The Economics of Border Crossings*

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Abstract

We estimate a structural model of the decision to cross an international border to purchase products. Using data from a travel survey filled out at the Canada-US border, we calculate the number of car crossings originating from each of the 250 Census Divisions in the seven border provinces in Canada, for the period 1990–2009. Our model predicts that a higher fraction of individuals in a Division will cross when the home exchange rate is strong. Because real appreciation expands the set of goods that are cheaper abroad, the model predicts a convex relationship between crossing benefits and the real exchange rate. Estimates using a fractional probit specification support this prediction. Distance from the border reduces crossing probabilities with about the same elasticity as foreign currency appreciation. Taking into account the geographic distribution of Canadian residents, counterfactual simulations show that a 10% appreciation of the Canadian dollar from its 2009 value would increase travel by Canadian residents by almost 20%.

1 Introduction

Residents of Canada and the United States made more than 64 million cross-border trips in 2009. The vast majority of these trips—more than 50 million—were made across the land border, primarily by car. This amounts to almost one trip per year for each person living within three hours of the border.[1] Indeed, Canadian residents travel more frequently to the US than they do to other provinces in Canada.[2] However, these

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[1]The combined population of Canada and the United States that lives within 300 km of the border is about 55 million.

[2]In 2004, the most recent year for which data are available from Statistics Canada, Canadian residents made 22 million inter-provincial trips, compared to 36 million trips to the US.

rates of cross-border travel vary substantially over time. Since 2001, US car trips have dropped by 55%, while Canadian car trips have barely changed. By contrast, during the period 1991–2000, Canadian travel dropped by 50% whereas US travel rose 37%.

The massive, but volatile, movement of people across the Canada-US border affects sales for businesses on both sides of the border and therefore also affects government revenues. Moreover, decisions to build infrastructure to accommodate such flows should be based on the predicted evolution of cross-border movement in the future. These considerations point to the importance of developing and estimating a quantitative model of the crossing decision.

This paper measures the effect that the US-Canada exchange rate has on the propensity of residents of the two countries to cross their common border. We start by establishing stylized facts on cross-border travel by relating monthly trips made across the land border for 38 years to the monthly average real exchange rates between Canada and the US. We then exploit detailed geographic data on the residence of Canadian cross-border travellers to estimate the parameters of a structural model of the crossing decision.

Understanding human travel is important for urban planning, traffic forecasting, and controlling the spread of infectious diseases. Recent studies of "human dynamics" have applied ideas from statistical physics to analyze data on movement of people. Using bank notes (Brockmann, Hufnagel, and Geisel, 2006) and mobile phones (González, Hidalgo, and Barabási, 2008) to track individuals, scientists have shown that most travel is over short ranges but the distribution of distances traveled has a very long tail (Levy flights). Song, Qu, Blumm, and Barabasi (2010) show that individual travel patterns are potentially highly predictable. While sharing the goal of improved modeling of human travel, our study differs starkly in terms of both data and modeling style. We measure human movement by taking advantage of the careful tracking of border crossings undertaken by the Canadian Border Services Agency. In contrast to the purely statistical models employed so far in the science literature, our model emphasizes the economic motivations that we hypothesize to underlie much human travel.

The Economics literature has also examined the phenomenon of individuals crossing borders in order to purchase goods that are more attractively priced than in their home jurisdictions. Studies of intra-national border crossings generally rely on differences in taxes, since products are priced in the same currency. They also tend to examine cross-border shopping for a single good. Chiou and Muehlegger (2010) examine the circumstances under which US residents cross state lines to take advantage of tax differences on the sale of cigarettes. Similar to our paper, they have access to survey data on the residence of individuals, which allows them to calculate the distance to the nearest state border, and thus permits them to estimate the relative importance of cigarette taxes and travel costs. However, other studies generally do not have data on the location of consumers: for example, Manuszak and Moul (2009) estimate how differences in gasoline taxes across US states create incentives for residents to cross state borders. Knight and Schiff (2010) exploit the varying payoffs offered by state lotteries, rather than tax differences, to estimate the extent to which consumers cross US state borders to purchase lottery tickets. They find that consumers are more likely to do so when they live closer to the border and when the jackpots of neighboring state lotteries are higher.

Studies of international border crossings generally exploit differences in prices created by exchange rate variation. Some of these studies explicitly measure cross-border travel: see Ford (1992), Di Matteo and Di Matteo (1993, 1996) and Ferris (2000).[3] Campbell and Lapham (2004) and Baggs et al (2010) examine the effect of exchange rate changes on firms located near the US-Canada border, with the implicit assumption that exchange rate changes have a direct effect on cross-border travel. Asplund, Friberg, and Wilander (2007) infer the extent of cross-border shopping by using data on taxes and sales of alcohol on both sides of the Sweden-Denmark border. While not measuring border crossings, Gopinath, Gourinchas, Hsieh, and Li (2010) examine grocery store products on both sides of the US-Canada border and find a large discontinuity in their prices at the border. Related papers by Engel and Rogers (1996), Gorodnichenko and Tesar (2009) and Broda and Weinstein (2008) measure the "width" of the US-Canada border by estimating the extent to which the presence of the international border generates differences in prices.

Our work contributes to these literatures in a number of ways. First, we use actual data on cross border travel, instead of inferring border-crossings from price differences or changes in retail sales. Second, rather than focus on the decision of where to purchase a single good, we model the endogenous decision of consumers

[3] These studies have relied on aggregate national data, with one exception that examines province level data. Moreover, these studies have only examined the behaviour of Canadian residents going to the US, not the reverse.

regarding the range of goods that they will purchase across the border. Our empirical results conform to the predictions of the model by establishing a convex elasticity of crossings with respect to the exchange rate, thus supporting the hypothesis that the set of goods purchased in the foreign country expands as the home currency appreciates. Third, we use data on the residence of cross-border travelers, instead of inferring this information from the geographical distribution of the population. In this way, our paper incorporates both the importance of distance, as well as the importance of relative prices on the decision to cross the border.

An initial set of reduced form regressions uncover a number of interesting patterns of cross-border travel and its relationship to the exchange rate. First, we find that US and Canadian residents respond differently to changes in the buying power of their home currency. In particular, while residents of both countries cross the border more when their currency appreciates, Canadian residents have a higher elasticity to exchange rate changes. Second, we find that, even within a country, residents of different regions respond differently to exchange rate shocks. And finally, we find that exchange rate elasticities vary over time because they depend on the level of the exchange rate. The elasticity of crossings with respect to the exchange rate increases in absolute value as the home currency strengthens.

We then develop a model to explain these patterns. Following Dornbusch, Fischer, and Samuelson (1977), our model assumes a continuum of goods available in both countries. Each country produces a subset of these goods more cheaply than the other. The model assumes that those travelers who cross the border will always purchase the set of goods in each country that is cheaper in that country. Travelers who do not cross must necessarily purchase all goods at home. The model naturally generates a prediction for a convex elasticity of crossings with respect to the exchange rate; that is, as the home currency strengthens, the elasticity of crossings rises in absolute value. This is for two reasons. First, the set of goods that are cheaper in the foreign country expands. Second, the goods that were already cheaper in the foreign country are even more attractive now.

We estimate the parameters of this model using a new dataset with information on the residence of cross-border travelers. Using this model we can estimate the effects of travel costs and the real exchange rate on the probability of crossing the border. We also conduct counterfactual experiments with respect to the key variables. We show that an appreciation of the real exchange rate would have similar effects across

Canadian provinces, but very different effects across time. Conversely, an increase in the costs associated with crossing the border would have different effects across provinces, but would have very stable effects across time.

An important component of our research is determining the extent to which the US and Canadian markets are segmented by the presence of the border. The costs generated by taxes, customs regulations and other barriers to trade may prevent consumers from arbitraging price differences. One of the goals of the next section will be to show that cross-border shopping really is an important motive for consumers. In Table 1, we present the commonly stated motives for crossing the border. The data are based on the International Travel Survey of visitors and returning residents to Canada; more details on the data are presented in Section 4.2. Approximately 50,000 travelers who cross the land border are asked to fill out these anonymous surveys each year. The responses in Table 1 indicate the main reason for the cross-border trip.

Table 1: Reasons for Crossing the Border 1990–2009 (survey responses)

Percent Answering:	Sameday		Overnight	
	US	Canada	US	Canada
Business Affairs	7.4	7.4	7.6	7.5
Visit friends/relatives	15.1	8.9	22.6	22.2
Pleasure or personal trip	43.3	53.1	62.4	64.5
Commuting to work	2.3	5.9	-	-
Other	21	15.4	7.4	5.6
Not stated	10.9	9.3	0.1	0.2
Total Respondents ('000s)	289	414	217	252

Source: Authors' calculations from the International Travel Survey

Trips for pleasure or personal reasons, which include shopping trips, are potentially the most likely to respond to exchange rates. The survey responses indicate that this is easily the largest category among stated reasons to cross the border. Trips for the purpose of business or driving to work, which are likely to be less sensitive to the exchange rate, account for under 10% of responses.

Canadian residents have a zero exemption from taxes and duties on goods purchased abroad when returning from a trip of less than 24 hours.[4] Despite this, we observe same day travel being extremely sensitive to exchange rates: we estimate the

[4]Under NAFTA, Canadian residents are not required to pay duties on most products that were manufactured in the US or Mexico. They are generally still required to pay taxes on these purchases.

elasticity of Canadian residents as well over 1. It may well be the case that some residents do not report their purchases truthfully, or that Border Agents do not bother to charge taxes for small amounts. Regardless, the fact that same-day purchases are not exempt from taxes or duties will tend to underestimate the true responsiveness of travelers to the exchange rate, and therefore serve as a lower bound for our estimate of the extent to which travelers respond to the exchange rate.

In the next section we present simple empirical results that establish patterns of cross-border travel and document the differing effects of exchange rate changes across and within countries as well as over time. In Section 3 we develop a model of cross-border travel that develops testable relationships between the variables of interest. In Section 4 we estimate this model, and use the estimated parameters to explain the observed patterns of cross-border travel. We then conduct counterfactual experiments with respect to the key variables. We conclude in Section 5.

2 Exchange rate elasticities, reduced form estimates

In this section we estimate the relationship between exchange rates and the propensity of residents of the US and Canada to cross the border. We first show that there is strong evidence that exchange rates influence travel behaviour, which is indicative of cross-border shopping. Additionally, we show that the relationship between exchange rates and travel depends on other factors as well. Notably, we find interesting variation in the response of travelers to currency fluctuations, both within and across countries.

2.1 Data

We obtained data on cross-border travel from Statistics Canada, using information collected by the Canadian Border Services Agency (CBSA) and made available by Statistics Canada as Cansim Table 427-0002. These data consist of counts of all vehicles entering Canada at all land crossings with the United States. When residents of either country make cross-border trips, they need to cross the frontier twice: once on the outbound journey and once on the return.[5] Therefore, US residents encounter the CBSA on their outbound journey and Canadian residents on their return journey.

[5]It is possible that some travelers cross the land border in one direction but travel by air or sea in the other direction. We ignore these cases, which comprise a very small fraction of cross-border trips.

Data on vehicle counts are available separately for each of the 127 ports of entry on the US-Canada border.[6] In this section we aggregate the port-level information to province-level data; we will return to port-level data in Section 4. Moreover, we restrict attention to the 7 Canadian provinces that share a land border with the United States: British Columbia, Alberta, Saskatchewan, Manitoba, Ontario, Quebec and New Brunswick.[7]

Data are available separately for passenger vehicles, commercial vehicles, trucks, motorcycles etc. We focus only on travel by passenger vehicles. The counts are separated by whether the vehicle has American or Canadian license plates, which allows us to identify the travelers' country of residence. Finally, the data are broken down by the length of the cross-border trip. In particular, US residents are asked how long they intend to stay in Canada, while Canadian residents are asked how long they were away.[8] We use data for the calendar years 1972–2009.

We obtained monthly data on the exchange rate between the US and Canadian currencies for the same years, from the Pacific Exchange Rate Service. We use these nominal exchange rates (E), along with monthly CPIs for both countries, to construct the Real Exchange Rate (e) for each month. We fixed the absolute level of e using relative price levels from OECD data for 2002.

Summary statistics of the data are presented in Table 2. Each observation is a calendar month in a given province. The data indicate that, on average, Canadian residents make more daytrips across the border than do US residents. The number of overnight trips is comparable for the two countries.

Figure 1 shows patterns in the data over time. Figure 1(a) shows monthly same-day trips by residents of the two countries over the entire 38-year period. Travel is highly seasonal, for residents of both countries. Canadian residents exhibited a sharp rise in same-day trips during the period 1988–1993. The decline in US travel in recent years appears to coincide with the period of heightened security concerns after September 2001, and stricter requirements regarding passports or other identification

[6]A small number of these are marine ports which admit vehicles arriving by ferry. There are a few other crossings that do not admit vehicles; these are primarily train crossings.

[7]Nova Scotia has a marine border with the US as it accepts ferry traffic from Maine. The Yukon Territory shares a border with Alaska. We omit these jurisdictions due to difficulties in ascertaining the corresponding US port from which vehicles enter Canada.

[8]Canadian residents are asked about their absences in order to calculate exemption limits for foreign purchases. US residents are asked about their intended period of stay so that the CBSA can make a decision on whether to approve entry.

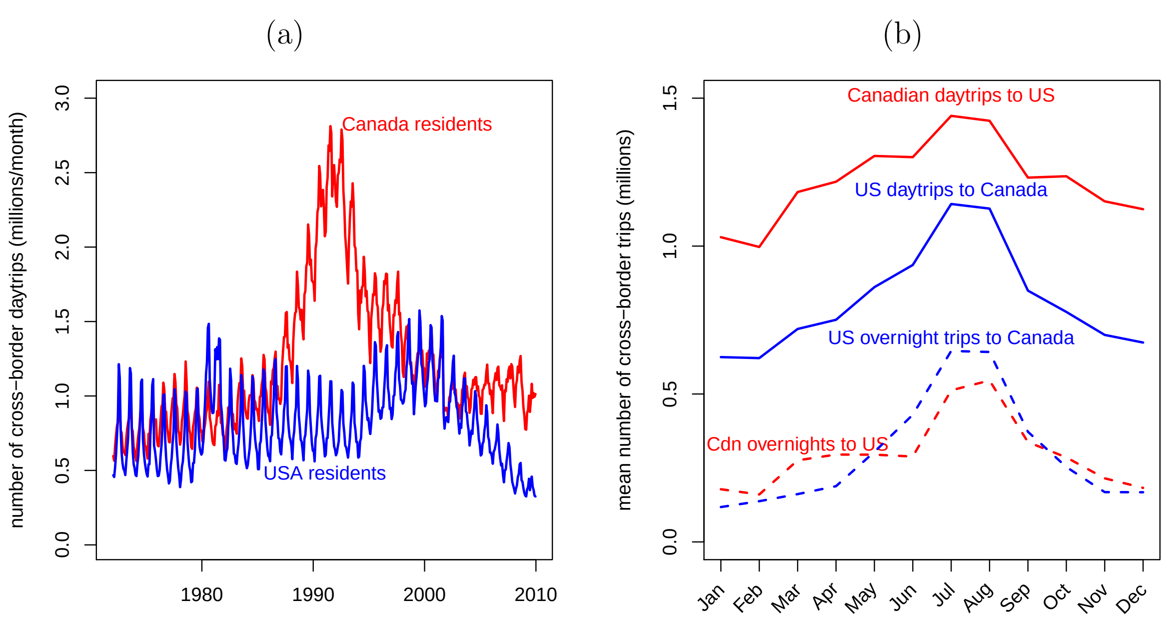
Table 2: Summary Statistics: 1972–2009 (3192 province-months)

	Mean	SD	Median	Min	Max
Day Trips (1000 vehicles):					
US	116.3	213.5	43.4	1	1224.8
CA	174.2	213.7	102.2	2.9	1192.9
Overnight Trips (1000 vehicles):					
US	41.9	72.2	14.4	0.5	519.1
CA	42.3	51.1	18.0	1.1	346.4
Nominal ER (CAD/USD)	1.241	0.165	1.225	0.962	1.600
Real ER	0.996	0.125	0.977	0.802	1.313

since 2008.

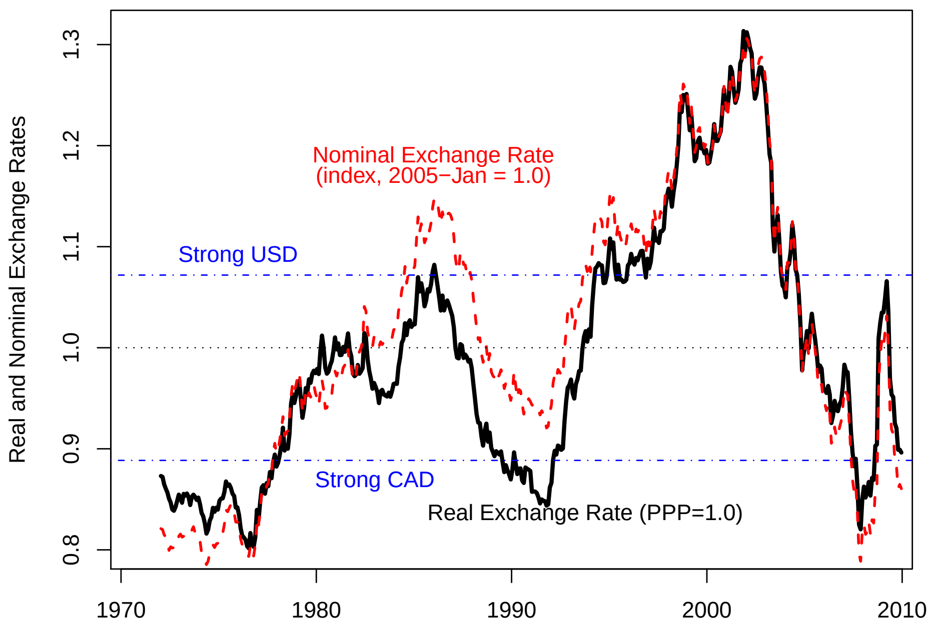
Figure 1(b) shows average travel over the 38-year period for each calendar month. We show this separately for each country of residence and for same-day versus overnight trips. Cross-border travel peaks in the summer months for all groups.

Figure 1: Annual and monthly variation in crossings



The solid line in Figure 2 shows the real exchange rate (e) starting in January 1972 and continuing to December, 2009. The dashed line shows the monthly nominal (or market) exchange rates expressed as an index of the January 2005 level (1.22 CAD per USD), when the RER was approximately one. Horizontal dot-dashed lines show the 25th and 75th percentiles of the real exchange rate—"strong USD" corresponds to $e > 1.072$ and "strong CAD" corresponds to $e < 0.888$.

Figure 2: Canada-US real and nominal exchange rates since 1972



2.2 Cross-Border Travel and the Exchange Rate

We are interested in determining the relationship between the US-Canada real exchange rate and the number of cross-border trips made by residents of the two countries. It is important to control for seasonal effects, as these are highly correlated with travel. We also control for secular trends in the propensity to cross the border. Our baseline regression specification is the following:

$$\log n_{it} = \eta_0 + \eta_1 \log \mathrm{e}_t + \eta_2 \mathrm{post911}_t + \mathrm{Month}_i(t) + \eta_3 t + \eta_4 t^2 + \varepsilon_{it} \qquad (1)$$

where i denotes a particular province or the entire country; and t denotes a calendar month. We include an indicator variable for the period following September 11, 2001 which reduced travel by residents of both countries and also led to increased border security measures. We include fixed effects for each calendar month to account for the strong seasonality in travel. Finally, we add a linear and quadratic trend to capture secular effects such as population changes. The data span the period 1972–2009. We estimate this equation separately for residents of each country.

We can also express this regression in a difference equation framework. We take 12-month differences of each observation in order to hold constant seasonal effects but also to address time-varying factors that may not be captured by the trend variables, such as the effects of free-trade agreements and changes to border security in recent years:

$$\begin{aligned} \ln n_{it} - \ln n_{i,t-12} &= \{12\eta_3 + 144\eta_4\} + \eta_1 [\ln \mathrm{e}_t - \ln \mathrm{e}_{t-12}] \\ &+ \eta_2 \left[\mathrm{post911}_t - \mathrm{post911}_{t-12}\right] + 24\eta_4 t + \varepsilon_{it} - \varepsilon_{i,t-12}. \end{aligned} \tag{2}$$

Note that the 12-month differences eliminate the linear trend and the month fixed-effects in equation 1, and transform the quadratic trend to a linear trend.

Table 3: Regression of log crossings

Method:	Levels (contemp.)				Year-on-year diffs.			
Length of stay:	Daytrip		Overnight		Daytrip		Overnight	
Residence:	US	CA	US	CA	US	CA	US	CA
$\ln e$	1.17^a	-1.61^a	0.41^a	-1.76^a	0.42^a	-1.14^a	0.14^b	-1.38^a
	(0.05)	(0.08)	(0.08)	(0.08)	(0.07)	(0.08)	(0.07)	(0.10)
N	3192	3192	3192	3192	3108	3108	3108	3108
r2	0.98	0.98	0.96	0.97	0.06	0.27	0.01	0.23
r2 (baseline)	0.98	0.97	0.96	0.95	0.03	0.06	0.00	0.00
rmse	0.25	0.22	0.30	0.23	0.14	0.13	0.14	0.15

Standard errors clustered by year-month. An observation is a province-year-month.
^{c} $p < 0.1$, ^{b} $p < 0.05$, ^{a} $p < 0.01$

The results of estimating these equations are presented in Table 3. We treat each province in a calendar month as a separate observation.[9] The first four columns present results using the contemporaneous specification described in equation 1 and the next four columns use the 12-month difference specification in equation 2. Coefficients on month fixed-effects, the post 9/11 indicator and the trend variables are

[9] In Appendix Table B.1 we present a similar regression using country-level data, instead of breaking up the data by provinces. The results in that table are similar to those presented here.

not reported. Standard errors are clustered by year-month, as exchange rates do not vary across provinces for a given month.

The results of both specifications indicate that travelers respond to the exchange rate, as represented in the negative elasticity of Canadian residents and the positive elasticity of US residents with respect to the real exchange rate. In addition, the elasticities of Canadian residents are bigger in magnitude than those of US residents, across both specifications and both categories of trip-length.

We investigate whether the elasticity with respect to exchange rates was constant over the entire 38-year period, or whether it varied according to the level of the exchange rate. In order to do this we divide the data according to quartiles of the real exchange rate. We interact e with indicator variables for whether it fell in the highest or lowest quartiles.

We present the results of these regressions in Table 4. The interactions reveal that travelers' elasticity with respect to exchange rate changes is itself dependent on the level of the exchange rate. In particular, the coefficient for the period when the US dollar was stronger than usual is generally positive, for residents of both countries. This has the effect of increasing the positive elasticity of US residents, and decreasing the negative elasticity of Canadian residents. In other words, US residents become more responsive to the exchange rate in periods when the US dollar is overvalued, while Canadian residents become less responsive. The opposite is observed during periods when the US dollar is considered undervalued.[10]

We also examine whether consumers' responses to the exchange rate vary within each country. To do this, we ran province-by-province regressions corresponding to equations 1 and 2. We report the estimated elasticities, by province, length of trip and country of residence, in Table 5.

The estimated elasticities have the expected sign for all provinces, groups of residents, and trip lengths. The results confirm our prior finding that Canadian residents have larger elasticities than US residents. This is the case for 26 out of 28 comparisons in Table 5; the two exceptions are day trips in Ontario and Alberta in the contemporaneous specification. The results also indicate that there is considerable variation in

[10] In Appendix Table B.2 we present corresponding regressions using country-level data. The results in that regression are similar to those presented here. We also conducted other robustness checks. Instead of using indicators for the top and bottom quartiles of the RER, we used a 10% cutoff above and below PPP values. We also included a second-order term for $\ln e$. All the results indicated the same pattern of exchange rate elasticities being sensitive to the level of the RER.

Table 4: Regression of log crossings using Quartiles of RER

Method:	Levels (contemp.)				Year-on-year diffs.			
Length of stay:	Daytrip		Overnight		Daytrip		Overnight	
Residence:	US	CA	US	CA	US	CA	US	CA
$\ln e$	0.95^{a}	-1.72^{a}	0.38^{b}	-2.10^{a}	0.62^{a}	-1.21^{a}	0.14	-1.54^{a}
(CAD/USD)	(0.16)	(0.17)	(0.17)	(0.19)	(0.10)	(0.09)	(0.10)	(0.14)
$\ln e \times \ [e > 1.07]$	0.88^{a}	0.61^{a}	0.76^{a}	0.75^{a}	0.09	0.36^{b}	0.30^{b}	0.42^{b}
(strong USD)	(0.21)	(0.20)	(0.21)	(0.25)	(0.11)	(0.17)	(0.12)	(0.18)
$\ln e \times \ [e < 0.89]$	-0.84^{a}	-0.73^{a}	-1.25^{a}	-0.24	-0.44^{a}	-0.09	-0.20^{b}	0.02
(strong CAD)	(0.15)	(0.17)	(0.19)	(0.19)	(0.08)	(0.09)	(0.09)	(0.12)
N	3192	3192	3192	3192	3108	3108	3108	3108
r2	0.98	0.98	0.97	0.97	0.09	0.27	0.02	0.24
rmse	0.24	0.22	0.28	0.23	0.14	0.13	0.14	0.15

Standard errors clustered by year-month. An observation is a province-month.
^{c} $p < 0.1$, ^{b} $p < 0.05$, ^{a} $p < 0.01$

Table 5: Province-specific elasticities of crossings w.r.t. exchange rates

Method:	Levels (contemp.)				Year-on-year diffs.			
Length of stay:	Daytrip		Overnight		Daytrip		Overnight	
Residence:	US	CA	US	CA	US	CA	US	CA
BC	1.26^{a}	-2.54^{a}	0.54^{a}	-1.38^{a}	0.51^{a}	-1.63^{a}	0.23^{b}	-1.16^{a}
	(0.07)	(0.08)	(0.10)	(0.06)	(0.10)	(0.11)	(0.12)	(0.09)
Alberta	1.36^{a}	-0.67^{a}	0.45^{a}	-1.79^{a}	0.57^{a}	-0.67^{a}	-0.02	-1.63^{a}
	(0.10)	(0.09)	(0.13)	(0.09)	(0.15)	(0.15)	(0.14)	(0.12)
Saskatchewan	1.07^{a}	-2.42^{a}	0.25^{a}	-2.48^{a}	0.21^{b}	-1.43^{a}	0.01	-1.98^{a}
	(0.07)	(0.08)	(0.08)	(0.09)	(0.10)	(0.11)	(0.10)	(0.14)
Manitoba	1.07^{a}	-1.42^{a}	0.56^{a}	-2.33^{a}	0.22^{b}	-1.07^{a}	0.15	-1.85^{a}
	(0.08)	(0.07)	(0.09)	(0.08)	(0.10)	(0.09)	(0.12)	(0.14)
Ontario	1.63^{a}	-1.21^{a}	0.27^{a}	-1.51^{a}	0.67^{a}	-0.97^{a}	0.29^{a}	-0.88^{a}
	(0.09)	(0.07)	(0.08)	(0.07)	(0.12)	(0.06)	(0.08)	(0.09)
Quebec	1.15^{a}	-1.46^{a}	0.50^{a}	-1.09^{a}	0.45^{a}	-1.12^{a}	0.23^{a}	-0.78^{a}
	(0.07)	(0.08)	(0.09)	(0.07)	(0.09)	(0.08)	(0.08)	(0.09)
New Brunswick	0.66^{a}	-1.59^{a}	0.31^{a}	-1.76^{a}	0.30^{a}	-1.09^{a}	0.08	-1.40^{a}
	(0.06)	(0.07)	(0.09)	(0.11)	(0.08)	(0.07)	(0.10)	(0.13)

Standard errors in parentheses.
^{c} $p < 0.1$, ^{b} $p < 0.05$, ^{a} $p < 0.01$

the behavior of travelers across provinces. Among Canadian residents who make day trips, travelers from BC and Saskatchewan generally have the largest elasticities with respect to the Real Exchange Rate. Among US residents making day trips, travelers entering Ontario — who are likely to be residents of either Michigan or New York State — have the largest elasticities.

3 Model of the crossing decision

The previous section has uncovered several stylized facts of cross-border travel that need to be incorporated in a model. First, while there is always two-way movement across the border, there are large within- and between-year fluctuations. Second, there is a robust relationship between exchange rates and travel: the stronger the currency in the country of residence, the more trips. Third, elasticities are asymmetric: In absolute value Canadian residents have higher percentage responses to changes in the exchange rate. Fourth, exchange rate elasticities are higher (in absolute value) when the home currency is stronger. Fifth, within–country elasticities differ substantially across provinces.

The model specifies the relationship between exchange rates and border crossings. We formalize a mechanism for a non-linear relationship between the crossings and the exchange rate. That is, we show why exchange rate responsiveness itself depends on the exchange rate. The model also rationalizes the large elasticities for residents of Canada and potentially explains differences in exchange rate elasticities across locations within the same country. The model is inspired by Dornbusch, Fischer, and Samuelson (1977), hereafter DFS, so we retain that paper's notation where convenient.

3.1 The "supply side"

There is a continuum of goods indexed z on the interval from zero to one. In contrast to Armington style models in which goods in one country are differentiated from those in another, this model follows the Eaton and Kortum (2002) approach in which all goods differ only in their prices. Unit labour requirements are $a(z)$ at home and $a^*(z)$ abroad. Defining $A(z) \equiv a^*(z)/a(z)$, goods are ordered such that $A'(z) < 0$. Following DFS we assume competitive product markets. Consequently, goods at home have prices given by $P(z) = Wa(z)$ where W are wages at home. Correspondingly, prices in the foreign country are $P^*(z) = W^* a^*(z)$. Both prices and wages are expressed in terms of the respective local currency units. The relative price of *foreign goods* in local units is defined as

$$p(z) \equiv P^*(z)/P(z) = A(z)(W^*/W).$$

Since $A'(z) < 0$ and we take relative wages as given, we have $p' < 0$.

Let the domestic currency price of foreign currency be E. The relative price of foreign goods is therefore $Ep(z) = A(z)/\omega$ where $\omega \equiv W/EW^*$ is the relative wage of domestic workers expressed in a common currency. Again following DFS, we define $\tilde{z}$ as the borderline good for which prices in a common currency are equal, that is

$$P(\tilde{z}) = EP^*(\tilde{z}) \quad \Leftrightarrow \quad Ep(\tilde{z}) = 1. \tag{3}$$

For $0 \leq z \leq \tilde{z}$, goods are cheaper at home and the remaining goods $\tilde{z} \leq z \leq 1$ are cheaper in the foreign country. Given wages and the exchange rate, we can solve for the borderline good as

$$\tilde{z} = A^{-1}(\omega) = p^{-1}(1/E). \tag{4}$$

Since $A' < 0$ we know that $\tilde{z}$ is decreasing in ω, the relative wage of domestic workers. Using relative price notation, the expression following the second equality shows that $\tilde{z}$ is increasing in E. Thus a nominal appreciation of the foreign currency, holding prices constant in local currency units, contracts the range of goods that are cheaper in the foreign country.

We illustrate the model in Figure 3 using data from Porter (2009). The author reports prices for 19 well-defined goods available on both sides of the border. Calculating $p(z)$ as the ratio of the US price (in USD) to the Canadian price (in CAD), we sort z in decreasing order and plot. With the lone exception of ice cream obtained at a Cold Stone Creamery and a 32G iTouch, Canadian prices were higher (expressed in local units). At the time the article was written the exchange rate was 1.09 CAD/USD. With such a strong Canadian dollar, it is not surprising that 15 out of 19 goods were less expensive in the US after converting prices to a common currency. The figure shows that dramatic changes would arise if the USD were to revert to the 1972–2009 mean and appreciate by about 15% to 1.25 CAD/USD. This would lead to a rise in $\tilde{z}$, i.e. a contraction in the set of goods that are cheaper in the

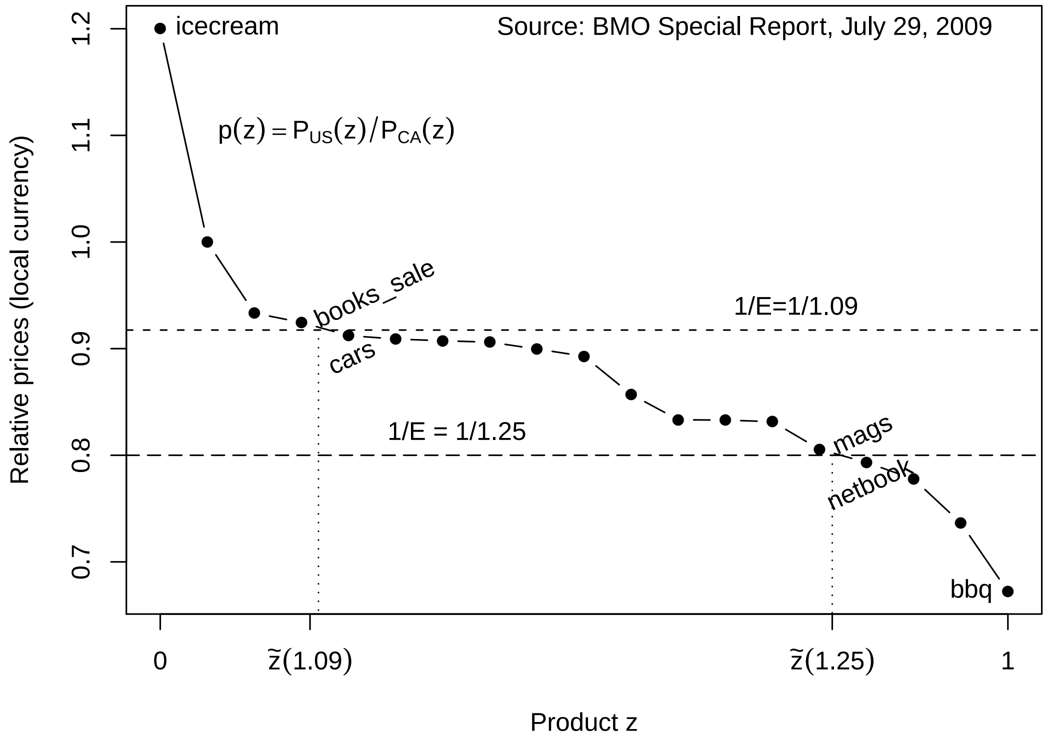


Figure 3: Exchange rates and relative prices: 19 products

US.

3.2 Consumer problem

Consumers have Cobb-Douglas utility and spend $b(z)$ on each good z.[11] Utility (subject to a monotone transformation) can be expressed as

$$\ln U = \int_0^1 b(z) \ln C(z) dz,$$

where $C(z)$ denotes consumption.

Up to this point there has not been any deviation from DFS. We now make a major departure. The purchase of products made in the foreign country requires the consumer to engage in cross-border shopping. This assumption works well for most services (e.g. hotels, restaurant meals, concerts). Goods purchased at stores have a

[11]Under Cobb-Douglas utility, the $b(z)$ are utility function parameters.

substantial non-traded retail service component embedded in their prices. The model does not apply to goods bought by phone or Internet.[12] The cost of the cross border trip enters as a reduction in income. Thus, the income of workers who stay at home is W whereas those who travel have income reduced to W/τ where $\tau > 1$ is the travel cost. This is like the Samuelson "iceberg" form for transport costs in the sense that a share $1/\tau$ of income "melts away" in the trip across the border. The idea is that travel to and across the border consumes time and represents a loss in income. The stayer buys all goods at home and has indirect utility v_S where the mnemonic for S is "stay":

$$v_S = \ln W - \int_0^1 b(z) \ln P(z) dz.$$

Crossers buy goods $\tilde{z} \leq z \leq 1$ in the foreign country. Their indirect utility is therefore given by

$$v_X = \ln W/\tau - \int_0^{\tilde{z}} b(z) \ln P(z) dz - \int_{\tilde{z}}^1 b(z) \ln EP^*(z) dz,$$

where the X subscript is a mnemonic for "cross." Define B as the integral of savings made by buying goods in the foreign country instead of domestically:

$$B \equiv \int_{\tilde{z}}^1 b(z)[\ln P(z) - \ln EP^*(z)] dz. \tag{5}$$

The increase in utility obtained from crossing is therefore

$$v_X - v_S = B - \ln \tau.$$

The mnemonic for B is that it represents the (gross) benefits of crossing as a function of the exchange rate. For any interior value of $\tilde{z}$, B is positive since $P(z) > EP^*(z)$ for all $z > \tilde{z}$.

To specify benefits of crossing in terms of the real exchange rate, we now make use of the supply-side assumptions from subsection 3.1. Replacing $\ln P(z)$ with $\ln a(z) + \ln W$ and $\ln P^*(z)$ with $\ln a^*(z) + \ln W^*$, recalling that $A(z) \equiv a^*(z)/a(z)$,

[12] The model could be extended to allow for goods that are traded. For the set of imported goods, consumers at home pay $P(z) = EP^*(z)(1+s)$ where s is the advalorem shipping cost for goods from the foreign country. In the foreign country these goods cost $EP^*(z)$. Thus there is a potential saving to be made by crossing the border to purchase these goods. Its contribution to the benefit of crossing, B, would be $\ln(1+s)b(z)$. This term does not depend on the exchange rate and therefore would not affect the predictions of the model.

and rearranging we obtain

$$B = -\int_{\tilde{z}}^{1} b(z)\ln A(z)dz - \ln\frac{EW^*}{W}\int_{\tilde{z}}^{1} b(z)dz. \tag{6}$$

While EW^*/W would be one way to define the real exchange rate, it is more customary to do so in terms of price indexes. The model implies a simple relationship between relative price indexes and relative wages. With Cobb-Douglas preferences the natural definition of the price indexes are $\mathbf{P} = \exp(\int_0^1 b(z)\ln P(z)dz)$ and $\mathbf{P}^* = \exp(\int_0^1 b^*(z)\ln P^*(z)dz)$. Substituting in the expressions for prices we obtain

$$\ln \mathbf{P}^*/\mathbf{P} = \ln W^*/W + \ln\kappa,$$

where $\kappa \equiv \exp\left(\int_0^1 [b^*(z)\ln a^*(z) - b(z)\ln a(z)]dz\right)$ is a constant if budget shares and relative productivities across goods do not change over time. We can now express the real exchange rate as a function of the relative wage:

$$e \equiv \frac{E\mathbf{P}^*}{\mathbf{P}} = \kappa\frac{EW^*}{W}. \tag{7}$$

It is also useful to follow DFS in defining $\vartheta(\tilde{z}) = \int_0^{\tilde{z}} b(z)dz$ as the share of expenditures on all goods that fall on goods for which the home country is the low-price supplier. Making these substitutions in equation (6), we can express the benefits of crossing as a function of the log real exchange rate:

$$B(\ln e) = -\int_{\tilde{z}}^{1} b(z)\ln A(z)dz - (1-\vartheta(\tilde{z}))(\ln e - \ln\kappa). \tag{8}$$

Taking the derivative of (8) with respect to $\ln e$ we obtain

$$B' \equiv \frac{\partial B}{\partial \ln e} = -(1-\vartheta(\tilde{z})) < 0. \tag{9}$$

The benefits from crossing are a negative function of the real exchange rate.[13] The derivative is just minus the share of consumer income spent on the goods that are

[13] Consistent with our model, we assume that changes in $\ln e$ are generated by either the nominal exchange rate E or the ratio of relative wages (through its effect on relative prices $\mathbf{P}^*/\mathbf{P}$). We do not consider the changes in $\ln e$ generated by adjustments in $A(z)$ as it would require a reordering of goods in the $[0,1]$ interval.

cheaper in the foreign country. Thus the benefits of crossing the border respond more to a given percentage change in the exchange rate when this budget share is high. The slope of the benefits of crossing B is partly determined by the effect of $\ln e$ on the intensive margin (goods that would be bought in the foreign country anyway but are now cheaper). But the benefits are augmented by the fact that the basket of goods that are cheaper abroad also changes with the exchange rate. This leads the benefit function to be convex in the real exchange rate. Using the fact that $\vartheta'(\tilde{z}) = b(\tilde{z})$, we obtain a second derivative of

$$B'' \equiv \frac{\partial^2 B}{\partial \ln e^2} = b\left(\tilde{z}\right) \frac{\partial \tilde{z}}{\partial \ln e} = -b\left(\tilde{z}\right) \frac{A\left(\tilde{z}\right)}{A'\left(\tilde{z}\right)} > 0. \tag{10}$$

These results show that convexity of the $B(\ln e)$ function arises under general functional form assumptions for preferences, $b(z)$, and technology $A(z)$. However, it is also useful to consider a special case where the integrals have closed form solutions. Suppose equal expenditure on all goods, i.e. $b(z) = 1 \forall z$ and $\ln A(z) = \alpha_0 - \alpha_1 z$ with $\alpha_0, \alpha_1 > 0$. Parameter α_0 shifts the relative costs of the foreign country in all goods whereas α_1 measures the extent of cross-good heterogeneity in relative costs. Imposing these functional forms simplifies $\ln \kappa$ to $\alpha_0 - \alpha_1/2$ and $\vartheta(\tilde{z})$ to $\tilde{z}$. Solving for the critical good defined in equation 4) yields

$$\tilde{z} = \frac{1}{2} + \frac{1}{\alpha_1} \ln e. \tag{11}$$

To obtain an interior solution the real exchange rate must lie in the range $\exp(-\alpha_1/2) < e < \exp(\alpha_1/2)$.

Plugging in these equations and integrating the benefit function shown in equation (8) yields a quadratic formula in which the coefficients are simple functions of α_1, the key cost heterogeneity parameter,

$$B(\ln e) = \beta_0 + \beta_1 \ln e + \beta_2 [\ln e]^2, \tag{12}$$

where $\beta_0 \equiv \alpha_1/8$, $\beta_1 \equiv -1/2 < 0$, and $\beta_2 \equiv \frac{1}{2\alpha_1} > 0$. Note that α_0 does not appear, because it only matters through changes in the real exchange rate. Increasing α_0 is equivalent to increasing $\ln W^*/W$. On the other hand, α_1, the measure of dispersion in productivities, determines both the size of savings for a given basket of goods to be bought in the foreign country and the extent of that shopping basket.

We use the quadratic form shown in equation (12) in our empirical specification. It can be thought of either as a second-order approximation of a general B or as the exact solution under the uniform budgeting and log-linear $A(z)$.

Foreign crossings into the home country depend on a similar benefit function:

$$B^*(\ln e) \equiv \int_0^{\tilde{z}} b^*(z)[\ln EP^*(z) - \ln P(z)]dz \tag{13}$$

$$= \int_0^{\tilde{z}} b^*(z) \ln A(z)\, dz + \vartheta^*(\tilde{z})(\ln e - \ln \kappa), \tag{14}$$

where $\vartheta^*(\tilde{z}) = \int_0^{\tilde{z}} b^*(z)dz$ is the share of expenditures that foreign consumers allocate to goods that are less expensive in the home country. The derivative of B^* with respect to the log real exchange rate is given by

$$\frac{\partial B^*}{\partial \ln e} = \vartheta^*(\tilde{z}). \tag{15}$$

The derivatives of the benefits of crossing can only be equal in absolute value if $1 - \vartheta(\tilde{z}) = \vartheta^*(\tilde{z})$. [14] There is only a single value of the real exchange rate that meets this condition.

3.3 A continuum of individuals

The DFS model has representative agents and needs to be amended to take into account the fact that most people do not cross in a given month. There are two important forms of heterogeneity. The first, which we measure, is proximity to the border. The second, which is unobserved, explains why even in the same community some people cross and some do not. We formalize these ideas by generalizing the setup to allow for a continuum of individuals i in each community c. The travel cost comprises a term reflecting the geographic location of the community and an unobserved variable reflecting individual-specific shocks:

$$\ln \tau_c(i) = g(d_c(i)) + \zeta(i),$$

[14] With identical preferences, $b(z) = b^*(z)$, this requires $\vartheta^*(\tilde{z}) = 1/2$.

where $d_c(i)$ is the geographic distance of individual i from the border. The idiosyncratic cost of individual i is distributed with a CDF denoted $F(\zeta)$.[15] The function $g()$ maps geographic distance into indirect utility costs. We denote the probability of crossing for residents of community c as x_c. With a continuum of individuals, x_c also measures the fraction who cross:

$$x_c = \mathbb{P}(v_X > v_S) = \mathbb{P}(B(\ln e) > \ln \tau_{ic}) = \mathbb{P}(\zeta(i) < B(\ln e) - g(d_c)) = F(B(\ln e) - g(d_c)). \tag{16}$$

That is, residents of distant communities require a lower idiosyncratic cost to benefit from crossing the border.

The elasticity of crossing probability with respect to the real exchange rate is

$$\frac{\partial \ln x_c}{\partial \ln e} = \frac{F'}{F} B' < 0. \tag{17}$$

Since crossing costs are increasing in distance, $g' > 0$, the elasticity of crossing with respect to distance is negative:

$$\frac{\partial \ln x_c}{\partial \ln d_c} = -\frac{F'}{F} g' d_c < 0. \tag{18}$$

While these elasticities can be signed for general distributions on individual heterogeneity, the second derivative with respect to $\ln e$ and the cross-partial of how distance affects the exchange rate elasticity cannot be signed without determining the shape of F'/F. The marginal crosser in a given community (i.e., holding distance to the border fixed) is the resident that is indifferent between crossing and shopping at home. A change in the real exchange rate level determines (through B) a shift in the location of the marginal crosser in $F's$ domain and translates into a new crossing probability for the community. The rate of such change depends on the initial location of the marginal crosser as well as the curvature of F.

How does exchange rate responsiveness change in response to foreign appreciation? Differentiating equation (17) we obtain

$$\frac{\partial^2 \ln x_c}{\partial \ln e^2} = \frac{[FF'' - (F')^2]}{F^2} (B')^2 + \frac{F'}{F} B''. \tag{19}$$

[15]We introduce $\zeta(i)$ as an idiosyncratic cost although it is intended to capture heterogeneity in the crossing costs and benefits. That is $\zeta(i) \in (-\infty, \infty)$.

Examination of this expression leads to two important results. First, once heterogeneity is added into the model, the positive second derivative of the individual benefit function (B'') shown in (10) will not translate into a positive second derivative for aggregate log crossings if the term in brackets is sufficiently negative. Second, we see that in models with a constant elasticity at the individual level ($B'' = 0$), convexity of log crossings requires the term in square brackets to be positive. For commonly used distributions of individual heterogeneity, the factor in brackets has a negative sign.[16] In our empirical work we find that the convexity we estimate in the $B()$ function is retained in the relationship between log of predicted crossings and the log real exchange rate.

3.4 Geography

Consider the consequences of aggregation of multiple communities c, of size N_c into a single "region" R of size $N_R = \sum_{c \in R} N_c$ which could be a province, state or country.

$$x_R = \sum_{c \in R} \frac{N_c}{N_R} x_c$$

$$\frac{\partial \ln x_R}{\partial \ln e} = \sum_{c \in R} \frac{N_c}{N_R} \frac{x_c}{x_R} \frac{\partial \ln x_c}{\partial \ln e}. \tag{20}$$

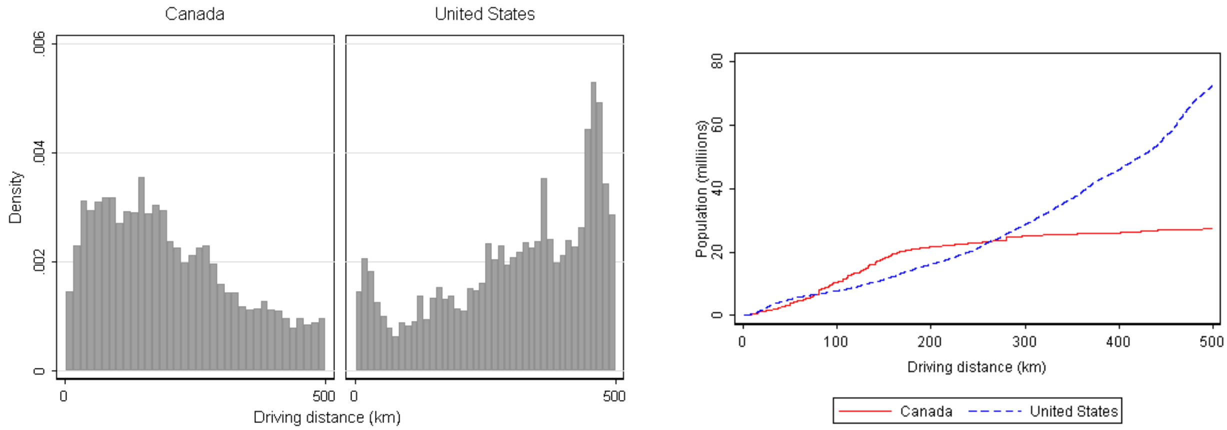
From (17) we know that (holding the marginal benefit of crossing constant) elasticities are highest when crossing rates are near zero. We can therefore sign the cross partial effect of distance and the log exchange rate:

$$\frac{\partial^2 \ln x_c}{\partial \ln e \partial d_c} = -\frac{[FF'' - (F')^2]}{F^2} B' g'. \tag{21}$$

Since commonly used $F()$ distributions imply that the term in brackets is negative, equation (21) leads to a somewhat counterintuitive prediction: As distance to the border increases, exchange rate responsiveness—as measured by the absolute value of the elasticity—becomes stronger.[17] Consider both equation (20) and (21 we can infer

[16] F'/F is globally decreasing for uniform, normal, logit, gumbel. Even the highly convex power distributions, $F(\zeta) = (\zeta/\bar{\zeta})^\lambda$ for $0 < \zeta < \bar{\zeta}$ has F'/F decreasing. Although certain parameterizations of beta distributions can have upward sloping regions in the right tail, our numerical analysis suggests F'/F is decreasing over most of the support.

[17] The response in *levels* of crossings shrinks with distance from the border.



(a) (b)

Figure 4: Population and Distance to the Border

that the regions where population is clustered farthest from the border should have the largest (in absolute value) estimated crossing elasticities, in large part because they have the lowest crossing rates.

The model presented in this section strongly suggests that geography has an important role to explain the differences in crossing elasticities from Tables 3 and 5. Population and geography together determine the size of the potential number of crossers as well as its likelihood of crossing. Due to differences in crossing costs (distance to the border) and population density, it is very unlikely for two regions to have the same crossing elasticity.

To understand the role of geography it is instructive to ignore the idiosyncratic shocks for a moment. Assume that all communities in a given region receive the same idiosyncratic shock. In such case, there is no heterogeneity within a community and either all residents cross the border or no one does. The marginal community is located at $\bar{d}$ km from the border and is determined by $B\left(e\right) = g\left(\bar{d}\right)$. A stronger domestic exchange rate increases the critical distance and more communities engage in cross-border shopping. In terms of elasticities, the result of a change in $\ln e$ depends on the population size of the new communities that now cross the border with respect to the total number of residents that were engaged in cross-border shopping.

Consider Canada and United States as the two regions. Figure 4 show the differences in terms of population density and distance to the border.[18] Panel (a) shows

[18]The Figures were constructed by calculating the driving distance from each census tract to the

that a higher proportion of Canadians live near the border relative to the United States. Panel (b) shows the accumulated population as we move farther from the border. The logic described above would suggest that the crossing elasticity with respect to the real exchange rate is similar for the two regions if the critical distance $\bar{d}$ falls within 100km from the border (both countries have roughly the same total population at any given point in the range). The elasticity for Canadians would be higher than for Americans in the 100–170km range since, as the critical distance increases, Canadian communities that start crossing are larger than Americans. The opposite is true for distances larger than 200km, as the number of American crossers expands significantly while it does not change for Canada.

Figure B.1 in the Appendix shows significant variation across regions in total population and distances to the border.[19] In some regions, more Canadians than Americans live near the border (Alberta, Manitoba, Saskatchewan), the reverse is true in Ontario, and in other cases (Quebec and British Columbia) it depends on the distance to the border. Additionally, the pool of potential crossers is different across regions of the same country. Differences in size and population distribution may contribute to the differences in the estimated elasticities found in Table 5. In the next section, we exploit the variation in distances traveled by actual crossers to estimate the elasticities with respect to the exchange rate and the effect of distance to the border.

4 Estimation of the model

In this section we take the model of the previous section to the data. We combine our dataset on cross-border travel with detailed information on the location of Canadian travelers and their distance from the US border. We use our estimates to calculate the relative importance of exchange rates and distance effects in determining the decision to cross the border. Our results provide an estimate of the extent of market segmentation that is created by the presence of the US–Canada border.

closest land border.

[19] A census tract was assigned to a Canadian province based on the location of the nearest crossing port.

4.1 Regression Specification

In order to estimate the crossing fraction equation shown in equation (16), we need to parameterize the crossing benefit and cost functions (B and g) as well as specify the distribution of individual heterogeneity (F). We make use of the quadratic form for $B(e)$ shown in equation (12). It can be thought of either as the exact solution under uniform budgeting and log-linear $A(z)$ or can be considered a second order approximation for $B(e)$. This is the simplest form that allows us to test for the convexity which is a distinguishing feature of our model of cross-border shopping.

Next we parameterize the border crossing costs that apply to all individuals as

$$g(d_c) = \gamma_0 + \gamma_1 \ln(1 + d_c) \tag{22}$$

The γ_0 parameter represents the $\ln \tau$ for a hypothetical individual i with no idiosyncratic crossing costs ($\zeta(i) = 0$) located at the border ($d_c(i) = 0$). Thus γ_0 captures the pure cost of crossing the border. If d_c is measured in terms of distance, γ_0 includes the disutility of waiting in border line-ups. We also consider a specification where d_c measures time from the census division to the border crossing. In that specification we add an estimated wait time to the drive time. In both specifications, γ_0 captures border formality compliance costs.[20]

Substituting these B and g functions into equation 16, we can express the crossing fraction as

$$x_c = F[\beta_0 - \gamma_0 + \beta_1 \ln e + \beta_2 (\ln e)^2 - \gamma_1 \ln(1 + d_c)]. \tag{23}$$

Next we need to impose a particular functional form for $F(\zeta)$. Idiosyncratic crossing costs $\zeta(i)$ are likely to depend on the sum of a large number of at least partially independent factors. The central limit theorem would therefore lead ζ to be distributed normally. Assuming ζ has expectation μ and variance σ^2, $F(\zeta) = \Phi([\zeta - \mu]/\sigma)$, where $\Phi()$ denotes the standard normal CDF. Substituting these parameterizations into equation (23) and adding time subscripts we obtain

$$x_{ct} = \Phi[\theta_0 + \theta_1 \ln e_t + \theta_2 [\ln e_t]^2 + \theta_3 \ln(1 + d_c)], \tag{24}$$

where Table 6 shows the mapping between the θ and the structural parameters as

[20] Since these costs are thought to have risen following September 11th, 2001, we include a post911 dummy in most specifications.

well as the expected signs for each coefficient.

Table 6: Interpretation of coefficients

Parameter	Covariate	Structure	Sign
θ_0	constant,	$(\beta_0 - \gamma_0 - \mu)/\sigma = (\alpha_1/8 - \gamma_0 - \mu)/\sigma$	+ or −
θ_1	$\ln e_t$ (RER)	$\beta_1/\sigma = -1/(2\sigma)$	−
θ_2	$(\ln e_t)^2$	$\beta_2/\sigma = 1/(2\alpha_1\sigma)$	+
θ_3	$\ln(1 + d_c)$	$-\gamma_1/\sigma$	−

Equation 24 is not yet suitable for estimation purposes because it does not allow for deviations between observed crossing fractions and those predicted by the model. Such deviations would arise from at least three sources. First, the continuum assumption is only an approximation, so the actual crossing share would only be equal to the crossing probability in expectation. Second, our data is based on a survey given out to a subset of the actual population of crossers. We elaborate on this point in section 4.2 and Appendix A. Third, the parameterizations of $B(e)$ and $g(d)$ are approximations and therefore introduce approximation error. We restate equation (24) in the form of a conditional expectation:

$$\mathbb{E}[x_{ct} \mid e_t, d_c] = \Phi[\theta_0 + \theta_1 \ln e_t + \theta_2 [\ln e_t]^2 + \theta_3 \ln(1 + d_c)]. \tag{25}$$

Papke and Wooldridge (1996) point out that quasi-likelihood estimation will yield consistent estimates of the model parameters so as long as the conditional expectation shown in (25) is correctly specified. This method is easy to estimate using Stata's "glm" command with "family(binomial)" and "link(probit)". Following the suggestion of Papke and Wooldridge (2008) we estimate robust standard errors clustering at the census division (c) level to allow for arbitrary serial correlation within panels.

We also consider two more commonly used methods for estimating models where the dependent variable is a fraction. The first is to (implicitly) assume F is uniform which makes the crossing share, x_{ct}, linear in the parameters and therefore estimable using OLS. The second method assumes F is logistic with location and scale parameters μ and σ and applies the log odds transformation ($\ln[x/(1 - x)]$) to obtain an equation that is linear in the parameters. The log-odds method is often preferred because it forces predictions for x_{ct} to lie between zero and one. Replacing F with the logistic function, transforming the dependent variable, and grafting on an error term denoted ε, yields

$$\ln\left(\frac{x_{ct}}{1-x_{ct}}\right) = \theta_0 + \theta_1 \ln e_t + \theta_2[\ln e_t]^2 + \theta_3 \ln(1+d_c) + \varepsilon_{ct}. \tag{26}$$

While this method has the virtue of being estimable using OLS, Papke and Wooldridge (1996) identify two critical defects. First, the dependent variable is undefined for $x_{ct} = 0$ and $x_{ct} = 1$. As we discuss in section (4.2), over half the ct combinations in our data have $x_{ct} = 0$ and these tend to occur in divisions that are far from the border, implying that the log-odds procedure is likely to induce selection bias. A second problem with the log odds specification is that it yields the conditional expectation of the log odds ratio, a variable that is not of direct interest. As Papke and Wooldridge (1996) show, one cannot simply plug the estimated θ estimated using specification (26) into the logistic function to recover the conditional expectation of x_{ct}.[21] Based on these arguments, we only report estimates from the log-odds method as a robustness check.

4.2 Data

The dependent variable is the crossing fraction, x_{ct}, which is measured as the number of car crossings, n_{ct}, from Census division c in month t, divided by estimate of the number of potential crossings, denoted N_{ct}. Potential trips, N_{ct}, are defined as the population of the census division (Pop_{ct}), multiplied by the number of cars per capita (CPC) in the province multiplied by the number of days in the month. Thus, the crossing fraction is given by

$$x_{ct} = \frac{n_{ct}}{N_{ct}} \approx \frac{\hat{n}_{ct}}{\text{Pop}_{ct} \times \text{CPC}_c \times 30}. \tag{27}$$

We estimate $\hat{n}_{ct}$ using data from the International Travel Survey (ITS), which is filled out by travelers returning to Canada from trips abroad. The data consist of questionnaires that collect information on the nature and purpose of the trip, the dates on which travelers exited and entered Canada, and information on the Census Division (CD) in which the travelers reside and ports used to cross to the US. We

[21] Intuitively, this is because the log of the expectation is not equal to the expectation of the log.

keep data on Canadian residents returning from the United States by car during the period 1990–2009.[22] The Appendix details how we construct $\hat{n}_{ct}$ by weighting the ITS responses using the aggregate crossing data so as to make the sample representative at the monthly level as well as representative at each port of entry. Census division populations, Pop_{ct}, are available annually from Cansim Table 051-0034, provided by Statistics Canada. Car registration data come from Statistics Canada publication 53-219-XIB ("Road Motor Vehicle Registrations 1998").

We measure d_c, the distance from census division c to the border, in three ways. Our preferred form is the weighted average of the driving distances to frequently used ports (those that account for more than 5 percent of the division's total crossings).[23] In robustness checks we also measure d_c as driving distance to the nearest port and as average driving time. We obtained the driving distances and times from the CD's centroid to its associated crossing ports using centroid information from the Standard Geographical Classification of 2001 and Google's application for determining driving directions.

We present summary statistics of the data in Table 7. An observation is a combination of a census division in a given month. The top panel presents statistics for all possible combinations of Census-Divisions, months, and length of trip. Many of these combinations do not have corresponding cross-border trips. The lower panel presents data for the sample with positive cross-border trips in that month. Conditioning on positive trips, Census Divisions tend to be closer to the border, and more populated.

4.3 Baseline Estimation

In this section we estimate the structural model implied by equation (25). The results using the fractional probit method of estimation are presented in Table 8. The first three columns use daytrips to construct the dependent variable, while the next three use overnight trips. Coefficients for month fixed-effects and other control variables are

[22] The survey began in 1990. We do not use information on US residents since the only information on their place of residence within the US is the state in which they live. This level of aggregation is too coarse to provide meaningful information on their distance to the border.

[23] Note that residents of a CD use many different ports to enter the US and return to Canada. The ITS data records crossings from 250 CDs to 102 ports located in the seven provinces that share a land border with the US. The weights are the fraction of crossings from the CD over the 1990–2009 period.

Table 7: Summary Statistics: Census Divisions-months

Sample	Variable	Mean	SD	Median	Min	Max
All CD-months	Driving Distance (km)	381	417.5	226.2	23.8	3421.2
N=118500	Driving time (hrs)	5.2	5.7	3.1	0.5	48.6
	Population (1000)	115.4	271.5	40.7	1.2	2662.5
	Cross-border trips (cars):					
	Same-day	4116	20403	0	0	456542
	Overnight	1323	4161	81	0	90662
CD-months	Driving Distance (km)	232.5	250.8	157.4	23.8	3421.2
with trips>0	Driving time (hrs)	3.2	3.4	2.2	0.5	48.6
N=54722	Population (1000)	196.3	377	76.4	1.2	2662.5
	Cross-border trips (cars):					
	Same-day	10695	31802	1493	2	456542
	Overnight	2456	5418	637	2	90662

not reported.[24] Columns 1 and 4 use the specification of equation 25, adding month dummies to allow μ to change reflecting the seasonal pattern shown in Figure 1(b). Columns 2 and 5 add controls that are expected to shift the net benefits of crossing at any given exchange rate and distance, that is variables that affect $\gamma_0 + \mu$ in the model: an indicator for travel after September 2001, and province fixed-effects. Columns 3 and 6 add linear and quadratic time trends.

The results show that driving distance creates a strong disincentive to cross the border. This is especially the case for daytrips; distance is a weaker disincentive for those planning trips of a longer duration. The coefficient on the exchange rate variables indicate that a higher value of the real exchange rate (implying a weaker CAD) reduces the probability of cross-border trips. The coefficient on the second order term is positive for daytrips, implying that travelers' responsiveness to the real exchange rate decreases as its level rises. This is in accordance with the predictions of our model and is also consistent with the reduced form results of Table 4. We do not observe this result for overnight trips; in column 6 we actually get the opposite result for the second-order term.[25]

[24] R-squares are calculated as the square of the correlation between the dependent variable and its predicted value.

[25] Note, though, that we include linear and quadratic trends in columns 3 and 6 in order to test the robustness of our results to the specification in equation 1. However, the structural model does not imply that these terms should be included.

Table 8: Fractional Probit estimation of crossing fractions (x_{ct})

Length of stay:	Daytrip			Overnight		
θ_0: constant	0.30	0.29	0.34	-2.58^a	-2.50^a	-2.37^a
	(0.47)	(0.31)	(0.31)	(0.09)	(0.10)	(0.10)
θ_1: $\ln e$ [RER]	-0.52^a	-0.59^a	-0.47^a	-0.68^a	-0.73^a	-0.30^a
	(0.11)	(0.12)	(0.09)	(0.13)	(0.13)	(0.10)
θ_2: $(\ln e)^2$	0.80^b	0.70^b	0.54^c	0.13	-0.21	-0.57^b
	(0.33)	(0.33)	(0.30)	(0.36)	(0.29)	(0.29)
θ_3: $\ln(1+d_c)$ [distance]	-0.63^a	-0.58^a	-0.58^a	-0.15^a	-0.14^a	-0.14^a
	(0.09)	(0.05)	(0.05)	(0.02)	(0.02)	(0.02)
New Brunswick		0.40^b	0.40^b		-0.06	-0.06
		(0.17)	(0.17)		(0.06)	(0.06)
Quebec		-0.57^a	-0.57^a		-0.19^a	-0.19^a
		(0.14)	(0.14)		(0.07)	(0.07)
Ontario		-0.12	-0.12		0.12^a	0.12^a
		(0.19)	(0.19)		(0.04)	(0.04)
Manitoba		-0.34^b	-0.34^b		-0.03	-0.03
		(0.15)	(0.15)		(0.04)	(0.04)
Saskatchewan		-0.35^b	-0.35^b		-0.19^a	-0.19^a
		(0.18)	(0.18)		(0.04)	(0.04)
Alberta		-0.48^a	-0.48^a		-0.15^a	-0.15^a
		(0.13)	(0.13)		(0.05)	(0.05)
Post-911		-0.16^a	-0.05^a		-0.16^a	-0.00
		(0.03)	(0.02)		(0.03)	(0.01)
Trend variables	No	No	Yes	No	No	Yes
N	59250	59250	59250	59250	59250	59250
$R^2 = [\mathrm{cor}(x_{ct}, \hat{x}_{ct})]^2$	0.18	0.49	0.49	0.04	0.07	0.07

Standard errors clustered by census-division. BC is the omitted province.
Regressions include month fixed-effects. ^{c} $p<0.1$, ^{b} $p<0.05$, ^{a} $p<0.01$

This result suggests that residents making daytrips are more likely to expand the bundle of goods that they purchase in the US when the exchange rate becomes more favorable, as our model predicts. It also indicates that overnight travelers do not exhibit the same behavior. This may be because the overnight travelers are less likely to cross the border in order to shop for goods to bring home, since they purchase a standard bundle of goods in the US: hotel stays, vacations, restaurant meals etc.

4.4 Robustness to specification changes

In this section we examine the robustness of our results to different specifications and variable definitions. The results are in Table 9. We use the set of controls corresponding to columns 2 and 5 of Table 8.

The first two columns present results using the log-odds model depicted in equation (26). The remaining columns use the fractional probit model, but use different measures of the costs of travel. In columns 3 and 4 we use the average driving time to the border from each Census Division, instead of the driving distance, since the former may be of greater relevance to travelers deciding whether to cross the border. We add 26 minutes to the driving time to account for border wait times.[26] In columns 5 and 6 we use the distance of each Census Division to the closest border crossing, rather than the average distance that was used in Table 8.

Our results are generally robust in all specifications. The positive second-order effect for exchange rates continues to hold for daytrips but not for overnight trips. The cost of traveling to the border, whether measured in terms of distance or time, has a negative and strongly significant effect on the probability of crossing the border; more so for daytrips than overnight ones.

4.5 Quantification of the exchange-rate responsiveness of travel

In this section we show how travel costs and the real exchange rate affect the likelihood of crossing the border. We quantify the effect of these variables through

[26]This is the median wait time for all travelers entering the United States during the hours of 7 AM and 12 PM at the two largest ports in British Columbia, using daily data from 2006 to 2010. Data on wait times were obtained from the Whatcom Council of Governments. While the wait times are only calculated for British Columbia, we use these estimates for all provinces under the assumption that border agencies allocate staff in order to generally equalize wait times across ports. The results are not sensitive to including wait times.

Table 9: Alternative specifications of regression and travel costs

Method:	Log Odds (OLS)		Fractional Probit			
Length of stay:	Daytrip	Overnight	Daytrip	Overnight	Daytrip	Overnight
θ_0: constant	-0.15	-5.81^a	-1.72^a	-2.98^a	0.01	-2.59^a
	(0.66)	(0.25)	(0.18)	(0.04)	(0.25)	(0.09)
θ_1: $\ln e$	-1.30^a	-1.76^a	-0.58^a	-0.73^a	-0.59^a	-0.72^a
	(0.23)	(0.13)	(0.12)	(0.13)	(0.12)	(0.13)
θ_2: $(\ln e)^2$	2.85^a	-0.36	0.69^b	-0.21	0.71^b	-0.21
	(0.71)	(0.46)	(0.33)	(0.29)	(0.33)	(0.29)
θ_3: ln travel cost	-1.19^a	-0.30^a	-0.83^a	-0.19^a	-0.56^a	-0.13^a
	(0.10)	(0.04)	(0.09)	(0.02)	(0.04)	(0.02)
N	22802	31920	59250	59250	59250	59250
$R^2 = [\text{cor}(x_{ct},\hat{x}_{ct})]^2$	0.37	0.28	0.45	0.08	0.47	0.07

Standard errors clustered by census-division. Regressions include controls from Table 8. Travel cost is measured as average driving distance in cols. 1 and 2; average driving time in cols. 3 and 4; driving distance to closest port in columns 5 and 6. ^{c} p<0.1, ^{b} p<0.05, ^{a} p<0.01

counterfactual experiments. First, however, we show the relationship between the crossing fraction and the real exchange rate for specific distances from the border in Figure 5. This figure is based on the specification in column 2 for Table 8 (adjusting using the coefficients on the Ontario, post 9/11, and July dummy variables). Each line corresponds to a census division in Southern Ontario. The curves show that the convexity in the B function carries over to the log crossing function. Thus, the elasticity of crossing is larger in absolute value when the home currency is strong. This effect is large: In the case of Essex county (where Windsor is located), the average elasticity over the full range of elasticities is -1.25. For $0.8 < e < 0.9$ the elasticity rises (in absolute value) to -1.82. Furthermore, the elasticity of crossing implied by the model is larger at greater distances from the border. We can see this in the figure as the curve for Toronto is steeper (which corresponds to greater elasticity since both axes are drawn on a log scale) than that for Essex.

To quantify the aggregate effect of policy changes, it is necessary to aggregate over the effects at each census division. In Table 10 we show the effect of two possible changes. The first two columns show the effect, in two different years, on the number of cross-border trips from a decrease in the Real Exchange Rate of 10%. This is

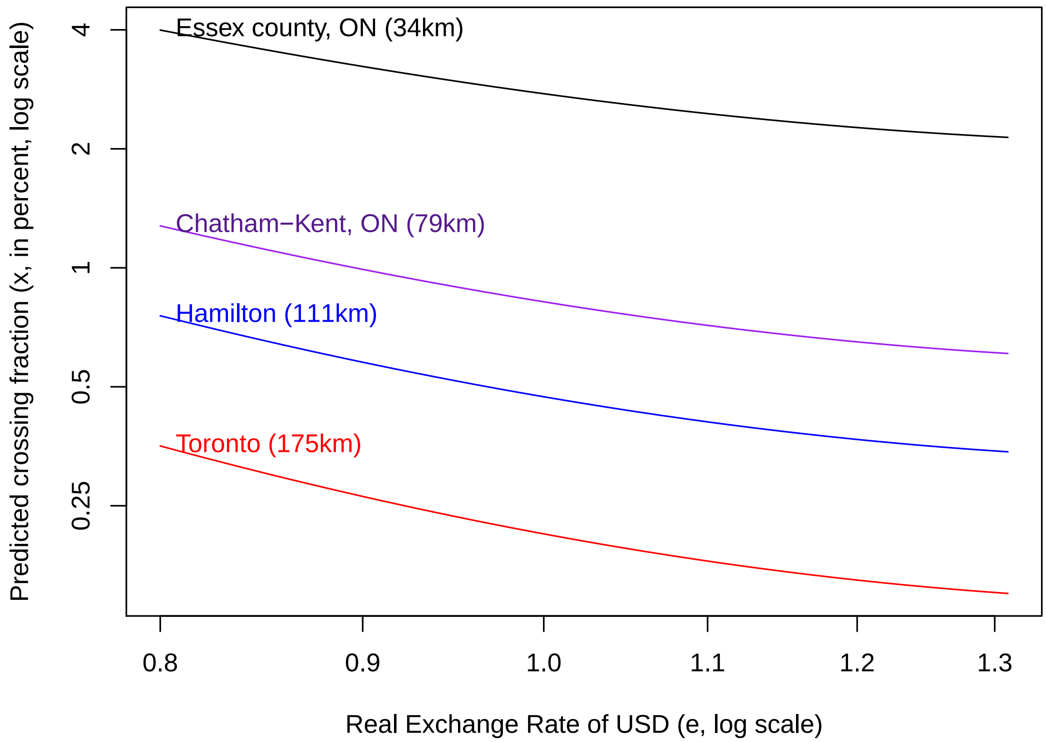


Figure 5: How the probability of crossing varies with the real exchange rate

equivalent to a strengthening of the Canadian Dollar. These estimates were derived by calculating, for each month in the corresponding year, the number of car trips from each Census Division had the RER in that month been 10% lower than its actual value. These counterfactual values were then aggregated across all census-divisions in the province and compared to the predicted values using the specification of Column 2 in Table 8.

As expected, cross-border trips by Canadian residents increase when the RER (defined in CAD per USD) decreases. The differences across time are far larger than differences across provinces for a given year. This is due to the structure of the model. At a given point in time, a change in the RER has a similar proportional effect on all census-divisions and therefore on all provinces. The differences over time come from the convex effect of the RER. The effects in 2009 are much larger than those in 2000; this is because the Canadian dollar was much stronger in 2009, and so the same proportional decrease in the RER (or equivalently, increase in the CAD) had a larger effect in that year compared to when the currency was weaker.

Table 10: Counterfactual scenarios: Percent change in crossings

Counterfactual:	RER +10%		Wait time +100%	
Year:	2000	2009	2000	2009
Canada	9.87	19.36	-41.95	-42.31
New Brunswick	8.28	16.14	-35.62	-35.95
Quebec	12.12	24.05	-48.25	-48.6
Ontario	10.33	20.4	-41.9	-42
Essex (34 km)	8.78	17.09	-50.4	-50.89
Chatham-Kent (79 km)	10.51	20.66	-40.52	-40.91
Hamilton (111 km)	11.23	22.16	-40.31	-40.69
Toronto (175 km)	12.23	24.26	-33.9	-34.21
Manitoba	11.86	23.43	-40.36	-40.84
Saskatchewan	13.34	26.73	-32.38	-32.49
Alberta	14.48	29.18	-29.82	-29.84
BC	9.21	17.94	-42.53	-43.06

In columns 3 and 4 we show the effect of increasing wait times at the border. We use the specification from Column 3 of Table 9. This specification had assumed a wait time of 26 minutes at the border. In our counterfactual experiment we double this to 52 minutes.[27] This naturally decreases the likelihood of cross-border trips by Canadians. However, now there are significant differences across provinces, and almost no variation over time. The smallest effect of the increased wait times is in the provinces of Alberta and Saskatchewan. These provinces do not have large cities close to the border. Since the wait time is incurred by all travelers, those driving longer distances have to pay a proportionately lower cost. By contrast a province like BC has a large population located very close to the border and therefore our model predicts a very large decrease in trips for a given increase in wait times. The predicted values do not change much over time since the effect of travel costs is independent of the value of the RER.

[27]Note that this increase in wait times needs to occur for exogenous reasons such as reduced staffing at the border or an increase in the time taken to process vehicles. Increases in wait times due to an increased number of cars arriving at the border will confound our predictions.

5 Conclusion

Using monthly data on the number of car trips across the border from 1972–2009, we establish that residents of both countries respond strongly to the real exchange rate, with estimated absolute elasticities for day trips of 0.95 for US residents and 1.72 for Canadians. Our model shows that these elasticities are not structural parameters in the sense that they could be used to evaluate counterfactuals. Rather, the exchange rate elasticity of travel is determined along several non-linear margins. First, it depends on the endogenously determined extent of the bundle of goods that are priced lower in the foreign country. Second, it depends on the geographic location of the prospective crosser. Finally, it depends on the distribution of idiosyncratic heterogeneity in crossing costs. Changing exchange rates move these margins, leading to new crossing elasticities. Our model delivers an estimable equation that permits estimation of the underlying structure needed to conduct counterfactuals. Estimation proceeds with standard software and no customized programming, thus removing a common impediment to structural estimation.

Structural estimates based on geographically detailed data for Canada from 1990 to 2009 support a key hypothesis generated by the model: the exchange rate elasticity grows with the strength of the home currency. This result is driven not by the unobserved heterogeneity, which tends to push in the opposite direction, but by the convexity in the (common) benefits of crossing function. A second key empirical result is that distance strongly inhibits crossing, with almost the same impact on net benefits of crossing as an equivalent percentage change in the real exchange rate. The estimated parameters and the detailed geographic data allow us to conduct counterfactual experiments with respect to key variables of the model. We show that an appreciation of the Canadian Dollar by 10% would have would increased cross-border travel by almost 20% in 2009, when the currency was relatively strong, but by only around 10% in 2000, when it was quite weak. An increase in border wait times would disproportionately reduce travel in provinces with large populations located near the border.

The strong effect of exchange rates on same-day travel poses interesting policy dilemmas. The convexity in the benefit function suggests a widening of the range of products purchased on trips. This seems more likely to apply to goods being brought back rather than services consumed during the day trip. The hypothesis that travellers are returning with purchased goods also finds support in news reports of Canadian license plates in the parking lots of shopping malls just south of the border. Since there is no exemption for goods purchased on daytrips, travelers who declared their purchases would be subject to sales taxes for all goods (12–13% for most crossers) and duties for non-North American goods. Alternatively, there may be substantial amounts of smuggling underlying the travel patterns we have estimated. A question for policy is whether the border agencies should devote more resources to inspections, especially during periods when the home exchange rate is strong. This would likely impose higher waiting costs on all travelers. Alternatively, exemptions could be increased so that small-scale undeclared purchases would no longer violate the law.

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Appendices

A Data construction

This Appendix describes how we select our sample from the ITS data and apply weights in order to make the sample representative. It also describes our method for calculating the distance to the border for residents of each Census Division in Canada.

Each observation in the ITS data is a questionnaire filled out by a Canadian resident returning to Canada from a trip to the US. This includes people who enter by car, bus, train, air, foot, boat etc. A maximum of one questionnaire is given to each traveling party. We keep only those observations where the traveling party exited and re-entered Canada by car. We also restricted the sample to people who reside in one of the 7 provinces that share a land border with the United States: New Brunswick, Quebec, Ontario, Manitoba, Saskatchewan, Alberta and British Columbia. This leaves us with 646,223 questionnaires over 20 years (1990–2009).

These questionnaires are handed out at the various border crossing ports, but not in a representative manner (either across ports, or across months of the year for a given port). Therefore, Statscan has assigned weights to each questionnaire in order to address non-representative sampling and non-response. Applying these weights makes the data representative at the annual level for each port-factor-group

(PFG).[28] However, we also want to exploit within-year variation in the exchange rate, and therefore require representative data on monthly travel. More importantly, we also require representative data at the level of each Census Division (CD) in order to examine the effect of the geographic distribution of residents on their propensity to travel. In order to construct data that are representative for each CD in each month, we construct our own weights.

Each questionnaire is associated with a particular CD and a port of entry into Canada. It also provides the month of travel and the length of the trip.[29] Therefore, each observation is CD–port–month–trip length combination. For notational clarity, we suppress subscripts for month and trip length. Define r_{cp} as the number of respondents from census division c passing through port of entry p. Define r_c as total respondents (across all CDs) at port p: $r_p = \sum_c r_{cp}$. Let n_p be the true number of crossers at port p which we obtain on a monthly basis from Cansim Table 427-0002. To estimate crossings by census division, $\hat{n}_c$, we first allocate n_p across census divisions using shares of response counts: $\hat{n}_{cp} = (r_{cp}/r_p)n_p$. Alternatively, one can think of this as the weighted sum of questionnaire respondents, r_{cp}, where weights are given by n_p/r_p, the number of actual crossers per respondent at a given port-month. Summing over all p for a given c we obtain $\hat{n}_c = \sum_p r_{cp} n_p / r_p$. The estimated crossing fraction is given by dividing $\hat{n}_c$ by our estimate of cars at risk, N_c.

B Additional Tables and Figures

[28] A PFG is a combination of a port of entry, length of stay, and mode of travel. For example, the PFG defined as Blaine–1 night–automobile is the set of traveling parties that entered Canada at the Blaine, BC port, having claimed to have spent one night in the US.

[29] We construct the length of trip from the reported dates of exit and entry. We assign the month of travel as the calendar month in which the vehicle entered Canada.

Table B.1: Regression of log crossings

Method:	Levels (contemp.)				Year-on-year diffs.			
Length of stay:	Daytrip		Overnight		Daytrip		Overnight	
Residence:	US	CA	US	CA	US	CA	US	CA
$\ln e$	1.47^{a}	-1.62^{a}	0.39^{a}	-1.47^{a}	0.60^{a}	-1.15^{a}	0.26^{a}	-1.04^{a}
	(0.08)	(0.07)	(0.08)	(0.06)	(0.10)	(0.06)	(0.07)	(0.08)
N	456	456	456	456	444	444	444	444
r2	0.79	0.86	0.93	0.93	0.18	0.52	0.04	0.26
r2 (baseline)	0.62	0.67	0.93	0.83	0.12	0.14	0.01	0.00
rmse	0.16	0.14	0.16	0.13	0.13	0.08	0.08	0.11

Standard errors in parentheses. An observation is a country-month.

^{c} $p < 0.1$, ^{b} $p < 0.05$, ^{a} $p < 0.01$

Table B.2: Regression of log crossings using Quartiles of RER

Method:	Levels (contemp.)				Year-on-year diffs.			
Length of stay:	Daytrip		Overnight		Daytrip		Overnight	
Residence:	US	CA	US	CA	US	CA	US	CA
$\ln e$	1.38^{a}	-1.90^{a}	0.45^{b}	-1.71^{a}	0.82^{a}	-1.32^{a}	0.17^{c}	-1.26^{a}
(CAD/USD)	(0.19)	(0.16)	(0.20)	(0.16)	(0.16)	(0.10)	(0.10)	(0.13)
$\ln e \times$ High RER	0.61^{b}	0.90^{a}	0.45^{c}	0.53^{a}	-0.05	0.31^{b}	0.36^{a}	0.44^{a}
(CA cheap)	(0.24)	(0.20)	(0.24)	(0.20)	(0.20)	(0.12)	(0.13)	(0.17)
$\ln e \times$ Low RER	-0.76^{a}	-0.66^{a}	-0.99^{a}	-0.18	-0.38^{a}	0.13	-0.07	0.14
(US Cheap)	(0.20)	(0.17)	(0.21)	(0.17)	(0.14)	(0.09)	(0.09)	(0.12)
N	456	456	456	456	444	444	444	444
r2	0.82	0.88	0.94	0.93	0.20	0.53	0.07	0.28
rmse	0.15	0.12	0.15	0.12	0.13	0.08	0.08	0.11

Standard errors in parentheses. An observation is a country-month.

^{c} $p < 0.1$, ^{b} $p < 0.05$, ^{a} $p < 0.01$

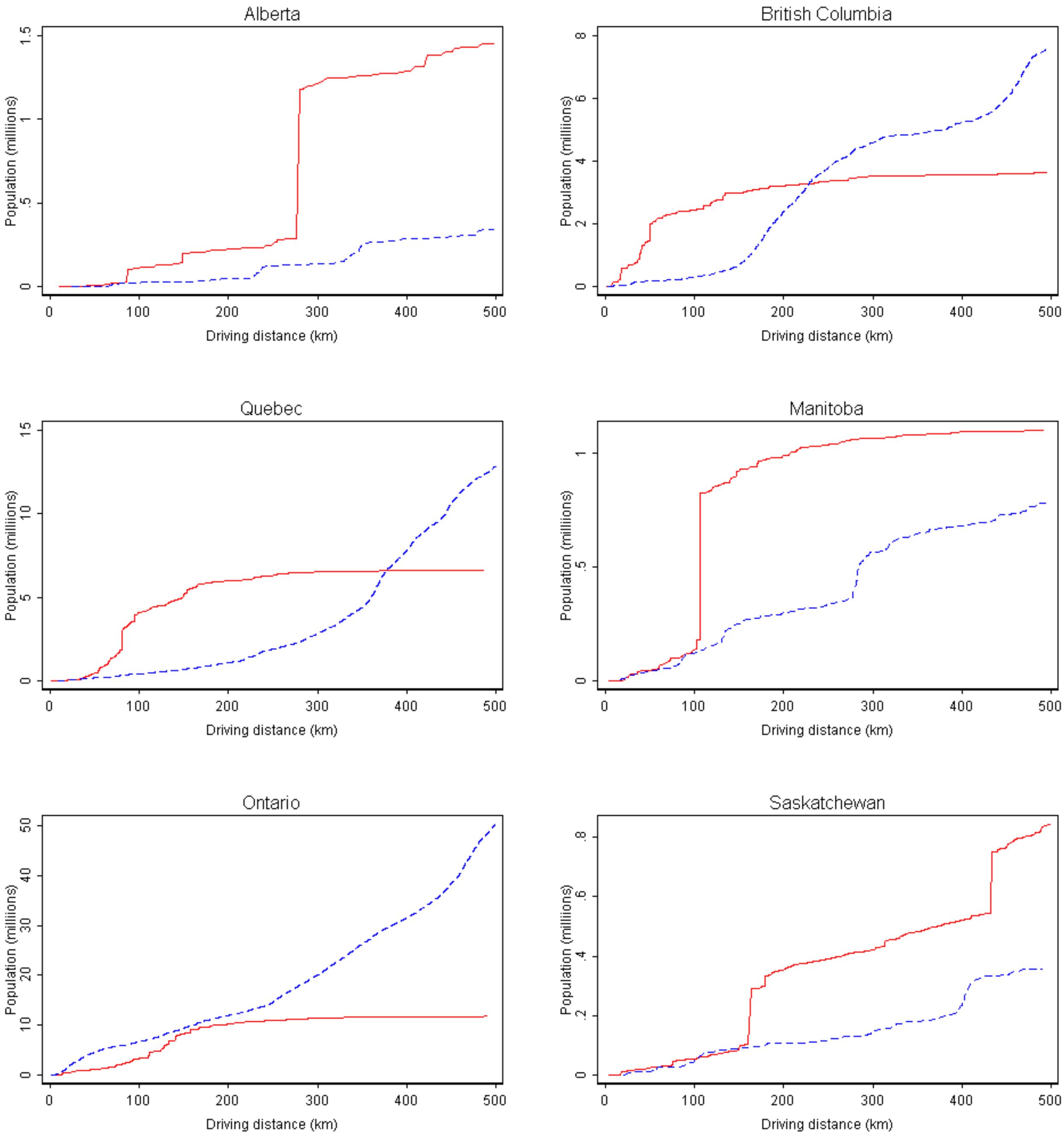

Figure B.1: Accumulated Population and Distance to the Border