savings that encourage or discourage cross border shopping, differences with respect to the qualitative dimension of shopping—the hours and days on which individuals can shop at home versus abroad and/or the level of screening of goods and individuals at the border—will also affect the net benefit of cross-border shopping and so change trading patterns.

Because many of the costs of cross border shopping are independent of the number of goods purchased, shopping will take place discretely in time such that same-day border crossings by automobile are often used to proxy the number of cross-border shopping trips (SAME DAY).³

³ There is no tax import exemption for one day travel into the U.S. Canadians must be out of the country at least 48 hours before a sliding (time) scale of exemptions begins.
Similarly because most cross border shopping is done by car, the goods that are targeted by cross border shoppers are typically those with a meaningful price advantage that are easily transported by car and are hard to detect by cursory border inspection. Hence the types of goods considered suitable for smuggling consist of goods with persistent cross-country price differences (products that are taxed less highly or regulated more leniently) that are also purchased frequently in both countries (so that country origin is ambiguous) and that are small in size relative to market value (easily disguised or hidden). Products that fit this bill include: groceries, clothing, gasoline, tobacco, and liquor. The discreteness of the shopping trip implies that cross-border shopping will be motivated more by the desire to purchase a bundle of these goods than the return to any individual item (PRBUNDLE). Combining this with the quantity restrictions imposed by car travel means that the feasibility of a shopping trip and hence the number of shopping trips made per period will depend on the savings that can be realized on the bundle. Last, because cross border shopping involves distance, the cost of the trip will depends both on the opportunity cost of the shopper’s time as well as the direct cost of travel by automobile. The former is metered (inversely) by the unemployment rate (URATE) and the later by gas prices (PR_CANGAS).

Because time costs figure heavily in cross-border shopping, the administrative hassle at the border becomes a factor that will be subtracted from pecuniary savings when comparing the purchase of the same goods either by direct mail or through Canadian retailers. And in the period between 1972 and 2009, a number of legislative changes affected the time/convenience cost of

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4 A product that is often mentioned for U.S. shoppers in Canada is prescription drugs. A separate CPI category for prescription drug prices was available only from 1985 onwards; however, experimentation by adding drug prices individually and as part of the shopping bundle over this shorter time interval suggest that prescription drug prices may be another part of the cross-border shopping story for Canadian cross border shoppers as well.

5 The time hypothesis is that the time cost of shopping is higher when employed than when unemployed. This is also when the same cost savings from cross border shopping are more highly valued.
using these alternatives relative to cross border shopping.\textsuperscript{6} They include: the liberalization of shopping hour restriction during the week and particularly on Sundays in Canada [\text{QUEBEC}, August 1990 and \text{SUNDAY} Ontario, July 1992] and the US/Canada Free Trade Agreement (\text{USFTA} in December 1988). The latter lowered the cost to individuals (relative to goods) of crossing the border. On the other hand, the successive stages of tariff reduction in the implementation of the North American Free Trade Agreement (\text{NAFTA}, January 1994 and \text{NAFTA2} January 1998) lowered the cost of taking goods to individuals. The final organizational change considered, and the particular focus of this paper, is the dramatic tightening of security arrangements for individuals attempting to cross from Canada into the U.S. following 9/11. The lengthy line-up of cars now crossing the border on weekends and holidays signals both a higher time cost and a greater probability of smuggling detection (due to the more intense inspection of cars and their content) for cross border shoppers.

In Table 1 the descriptive statistics of the variables used to test the hypotheses discussed above are presented. Note in particular that all time series variables are nonstationary, i.e., are I(1), and become stationary only after being first differenced. The data is described in greater detail in the paper’s appendix where the sources are also provided. Using these variables, the model’s comparative statics lead to the following estimation equation as a joint test of the set of proposed cross border shopping hypotheses:

\begin{equation}
\text{SAME DAY} = c_0 + c_1 \text{Pr Bundle} + c_2 \text{Pr Cangas} + c_3 \text{URATE} + c_4 \text{SUNDAY} + c_5 \text{QUEBEC} + c_6 \text{USFTA} + c_7 \text{NAFTA} + c_8 \text{NAFTA2}
\end{equation}

\textsuperscript{6} These changes in institutional arrangements of shopping are treated as known breaks in the time series connecting cross border travel and cross border relative prices and travel costs. They are all 0/1 dummy variables.
The comparative statics predict that the sign of coefficient estimates of $c_1$, $c_3$ and $c_6$ will be positive and the sign of $c_2$, $c_4$, $c_5$, $c_7$, and $c_8$ will be negative. This equation, estimated over the 1972/01 to 2001/08 time period, becomes the counterfactual model of cross border shopping used to describe shopping behavior in the absence of the 9/11 security arrangements.

---inset Table 1 here---

### 3. The effect of 9/11 on Cross Border Shopping

As is well known, a series of I(1) variables can lead to estimation results that are spurious by suggesting the presence of a significant relationship among variables that have only a time trend in common and no other statistical interrelationship. This is the potential problem presented by estimating equation (1) above. However, if the residuals of a linear combination of I(1) variables are stationary, then the group of variables is said to be cointegrated and the linear estimate describes a long run equilibrium relationship among these variables (Engle and Granger, 1987). Hence testing the hypotheses in the manner suggested by equation (1) first requires evidence of cointegration before we can treat the relationship seriously. In Table 2 below, the OLS estimates of the simplified shopping model over three different time periods are presented. The last row in the table reports the results of the Adjusted Dickey Fuller test for the stationarity of the equations’ residuals. The results are all consistent with stationarity in the residuals and thus with the hypothesis that estimation results for equation (1) describe a long run equilibrium relationship.

Column (1) of Table 2 presents the results of running the simplified model on data over the 1972:01-1997:12 time period. This is done simply to compare the simplified model first with the results found in the more detailed cross border shopping model used originally by Ferris
(2000). In column (2) the time period is extended through 2001:08, to cover the full period immediately prior to the World Trade Centre attack. This performs a robustness test of the underlying model and generates the empirical model that we can project into the post 9/11 time period as our counterfactual. Column (3) then presents the estimate of the same model over the entire 1972:01 – 2009:05 period using a 0/1 dummy for the months following 9/11. This is our complementary test of the strength of the new security restrictions on trade.

As expected, the results presented in column (1) are broadly consistent with the results presented in Ferris (2000). Hence an increase in the bundle price of goods that can be smuggled back into Canada relative to the same U.S. priced bundle increases significantly Canadian same day cross border travel. A fall in the opportunity cost of the time used in cross border shopping, as represented by a rise in the Canadian unemployment rate, also significantly increases cross border travel. On the other hand an increase in the direct travel cost of cross border shopping, a rise in the Canadian price of gasoline, has had no significant effect on cross-border travel over this time period. The results found for the effects for the institutional shopping dummies also mimic those found in the earlier study. In particular, the passing of the USFTTA is associated with a rise in same day cross border travel, consist with institutional arrangements lowering the obstacles to individual travel relative to the movement of goods and services. The liberalization of shopping hour restrictions on week nights and Sundays in Ontario and the tariff reducing

---insert Table 2 about here---

To assess the significance of the individual coefficients properly, I followed Saikkonen (1991) to adjust the equations and their standard errors. This requires re-estimation of the models in Table 2 by adding first differences of all left and right-hand side continuous variables together with their leads and lags and then adjusting the resulting standard errors (for correlations among the innovations). In this case the resulting equations appeared very similar to those presented in Table 2 but with a Saikkonen adjustment factor for standard errors of .76 for the equations in columns (2) and (3). Hence while the significance of the coefficient estimates in Table 2 is overstated, the Saikkonen adjustment does not change the significance of any variable. This work is available upon request.
features of NAFTA had their expected effect of decreasing cross border shopping. As in the earlier study, the relaxation of shopping restrictions in 1990 in Quebec stands as the exception to the general rule (increasing rather than decreasing cross border shopping). In this case, however, there is a suggestion in the literature that the extension of shopping hours in Quebec increased rather than decreased domestic retail prices and this rise could explain the otherwise perverse finding.8

Re-estimation of the model over the longer time period leading into 9/11 illustrates the robustness of these early estimates. After adding a second dummy variable to account for the second stage of NAFTA implementation in 1998, the estimated results in column (2) appear virtually identical to those presented in column (1). Only in the (insignificant) case of Canadian gas prices is there any discernable change in an estimated coefficient size. The constancy of the significant coefficient estimates increase our confidence that the empirical model of column (2) is a stable relationship that can serve as a useful empirical of cross border shopping that can be projected into the future. Using the estimated model from the pre 9/11 time period and the actual explanatory variables post 9/11, we project same day border crossings through 2009:04 as the counterfactual representation of what cross border trade flows would have been in the absence of 9/11. The model forecast is shown as the dashed line in Figure 1. The solid line on the diagram represents actual same day cross border travel.

---insert Figure 1 around here---

By inspection it can be seen that the estimated model does a good job of matching actual travel flows through September 2001. What is quite striking, however, is that immediately following 9/11 actual crossings fall off dramatically while projected crossings do not. The

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8 Tanguay, Valee and Lanoie (1995) find retail prices in the large retail stores that benefit best from extended hours rise following shopping hour liberalization (as the consumer price of receiving greater shopping convenience).
difference is sudden and sizeable, consistent with a large immediate response to the new security arrangements.\footnote{Same day cross border travel fell from 2,221,530 individuals per month in August 2001 to 1,501,864 per month in October of 2001, a change of over 700,000 and a drop of almost one third (32.4%).} It is also apparent from the diagram, however, that not only is there a large discrete fall in the level of cross border travel at 9/11, but the size of the gap has been maintained and perhaps even increased in the time that has followed. This suggests that the new security measures, and shoppers’ reaction to these stricter arrangements, are not simply transitory but have become a permanent deterrent to cross border travel.

An alternative way of assessing the significance of 9/11 for Canadian cross border shopping is represented in the regression results of Table 2, column (3). Under this approach the shopping equation is run over the entire 1972:01 – 2009:04 time period, with a dummy variable added for the post 9/11 period (1 for September 2001 and the months following, 0 for the months earlier). By inspection it can be seen that the overall equation estimate has a remarkable degree of conformity with the two earlier models estimated on pre 9/11 data. All coefficient estimates have the same sign, with no evidence of diminished significance. If anything, the results of column (3) indicate that greater significance should be given to some shopping hypotheses (compared to what was found earlier). In particular, travel costs as represented both by Canadian gas prices and unemployment rates have become significant determinants of same day cross border travel.

In column (3), the size and significance of the 9/11 coefficient stand out. The coefficient itself is significantly negative and its size indicates a large downward shift in the level of cross border travel from 9/11 onwards. The coefficient estimate implies that over 800,000 individual shopping trips per month were lost over the period following 9/11, a loss of well over 1/3 of the monthly trips that were being made in the months immediately prior. The dashed line in Figure
2 shows the outcome predicted by the model for the 1972:01-2009:04 time period. This is contrasted with the solid line reflecting actual travel. While comparison illustrates that the model fits the data quite well, the immediate post 9/11 prediction overestimates somewhat the size of immediate negative effect on cross border travel and underestimates somewhat the persistence of the effect over the longer run.

--insert Figure 2 here--

More generally, the two methods used to assess the impact of 9/11 on cross border travel are complementary in suggesting a strong persistent negative effect on cross border shopping. The side effect of the desire for greater border security has been a significant reduction in this dimension of cross border trade. What is also of interest is that while the size of this non-price trading impediment is quantitatively large, the effect produced is not significantly different in absolute size from some of the other institutional effects on cross border travel. The effects on cross border shopping produced by such changes as the adoption of the USFTA, the relaxation of shopping restrictions in Quebec and Sunday shopping in Ontario, and the first two stages of NAFTA appear to be quantitatively similar. While one could argue that the levels of monthly crossing were much higher in the earlier periods so that the relative impact of these changes was smaller, the more important point is that there have been a number of changes in the shopping arrangements that have affected shopping convenience that have had a significant statistical effect on the level of cross border shopping. So that while pecuniary incentives may motivate cross border shopping and direct its evolution through time, both legislative, organizational, and regulatory changes that affect the convenience (in addition to the money savings) of crossing the

10 The introduction of NAFTA in January 1994, for example, had a similar sized estimated effect on the number of same day cross border travelers (roughly .744 versus .823 million), but with cross border traveling averaging 4 million per month in the period before NAFTA’s introduction in 1994 (versus 2 million in 2001), the relative impact of NAFTA on travel was much smaller (an 18.6% change versus 41.1%).
border can be evaluated quantitatively for their effects on the volume as well as the cost of cross border trade.

4. Shorter run adjustment: the implied error correction models

In Table 3 the error correction models of short run adjustment about the long run equilibriums (represented by the cointegration equations in Table 2) are presented. In each equation the error correction coefficient is found to be significantly negative—as required to produce the reversal of temporary deviations from the long run equilibrium. The negative error correction term then reinforces the hypothesis that the linear equation in levels represents an equilibrium relationship. The size of the coefficient estimates indicate a relatively long period of adjustment, with transitory deviations taking as long as a year before affecting convergence back to long run equilibrium.

The equation results also reinforce earlier individual findings in first indicating that cross border travel does respond significantly to short run changes in the price of the bundle of smuggled goods. The estimates indicate that travel response is at least as strong in the second month of a change as in the first, implying that price effects build in strength through time before producing their full long run effect on cross border travel. The short term results also mirror long term findings in suggesting that it is only over the latest time period that time considerations have become significant predictors of cross border travel. In this case only changes in the unemployment rate signal changes in cross border travel and only in the latest time period. Finally, the institutional and regulatory changes that indicated significant breaks in the stochastic process describing the long run also indicate breaks in the stochastic process describing short run adjustment. However, while all of the dummy variables enter with their same sign, only the USFTA and the two NAFTA dummies consistently retain their significance. In this context, the
size and significance of the 9/11 dummy variable stands out even more clearly than it did in the long run equation. The size of the contemporaneous 9/11 coefficient indicates that the change brought about by 9/11 is the largest of the institutional shocks on short run same day cross border travel, roughly double the size of the USFTA and NAFTA tariff effects. In addition, the data suggest that the effect of 9/11 has increased over time, with the 9/11 dummy producing a second significantly negative effect in the period immediately following.

--Insert Table 3 about here--

5. Conclusions

In this paper I have illustrated the idea that non-tariff trade barriers such as the tightening of the security arrangements in the period following 9/11 will produce reductions in the volume (as well as generating higher real costs) of cross border trade that can be measured empirically. By focusing on the case of cross border shopping by Canadians, an empirical counterfactual is constructed and used to determine the magnitude of the fall in same day cross border travel in the period following 9/11. The robustness of these findings was assessed by rerunning the shopping model over the entire sample period with a dummy variable for the post 9/11 time period and by investigating the characteristics of the implied error correction model. The results found are consistent and indicate that the number of same day cross border shoppers has fallen by as many as 800,000 individuals per month in the post 9/11 period for reasons unrelated to cross country price differences and usual travel cost considerations. This represents roughly one third of the level of shoppers that had crossed the border to shop in the period immediately before the imposition of the new security arrangements.
Table 1
Descriptive Statistics for Monthly Cross Border Shopping Data:
1972:01 – 2009:05

<table>
<thead>
<tr>
<th>Variable Name</th>
<th>Mean</th>
<th>Maximum</th>
<th>Minimum</th>
<th>Standard Deviation</th>
<th>ADF (level)</th>
<th>ADF (difference)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Sameday border crossings (seasonally adjusted)</td>
<td>2.43 million per month</td>
<td>5.350</td>
<td>1.502</td>
<td>0.881</td>
<td>-1.22</td>
<td>-23.7</td>
</tr>
<tr>
<td>Pr_Bundle</td>
<td>0.426</td>
<td>0.584</td>
<td>0.339</td>
<td>0.055</td>
<td>-1.33</td>
<td>-4.01</td>
</tr>
<tr>
<td>Relative_Cangas</td>
<td>0.984</td>
<td>1.70</td>
<td>0.680</td>
<td>0.180</td>
<td>-2.08</td>
<td>-6.64</td>
</tr>
<tr>
<td>URate (seasonally adjusted)</td>
<td>8.28%</td>
<td>13.00</td>
<td>4.70</td>
<td>1.83</td>
<td>-1.85</td>
<td>-18.67</td>
</tr>
<tr>
<td>Sunday Shopping</td>
<td>1 July 1992 onwards</td>
<td>0</td>
<td>0 otherwise</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>USFTA</td>
<td>1 December 1988 onward</td>
<td>0</td>
<td>0 otherwise</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>QUEBEC</td>
<td>1 August 1990 onwards</td>
<td>0</td>
<td>0 otherwise</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>NAFTA</td>
<td>1 January 1994 and onwards</td>
<td>0</td>
<td>0 otherwise</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>NAFTA2</td>
<td>1 January 1998 and onwards</td>
<td>0</td>
<td>0 otherwise</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>9/11</td>
<td>1 September 2001 onwards</td>
<td>0</td>
<td>0 otherwise</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

MacKinnon Critical values at 1% (5%), -3.45 (-2.87)
Table 2
(absolute value of t-statistic in brackets)

<table>
<thead>
<tr>
<th></th>
<th></th>
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</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>-1.61 (9.47)</td>
<td>-1.62 (10.14)</td>
<td>-0.288 (2.42)</td>
</tr>
<tr>
<td>Pr_Bundle</td>
<td>8.83 (19.50)</td>
<td>8.78 (20.93)</td>
<td>5.25 (16.58)</td>
</tr>
<tr>
<td>Relative_Cangas</td>
<td>-0.106 (0.804)</td>
<td>-0.023 (0.220)</td>
<td>-0.267 (3.08)</td>
</tr>
<tr>
<td>Urate</td>
<td>0.022 (2.11)</td>
<td>0.017 (1.93)</td>
<td>0.053 (7.08)</td>
</tr>
<tr>
<td>Sunday</td>
<td>-0.343 (5.09)</td>
<td>-0.332 (5.26)</td>
<td>-0.564 (8.88)</td>
</tr>
<tr>
<td>Quebec</td>
<td>0.904 (14.04)</td>
<td>0.918 (15.38)</td>
<td>0.889 (14.07)</td>
</tr>
<tr>
<td>USFTA</td>
<td>0.902 (13.87)</td>
<td>0.898 (14.66)</td>
<td>1.300 (24.03)</td>
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<tr>
<td>NAFTA</td>
<td>-0.600 (9.72)</td>
<td>-0.614 (10.76)</td>
<td>-0.744 (12.50)</td>
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<tr>
<td>NAFTA2</td>
<td></td>
<td>-0.426 (9.27)</td>
<td>-0.452 (9.50)</td>
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<tr>
<td>9/11</td>
<td></td>
<td></td>
<td>-0.823 (18.74)</td>
</tr>
<tr>
<td>Statistics</td>
<td></td>
<td></td>
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</tr>
<tr>
<td>Adj. R²</td>
<td>0.962</td>
<td>0.962</td>
<td>0.951</td>
</tr>
<tr>
<td>DW</td>
<td>0.445</td>
<td>0.466</td>
<td>0.402</td>
</tr>
<tr>
<td>ADF on residuals</td>
<td>-6.27</td>
<td>-6.91</td>
<td>-7.11</td>
</tr>
</tbody>
</table>

MacKinnon Adjusted Critical Values at 5 percent for 7(8) and [9] variables: -4.83 (-5.12) [-5.40].
Table 3
(absolute value of t-statistic in brackets)

<table>
<thead>
<tr>
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</thead>
<tbody>
<tr>
<td>Error Correction Term</td>
<td>-0.124* (4.01)</td>
<td>-0.132* (4.49)</td>
<td>-0.078* (3.35)</td>
</tr>
<tr>
<td>D(Pr_Bundle)</td>
<td>2.18** (2.26)</td>
<td>2.21* (2.58)</td>
<td>1.92* (3.30)</td>
</tr>
<tr>
<td>D(Pr_Bundle(-1))</td>
<td>2.28** (2.38)</td>
<td>2.53* (2.88)</td>
<td>2.42* (4.12)</td>
</tr>
<tr>
<td>D(Canadian Gas Price)</td>
<td>0.109 (0.494)</td>
<td>0.153 (0.827)</td>
<td>0.128 (1.27)</td>
</tr>
<tr>
<td>D(Urate)</td>
<td>0.019 (1.03)</td>
<td>0.025 (1.43)</td>
<td>0.029*** (1.82)</td>
</tr>
<tr>
<td>D(Sunday)</td>
<td>-0.147 (1.52)</td>
<td>-0.146 (1.57)</td>
<td>-0.155*** (1.76)</td>
</tr>
<tr>
<td>D(Quebec)</td>
<td>0.094 (0.958)</td>
<td>0.094 (1.00)</td>
<td>0.080 (0.899)</td>
</tr>
<tr>
<td>D(USFTA)</td>
<td>0.245* (2.48)</td>
<td>0.255* (2.69)</td>
<td>0.250* (2.74)</td>
</tr>
<tr>
<td>D(NAFTA)</td>
<td>-0.315* (3.24)</td>
<td>-0.317* (3.40)</td>
<td>-0.307* (3.45)</td>
</tr>
<tr>
<td>D(NAFTA2)</td>
<td></td>
<td>-0.211** (2.27)</td>
<td>-0.204** (2.30)</td>
</tr>
<tr>
<td>D(9/11)</td>
<td></td>
<td></td>
<td>-0.582* (6.54)</td>
</tr>
<tr>
<td>D(9/11(-1))</td>
<td></td>
<td></td>
<td>-0.182** (1.95)</td>
</tr>
<tr>
<td>D(Sameday(-1))</td>
<td>-0.150* (2.72)</td>
<td>-0.166* (3.28)</td>
<td>-0.175* (3.81)</td>
</tr>
<tr>
<td>Statistics</td>
<td></td>
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<tr>
<td>Adj. R²</td>
<td>0.138</td>
<td>0.165</td>
<td>0.214</td>
</tr>
<tr>
<td>DW</td>
<td>1.97</td>
<td>1.98</td>
<td>1.99</td>
</tr>
</tbody>
</table>

Significantly different from zero at 1% (*), 5% (**) and 10% (**).
Data Sources (all monthly)

SAME DAY = Canadian same day border crossings by automobile, seasonally adjusted: Cansim II V129491

CANCPI = Canadian Consumer Price Index, All items: Cansim II V41690973
CANALCOHOL = CPI price index for Alcoholic Beverages: Cansim II V41691207
CANAPPAREL = CPI index for Apparel: Cansim II V41691123
CANFOOD = CPI Index for Food: Cansim II V41690974
CANGAS = CPI Index for Gasoline: Cansim II V41691136
CANTOBACCO = CPI Index for Tobacco products: Cansim II V41691216

EXCH_RATE = Canadian/US Exchange rate = # of Can$/US$; Cansim II V37426

USCPI = U.S. Consumer Price Index, Urban all items: Cansim II V11123
USALCOHOL = CPI Index for Alcoholic Beverages: US Bureau of Labor Statistics Series Id: CUUR0000SAF1,CUUS0000SAF1
USAPPAREL = CPI Index for Apparel: BLS Series Id: CUUR0000SAA,CUUS0000SAA
USFOOD = CPI Index for Food: BLS Series Id: CUUR0000SAF1,CUUS0000SAF1
USGAS = CPI Index for Gasoline: US BLS Series Id: CUUR0000SETB01,CUUS0000SETB01
USTOBACCO = CPI Index for Tobacco Products: US BLS Series Id: CUUR0000SEGA,CUUS0000SEGA

PR_ALCOHOL = (CANALCOHOL/EXCH_RATE)/USALCOHOL
PR_APPAREL = (CANAPPAREL/EXCH_RATE)
PR_FOOD = (CANFOOD/EXCH_RATE)/USFOOD
PR_GAS = (CANGAS/EXCH_RATE)/USGAS
PR_TOBACCO = CANTOBACCO/EXCH_RATE)/USTOBACCO

PRBUNDLE = .2*PR_ALCOHOL + .2*PR_APPAREL + .2*PR_FOOD + .2*PR_GAS + .2*PR_TOBACCO

URATE = unemployment rate (both sexes 16 plus) seasonally adjusted: Cansim II v2062815
USFTA = 0 for the months before December 1988, 1 afterwards
QUEBEC = 0 for the months before August 1990, 1 afterwards
SUNDAY = 0 for the months before July 1992, 1 afterwards
NAFTA = 0 for the months before January 1994, 1 afterwards
NAFTA2 = 0 for the months before January 1998, 1 afterwards
US_9/11 = 0 for the months before September 2001, 1 afterwards
References


