

The Economics of Cross-Border Travel*

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Abstract

We model the decision to travel across an international border as a trade-off between benefits derived from buying a range of products at lower prices and the costs of travel. Using micro-data on Canada-US travel, we structurally estimate this model. Price differences motivate cross-border travel; our estimates indicate that a 10% home appreciation raises the frequency of cross-border day trips by 8% to 26%. The larger elasticity arises when the home currency is strong, a result predicted by the model. Distance to the border strongly inhibits crossings, with an implied cost of 87 cents/mile. Geographic differences can partially explain why American travel is less exchange-rate responsive.

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1 Introduction

International border crossings retain a vital influence on the economy since they are the sites where governments control the movement of goods and people between nations. Most economic research on people crossing borders considers permanent migration. This focus is natural given public concern over potential impacts on labor markets. Much less is known about the causes and consequences of short-term movement. Nevertheless, at many borders travel flows dwarf permanent movement. For example, 50 million car trips were made across the US-Canada border in 2010, about 250 times the number of permanent migrants arriving from Canada or the US in that year.¹ Understanding travel patterns has important implications for traffic forecasting, infrastructure planning, taxation, preventing terrorism, and controlling the spread of infectious diseases.

Discrete choice models have been successfully applied to cross-border migration² but this is the first paper to apply them to cross-border travel. One possible reason for lack of attention to cross-border travel is the perception that its motivations are non-economic. Whereas economic migrants seek out better jobs, travellers by definition return to their country of residence. Since this normally precludes earning income in the visited country, we must therefore look to motivations other than labor supply to explain most international travel.³ From an economic standpoint, the consumption motive is the natural explanation to consider.

This paper estimates a model in which the consumption motive drives short-term travel between Canada and the US. Drawing on evidence from descriptive statistics and reduced form regressions, we develop the first model of the decision by residents of one country to cross the border and purchase a cheaper bundle of goods in the other country. The model combines the decision of *whether to cross* with that of *what to buy* if one crosses. Because a stronger home currency expands the set of goods that are cheaper in the foreign country, the benefits of crossing are shown to be a convex function of the real exchange rate. Estimates of the model's parameters provide robust support for this hypothesis. Evaluated at 2010 exchange rates, the crossing elasticity is 2.6, three times the elasticity observed when the currency is weak, and higher than the Blonigen and Wilson (1999) estimates for the responsiveness of US-Canada trade in goods. This quantity response to price differences shows that the border does not fully segment the two markets as it appeared to do in the price shock tests of Gopinath et al. (2011). On the other hand, travelers' responsiveness is too limited to eliminate price differences, which points to the importance of significant travel costs. We find that a one percent increase in residents' distance to the border sharply reduces crossings. The consumption motive for cross-border travel is visible

¹Source: Statistics Canada and Migration Policy Institute

²See Grogger and Hanson (2011) for a estimated model in which income maximization drives the decision to emigrate.

³For the US and Canada commuters constitute only 5% of daytrips.

in the behavior of same-day travelers but it is concentrated among residents near the border. The median day tripper in Canada lives 18 miles from the border whereas the median Canadian lives 81 miles away.

The consumption motive for travel is predicated on the existence of price differences. Evidence of such differences on either side of the Canada-US border has been convincingly demonstrated by a series of papers. Engel and Rogers (1996) study of price dispersion between cities in Canada and the US reports that crossing the border is equivalent to a distance of 1,780 miles. While the Engel and Rogers estimate of the border's width has been challenged by Gorodnichenko and Tesar (2009), recent studies that examine disaggregated price data for identical goods on both sides of the border confirm the importance of price gaps. Notably, Gopinath et al. (2011) find a discontinuity in grocery prices of 24% at the Canada-US border. They also find that shocks to wholesale prices for supermarkets on one side of the border do not affect prices on the other side, a result which shows the segmentation of the two markets. Similarly, Burstein and Jaimovich (2009) find substantial amounts of pricing to market using scanner data from both sides of the border.⁴

Prior studies of cross-border travel for consumption purposes have tended to examine price or tax differences across jurisdictions for specific goods, and have inferred travel from measures of sales or retail activity. Asplund et al. (2007) infer cross-border shopping for alcohol between Sweden and Denmark by observing the response of retail sales to changes in taxes and exchange rates. Manuszak and Moul (2009) examine how differences in gasoline and cigarette taxes create incentives for residents to cross state borders, and thereby calculate consumers' travel costs. Knight and Schiff (2010) exploit the varying payoffs offered by state lotteries to infer that cross-border shopping has a large effect on the lotteries run by small states with densely populated border regions. To our knowledge, only one paper uses data on actual travel across borders in response to price changes: Chiou and Muehlegger (2008) examine how differences in cigarette taxes motivate residents to cross state borders, and estimate the relative importance of cigarette taxes and travel costs. While prior research has not examined the consumption of a general basket of goods by cross-border shoppers, two studies have provided indirect evidence of how exchange rate changes, which potentially alter prices of all goods, can affect cross-border travel. Campbell and Lapham (2004) and Baggs et al. (2010) find that exchange rate changes affect the employment and exit of retail firms located near the US-Canada border.

The paper proceeds in three steps. We begin by using graphs and reduced form regressions to uncover stylized facts of travel. We then develop a discrete choice model of travel consistent these patterns. The third step is to estimate the parameters of the model and use those estimates to quantify the roles of the key determinants of

⁴Goldberg and Knetter (1997), summarizing the earlier literature, point out that studies consistently find significant pricing to market. Boivin et al. (2011) show that even online book prices differ greatly between the US and Canada, and that their prices do not respond to exchange rate movements.

cross-border travel.

The reduced form exercises in Section 2 establish that travelers respond strongly to the economic incentives created by fluctuations in the exchange rate, suggesting that cross-border shopping is an important economic phenomenon. This finding corroborates results from reduced form estimations conducted by Ford (1992), Di Matteo and Di Matteo (1993, 1996) and Ferris (2000, 2010). We also find that travel by Canadian residents has a higher elasticity with respect to exchange rate changes than US travel. Moreover, for residents of both countries, the elasticity of crossings with respect to the exchange rate increases in absolute value as the home currency strengthens.

To make sense of these findings and to allow investigations of counterfactuals, Section 3 presents a model the decision to cross based on the premise that travellers seek bargains on the other side of the border. At any observed exchange rate, neither country has lower prices for all products. Our model therefore assumes a continuum of goods available in both countries at heterogeneous relative prices. Travelers who cross the border purchase the set of goods in each country that is cheaper in that country. Travelers who do not cross purchase all goods at home. The model naturally generates the prediction that as the home currency strengthens, the elasticity of crossings rises in absolute value. However, this is not because of heterogeneity in travel costs across residents, which tends to work in the opposite direction. Instead, the result is for two reasons: first, goods that were already cheaper in the foreign country are even more attractive now. Second, the set of goods that are cheaper in the foreign country expands.

Using a new dataset with information on the residence of consumers and their date of crossing, we estimate the parameters of this model in Section 4. Our estimated coefficients imply that the median crosser requires savings of almost \$30 per hour of travel time. The model also permits counterfactual experiments with respect to the key variables. We show that a 10% appreciation of the real exchange rate would increase cross-border travel frequencies by about 8% when the Canadian dollar is weak, but by 26% when it is strong. On the other hand, an exogenous doubling of border wait times would lower crossing frequencies by 50–60%, depending on the province. We estimate that travel has fallen by 32% since September 11, 2001, compared with the otherwise expected level of travel given the realized values of the exchange rate, gasoline prices, income, and population. The model provides a natural way to calculate the average crosser’s welfare gains in response to these changes. We find the 10% appreciation yields average crosser gains of 2.1% whereas the consequences of 9/11 have lowered average crosser gains by 3.4%. We also show that differences in the geographic distribution of residents partially explains the observation from the reduced-form regressions that Canadian elasticities with respect to the exchange rate are greater than US elasticities. We achieve this by simulating travel by Canadian residents in the event that their population distribution resembled that of the northern United States.

2 Stylized facts of border crossings

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 behavior in a manner that is consistent with cross-border shopping. Additionally, we find interesting variation in the response of travelers to currency fluctuations, both across countries and over time.

Surveys of travelers crossing the US-Canada border indicate that a majority of cross-border trips are made for pleasure or personal reasons, which include shopping trips. These sorts of trips are potentially the most likely to respond to exchange rates. 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.⁵ This information suggests that the exchange rate potentially plays an important role in the decision to cross the border, for residents of both countries. We now attempt to quantify the relationship between exchange rates and cross-border travel.

2.1 Data

We obtained data on cross-border travel from Statistics Canada, using information collected by the Canadian Border Services Agency (CBSA).⁶ These data consist of counts of all vehicles entering Canada at all land crossings with the United States. US residents encounter the CBSA on their outbound journey and Canadian residents on their return journey.

We use these data on vehicle counts for the 7 Canadian provinces that share a land border with the United States: British Columbia, Alberta, Saskatchewan, Manitoba, Ontario, Quebec and New Brunswick.⁷ We use monthly data for the calendar years 1972–2010. 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 travelers’ country of residence, which is determined by whether the vehicle has US or Canadian license plates. Finally, the data are broken down by the length of the cross-border trip. We analyze same-day and overnight trips separately.⁸

We obtained monthly average data on the spot market exchange rate between the US and Canadian currencies. Multiplying the nominal exchange rate by the ratio of monthly CPIs for both countries we construct the Real Exchange Rate (RER) for

⁵In Table 1 in the supplementary file, we tabulate the commonly stated motives for crossing the border. The data are based on the International Travel Survey of visitors and returning residents to Canada.

⁶See Cansim Table 427-0002.

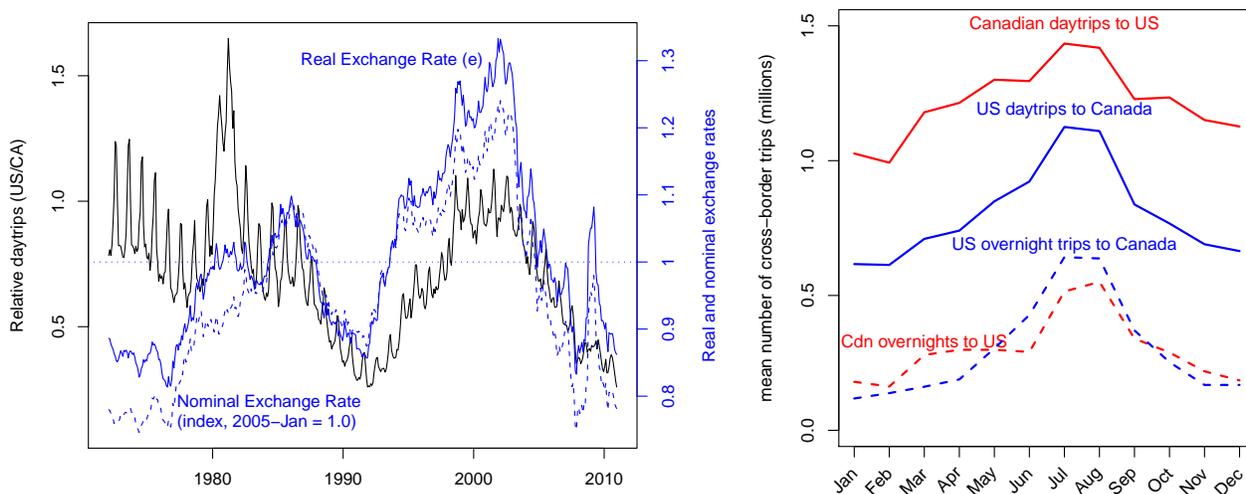
⁷Nova Scotia has a marine border with the US as it accepts ferry traffic from Maine. The Yukon Territory shares a border with Alaska. We omit these jurisdictions due to difficulties in ascertaining the corresponding US port from which vehicles enter Canada.

⁸“Overnight” is a short-hand to refer to trips spanning two or more days.

each month.⁹ It is defined with US prices in the numerator such that RER increases correspond to Canadian dollar depreciations. The RER incorporates relative taxes on goods and services in the two countries because the consumer price indexes in both countries are based on after-tax prices. Thus the 1991 introduction of the 7% goods and service tax (GST) in Canada is built into the RER. We fixed the absolute level of the RER using relative price levels from OECD data.

Figure 1 displays key temporal patterns in the data.¹⁰ Figure 1(a) shows the ratio of monthly same-day trips by residents of the two countries from 1972 through 2010. The solid blue line shows the real exchange rate. The dashed blue line shows the monthly nominal exchange rates, expressed in the figure as an index of the July 1993 level (1.29 CAD per USD), when the RER was approximately one (that is, prices of the consumer bundle expressed in a common currency were approximately equal). Because both countries have mainly had similar inflation rates, the primary source of real exchange rate variation is nominal variation.¹¹ US trips rise relative to Canadian trips when the US has relatively high price levels. Since the 1980s the relationship between relative travel and the exchange rate has become very strong.

Figure 1: Annual and monthly variation in crossings
(a) (b)



Travel is highly seasonal, for residents of both countries. Figure 1(b) shows average travel over the 38-year period for each calendar month. On average, Canadian residents make about 50% more daytrips across the border than do US residents. The

⁹Data sources and other details are provided in Appendix A.3.

¹⁰A table of summary statistics for these data is available in the supplementary file to this paper.

¹¹Put more precisely, log first differences of the nominal exchange rate can explain 94% of the variation in log first differences of the real exchange rate over the period 1972–2010. In levels the R^2 is 0.89.

number of overnight trips for the two countries is approximately the same. Cross-border travel peaks in the summer months for all groups.

We employ the CBSA data in the reduced form regressions that follow. However, for the structural model that we estimate in Section 4, we require information on the geographic distribution of crossers and the distance they travel to and from the border. This information is not available in the CBSA data, so we use a second source of data on cross-border travel: the International Travel Survey (ITS), which is also made available by Statistics Canada. This survey is filled out by travelers returning to Canada from trips abroad. The data were derived from questionnaires distributed from 1990 to 2010 that collected information on the nature and purpose of the trip, the dates on which travelers exited and entered Canada, and information on the Census Division in which the travelers reside and the ports used to cross to the US. We retain data on Canadian residents returning from the United States by car.¹²

We present summary statistics of the ITS data in Table 1. There are 63000 observations, each corresponding to a census division in a given month. The first column presents variable means across all observations, while the second column does so only for the subset of observations (39088) in which there was at least one car trip across the border in the given month. Conditioning on positive trips, Census Divisions tend to be closer to the border, and more populated. The large standard deviation for gas prices is mainly driven by temporal variation, whereas there is substantial cross-CD variation in household incomes, with the richest having incomes that are several times larger than the poorest.

Table 1: Summary Statistics: 63000 Census Divisions-months

Variable	Mean	Mean trips>0 ^a	SD	Median	Min	Max
Driving Distance (km)	263.0	187.0	281.2	161.9	6.8	1877.1
Driving time (hrs)	3.7	2.6	3.9	2.2	0.2	26.7
Population (1000)	116.2	165.8	273.8	40.8	1.2	2667.9
Gasoline Price (c/L)	73.5	72.5	21.1	66.5	39.5	146.6
Median HH Income (\$1000)	42.8	44.1	11.3	41.2	15.2	157.7
Cross-border trips (cars):						
Same-day	4093	6597	20229	0	0	456542
Overnight	1319	2126	4146	80	0	90662

^a 39088 CD-months with at least one car trip across the border.

¹²The 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.

2.2 The Exchange Rate Elasticity of Cross-Border Travel

Our first regression exercise is to determine the elasticity of cross-border trips with respect to the real exchange rate. Our main goal is establish simple data relationships to motivate the development of a model in the subsequent section of the paper. We therefore work with a minimal specification. Denoting the number of cars that cross the border by n , and the real exchange rate by e , our specification is:

$$\ln n_{it} = \text{Month}_t + \text{Province}_i + \eta_1 \ln e_t + \eta_2 \text{post911}_t + \eta_3 t + \eta_4 t^2 + \varepsilon_{it}, \quad (1)$$

where i denotes a province and t denotes time (in months since January 1972). The month effects account for the strong seasonality in travel. We add province fixed-effects, as well as an indicator variable for the period following September 11, 2001 when border security was increased. Finally, we add a linear and quadratic trend to capture secular effects such as population changes. We estimate this equation separately for residents of each country. Therefore, for Canada, this regression models the number of cars returning from the US in a given province and month. For the US, it represents the cars that enter the corresponding Canadian province.

Implicit in the estimation of equation 1 is the assumption that causation runs only from the real exchange rate to crossing decisions. This assumption is defensible because demand for foreign currency created by US and Canadian cross-border shoppers is unlikely to be large enough to move the global foreign exchange markets. To gain some perspective on relative magnitudes, Canadians spent \$4.2bn in the US while Americans spent \$1.8b in Canada during the first quarter of 2010.¹³ This represents a mere 0.04% of the foreign exchange turnover involving the Canadian Dollar.¹⁴

To establish the robustness of the stylized facts, we also estimate using year-on-year differences of equation 1. That is, we subtract from each variable the value it had twelve months before. This holds constant season and province effects and also removes time-varying factors that may not have been well captured by the trend variables:

$$\begin{aligned} \ln n_{it} - \ln n_{i,t-12} &= \{12\eta_3 + 144\eta_4\} + \eta_1 [\ln e_t - \ln e_{t-12}] \\ &+ \eta_2 [\text{post911}_t - \text{post911}_{t-12}] + 24\eta_4 t + \varepsilon_{it} - \varepsilon_{i,t-12}. \end{aligned} \quad (2)$$

The 12-month differences transform the linear trend into the constant term and the quadratic trend to a linear trend.

The results of estimating these equations are presented in Table 2. We treat each province in a calendar month as a separate observation.¹⁵ Since monthly crossing data

¹³This includes expenditures by air travelers. Source: International Travel Account Receipts and Payments (<http://statcan.gc.ca/daily-quotidien/100827/dq100827-eng.pdf>)

¹⁴Source: Authors' calculations from the BIS Central Bank Survey of Foreign Exchange and Derivatives Market Activity, 2010 (<http://www.bis.org/publ/rpfx10t.htm>)

¹⁵In Table 2 in the supplementary file we present corresponding regressions using country-level data, instead of breaking up the data by provinces. The results in that regression are similar to those presented here.

Table 2: Regression of log crossings, 1972–2010.

Method:	Levels (contemp.)				Year-on-year diffs.			
Length of stay:	Daytrip		Overnight		Daytrip		Overnight	
Residence:	US	CA	US	CA	US	CA	US	CA
$\ln e$	1.24 ^a	-1.62 ^a	0.47 ^a	-1.78 ^a	0.38 ^a	-1.16 ^a	0.12 ^c	-1.36 ^a
(CAD/USD)	(0.17)	(0.24)	(0.17)	(0.17)	(0.09)	(0.12)	(0.07)	(0.16)
N	3276	3276	3276	3276	3192	3192	3192	3192
R^2	0.98	0.98	0.96	0.97	0.06	0.27	0.01	0.24
R^2 (excl. $\ln e$)	0.98	0.97	0.96	0.95	0.03	0.05	0.00	0.00
RMSE	0.25	0.22	0.30	0.23	0.14	0.13	0.14	0.15

Newey-West standard errors in parentheses are robust to serial correlation out to 60 months. Significance indicated by ^c $p < 0.1$, ^b $p < 0.05$, ^a $p < 0.01$. An observation is a province-year-month. Coefficients on month and province fixed-effects, the post 9/11 indicator, and the trend variables are not reported.

are serially correlated, we use Newey-West standard errors.¹⁶ 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. 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 than those of US residents, across both specifications and both categories of trip-length.¹⁷

We investigate whether the crossing elasticity with respect to exchange rates varies with the level of the exchange rate in Table 3. We find significant interactions between the log of the RER and indicators for the highest and lowest quartiles of the RER over the 38-year period. In particular, the coefficient for the period when the US dollar was strong 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 strong, while Canadian residents become less responsive. We observe the opposite pattern during periods when the US dollar is in its lowest quartile.¹⁸

¹⁶There are too few provinces (7) for clustering at the province level to work.

¹⁷Adding economic indicators, such as unemployment and GDP, to the regressions has a modest effect on the coefficient of interest, and does not affect the general pattern of results. See Table 4 in the supplementary file for details. We do not include these variables here in order to maintain a minimal specification.

¹⁸In Table 3 in the supplementary file 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

Table 3: Regression of log crossings using Quartiles of RER, 1972–2010.

Method:	Levels (contemp.)				Year-on-year diffs.			
	Daytrip		Overnight		Daytrip		Overnight	
Residence:	US	CA	US	CA	US	CA	US	CA
$\ln e$ (CAD/USD)	0.93 ^a (0.28)	-1.71 ^a (0.28)	0.32 (0.23)	-2.08 ^a (0.21)	0.55 ^a (0.11)	-1.13 ^a (0.12)	0.14 ^c (0.08)	-1.42 ^a (0.15)
$\ln e \times [e > 1.09]$ (strong USD)	0.90 ^b (0.37)	0.54 ^c (0.33)	0.83 ^a (0.31)	0.65 ^b (0.29)	0.14 (0.11)	0.25 (0.18)	0.26 ^a (0.07)	0.27 (0.16)
$\ln e \times [e < 0.90]$ (strong CAD)	-0.87 ^b (0.34)	-0.87 ^a (0.24)	-1.25 ^a (0.32)	-0.31 (0.22)	-0.44 ^a (0.11)	-0.25 ^b (0.11)	-0.22 ^b (0.09)	-0.06 (0.12)
N	3276	3276	3276	3276	3192	3192	3192	3192
R^2	0.98	0.98	0.97	0.97	0.08	0.28	0.02	0.24
RMSE	0.24	0.21	0.29	0.23	0.14	0.13	0.14	0.15

Newey-West standard errors in parentheses are robust to serial correlation out to 60 months. Significance indicated by ^c $p < 0.1$, ^b $p < 0.05$, ^a $p < 0.01$. An observation is a province-year-month. Coefficients on month and province fixed-effects, the post 9/11 indicator, and the trend variables are not reported.

Canadian residents have zero exemptions from taxes and duties on goods purchased abroad when returning from a trip of less than 24 hours.¹⁹ Despite this, we observe same day travel being extremely sensitive to exchange rates: we estimate the 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.

This section has uncovered four stylized facts of cross-border travel that should be features of a quantitative model of crossing decisions. 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.

RER.

¹⁹Under 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. US residents generally have a \$200 exemption when returning from a same-day trip to Canada.

3 Model of the crossing decision

In this section we model the trade-offs faced by potential cross-border shoppers. The benefits from crossings are modeled using a continuum of goods structure similar to Dornbusch, Fischer, and Samuelson (1977). To focus on the crossing decision we omit the supply-side of that model. We show that the model generates a convex relationship between the savings obtained from cross-border shopping and real exchange rates that rationalizes the findings of the previous section.

Consumers purchase a continuum of goods on the unit interval. Good z has price $P(z)$ in the home country and a price $P^*(z)$ in the foreign country. Both prices are expressed in terms of the respective local currency units. Let E represent the nominal exchange rate defined in local currency units per foreign currency unit. Define \bar{P} and \bar{P}^* as the domestic and foreign consumer price indexes. The real exchange rate, which indicates the relative price of the foreign consumption bundle expressed in a common currency is given by $e = E\bar{P}^*/\bar{P}$. Lastly, we define $\delta(z)$ as the relative price deviation of good z :

$$\delta(z) = \frac{P(z)/\bar{P}}{P^*(z)/\bar{P}^*}. \quad (3)$$

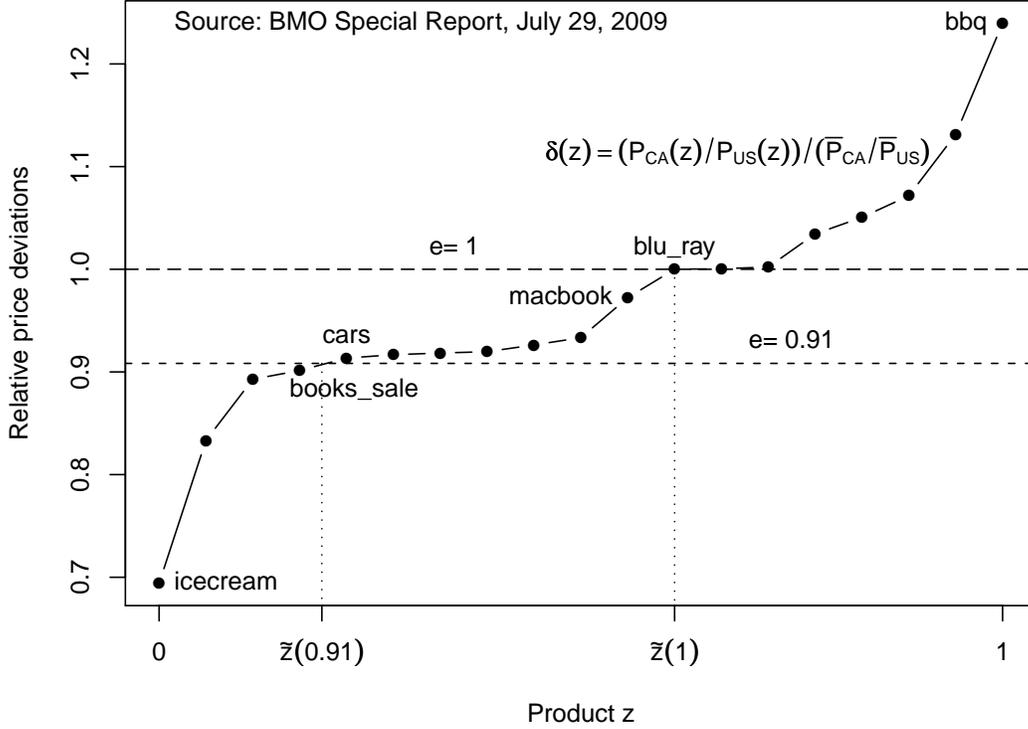
We order goods such that $\delta'(z) > 0$ and assume that *relative price deviations are invariant to the real exchange rate*, that is $\partial\delta(z)/\partial e = 0$. In Appendix A we show that this assumption is implied by the full DFS model.

The borderline good for which prices are equal after converting currency, is denoted \tilde{z} and defined implicitly as $P(\tilde{z}) = EP^*(\tilde{z})$. Substituting this relationship and the definition of the real exchange rate back into equation (3), it follows that $\delta(\tilde{z}) = e$. Goods $z < \tilde{z}$ are cheaper at home and the remaining goods are cheaper abroad. Inverting $\delta(z)$, we find $\tilde{z} = \delta^{-1}(e)$. Using the implicit function theorem, $\partial\tilde{z}/\partial e = 1/\delta'(\tilde{z}) > 0$. Thus a real appreciation of the foreign currency contracts the range of goods that are cheaper in the foreign country.

We illustrate the model in Figure 2 using data from Porter (2009). The author reports prices for 19 goods available on both sides of the border. Calculating $\delta(z)$ as the ratio of the Canadian price (in CAD) to the US price (in USD), all divided by the relative price levels (1.2, obtained from the OECD), we sort z in increasing order and plot relative price deviations, $\delta(z)$, against $z = (i - 1)/18$. At the time the article was written the exchange rate was 1.09 CAD/USD, leading to a real exchange rate of $e = 0.91$. With a Canadian dollar at this strength, 15 out of 19 goods (from cars to BBQ grills) were less expensive in the US after converting prices to a common currency. The figure shows that seven goods—from cars to MacBooks—would switch to being cheaper in Canada if the USD appreciated by 10% to $e = 1$ (purchasing power parity). Thus holding nominal price deviations constant, real exchange rate changes can produce dramatic shifts in \tilde{z} .

To take advantage of lower prices in foreign retail stores, the consumer engages in cross-border shopping. Thus wholesalers can trade goods across borders but, due to

Figure 2: Exchange rates and relative prices: 19 products



pricing-to-market by home retailers, consumers can only obtain the foreign price by travelling to the foreign retail store.²⁰ Individuals decide whether to stay at home or cross by comparing the indirect utility associated with each option. Consumers have Cobb-Douglas utility with expenditure share parameters $b(z)$. Stayers spend their entire income, W , in the home country. This implies an indirect utility, v_S , given by

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

Crossers buy goods $\tilde{z} \leq z \leq 1$ in the foreign country but make the rest of their purchases at home. Travel costs take the “iceberg” form: $1 - 1/\tau$ is the fraction of income that “melts away” in the trip across the border.²¹ Neglecting any home government taxes on the goods travelers bring back, the price paid for foreign goods is

²⁰An implicit assumption is that the proportion of cross-border shoppers is not large enough to have a material effect on pricing decisions by firms on either side of the border. This assumption is consistent with the price differences shown in Figure 2 and the discontinuities at the border documented by Gopinath et al. (2011).

²¹This assumption is made here for expositional purposes. In the implementation we specify travel costs as a function of distance to the border, gas prices, and the opportunity cost of time.

$EP^*(z)$ in domestic currency.²² Finally we assume a non-pecuniary benefit (or cost, if negative) of travel given by ζ . The indirect utility of crossers is therefore given by

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

The model should not be taken literally since cars cannot physically accommodate all the products that are cheaper in the foreign country.²³ The important idea is that the indirect utility of a cross-border trip depends on the prices of the goods that a consumer would actually choose to buy in the foreign country.

The net benefits of crossing is obtained by subtracting v_S from v_X , yielding

$$v_X - v_S = B - \ln \tau + \zeta, \quad (4)$$

where $B \equiv \int_{\tilde{z}}^1 b(z) [\ln P(z) - \ln EP^*(z)] dz$, the gross benefit of crossing, is the savings from buying goods in the foreign country instead of domestically. For any interior value of \tilde{z} , B is positive since $P(z) > EP^*(z)$ for all $z > \tilde{z}$.

Using the notation of DFS we also define $\vartheta(\tilde{z}) = \int_0^{\tilde{z}} b(z) dz$ as the share of expenditures on goods for which the home country is the low-price supplier. Inserting the definitions of e , $\delta(z)$ and ϑ into B , we express the benefits of crossing as a function of the log real exchange rate:

$$B(\ln e) = -(1 - \vartheta(\tilde{z})) \ln e + \int_{\tilde{z}}^1 b(z) \ln \delta(z) dz. \quad (5)$$

The first term shows that holding \tilde{z} constant, a stronger foreign currency lowers the benefit of crossing. The second term can be thought of as the correlation between budget shares and price deviations for the set of goods $z > \tilde{z}$. It says that the benefits of crossing are higher if consumers happen to particularly like the goods that are relatively expensive at home.

Noting that $\vartheta'(\tilde{z}) = b(\tilde{z})$, the derivative of (5) with respect to $\ln e$, while holding $\delta(z)$ constant,²⁴ can be expressed as

$$B' = -(1 - \vartheta(\tilde{z})) + b(\tilde{z})(\ln \delta(\tilde{z}) - \ln e) \frac{\partial \tilde{z}}{\partial \ln e} = -(1 - \vartheta(\tilde{z})) < 0,$$

where the second term equals zero because $\delta(\tilde{z}) = e$. This reveals that the impact of the exchange rate on the benefits of crossing depends on the share of goods that are

²²This assumption is grounded in anecdotal accounts of undeclared purchases and *de facto* exemptions for small amounts of declared spending. Adding a tax would just be a scalar multiplying the real exchange rate.

²³A more realistic approach would be to consider a model of random replacement of durable goods. The $b(z)$ would measure the probability that a particular good needed to be replaced. The v_S and v_X would become expected utilities.

²⁴We assume that changes in $\ln e$ are generated by changes in the nominal exchange rate E or by proportional shocks to all prices such as ad valorem taxes or factor price increases.

cheaper abroad. Foreign appreciation contracts the basket of goods that are cheaper abroad, i.e. rising e decreases $1 - \vartheta(\tilde{z})$. This leads the benefit function to be convex in the real exchange rate:

$$B'' \equiv \frac{\partial^2 B}{\partial \ln e^2} = b(\tilde{z}) \frac{\partial \tilde{z}}{\partial \ln e} = b(\tilde{z}) \frac{\delta(\tilde{z})}{\delta'(\tilde{z})} > 0. \quad (6)$$

The convexity of the $B(\ln e)$ function arises under general functional form assumptions for preferences, $b(z)$, and relative price deviations $\delta(z)$. We use a quadratic form for $B(\ln e)$ in our empirical specification. It is the simplest way to capture and test for the convexity predicted by the model and can be thought of as a second-order approximation of a general B .

$$B(\ln e) = \beta_0 + \beta_1 \ln e + \beta_2 [\ln e]^2. \quad (7)$$

The model predicts $B' = \beta_1 + 2\beta_2 \ln e < 0$ for the observed range of e and $B'' = 2\beta_2 > 0$. The quadratic is the exact solution under the assumptions of uniform budgeting and exponential relative price deviations, that is $b(z) = 1$ and $\delta(z) = \exp[\lambda(z-1/2)]$.²⁵ Under these assumptions, the solution for the borderline good is linear in the log real exchange rate ($\tilde{z} = \frac{1}{2} + \frac{1}{\lambda} \ln e$) and parameters of equation (7) have structural interpretations, with $\beta_0 = \lambda/8$, $\beta_1 = -1/2 < 0$, and $\beta_2 = 1/(2\lambda) > 0$.

We introduce individual heterogeneity to the net benefits of crossings in two ways. First, we add community c subscripts to the determinants of travel costs to reflect differences in distance to the border and wages (opportunity cost of time). Second, we allow for individual-specific heterogeneity in the unobserved non-pecuniary benefits of crossing, distributed with a CDF denoted $F(\zeta)$. Within each community c there is a marginal individual who is indifferent between crossing and staying. This ζ_c^* is defined by setting $v_X = v_S$, yielding $\zeta_c^* = -B(\ln e) + \ln \tau_c$. Thus, residents of distant communities (high τ_c) require a higher idiosyncratic shock to justify crossing the border. With a continuum of individuals, the fraction of crossers, denoted x_c , will be equal to the probability that a potential crosser has $v_X > v_S$:

$$x_c = \mathbb{P}(\zeta_c^* < \zeta) = F(B(\ln e) - \ln \tau_c). \quad (8)$$

The model's predicted travel elasticities with respect to the real exchange rate and travel costs depend on the curvature of the CDF, but both are unambiguously negative:

$$\frac{\partial \ln x_c}{\partial \ln e} = \frac{F'}{F} B' = -\frac{F'}{F} [1 - \vartheta(\tilde{z})] < 0, \quad \frac{\partial \ln x_c}{\partial \ln \tau_c} = -\frac{F'}{F} < 0. \quad (9)$$

²⁵The exponential deviations assumption is not as arbitrary as it might seem. Since z spans the unit interval and $\delta(z)$ is sorted in increasing order, the $\delta(z)$ function is actually the inverse of the the cumulative distribution function (CDF) of relative nominal prices. Hence an exponential form implies that the CDF of relative prices is linearly related to the log of $\delta(z)$. Strictly positive variables are often distributed log-normally in practice and this distribution has the feature that for most of the data except the tails, there is a close-to-linear relationship between the CDF and the log of the variable.

Since travel costs increase with distance to the border, incomes, and gas prices, the model predicts negative effects of these variables on the propensity to cross.

A key relationship uncovered in our reduced form regressions in section 2 was that the exchange rate elasticity of travel diminishes in periods when the foreign currency is strong. To see whether this effect is predicted by our model we differentiate the first equation in (9) to obtain

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

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 (6) will not translate into a positive second derivative for aggregate log crossings if the term in brackets is sufficiently negative. Second, we see that if we eliminated the convexity in the benefit function and set $B'' = 0$, as in the single-product model sketched in Appendix A, the term in square brackets would have to be positive to yield convexity of log crossings. For commonly used distributions of individual heterogeneity, the factor in brackets has a negative sign.²⁶

A second reduced-form finding we would like to reconcile with the model is that crossers from Canadian provinces into US states exhibit higher exchange rate elasticities than residents of the US states on the other side of the border. To think about why a province (or country) might have a higher elasticity, we need to aggregate multiple communities c , of size N_c into a single region R of size N_R . The crossing rate of the aggregate is $x_R = \sum_{c \in R} \frac{N_c}{N_R} x_c$. The elasticity of crossings of this region with respect to e is given by

$$\frac{\partial \ln x_R}{\partial \ln e} = \sum_{c \in R} \frac{N_c}{N_R} \frac{x_c}{x_R} \frac{\partial \ln x_c}{\partial \ln e} = -\frac{[1 - \vartheta(\tilde{z})]}{x_R} \sum_{c \in R} \frac{N_c}{N_R} F'. \quad (10)$$

Inspection of equation (10) suggests various ways in which crossing elasticities can differ between regions. One way US elasticities could be smaller is if $\vartheta(\tilde{z})$, the expenditure share of goods that are cheaper at home, were sufficiently lower in Canada than its counterpart for the US. In the model this would occur if $b(z)$ and $b^*(z)$ are positively correlated with $\delta(z)$, that is if both countries tend to spend high shares of their incomes on goods that are relatively expensive in Canada.

Equation (10) also reveals that differences in regional elasticities can arise from differences in the geographic distribution of the potential crossers in each region. If cities in one region all have higher τ_c , x_R decreases and the absolute value of the crossing elasticity in equation (10) becomes larger. There is a secondary impact of higher τ_c via changes in F' . The elasticity is only certain to rise (in absolute value)

²⁶ F'/F is globally decreasing for uniform, normal, logit, gumbel, and power-law distributions. 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.

if $F'' < 0$ for all communities c . The analysis is further complicated when taking into account difference in the weights of potential crossers, N_c/N_R . In general, the relationship between geography and the regional exchange rate crossing elasticity must be addressed numerically. After estimating the model's parameters, we will investigate whether geographic differences contribute to explaining the stylized fact that US residents have a lower crossing elasticity than Canadians.

A final point to note is that, because of the aggregation issues revealed in equation (10), the model should be estimated using geographically disaggregated data. In particular, it is important to use data on the distance to the border from each community from which travellers originate.

4 Estimation of the model

In this section we take the model of the previous section to the data. We use our estimates to calculate implied travel costs and to conduct counterfactual welfare analysis.

4.1 Regression Specification

In order to estimate the crossing fraction equation shown in (8), we need to parameterize the crossing benefit and cost functions (B and $\ln \tau_c$) as well as specify the distribution of individual heterogeneity ($F(\zeta)$). We make use of the quadratic form for $B(\ln \epsilon)$ shown in equation 7.

The next step is to parameterize τ_c in terms of its underlying observable determinants. The cost of the cross-border trip consists of the sum of the opportunity cost of driving time and fuel costs. Letting parameters ψ equal speed (kilometer per hour), ϕ equal fuel efficiency (kilometers per liter), and H equal the endowment of hours, the total crossing cost is $D_c[W_c/(\psi H) + P(g)_c/\phi]$, where $P(g)_c$ is the price of gasoline (per liter) and D_c is driving distance (in kilometers). Expressing travel costs in iceberg form, that is the ratio of initial income to net income after incurring travels yields

$$\tau_c = \left[1 - D_c \left(\frac{1}{\psi H} + \frac{P_c(g)}{\phi W_c} \right) \right]^{-1}. \quad (11)$$

We see that the strict iceberg assumption of a constant fraction of income lost from travel is only met in the limit as the gas price to income ratio goes to zero. To facilitate estimation, we use a linear-in-logs approximation of equation 11:

$$\ln \tau_c = \gamma_0 + \gamma_1 \ln D_c + \gamma_2 \ln [P(g)_c/W_c]. \quad (12)$$

The γ_0 parameter shifts travel costs at all distances. One such shifter would be

border formality compliance costs.²⁷ The γ_1 parameter represents the elasticity of travel costs with respect to distance.²⁸

Substituting the B and $\ln \tau$ functions into equation 8, we can express the crossing fraction as

$$x_c = F[\beta_0 - \gamma_0 + \beta_1 \ln e + \beta_2 (\ln e)^2 - \gamma_1 \ln D_c - \gamma_2 \ln (P(g)_c/W_c)]. \quad (13)$$

We assume that ζ follows a normal distribution, with expectation μ and variance σ^2 . Thus, $F(\zeta) = \Phi([\zeta - \mu]/\sigma)$, where $\Phi()$ denotes the standard normal CDF.²⁹ Substituting these parameterizations into equation 13 and adding time subscripts we obtain

$$x_{ct} = \Phi[\theta_0 + \theta_1 \ln e_t + \theta_2 (\ln e_t)^2 + \theta_3 \ln D_c + \theta_4 \ln (P(g)_{ct}/W_{ct})], \quad (14)$$

where Table 4 shows the mapping between the θ and the structural parameters as well as the expected signs for each coefficient.

Table 4: Interpretation of coefficients

Parameter	Covariate	Structure	Sign
θ_0	constant,	$(\beta_0 - \gamma_0 + \mu)/\sigma = (\lambda/8 - \gamma_0 + \mu)/\sigma$	+ or -
θ_1	$\ln e_t$ (RER)	$\beta_1/\sigma = -1/(2\sigma)$	-
θ_2	$(\ln e_t)^2$	$\beta_2/\sigma = 1/(2\lambda\sigma)$	+
θ_3	$\ln D_c$	$-\gamma_1/\sigma$	-
θ_4	$\ln [P(g)_{ct}/W_{ct}]$	$-\gamma_2/\sigma$	-

Equation 14 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 two 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, the data that we use for estimation are the ITS data, described in Section 2.1, which are based on a survey given out to a subset of the actual population of crossers. We restate equation 14 in the form of a conditional expectation:

$$\mathbb{E}[x_{ct} | e_t, D_c, P(g)_{ct}, W_{ct}] = \Phi[\theta_0 + \theta_1 \ln e_t + \theta_2 (\ln e_t)^2 + \theta_3 \ln D_c + \theta_4 \ln (P(g)_{ct}/W_{ct})]. \quad (15)$$

²⁷Since these costs are thought to have risen following September 11th, 2001, we include a Post-9/11 dummy in most specifications.

²⁸The empirical trade literature routinely assumes a constant elasticity of distance. We report estimation results using quadratic distance functions in the supplementary file. We also estimated a semi-parametric step function. Neither generalization improves the fit enough to justify the loss in parsimony.

²⁹We show later that alternative distributional assumptions for ζ (logistic, Gumbel) lead to very similar ratios of coefficients.

When estimating this equation, we need to recognize that the dependent variable is a fractional response bounded between 0 and 1. As Table 1 shows, in many census-divisions the number of cars crossing the border is zero. It is important to employ an estimation method that (a) incorporates zeros into the estimation and (b) does not yield out-of-bounds predictions. We therefore estimate equation (15) using as a fractional probit model using quasi-likelihood methods. Fractional probit yields consistent estimates of the model parameters so long as the conditional expectation shown in Equation 15 is correctly specified. Had we assumed that F is logistic, it would have been possible to take the log of the odds, $x/(1-x)$, and obtain a right hand side that is linear in the parameters and therefore estimable using OLS. We show the results for this approach in the robustness section but do not adopt it as the main method for two reasons. First, the log odds is undefined at the limit values and thus can induce selection bias by dropping observations with zero crossings. Second, it does not estimate the conditional expectation of x_{ct} consistently.³⁰ Assuming a uniform distribution for F would lead to a linear probability model. This method predicts values outside of the $[0, 1]$ interval, rendering it unsuitable for counterfactuals.

The dependent variable is the crossing fraction, x_{ct} , which is defined as the number of car crossings, n_{ct} , from Census Division (CD) c in month t , divided by the number of potential crossings, N_{ct} . Potential crossings are approximated as the population of the census division (Pop), 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

$$x_{ct} = \frac{n_{ct}}{N_{ct}} \approx \frac{\hat{n}_{ct}}{\text{Pop}_{ct} \times \text{CPC}_c \times 30}. \quad (16)$$

We estimate \hat{n}_{ct} using data from the International Travel Survey (ITS), which was described in Section 2.1. Appendix B.1 shows the sources for the variables in equation 16 and details how we construct \hat{n}_{ct} by weighting the ITS responses using the port-level counts of all crossers, so as to make the sample representative at the monthly level as well as representative at each port of entry.

We measure D_c , the distance from census division c to the border, in two ways described in Appendix B.2. Our preferred form is the population-weighted median of the driving distances of all the subdivisions within a given CD.³¹ In robustness checks we also measure D_c as the median driving time to these ports and as the average of driving distances to the five most-used ports. Gas prices, $P(g)$, are obtained for the largest city in each province. Median household income, our proxy for W_c , is available

³⁰Papke and Wooldridge (1996) explain this defect and also elaborate on the advantages of the fractional probit method.

³¹Figure C.1 contains a map of a few CDs in south-eastern Ontario and shows the subdivisions (with thin borders) within each CD (with thick borders). Note the importance of using driving distance, as opposed to, say, great circle distance, given that there are a number of large lakes near the US–Canada border, as well as given the actual network of highways. Using a Euclidean distance metric would greatly understate the distance of a city such as Toronto from the border.

at the CD-level from the Canadian census in five year intervals.³²

4.2 Baseline Estimation

In this section we estimate the structural model implied by equation 15. We estimate the model separately for travelers making same-day and overnight trips; the latter are defined as stays of two or more days. We do this because travelers whose main reason for crossing the border is to shop are far more likely to make same-day trips, and these are travelers whose behavior is represented in the model of Section 3. By contrast, those making overnight trips may have purposes other than just shopping for goods to bring home: vacations, recreation spanning multiple days, visiting acquaintances etc. For these travelers, the single-good model sketched in Appendix A may be more appropriate. On a related note, same-day and overnight travelers may respond differently to gasoline prices and other travel cost shocks, as we discuss below.

The results using the fractional probit method of estimation are presented in Table 5. The first three columns use daytrips to construct the dependent variable, while the next three use overnight trips. All estimated specifications include (unreported) month dummies to allow shocks to the mean of the $\zeta(i)$ distribution reflecting the seasonal pattern shown in Figure 1(b). Standard errors are clustered at the census division (*c*) level to allow for arbitrary serial correlation within divisions.³³ The initial specification, shown in columns 1 and 4, assumes that travel costs are constant across time and depend only on the distance of the traveler’s origin to the border. Columns 2 and 5 estimate the influence of gas prices and incomes. We do not report the specification imposing equal and opposite coefficients on $\ln P(g)$ and $\ln W$ because we found that the same day travel data strongly reject this constraint. Our preferred specification, shown in columns 3 and 6, adds fixed effects (FE) for each province and a dummy for travel after September 11, 2001. The province FEs capture differences in $B(\ln e)$ that result from unmeasured cross-state differences in product prices.³⁴ The post 9/11 dummy is designed to capture real and perceived increases in the cost of crossing the border following heightened security measures.

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

³²Data details and sources are provided in Appendix B.3.

³³Ideally we would use two-way clustering of standard errors, to account for each census division in month t having the same real exchange rate. While this is not currently feasible using fractional probit, it can be done in the log-odds estimation which is another advantage of using that method as a robustness check.

³⁴They can also account for differences in the mean idiosyncratic shocks due to different population densities on the US side of the border which affect the likelihood of visiting friends and relatives.

Table 5: Fractional Probit estimation of crossing fractions (x_{ct})

Length of stay:	Daytrip			Overnight		
θ_0 : constant	-0.23 (0.31)	9.80 ^a (2.94)	4.42 ^a (1.52)	-2.68 ^a (0.07)	-4.59 ^a (0.57)	-5.20 ^a (0.99)
θ_1 : $\ln e_t$ [RER]	-0.44 ^a (0.10)	-0.77 ^a (0.14)	-0.65 ^a (0.13)	-0.61 ^a (0.12)	-0.92 ^a (0.13)	-0.75 ^a (0.12)
θ_2 : $(\ln e_t)^2$	0.39 (0.34)	1.24 ^a (0.33)	0.82 ^b (0.33)	-0.09 (0.30)	0.27 (0.28)	-0.17 (0.24)
θ_3 : $\ln D_c$ [distance]	-0.58 ^a (0.06)	-0.58 ^a (0.06)	-0.52 ^a (0.04)	-0.14 ^a (0.01)	-0.14 ^a (0.01)	-0.12 ^a (0.01)
$\ln P(g)_{ct}$ [gas price]		-0.35 ^a (0.09)	-0.07 (0.05)		-0.56 ^a (0.04)	-0.13 ^a (0.02)
$\ln W_{ct}$ [income]		-0.80 ^a (0.27)	-0.42 ^a (0.14)		0.40 ^a (0.06)	0.29 ^a (0.09)
New Brunswick			0.40 ^a (0.14)			0.00 (0.06)
Quebec			-0.46 ^a (0.08)			-0.15 ^b (0.07)
Ontario			-0.23 ^b (0.11)			0.07 ^b (0.03)
Manitoba			-0.33 ^a (0.13)			0.03 (0.04)
Saskatchewan			-0.45 ^a (0.10)			-0.15 ^a (0.04)
Alberta			-0.48 ^a (0.11)			-0.18 ^a (0.05)
Post-911			-0.14 ^a (0.03)			-0.14 ^a (0.03)
R^2	0.24	0.29	0.53	0.05	0.07	0.08
AIC	1935.18	1908.66	1778.11	629.59	626.92	636.59

Standard errors clustered by census-division. British Columbia is the omitted province. Regressions include month fixed-effects. ^c $p < 0.1$, ^b $p < 0.05$, ^a $p < 0.01$. $N = 63000$

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 3. 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.

We do not observe the same behavior by overnight travelers: the coefficients on $[\ln e_t]^2$ are small and statistically insignificant in columns 4–6. This may be because overnight travelers are more likely to purchase a standard bundle of goods in the US (hotel stays, vacations, restaurant meals etc.) without adjusting the scope of the bundle in accordance with relative prices. This still implies a positive elasticity of overnight travel with respect to the home currency, but does not imply that the elasticity changes with the RER. In other words, day trips are consistent with the multi-product shopping motive, whereas overnight trips instead appear to better fit with a single-good model such as the one in Appendix A.

Examining expenditure data provides additional support for this hypothesis. The International Travel Survey asks returning residents about their purchases made outside the country. These figures are subject to travelers' accurate recollection and truthful reporting of these amounts, and therefore susceptible to bias. Nevertheless, our examination of reported expenditures shows that same-day travelers have a positive elasticity of spending with respect to the home currency, while overnight travelers exhibit no effect of the exchange rate on their spending.³⁵

High gasoline prices should increase travel costs and reduce the propensity of border crossings. But at the same time gasoline is a commodity that is starkly cheaper in the US, due to tax differences. Canadians who cross to the US are overwhelmingly likely to fill up their cars, and some Canadians explicitly cross the border simply to buy gas. Media reports indicate that high gasoline prices motivate a significant number of Canadian residents to cross the border to purchase gasoline. A comparison of gas prices suggests that the absolute savings on gas purchases in the US tend to increase as gas prices rise.³⁶

Those travelers whose primary motive to cross the border is to buy gas must almost surely be making same-day trips. Therefore, it is not clear what the net effect of gas prices will be on the behavior of same-day travelers. By contrast, overnight travelers should only face a disincentive effect on travel as gas prices rise. This is especially since people making overnight trips tend to drive longer distances than those making same day trips. The results of Table 6 bear this out. Gas prices do not have a significant effect on day trips but have a negative and significant effect on overnight trips.

In separate regressions, we examined how weather affects travel. We find that rain and snow reduce the propensity of Canadians to make cross-border trips. This effect is observed only in regressions without month fixed-effects. When month effects

³⁵See Tables 7 and 8 in the supplementary file for these results.

³⁶This is partly due to higher ad valorem taxes in Canada. This behavior may be especially driven by household sub-budgets for gasoline, as documented by Hastings and Shapiro (2012).

are included, weather does not have a significant effect on travel, indicating that the regular pattern of the seasons explains travel behaviour more so than idiosyncratic deviations from it. Overall, our findings with regard to gas prices and weather are consistent with our results regarding exchange rates, and show that travelers respond appropriately to changes in the costs or benefits of travel

The coefficient in column 6 is about the same as the distance coefficient.³⁷ Income effects are strongly negative for day-trippers. This runs counter to what would be expected if income mattered just because it affects the fuel cost to income ratio in τ . Our model assumes a constant marginal utility of income across all individuals. One interpretation of the results is that richer households are less motivated by the savings to be had from cross-border shopping. For overnight trips income effects are positive. In column 6 the regression does not reject the restriction of equal and opposite effects for gas prices and incomes that is predicted by the transport cost function shown in equation 11.

The province fixed effects capture the underlying propensity of travelers from each province to cross the border, after accounting for exchange rate, distance, and income effects. They may reflect the presence of large cities, and the provision of goods and services that may be sought by Canadians, such as gasoline, outlet malls, casinos, airports etc. It is not surprising that British Columbia (the omitted category) and Ontario have higher fixed effects than Alberta and Saskatchewan. There are population centers near the border in Washington, Michigan, and New York but not in Montana and North Dakota.

The downward shift in travel to the US following September 2011 corroborates the finding of Ferris (2010) who estimates a linear reduced form regression using aggregate monthly same-day travel for Canada from 1972 to April 2009. The distance equivalent of 9/11 is given by $\exp(0.14/0.52) - 1 = 0.31$. Thus, the extra costs of crossing the border in the years since 9/11 corresponds to a 31% increase in distance. Alternatively, using a counterfactual calculation of the kind described in Section 4.5, we find a total reduction of 32% in travel attributable to 9/11. Remarkably, given the many differences in method, Ferris (2010) reports a 29% annual reduction.

We illustrate the magnitudes of the effects we have estimated by showing how predicted crossing shares respond to changes in our key explanatory variables. This is important since the estimated coefficients are scaled by the unobserved σ parameter. Moreover, the effects of the RER and distance have to pass through the nonlinear $\Phi()$ function to determine the predicted crossing share.

We show the relationship between the crossing fraction and the real exchange rate for specific distances from the border in Figure 3. This figure is based on the specification in column 3 for Table 5 (adjusting using the coefficients on the Ontario, post 9/11, and April dummy variables). Each curve corresponds to a census division in Southern Ontario.³⁸ The curves show that the convexity in the B function carries

³⁷The larger negative effect in column 5 is mainly attributable to the absence of the 9/11 dummy.

³⁸These census divisions—Niagara, Hamilton and Toronto—are CDs 26, 25 and 20 respectively in

over to the log crossing function. Thus, the elasticity of crossing is larger in absolute value when the home currency is strong. 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 Niagara.

The main determinant of travel costs is distance to the border. Figure 4 shows the steep decline of crossing fractions associated with increased driving distances. The curve graphs the average of the predicted shares (in percent) that would cross from each Ontario census division during the sample period (1990–2010). The circles show actual crossing fractions averaged over the same period. The model fits the data well, further supporting the validity of the linear-in-logs approximation of the travel cost function.

Divisions further from the border than Toronto (about 90 miles) have predicted and actual crossing rates below 0.1%. This means that on any given day there is a less than 1 in 1000 chance for a car to be driven across the US border on a daytrip. In contrast, communities closer than Niagara (15 miles) have crossing rates that are more than an order of magnitude higher. The evidence of porous borders is consistent with market segmentation because of the combination of strong distance effects and the fact that the majority live more than 80 miles from the border.

4.3 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 6. We use the set of controls corresponding to columns 3 and 6 of Table 5. We prefer this specification since adding province fixed-effects improves the fit of the model considerably, compared with columns 2 and 5.

The first two columns of Table 6 present results using the log of the odds of travel ($x_{ct}/(1 - x_{ct})$) as the dependent variable and estimating with OLS. The remaining columns return to the fractional probit model, but use different measures of the costs of travel. In columns 3 and 4 we use the driving *time* to the border from each Census Division, instead of the driving distance. This exploits the information Google keeps about differences in average driving speeds relevant for different subdivisions.³⁹ We add 26 minutes to the driving time to account (very roughly) for border wait times.⁴⁰ In columns 5 and 6 we use our secondary measure of distance (detailed in the Appendix). Relative to the primary measure used in Table 5, it has the advantage of taking into account not just the nearest port but the five ports that residents of the

Figure B.1. The nearest border crossings are on the Niagara river at the bottom right of the map.

³⁹The average speed is 70 km/hour with a 5%–95% range of 51–84 km/hour.

⁴⁰Wait time data is not systematically available across Canada during our estimation period. The 26 minutes figure is the median wait 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.

Figure 3: The crossing fraction declines with strength of foreign currency

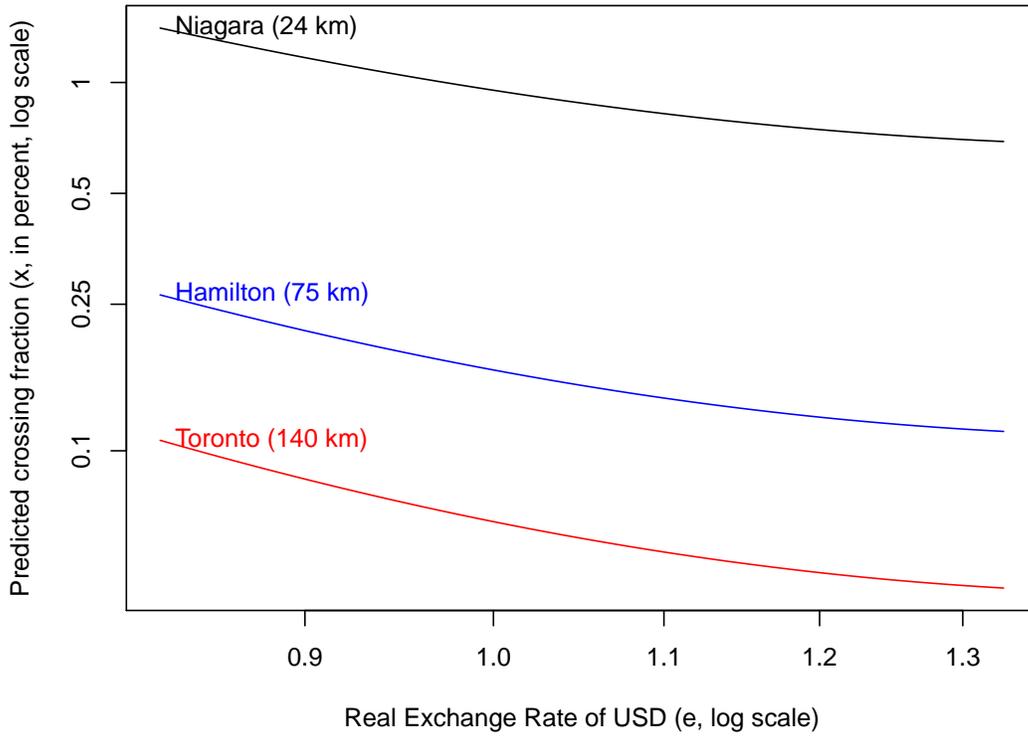
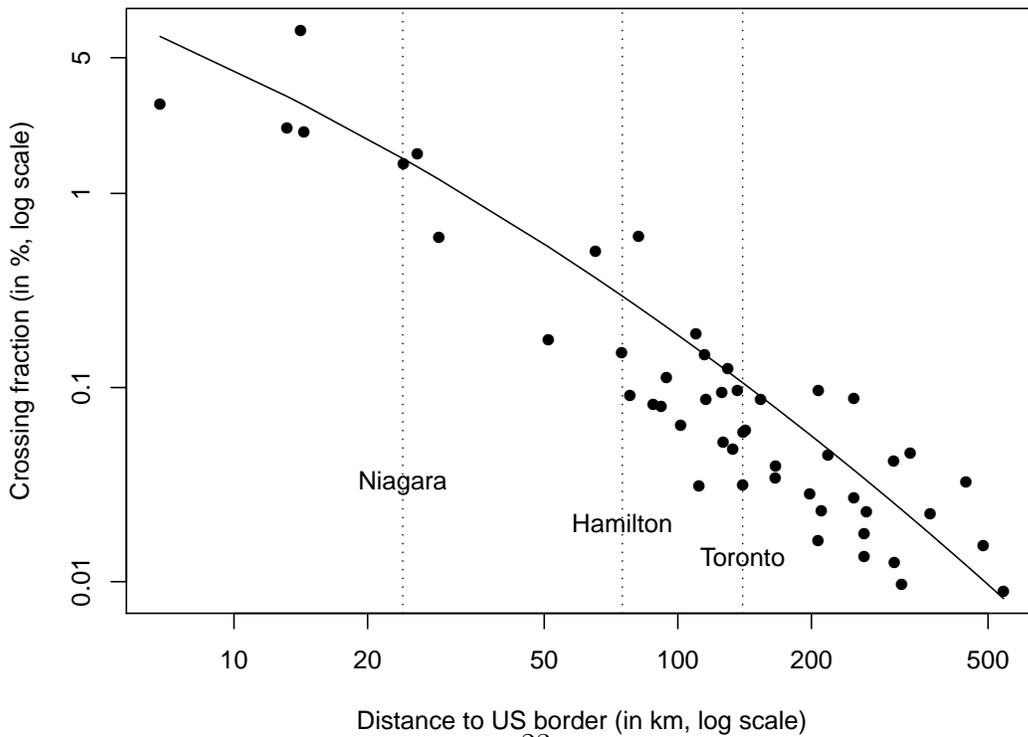


Figure 4: The crossing fraction declines with distance to the border



CD use most frequently. It has the disadvantage of using the geographic center of the CD as the origin point, which exaggerate distances severely for some large Divisions.

Table 6: Alternative specifications of regression and travel costs

Method: Stay:	Log Odds (OLS)		Fractional Probit			
	Daytrip	Overnight	Daytrip	Overnight	Daytrip	Overnight
θ_0 : constant	25.40 ^a (3.14)	-2.28 (1.87)	5.07 ^a (1.82)	-5.22 ^a (1.01)	10.33 ^a (2.47)	-4.61 ^a (1.08)
θ_1 : $\ln e_t$	-1.55 ^a (0.25)	-2.00 ^a (0.15)	-0.65 ^a (0.13)	-0.75 ^a (0.12)	-0.65 ^a (0.13)	-0.75 ^a (0.12)
θ_2 : $(\ln e_t)^2$	3.73 ^a (0.73)	0.20 (0.63)	0.93 ^a (0.33)	-0.15 (0.24)	1.03 ^a (0.32)	-0.16 (0.24)
θ_3 : \ln dist. or time	-1.14 ^a (0.07)	-0.28 ^a (0.04)	-0.89 ^a (0.06)	-0.19 ^a (0.02)	-0.56 ^a (0.05)	-0.14 ^a (0.02)
$\ln P(g)_{ct}$	-0.15 (0.13)	-0.42 ^a (0.08)	-0.05 (0.05)	-0.13 ^a (0.02)	-0.03 (0.05)	-0.13 ^a (0.02)
$\ln W_{ct}$	-2.41 ^a (0.30)	-0.19 (0.17)	-0.64 ^a (0.18)	0.26 ^a (0.10)	-0.94 ^a (0.24)	0.25 ^b (0.10)
Post-911	-0.25 ^a (0.06)	-0.18 ^a (0.04)	-0.13 ^a (0.03)	-0.14 ^a (0.03)	-0.12 ^a (0.03)	-0.14 ^a (0.03)
Observations	24232	33771	63000	63000	63000	63000
R^2	0.51	0.28	0.57	0.08	0.51	0.08
AIC	83374.53	93686.21	1792.09	635.49	1873.42	636.59

Standard errors clustered by census-division except cols. (1)–(2) where SEs also clustered by month-year. Regressions include month, province FEs. ^c $p < 0.1$, ^b $p < 0.05$, ^a $p < 0.01$. Driving time in cols. (3)–(4); port-use weighted average distances in cols. (5)–(6).

Our chief results on exchange rate and distance effects hold in all specifications. The positive second-order effect for exchange rates continues to hold for daytrips and is insignificant 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; much more so for daytrips than overnight ones.

There are a number of other robustness checks that we conducted, the results of which are contained in the supplementary file (see Table 5 in that document). We included a quadratic term for distance but it was not statistically significant nor did it

contribute significantly to the fit of the model. We also dropped observations where the drive times were extraordinarily long (more than 12 hours in one specification and more than 3 in another). The coefficients of the variables of interest in these specifications hardly change.

We examined whether commuters—residents of Canada who work in the United States—impact our results, since these travelers cross the border daily regardless of the exchange rate, and therefore are not the type of travelers that the model considers. Although commuters constitute less than 6% of travelers, they make up a disproportionate share of travelers in certain census divisions.⁴¹ We re-estimated the regressions dropping the census divisions where commuters made up 10% or more of travelers and found very similar results.

The real exchange rate and distance terms enter the crossing equation 13 additively. This suggests a simple falsification test. If the model is correctly specified, there should be no significant interaction between exchange rates and distance. When we add such an interaction term to the estimating equation, it is not statistically significant and it does not improve the R^2 relative to the equation implied by our model.

4.4 Implied travel cost estimates

One very useful way to evaluate our coefficients is to determine what they imply about travelers' willingness to trade off savings from cross-border shopping versus travel costs. Re-expressing the net benefits of crossing, $v_X - v_S$, using the parametric forms for $B(\ln e)$ and $\ln \tau(D)$ and setting $\zeta = 0$ we obtain

$$v_X - v_S = \beta_0 + \beta_1 \ln e + \beta_2 [\ln e]^2 - \gamma_0 - \gamma_1 \ln(D) - \gamma_2 \ln(P(g_e)/W_c).$$

Totally differentiating by e and D yields

$$d(v_X - v_S) = \frac{\partial(v_X - v_S)}{\partial e} de + \frac{\partial(v_X - v_S)}{\partial D} dD = 0.$$

Rearranging,

$$\frac{de/e}{dD/D} = \frac{\gamma_1}{\beta_1 + 2\beta_2 \ln e}$$

We do not observe β_1 , β_2 , or γ_1 but we do estimate $\theta_1 = \beta_1/\sigma$, $\theta_2 = \beta_2/\sigma$, and $\theta_3 = -\gamma_1/\sigma$. Plugging in these estimates, canceling out the σ , we obtain $(de/e)/(dD/D)$ as a function of the estimated parameters and the level of the real exchange rate. This calculation tells us the percent change in the real exchange rate required to compensate someone for a percentage increase in the distance or duration of the cross-border trip.

⁴¹The CD with the highest fraction commuters is Essex (35%), just across the border from Detroit. The next highest CD has just 13% commuters.

To obtain the change in expenditure, X , that would be required as compensation for the trip we note that expenditure in CAD is given by e times expenditure in USD. Holding USD-denominated expenditure constant, we have $dX/X = de/e$. We thereby arrive at the following formula for the travel cost:

$$\frac{dX}{dD} = \frac{-\hat{\theta}_3}{\hat{\theta}_1 + 2\hat{\theta}_2 \ln e} \left[\frac{X}{D} \right].$$

At the 2010 average real exchange rate of $e = 0.8846$, the first factor is given by -0.611 for distance (using $\hat{\theta}$ from column (3) of Table 5) and -1.02 for time (based on column (3) of Table 6). The second factor shown in brackets, X/D , is less straightforward to determine. We use the car-weighted median distance (or duration) of a round trip for daytrippers for D . This works out to 36 miles or 1.8 hours (including a 26 minute border wait in each direction). For X we use \$51, the 2010 median expenditure (in USD) of daytrippers who spent a positive amount, as calculated from the ITS.

Plugging in these values we obtain a travel cost of US \$0.87 per mile or \$29.69 per hour. These figures are in line with the \$0.89 per mile reimbursement rate for government travel within Ontario,⁴² and 2010 Canadian median hourly wages of US \$23.34 per hour.⁴³ Using means instead of medians for D (56 miles) and X (\$152) leads to travel cost estimates of \$1.66/mile and \$68.34/hour. These travel cost estimates are on the high end of the range reported in the literature on shopping within national markets.⁴⁴

Table 7: Travel cost estimates

Distribution	$d \ln e / d \ln D$	US \$/mile		$d \ln e / d \ln T$	US \$/hour	
		median	average		median	average
$\zeta(i) \sim \text{Normal}$	-0.611	0.87	1.66	-1.023	29.69	68.34
$\zeta(i) \sim \text{Logistic}$	-0.618	0.88	1.68	-1.124	32.63	75.10
$\zeta(i) \sim \text{Gumbel}$	-0.597	0.85	1.63	-0.946	27.47	63.23

The normality assumption for individual heterogeneity can be replaced with assumptions of logistic or Gumbel distributions. While each distributional assumption leads to different estimated coefficients, their relative values change very little. Thus,

⁴²See <http://www.njc-cnm.gc.ca/directive/travel-voyage/s-td-dv-a2-eng.php>. All Canadian dollar figures in this section were converted to USD using the 2010 average exchange rate of 1.03 CAD/USD.

⁴³See CANSIM Table 2820070.

⁴⁴Chiou and Muehlegger (2008) estimate that consumers would be willing to travel to a location 2.7 miles further away to save \$1 on cigarettes. This equates to a travel cost of 18.5 cents per mile. Manuszak and Moul (2009) estimate a marginal cost of around 50 cents per mile for consumers of gasoline in the Chicago area. Thomadsen (2005) estimates a travel cost of around \$1.50 per mile for consumers choosing fast food restaurants in Palo Alto.

we see in Table 7 that $-\hat{\theta}_3/(\hat{\theta}_1+2\hat{\theta}_2 \ln e)$ evaluated in 2010 ranges from -0.597 (Gumbel) to -0.618 (logistic), with the normal distribution in the middle. The monetary travel costs differ by only a few cents per mile. We are reassured that the results are not fragile to specific distributional assumptions. This is not to say that all distributions would yield the same results. However, it seems reasonable to infer that the travel cost results are unlikely to vary much so long as a bell-shaped distribution for heterogeneity is assumed.

4.5 Quantification: Crossing elasticities and crosser gains

An attractive property of structural estimation is the ability to conduct counterfactual exercises. Here we consider three scenarios: (i) a further 10% appreciation of the Canadian dollar, (ii) a doubling of wait times at the border, and (iii) re-play of history without the post-9/11 depression of travel. The first experiment is particularly useful because the fractional probit coefficients, like those in a binary probit, are not directly usable. The implied aggregate travel elasticities vary with the exchange rate and also depend on the geographic distribution of distances and incomes. Elasticities must therefore be obtained numerically as the aggregation of the predicted impacts in each census division-month combination.

Table 8: Counterfactual effects on same-day travel probabilities

Year:	RER +10%		Wait +100%	
	2002	2010	2002	2010
Canada	8.02	25.67	-57.08	-54.60
New Brunswick	6.33	19.92	-52.29	-49.10
Quebec	10.00	32.12	-55.77	-54.04
Ontario	7.94	25.47	-60.37	-57.33
Toronto (140 km)	10.78	34.35	-44.74	-42.84
Hamilton (75 km)	9.79	31.30	-53.72	-52.32
Niagara (24 km)	8.08	25.21	-64.16	-62.53
Manitoba	9.76	31.35	-53.42	-51.78
Saskatchewan	10.47	34.02	-53.31	-51.48
Alberta	11.41	37.81	-50.75	-49.23
British Columbia	8.31	25.88	-55.38	-53.48

Table 8 shows the effect of the first two experiments. Columns (1) and (2) show the impact, in two different years, on the number of cross-border trips from a decrease in a 10% decrease in e . This is 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 5. The years that we analyze are 2002 and 2010, when the Canadian Dollar was at its weakest ($e = 1.30$) and strongest ($e = 0.88$), respectively, against the US Dollar, in the last 50 years.

Table 8 reveals differences in the implied exchange rate elasticities across locations and time. Comparing the three Ontario census divisions, elasticities are larger for those communities located further from the border. It appears that the elasticities are also larger in provinces where most of the population is far from the border. As Figure C.2 in the Appendix shows, Alberta, Saskatchewan and Manitoba (three provinces with high elasticities) have relatively few inhabitants located at or very close to the border. This finding is consistent with our earlier discussion of equation 10. In other words, higher distances to the border lower the proportion of crossers in a region x_R and dominates any possible opposing effect arising from the curvature of the function F . At a given point in time, an appreciation of the RER shifts up the benefits of crossing for all census-divisions and therefore for all provinces, leading to proportional rises in the elasticities from 2002 to 2010. The elasticities rise due to the convex relationship between the crossing benefits and the log RER.

The implied crossing elasticities can be compared to those obtained in the trade literature to gain perspective on the extent of consumer arbitrage. When the Canadian dollar is at its weakest (2002), the Canada-wide elasticity of 0.80 (first row of Table 8) is almost the same as the average elasticity of 0.81 of that Blonigen and Wilson (1999) estimate for Canada-US trade in goods. At the strongest levels of the RER, elasticities for travel that are three times as large as those observed for goods. One reason why travel could be more elastic is that travelers can alter their border crossing decision immediately as relative prices change, whereas traders have to make various up-front investments in marketing, distribution, and logistics.

Columns (3) and (4) of Table 8 show the effect of increasing wait times at the border. We use the specification from Column 3 of Table 6. This specification had assumed a wait time of 26 minutes at the border. In our counterfactual experiment we double this to 52 minutes.⁴⁵ 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 effects of the increased wait times are in the provinces of Alberta and New Brunswick. These provinces do not have large cities close to the border. Since the wait time is incurred by all travelers, those driving longer distances pay a proportionately lower cost. By contrast, a province such as Ontario 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.⁴⁶ The welfare losses to the average crosser from increased wait times range from under

⁴⁵Note 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.

⁴⁶See Figure C.2 in the Appendix to understand the differences in the geographical distribution of population across Canadian provinces.

2% for Alberta to almost 10% for Niagara county. The predicted impacts of delay do not vary much over time since the effect of travel costs is independent of the value of the RER in the net benefits function.

The structural approach has the additional advantage that the impact of changes can be expressed in terms of percent changes in surplus accruing to the average traveler. Due to heterogeneity in travel costs, all crossers obtain weakly positive surplus. For a community with mass N_c of potential monthly crossers, aggregate surplus is the integral over individuals for whom $\zeta > \zeta_c^*$:

$$G_c = N_c \int_{\zeta_c^*}^{\infty} [B - \ln \tau_c + \zeta] dF(\zeta). \quad (17)$$

Integrating equation (17), G_c can be expressed as the product of two factors:

$$G_c = \underbrace{(B - \ln \tau_c + \mathbb{E}[\zeta \mid \zeta > -B + \ln \tau_c])}_{\text{Average crosser's gain}} \underbrace{F[B - \ln \tau_c] N_c}_{\text{Number of crossers}}. \quad (18)$$

To a first approximation, the percentage change in crosser welfare brought about by a change in the determinants of $B - \ln \tau_c$ will be given by the sum of the percentage changes in the number of crossers, n_c , and the average gain each crosser expects to obtain, G_c/n_c . We therefore quantify these components separately. The difference between their sum and the total welfare effect is negligible in the experiments we conduct.

With ζ distributed $\mathcal{N}(\mu, \sigma^2)$, we can compute the average crosser's gain as

$$G_c/n_c = (B - \ln \tau_c) + \mu + \sigma \frac{\phi[(\mu + B - \ln \tau_c)/\sigma]}{\Phi[(\mu + B - \ln \tau_c)/\sigma]} = \sigma \left(\mathbf{Z}\hat{\boldsymbol{\theta}}_c + \frac{\phi[\mathbf{Z}\hat{\boldsymbol{\theta}}_c]}{\Phi[\mathbf{Z}\hat{\boldsymbol{\theta}}_c]} \right),$$

where \mathbf{Z}_c is the vector of explanatory variables and $\boldsymbol{\theta}$ is the coefficient vector. The second equality comes from $(B - \ln \tau_c + \mu)/\sigma = \mathbf{Z}\hat{\boldsymbol{\theta}}_c$ (the prediction index obtained from the fractional probit regressions). Without being able to identify σ , levels of G_c/n_c cannot be determined but we can determine the percentage change resulting from any contemplated change in the \mathbf{Z}_c vector.⁴⁷ To quantify the aggregate effect of policy changes, it is necessary to aggregate over the effects at each census division, multiplying by N_c to give greater weight to larger divisions.

The model indicates that the home appreciation gives rise to aggregate gains of 28.20% in 2010. Most of this, 25.67%, comes from increased propensity to cross. Welfare changes for the average crosser contribute 2.22%.⁴⁸ The gains to the average crosser are approximately three times as high when the appreciation starts from an

⁴⁷This means that we cannot quantify the monthly welfare gains for community c relative to staying at home. All policy change exercises capture relative gains and can be applied to the daily or monthly as well as individual or collective welfare gains.

⁴⁸The remainder, 0.31%, is attributable to the weighted product of the changes.

already strong Canadian dollar.⁴⁹ The biggest percentage gains to the average crosser are obtained in census divisions close to the border, with Niagara crossers gaining 2.6% from the 10% home appreciation.. Increasing delays would lower average crosser gains by 4.6% in 2010 in Canada. Larger losses would occur at communities along the border where the wait constitutes a higher share of total trip length. In Niagara, for example, doubling wait times would lower average crosser gains by 9.7%.

One final counterfactual we consider is to “turn off” the estimated 9/11 effect. As we reported earlier, the post-9/11 period had a 32% reduction in same-day crossings relative to what the model would have predicted based on the evolution of the real exchange rate, gas prices and incomes. The average crosser incurs a 3.4% reduction in welfare.

4.6 Reconciling reduced-form and structural estimates

We now return to a key result obtained in Section 2: the elasticity of crossings with respect to the RER is about 25% lower for Americans than for Canadians. There may be a number of possible reasons for this asymmetry. First, there may be differences in the density of retail networks on each side of the border. Second, the variety of goods available on each side of the border may differ. Third, ...Fourth, the difference may arise from differences in the geographic distribution of US and Canadian residents.

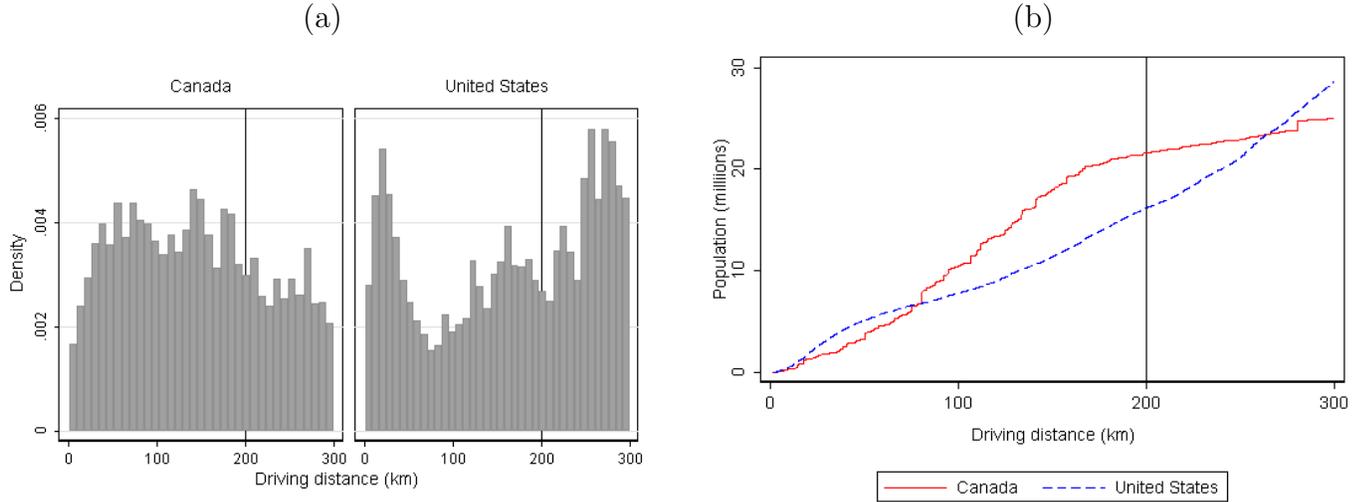
Some of these factors, and their effects on travel, are difficult to directly test, especially as they are outside of the model of Section 3. However, we can use our structural estimates to test the last of these possible factors, namely that differences in population distribution may be partially responsible for the observed lower crossing elasticities of US residents. Even though we do not have data on the geographical distribution of US crossers, we can use our structural estimates to simulate cross-border travel by Canadian residents in the event that their geographic distribution resembled that of the US population that is most likely to make cross-border trips for shopping purposes.

For this purpose, we use US population and driving distances at the census tract level. It is necessary to impose a cutoff distance of US census tracts to the border, in order for the set of included census tracts to generally resemble the Canadian census divisions that are likely to have same-day crossers. Otherwise, in principle, the simulation would include crossing elasticities for places as far from the Canadian border as Texas or Florida, which would greatly affect the results, despite the very low predicted crossing fractions at such distances. We therefore restrict the sample of US census tracts to those within 200 km of the border, since this distance bound contains about 97.5% of Canadian same-day crossers.⁵⁰ For each US census tract we can

⁴⁹As mentioned in Section 3, the benefits from crossing could be limited by car size constraints that prevent the crosser from taking full advantage of lower prices.

⁵⁰We do not use the distance corresponding to 100% of same-day crossers, since a small number of crossers report traveling implausibly large distances each way for a same-day trip. Canadian

Figure 5: Population and Distance to the Border



compute the predicted crossing probability, corresponding to estimating equation 15. We then conduct a counterfactual exercise similar to Section 4.5 where we increase the exchange rate by 10% in order to calculate elasticities.

Figure 5 shows the differences between the US and Canada in terms of population density and distance to the border.⁵¹ Panel (a) shows 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 figure shows that the northern US (within about 200 kms of the border) is generally less densely populated than a similar distance cutoff in Canada. This is the case for any distance cutoff greater than 70 kms.⁵² These different distributions and population densities can affect crossing elasticities as shown in Equation 10.

The comparison of Canadian and (simulated) US elasticities is shown in Table 9. We calculate these elasticities for 2002 and 2010, in order to correspond to Table 8, as well as for 2005, which had an intermediate value of the RER. We present elasticities corresponding to each of the three specifications from Table 5. Note that the elasticities for Canada in 2002 and 2010 using the column 3 specification are the same as those reported in Table 8. The elasticities for the US use the distribution of population across US census tracts, but applied to Canadian data on incomes and gas prices, and using the coefficients estimated on the Canadian population in Table

populations at this distance are low enough for this not to affect the results, but US populations at the same distance are very high. Note that using a lower distance cutoff strengthens the results that follow.

⁵¹The figures were constructed by calculating the driving distance from each census tract to the closest land border. Details are provided in Appendix A.2.

⁵²Although Canada generally has a larger population close to the border, two large US cities—Buffalo and Detroit—are located immediately on the border, unlike any similar sized city in Canada.

6. Even though the column 3 specification is preferred for the structural estimation, it is not necessarily the best specification to use for this exercise. This is because using either province fixed-effects or province level income and gas prices, which are included in the specifications of Columns 2 and 3, requires assigning US census tracts to Canadian provinces. It is not completely obvious how to do this, and therefore any mis-specification may affect the results.⁵³

Table 9: Counterfactual travel elasticities, with simulated US data

Table 6 Specification	2002 ($e = 1.3$)		2005 ($e = 1.01$)		2010 ($e = 0.88$)	
	Canada	US	Canada	US	Canada	US
Column 1	7.11	6.08	12.05	10.22	14.80	12.44
Column 2	6.46	5.65	23.82	20.55	34.23	29.13
Column 3	8.02	7.30	19.25	17.36	25.67	22.95

Canadian elasticities calculated using Census divisions as in Table 8. US elasticities simulated using Census tract populations, with estimated coefficients from Table 5.

The results of Table 9 suggest that changing the distribution of population in Canada to more closely reflect that of the (less densely populated) northern US would lower the elasticity of crossings with respect to the RER. In each year, and given any of the three specifications of Table 5, the elasticity using US population data is lower than using Canadian data.⁵⁴ In the most conservative estimate—that of Column 3—the simulated elasticities are about 10% lower using the US population distribution. In Table 2, (levels specification) Americans have 25% lower elasticities. In other words, using the US population distribution explains at least 40% of the difference in elasticities between Canadians and Americans.

5 Conclusion

On average, each person living within a three hour drive of the Canada-US border makes more than one cross-border car trip per year.⁵⁵ In this paper we develop and estimate a model of cross-border travel. In line with the shopping motive in our model, US-Canada border crossings are heavily influenced by exchange rate changes. Furthermore, the elasticity of crossings increases with the strength of the domestic currency, as predicted by expansion of the extensive margin of purchases. Consistent with the literature documenting pricing-to-market across borders, two forces prevent prices from fully converging in the two countries. First, consumers face large marginal

⁵³We assigned each US census tract to the Canadian province which is across the border from the closest port to that census tract.

⁵⁴We only present these three years in the Table. However, the simulated US elasticities are lower in each of the 21 years in our sample, and across all three specifications.

⁵⁵22 million Canadians and 24 million Americans reside in this region.

travel costs. Our estimates range between \$30 and \$68 per hour (or \$0.87 and \$1.66 per mile). Second, individuals are heterogeneous. While the majority of Canadian crossers live less than 18 miles from the border, the majority of Canadian residents are more than 81 miles away. We use our structural estimates to show that asymmetries in the geographical distribution of the population in the two countries can partially explain the differences in the sensitivity of US and Canadian travelers to the exchange rate.

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Appendices

A Theoretical derivations

Supply-side determination of price deviations

The Dornbusch, Fischer, and Samuelson (1977) model implies prices (in local currency) are given by $P(z) = a(z)W$ and $P^*(z) = a^*(z)W^*$. In DFS the $a(z)$ and $a^*(z)$ are unit labour requirements and product markets are perfectly competitive. For our purposes, the $a(z)$ could be the product of the cost parameter and a constant good-specific markup (such as would occur in the Dixit-Stiglitz monopolistic competition model).

Utility is $\ln U = \int_0^1 b(z) \ln C(z) dz$, where $C(z)$ denotes consumption of good z . With Cobb-Douglas preferences the natural definition of the price indexes are $\bar{P} = \exp(\int_0^1 b(z) \ln P(z) dz)$ and $\bar{P}^* = \exp(\int_0^1 b^*(z) \ln P^*(z) dz)$. The ratio of the domestic to foreign price index is given by $\bar{P}/\bar{P}^* = W/(W^*\kappa)$, 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. Relative price deviations are determined entirely in terms of exogenous parameters: $\delta(z) = \kappa a(z)/a^*(z)$. Hence under the DFS supply side assumptions, $\delta(z)$ is not influenced by the exchange rate.

Single-good model

Suppose instead of there being a continuum of goods which are available on both sides of the border there is only a single product that potential travellers are deciding where to buy. Maintaining Cobb-Douglas, consumers spend a fixed share b of their income on this product ($1 - b$ goes to items such as rent and taxes that are only purchased in the country of residence). This could be an all-inclusive holiday at a ski resort, for example. Let local currency prices be P and P^* . Let $F(\zeta)$ be the CDF of the difference in perceived quality of this good between the foreign and domestic version and, as before, τ_c is the iceberg travel cost. The indifferent potential crosser has $\zeta_c^* = b \ln P - b \ln(E P^*) - \ln \tau_c$. Assuming relative prices of this product are proportional to the ratio of CPIs ($P/P^* = a\bar{P}/\bar{P}^*$), the fraction who cross is

$$x_c = \mathbb{P}(\zeta > \zeta_c^*) = F(b \ln a - b \ln e - \ln \tau_c).$$

This model predicts a coefficient of zero for $[\ln e]^2$. Moreover, since F'/F is decreasing in its argument for the distributions of $F()$ used in fractional models (logit, probit, gumbel), the elasticity of crossings with respect to crossings will tend to *diminish* in absolute value with the strength of the home exchange rate, the opposite of our finding in section 2.

B Data construction

B.1 Crossing fractions

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 restrict 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 (1990–2010).

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, Statistics Canada 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).⁵⁶ 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.⁵⁷ 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, $N_c = \text{Pop}_{ct} \times \text{CPC}_c \times 30$. Census division populations, Pop_{ct} , are available annually from Cansim Table 051-0034, provided by Statistics Canada. Car registration

⁵⁶A 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.

⁵⁷We 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.

data used for generating CPC_c come from Statistics Canada publication 53-219-XIB (“Road Motor Vehicle Registrations 1998”).

B.2 Driving distances and times to the border

We calculate the distance from each Canadian Census Division (similar to a US county) to the nearest ports D_c using two methods. The primary method takes advantage of geographically detailed information at the level of Census Subdivisions (similar to US Census Tracts). The 250 CDs have an average of 20 subdivisions. We obtained Subdivision centroid information from the Standard Geographical Classification of 2001 and used Google’s driving distance application to measure the road distance and time from each centroid to the nearest crossing port. We obtained two measures: the median and the average distances for each CD. These two metrics are very similar for the majority of CDs except for two CDs in Ontario where the average distance is heavily influenced by outlier (low population and high distance to the border) subdivisions. We therefore used medians in our estimations. The results using averages do not differ much in terms of exchange rate or distance elasticities but the province and income effects are influenced by the two outliers.

The secondary method of calculating distances (employed in columns (5) and (6) of Table 6) takes into account the fact that crossers from a given census division do not always use the same port. At the CD level, we know shares of crossers from each CD that cross at 102 different ports. We use the average shares of the top 5 ports over the 1990 to 2010 period to construct weighted average distance and time from the CD’s geographic centroid. This measure generates several outliers in large CDs that have centroids that are far from the border but populations that are concentrated close to the border.

B.3 Prices, exchange rates, and incomes

Exchange rates obtained from Pacific Exchange Rate Service (fx.sauder.ubc.ca). The US Consumer Price Index is the US city average for all items and all urban consumers, not seasonally adjusted (Series ID CUUR0000SA0 from bls.gov/cpi#data). Canadian prices are from CANSIM Table 3260020, 2009 basket, all items. We choose July 1993 as the base period because in that month the nominal exchange rate was equal to the annual purchasing power parity rate provided by the OECD and thus the RER was approximately 1. Prices for regular unleaded gasoline at self service filling stations are obtained from CANSIM Table 3260009 for a major urban centre for each of the border provinces. We obtained median household income from the CHASS Canadian Census Analyser for the years 1991, 1996, 2001, and 2006. We linearly interpolated and extrapolated around July of each census year to obtain the monthly data from 1990 to 2010.

C Additional Figures

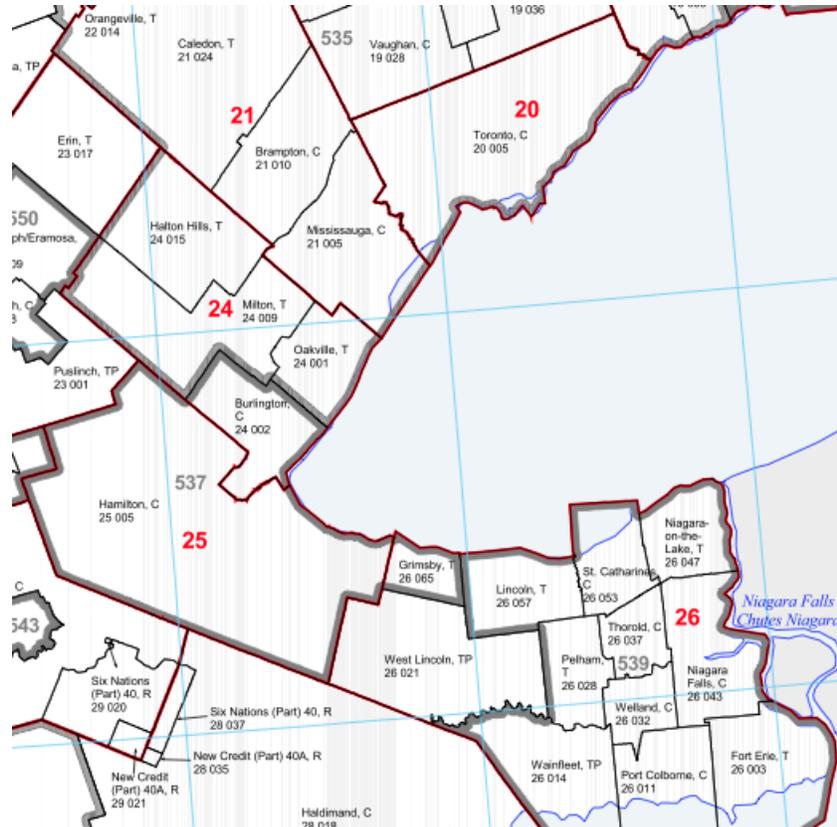


Figure C.1: Census Divisions in Southeastern Ontario

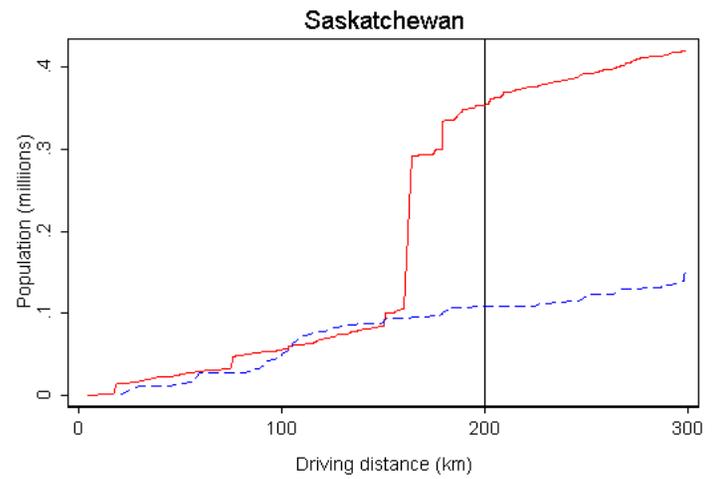
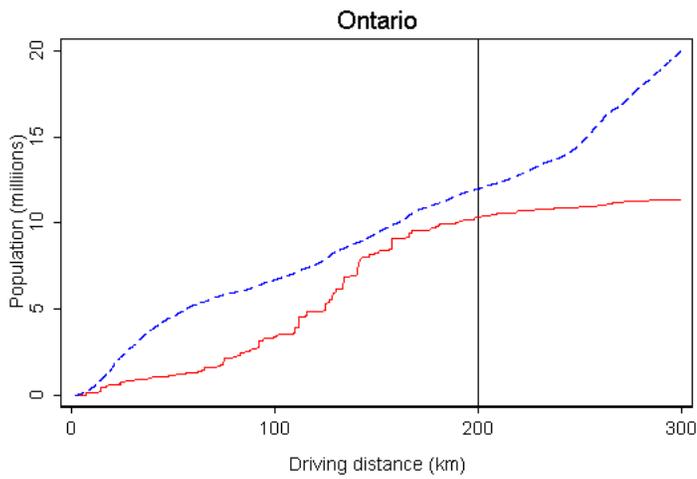
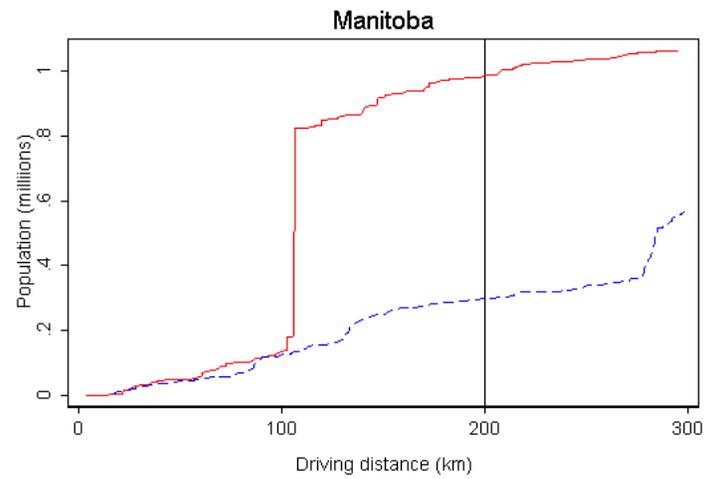
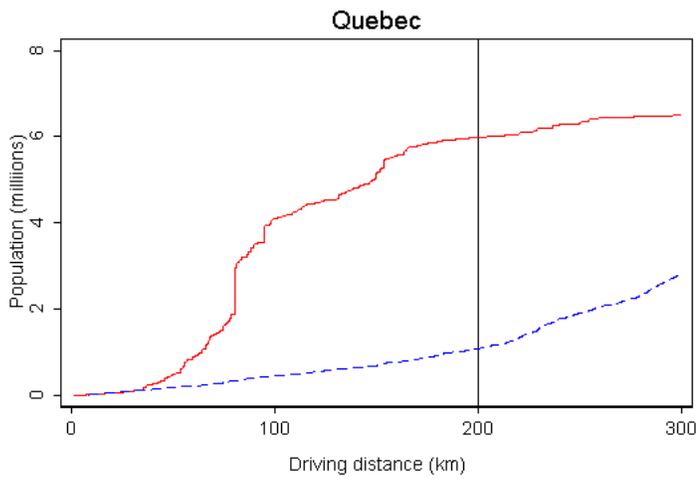
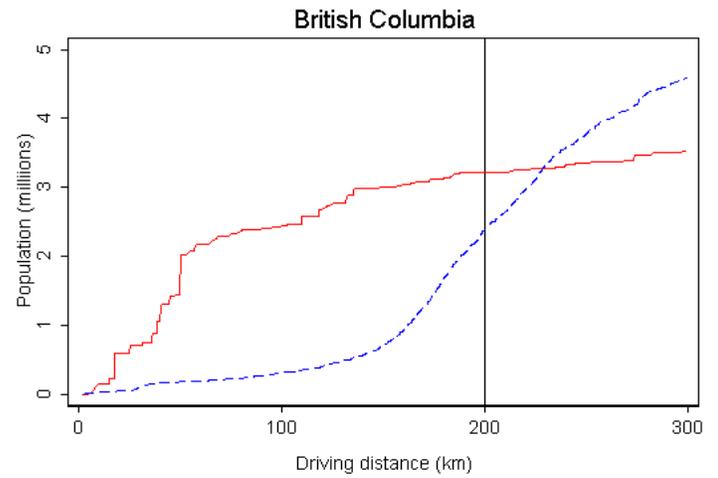
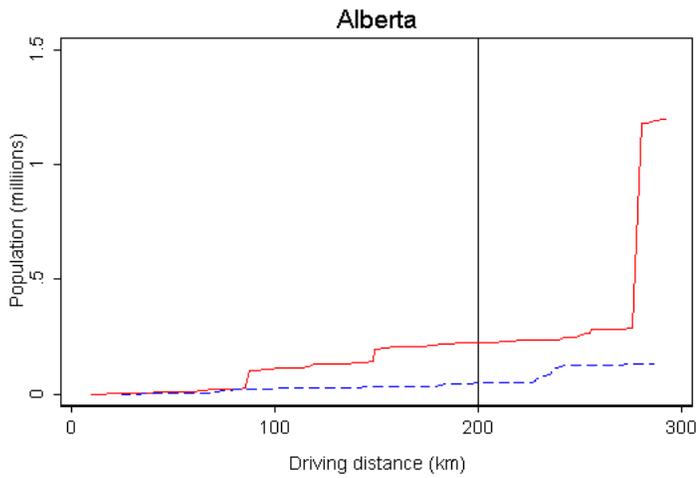


Figure C.2: Accumulated Population and Distance to the Border: Canada (solid) and US (dashed)