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. We estimate the model using microdata on Canada–United States travel. Price differences motivate cross-border travel; a 10% home appreciation raises the propensity to cross 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 per mile. Geographic differences can partially explain why American travel is less exchange rate responsive.

I. 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. Far 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 U.S.-Canada border in 2010, about 250 times the number of permanent migrants arriving from Canada or the United States in that year.[1] Understanding travel patterns has important implications for taxation, infrastructure planning, securing borders, and controlling the spread of infectious diseases.

Given the magnitude and policy relevance of cross-border travel, it has attracted surprisingly little formal analysis by economists. One possible explanation is the perception that travel motivations are noneconomic. Since travelers by definition return to their country of residence, this normally precludes earning income in the visited country. We must therefore look to motivations other than labor supply to explain most international travel.[2] This paper estimates a model in which the consumption motive drives short-term travel between Canada and the United States.

After establishing stylized facts using 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. Whereas models of the migration decision view potential crossers as maximizing earnings, we offer instead a model in which consumers seek to minimize expenditures.[3] 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 U.S.-Canada trade in goods.

The consumption motive for travel is predicated on the existence of price differences. Evidence of such differences on either side of the Canada-U.S. border has been convincingly demonstrated by a series of papers. Engel and Rogers's (1996) study of price dispersion between cities in Canada and the United States reports that crossing the border is equivalent to a distance of 1,780 miles. While this estimate of the border's width has been challenged by Gorodnichenko and Tesar (2009), empirical studies consistently find price differences.[4] Recent work has compared disaggregated price data for identical goods on both sides of the Canada-U.S. border at stores owned by the same large retailer. Burstein and Jaimovich (2009) find substantial amounts of pricing to market. Gopinath et al. (2011) also find evidence of market segmentation, including a 24% discontinuity in grocery prices at the border. Price differences on identical products are not unique to grocery products or to the Canada-U.S. border. Boivin, Clark, and Vincent (2012) show that even online book prices differ greatly between the United States and Canada and that their prices do not respond to exchange rate movements. Goldberg and Verboven (2004) compare prices for the same car model in different European countries and report price ranges of about 35% for the majority of models in the period before the euro was introduced.

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, Friberg, and Wilander (2007) infer cross-border shopping for alcohol between Sweden and Denmark by observing how retail sales respond to changes in taxes and exchange rates. Manuszak and Moul (2009) examine how differences in gasoline and cigarette taxes create incentives to cross U.S. state borders and thereby calculate consumers' travel costs. Inferring interstate travel from variation in lottery revenues

Received for publication December 9, 2011. Revision accepted for publication March 25, 2013. Editor: Gordon Hanson.

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An earlier version of this paper was titled "Consumer Arbitrage across a Porous Border." We thank participants at seminars at Cornell, Penn State, Ryerson, University of British Columbia, University of Toronto, UTDT, UdeSA, and the 2011 RMET conference for helpful comments. Suggestions by Pinelopi Goldberg, Gita Gopinath, Andy Neumeyer, and Andrés Rodríguez-Clare, Gordon Hanson, and three anonymous referees proved particulary useful. This research was supported by the Social Sciences and Humanities Research Council.

A supplemental appendix is available online at http://www.mitpressjournals.org/doi/suppl/10.1162/REST_a_00404.

[1] Statistics Canada and Migration Policy Institute.

[2] For the United States and, Canada commuters constitute only 5% of day trips.

[3] See Grogger and Hanson (2011) for an estimated model in which income maximization drives the decision to emigrate.

[4] Goldberg and Knetter (1997) summarize the earlier literature.

The Review of Economics and Statistics, October 2014, 96(4): 648–661

doi:10.1162/REST_a_00404

per resident, Knight and Schiff (2012) find that higher jackpots induce cross-border purchases in small, densely populated states. To our knowledge, Chiou and Muehlegger (2008) is the only paper that uses direct data on travel across borders to estimate responses to price differerences (caused by cigarette taxes). Moving beyond studies of individual goods, 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 U.S.-Canada border. Their results are consistent with cross-border shopping behavior, but they do not estimate travel responses directly.

Our paper proceeds in three steps. First, we use reduced-form regressions in section II to establish that travelers respond strongly to the economic incentives created by fluctuations in the exchange rate. This finding corroborates results from reduced-form estimations conducted by 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 the exchange rate than U.S. travel. Moreover, for residents of both countries, these elasticities increase in absolute value as the home currency strengthens. To make sense of these findings and allow investigations of counterfactuals, our second step is to develop a model of the decision to cross based on the premise that travelers seek bargains on the other side of the border. The model presented in section III naturally generates the prediction that the elasticity of crossings rises in absolute value as the home currency strengthens. The third step is to use a new data set with information on the residence of cross-border travelers to estimate the parameters of the model. The strong travel responses we estimate imply that the markets are not perfectly segmented. However, travel costs prevent the arbitraging away of all price differences. Our estimated coefficients imply that the median crosser requires savings of 87 cents per mile traveled. As a consequence, shopping-motivated travel is concentrated among the population living close to the border. Indeed, the median day tripper in Canada lives 18 miles from the border, whereas the median Canadian lives 81 miles away.

The model also permits counterfactual experiments. We show that an exogenous doubling of border wait times would lower crossing frequencies by 50% to 60%. 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 changes in the explanatory variables. We find that a 10% appreciation from current rates would yield average crosser gains of 2.1%, whereas the consequences of 9/11 have lowered these gains by 3.4%. We also show that differences in the geographic distribution of residents partially explain the difference in the Canadian and U.S. travel elasticities.

II. Stylized Facts of Border Crossings

In this section we describe our cross-border travel data and establish the main relationships between exchange rates and travel between Canada and the United States.

A. Data

Statistics Canada provides data on cross-border travel using information collected by the Canadian Border Services Agency (CBSA).[5] These data consist of counts of all vehicles entering Canada at all land crossings with the United States. U.S. residents encounter the CBSA on their outbound journey and Canadian residents on their return journey.

We use these data on vehicle counts for the seven Canadian provinces that share a land border with the United States: British Columbia, Alberta, Saskatchewan, Manitoba, Ontario, Quebec, and New Brunswick. We use monthly data on passenger vehicles for the calendar years 1972 to 2010. The counts are broken down by travelers' country of residence, which is determined by whether the vehicle has U.S. or Canadian license plates, and by the length of the cross-border trip. We analyze same-day and overnight (here defined as trips spanning two or more days) trips separately.

We obtained monthly average data on the spot market exchange rate between the U.S. and Canadian currencies. Using data on monthly CPIs for both countries, we construct the real exchange rate (RER) for each month.[6] It is defined with U.S. prices in the numerator such that RER increases correspond to Canadian dollar depreciations. We fixed the absolute level of the RER using relative price levels from OECD data.

Figure 1 displays key temporal patterns in the data.[7] Figure 1a shows the ratio of monthly same-day trips by residents of the two countries from 1972 through 2010. The solid black line shows the real exchange rate. The dashed black 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 1 (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 RER variation is nominal variation.[8] U.S. trips rise relative to Canadian trips (the thick gray line) when the United States has relatively high price levels. Since the 1980s, the relationship between relative travel and the exchange rate has become very strong.

Travel is highly seasonal for residents of both countries. Figure 1b shows average travel over the 38-year period for each calendar month. Cross-border travel rises with average

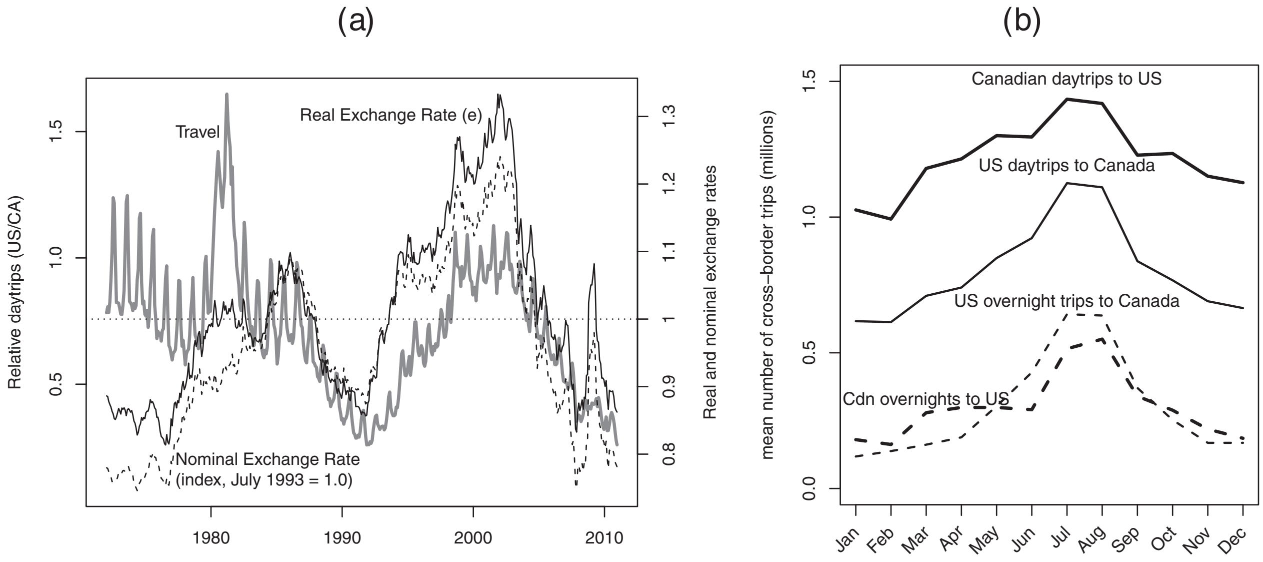
[5] See Cansim Table 427-0002.

[6] The consumer price indexes include sales taxes. Data sources and other details are provided in the online appendix.

[7] Table 1 in the online appendix to this paper presents summary statistics for these data.

[8] More precisely, log first differences of the nominal exchange rate explain 94% of the variation in log first differences of the real exchange rate over the period 1972 to 2010. In levels, the R^2 is 0.89.

FIGURE 1.—ANNUAL AND MONTHLY VARIATION IN CROSSINGS



temperatures, peaking in July and August, for all groups. One interpretation, to which we return when we estimate the model, is that cold weather raises travel costs. Figure 1b also serves to show relative magnitudes of the different travel categories. Over the 1972–2010 period, Canadian residents averaged about 50% more day trips across the border than U.S. residents. Overnight trips have similar monthly means for north- and south-bound travelers.

We employ the CBSA data in the reduced-form regressions that follow. However, for the model we estimate in section IV, 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 based on questionnaires filled out by travelers returning to Canada. It collects information on the nature and purpose of the trip, the dates on which travelers exited and entered Canada, the census division in which the travelers reside, and the ports used to cross to the United States. We retain data on Canadian residents returning from the United States by car between 1990 and 2010.[9] The ITS data contain 63,000 observations, each corresponding to a Canadian census division in a given month. Summary statistics of the ITS data are reported in table 2 in the online appendix.

The ITS data 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.[10] This suggests that the exchange rate potentially plays an important role in the decision to cross the border. We now attempt to quantify the relationship between exchange rates and cross-border travel.

B. 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 RER. Our goal is to 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

$$\ln n_{it} = Month_t + Province_i + \eta_1 \ln e_t + \eta_2 post911_t + \eta_3 t + \eta_4 t^2 + \varepsilon_{it}, \quad (1)$$

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 United States in a given province and month. For the United States, 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 RER to crossing decisions.

[9] We do not use information on U.S. residents since the only information on their place of residence within the United States is the state in which they live. This level of aggregation is too coarse to provide meaningful information on their distance to the border.

[10] In table 3 in the online appendix, we tabulate the commonly stated motives for crossing the border using ITS data on visitors and returning residents to Canada.

TABLE 1.—REGRESSION OF LOG CROSSINGS, 1972–2010

Length of Stay:	Day Trip		Overnight		Day Trip		Overnight	
Residence:	United States	Canada	United States	Canada	United States	Canada	United States	Canada
$\ln e$	1.24***	−1.62***	0.47***	−1.78***	0.93***	−1.71***	0.32	−2.08***
(CAD/USD)	(0.17)	(0.24)	(0.17)	(0.17)	(0.28)	(0.28)	(0.23)	(0.21)
$\ln e \times [e > 1.09]$					0.90**	0.54*	0.83***	0.65**
(strong USD)					(0.37)	(0.33)	(0.31)	(0.29)
$\ln e \times [e < 0.90]$					−0.87**	−0.87***	−1.25***	−0.31
(strong CAD)					(0.34)	(0.24)	(0.32)	(0.22)
R^2	0.98	0.98	0.96	0.97	0.98	0.98	0.97	0.97

Newey-West standard errors in parentheses are robust to serial correlation out to 60 months. Significant at *10%, **5%, ***1%. An observation is a province-year-month. $N = 3,276$. Regressions include month and province fixed effects, a post-9/11 indicator, and trend variables.

This assumption is defensible because demand for foreign currency created by U.S. 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 \$4.2 billion in the United States while Americans spent \$1.8 billion in Canada during the first quarter of 2010.[11] This represents a mere 0.04% of the foreign exchange turnover involving the Canadian dollar.[12]

The results of estimating this equation are presented in table 1. We treat each province in a calendar month as a separate observation. Since monthly crossing data are serially correlated, we use Newey-West standard errors. The results in the first four columns indicate that travelers respond to the exchange rate, as represented in the negative elasticity of Canadian residents and the positive elasticity of U.S. residents with respect to the RER. In addition, the elasticities of Canadian residents are bigger than those of U.S. residents for same-day and overnight trips.

In columns 5 to 8, we investigate whether the crossing elasticity with respect to exchange rates varies with the level of the exchange rate. 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 U.S. dollar was strong is generally positive, for residents of both countries. This has the effect of increasing the positive elasticity of U.S. residents and decreasing the negative elasticity of Canadian residents. In other words, U.S. residents become more responsive to the exchange rate in periods when the U.S. dollar is strong, while Canadian residents become less responsive. We observe the opposite pattern during periods when the U.S. dollar is in its lowest quartile.[13]

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 larger when the home currency is stronger.

III. Model of the Crossing Decision

Potential cross-border shoppers must decide whether it is worth incurring travel costs to obtain shopping benefits. The benefits 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 RERs 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, with both prices expressed in local currency units. Let E represent the nominal exchange rate defined in the home currency unit per foreign currency unit. Define $\bar{P}$ and $\bar{P}^*$ as the domestic and foreign consumer price indexes. The RER, which indicates the relative price of the foreign consumption bundle expressed in a common currency, is given by $e = E\bar{P}^*/\bar{P}$. Finally, we define $\delta(z)$ as the relative price deviation of good z:

$$\delta(z) = \frac{P(z)/\bar{P}}{P^*(z)/\bar{P}^*}. \tag{2}$$

We order goods such that $\delta'(z) > 0$ and assume that relative price deviations are invariant to the RER, that is, $\partial\delta(z)/\partial e = 0$.[14]

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 RER back into equation (2), 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 $\tilde{z} = \delta^{-1}(e)$,

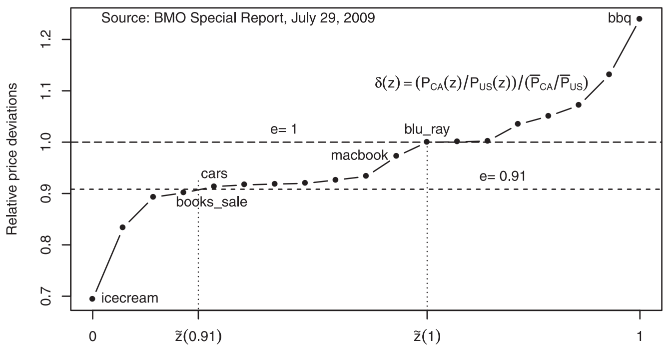
[11] This includes expenditures by air travelers. International Travel Account Receipts and Payments, http://statcan.gc.ca/daily-quotidien/100827/dq100827-eng.pdf.

[12] Authors' calculations from the BIS Central Bank Survey of Foreign Exchange and Derivatives Market Activity, 2010, http://www.bis.org/publ/rpfxf10t.htm.

[13] Tables 4, 5, and 6 in the online appendix show that these results are robust to using country-level data, taking first differences of equation (1) and adding economic indicators as regressors.

[14] The online appendix shows that this result is implied by the DFS supply-side assumptions.

FIGURE 2.—EXCHANGE RATES AND RELATIVE PRICES: NINETEEN PRODUCTS



with $\partial\tilde{z}/\partial e = 1/\delta'(\tilde{z}) > 0$. Thus, a real appreciation of the foreign currency contracts the range of goods that are cheaper in the foreign country.

Figure 2 illustrates the model using data from Porter (2009), who reports prices for nineteen goods available on both sides of the border. Calculating $\delta(z)$ as the ratio of the Canadian price (in CAD) to the U.S. price (in USD), all divided by the relative price level (1.2, obtained from the OECD), we plot it against z. The study was conducted when 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, fifteen of nineteen goods were less expensive in the United States 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 $e = 1$.

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 pricing-to-market by home retailers, consumers can obtain the foreign price only by traveling to the foreign retail store.[15] 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, who spend their whole income, W, in the home country, have indirect utility 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.[16] Neglecting any home government taxes on the goods travelers bring back, the price paid for foreign goods is $EP^*(z)$ in domestic currency.[17] Finally we assume a nonpecuniary 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. 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 benefit of crossing is obtained by subtracting v_S from v_X, yielding

$$v_X - v_S = B - \ln\tau + \zeta, \tag{3}$$

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 RER:

$$B(\ln e) = -(1 - \vartheta(\tilde{z})) \ln e + \int_{\tilde{z}}^1 b(z) \ln \delta(z) dz. \tag{4}$$

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 equation (4) with respect to $\ln e$, while holding $\delta(z)$ constant,[18] can be expressed as

$$\begin{aligned} B' &= -(1 - \vartheta(\tilde{z})) + b(\tilde{z})(\ln\delta(\tilde{z}) - \ln e)\frac{\partial\tilde{z}}{\partial \ln e} \\ &= -(1 - \vartheta(\tilde{z})) < 0. \end{aligned} \tag{5}$$

The impact of the exchange rate on the benefits of crossing depends on the share of goods that are cheaper abroad: $1 - \vartheta(\tilde{z})$. Foreign appreciation (rising e) contracts that share, leading to a benefit function that is convex in the RER:

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

[15] An 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.

[16] In the empirical work, travel costs are a function of distance to the border, gas prices, and time costs.

[17] Adding a tax would just be a scalar multiplying the real exchange rate.

[18] We 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.

The convexity of the $B(\ln e)$ function arises under general functional form assumptions for preferences, $b(z)$, and relative price deviations $\delta(z)$.

Individual heterogeneity enters the net benefits of crossings in two ways. First, the nonpecuniary benefits of crossing, ζ, assumed to be distributed with a CDF denoted $F(\zeta)$. Second, we add community c subscripts to the determinants of travel costs to reflect differences in distance to the border and wages (time costs). 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 $v_X > v_S$:

$$x_c = \mathbb{P}(\zeta_c^* < \zeta) = F(B(\ln e) - \ln \tau_c). \tag{7}$$

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

$$\begin{aligned} \frac{\partial \ln x_c}{\partial \ln e} &= \frac{F'}{F} B' = -\frac{F'}{F}[1 - \vartheta(\tilde{z})] < 0, \\ \frac{\partial \ln x_c}{\partial \ln \tau_c} &= -\frac{F'}{F} < 0. \end{aligned} \tag{8}$$

The regressions in section II showed that the exchange rate elasticity of travel diminishes in periods when the foreign currency is strong. Differentiating the first equation in equation (8) yields

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

This expression reveals that once heterogeneity is added into the model, the positive second derivative of the individual benefit function (B'') shown in equation (6) will not translate into a positive second derivative for aggregate log crossings if the term in brackets is sufficiently negative. However, convexity in the benefit function is almost a necessary condition for convexity in log crossings. This is because the term in square brackets is negative for most distributions of individual heterogeneity.[19]

A second reduced-form finding we would like to reconcile with the model is that crossers from Canadian provinces into U.S. states exhibit higher exchange rate elasticities than residents of the U.S. states on the other side of the border. In the context of our model, this can happen if Canadians spend a higher share of their income on goods that are cheaper in the United States than vice versa. As shown in the online appendix, the model also predicts elasticities to differ in response to different population distributions, a hypothesis we confirm after estimating the model.

IV. Estimation of the Model

In order to estimate the crossing fraction equation (equation [7]), we need to parameterize the crossing benefit and cost functions—B and $\ln \tau_c$ in equation (3)—as well as specify the distribution of individual heterogeneity ($F(\zeta)$). We use a quadratic form for $B(\ln e)$ in our empirical specification since it is the simplest way to capture and test for convexity:

$$B(\ln e) = \beta_0 + \beta_1 \ln e + \beta_2 [\ln e]^2. \tag{9}$$

The model predicts $B' = \beta_1 + 2\beta_2 \ln e < 0$ for the observed range of e and $B'' = 2\beta_2 > 0$. Equation (9) can be justified as a second-order approximation of a general B or, as shown in the online appendix, as the exact form implied by two additional assumptions on prices and budget shares.

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 (the ratio of initial income to income after deducting travel costs) yields

$$\tau_c = \left[1 - D_c\left(\frac{1}{\psi H} + \frac{P_c(g)}{\phi W_c}\right)\right]^{-1}. \tag{10}$$

We see that the strict iceberg assumption of a constant fraction of income lost from travel is met in the limit only as the gas price-to-income ratio goes to 0. To facilitate estimation, we use a linear-in-logs approximation of equation (10):

$$\ln \tau_c = \gamma_0 + \gamma_1 \ln D_c + \gamma_2 \ln \left[P(g)_c / W_c\right]. \tag{11}$$

The γ_0 parameter shifts travel costs at all distances. One such shifter would be border formality compliance costs.[20] The γ_1 parameter represents the elasticity of travel costs with respect to distance.

We assume that ζ in equation (7) follows a normal distribution, with expectation μ and variance σ^2. Adding time subscripts and substituting the functions $F(\zeta)$, B and $\ln \tau$ into equation (7), we can express the crossing fraction as

$$\begin{aligned} x_{ct} = \Phi[&\theta_0 + \theta_1 \ln e_t + \theta_2 (\ln e_t)^2 + \theta_3 \ln D_c \\ &+ \theta_4 \ln (P(g)_{ct}/W_{ct})], \end{aligned} \tag{12}$$

where the coefficients are the model parameters scaled by the standard deviation of individual heterogeneity (e.g., $\theta_1 = \beta_1/\sigma$). The parameter $\theta_0 = (\beta_0 - \gamma_0 + \mu)/\sigma$, is allowed to vary across months and provinces, reflecting seasonal and geographic influences on the average nonpecuniary benefits (μ) and costs (γ_0) of travel.

[19] F'/F is globally decreasing for uniform, normal, logit, gumbel, and power law distributions.

[20] Our specification includes a dummy for periods after September 2001 to capture changes in border impediments.

When taking equation (12) to the data, we should replace x_{ct} with its conditional expectation: $\mathbb{E}[x_{ct} \mid e_t, D_c, P(g)_{ct}, W_{ct}]$. Deviations between observed crossing fractions and those predicted by the model 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 IIA, which are based on a survey given out to a subset of the actual population of crossers.

When estimating this equation, we need to recognize that the dependent variable is a fractional response bounded between 0 and 1. As table 2 in the online appendix shows, in many census divisions, the number of cars crossing the border is 0. It is important to employ an estimation method that incorporates zeros into the estimation and does not yield out-of-bounds predictions. We therefore estimate the expected crossings equation as a fractional probit model. This method yields consistent estimates of the model parameters so long as the conditional expectation $\mathbb{E}[x_{ct} \mid \cdot]$ 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.[21]

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

$$x_{ct} = \frac{n_{ct}}{N_{ct}} \approx \frac{\hat{n}_{ct}}{\text{Pop}_{ct} \times \text{CPC}_c \times 30}. \tag{13}$$

We estimate $\hat{n}_{ct}$ using data from the International Travel Survey (ITS), which was described in section IIA. The online appendix shows the sources for the variables in equation (13) 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 the online appendix. Our preferred form is the population-weighted median of the driving distances of all the subdivisions within a given CD.[22] 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 at the CD level from the Canadian census every five years.[23]

A. *Baseline Estimation*

We estimate the model parameters in equation (12) separately for travelers making same-day and overnight (stays of two or more days) trips. Travelers whose main reason for crossing the border is to shop are much more likely to make same-day trips, and it is these travelers whose behavior is represented in the model. By contrast, those making overnight trips may have purposes other than just shopping for goods to bring home—vacations, recreation spanning multiple days, visiting friends, and so on. For these travelers, the single-good model sketched in the online appendix 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 2. The first three columns use day trips 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 1b. Standard errors are clustered at the census division (c) level to allow for arbitrary serial correlation within divisions.[24] 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 2001. The province FEs capture differences in $B(\ln e)$ that result from unmeasured cross-state differences in product prices.[25] We focus on the third specification since adding province fixed effects improves the fit of the model considerably.

The coefficients on the exchange rate variables indicate that a higher value of the real exchange rate (implying a weaker

[21] Papke and Wooldridge (1996) explain this defect and other advantages of the fractional probit method.

[22] Figure D.1 in the online appendix contains a map of a few CDs in southeastern Ontario, showing the subdivisions within each CD. The map shows that driving distances are needed (rather than great-circle distances) to take into account the Great Lakes.

[23] Data details and sources are provided in the online appendix.

[24] Ideally 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 is feasible in the log-odds estimation, another advantage of using that method as a robustness check.

[25] They can also account for differences in the mean idiosyncratic shocks due to different population densities on the U.S. side of the border that affect the likelihood of visiting friends and relatives.

TABLE 2.—FRACTIONAL PROBIT ESTIMATION OF CROSSING FRACTIONS (x_{ct})

Length of Stay:	Day Trip			Overnight		
θ_0: constant	−0.23	9.80***	4.42***	−2.68***	−4.59***	−5.20***
	(0.31)	(2.94)	(1.52)	(0.07)	(0.57)	(0.99)
θ_1: $\ln e_t$ [RER]	−0.44***	−0.77***	−0.65***	−0.61***	−0.92***	−0.75***
	(0.10)	(0.14)	(0.13)	(0.12)	(0.13)	(0.12)
θ_2: $(\ln e_t)^2$	0.39	1.24***	0.82**	−0.09	0.27	−0.17
	(0.34)	(0.33)	(0.33)	(0.30)	(0.28)	(0.24)
θ_3: $\ln D_c$ [distance]	−0.58***	−0.58***	−0.52***	−0.14***	−0.14***	−0.12***
	(0.06)	(0.06)	(0.04)	(0.01)	(0.01)	(0.01)
$\ln P(g)_{ct}$ [gas price]		−0.35***	−0.07		−0.56***	−0.13***
		(0.09)	(0.05)		(0.04)	(0.02)
$\ln W_{ct}$ [income]		−0.80***	−0.42***		0.40***	0.29***
		(0.27)	(0.14)		(0.06)	(0.09)
Post-9/11			−0.14***			−0.14***
			(0.03)			(0.03)
R^2	0.24	0.29	0.53	0.05	0.07	0.08

Standard errors clustered by census division. Significant at *10%, **5%, ***1%. Regressions include month FEs. Columns 3 and 6 include province FEs. $N = 63{,}000$.

CAD) reduces the probability of cross-border trips. The coefficient on the squared term is positive for day trips, implying that travelers' responsiveness to the real exchange rate decreases as its level rises. This accords with the predictions of our model and is also consistent with the reduced-form results of table 1. Residents making day trips are more likely to expand the bundle of goods that they purchase in the United States 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 to 6. This may be because overnight travelers are more likely to purchase a standard bundle of goods in the United States (e.g., hotel stays, vacations, restaurant meals) 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 multiproduct shopping motive, whereas overnight trips appear to fit better with a single-good model such as the one in the online appendix.

Examining expenditure data provides additional support for this hypothesis. The ITS 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 are therefore noisy and potentially biased. 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.[26]

The results on travel costs show that driving distance creates a strong disincentive to cross the border. This is especially the case for day trips; distance is a weaker disincentive for those planning trips of a longer duration. High gas prices lower overnight travel significantly as expected for variables that increase travel costs. The coefficient on log gas prices in column 6 is about the same as the distance coefficient. The negative effect on day-trippers is smaller and imprecisely measured. Its confidence intervals include the coefficient for overnight trips (−0.13), as well as 0. The effect of gas prices in Canada may be harder to discern for same-day travelers because the majority live less than 18 miles from the border and can fill up at U.S. prices at gas stations south of the border.

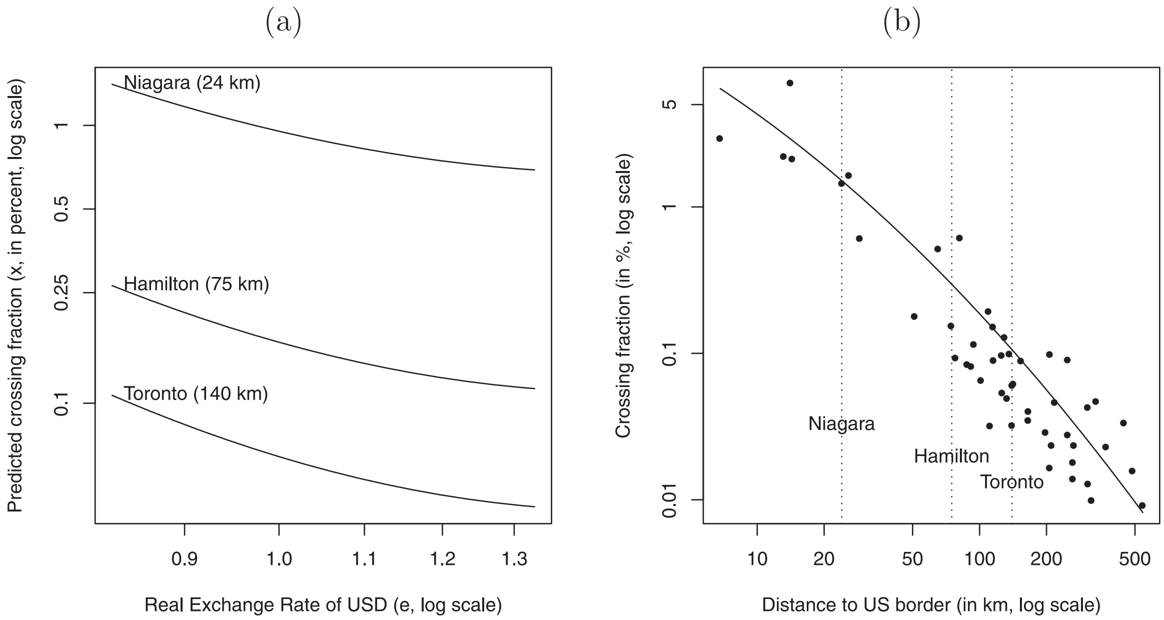
Income effects are strongly negative for day-trippers. This runs counter to what would be expected if income mattered just because it affects the fuel cost-to-income ratio in τ. Our model assumes a constant marginal utility of income across all individuals. One interpretation of the results is that richer 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 (10).

The downward shift in travel to the United States following September 2010 has a distance equivalent given by $\exp(0.14/0.52) - 1 = 0.31$. Thus, the extra costs of crossing the border in the years since 9/11 correspond to a 31% increase in distance. Alternatively, using a counterfactual calculation of the kind described in section IVD, 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.

Figure 3 illustrates the magnitudes of the estimated effects graphing predicted crossing shares as functions of 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 panel a. It is based on the specification in column 3 of table 2 (adjusting using the coefficients on the Ontario, post 9/11, and April dummy variables). Each curve corresponds to a census division in southern Ontario. The curves show that the convexity in the B function carries over

[26] See table 8 in the online appendix for these results.

Figure 3.—Crossing Declines with Foreign Appreciation and Distance to the Border



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 3b 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 farther 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 1,000 chance for a car to be driven across the U.S. border on a day trip. By contrast, communities closer than Niagara (15 miles) have crossing rates that are more than an order of magnitude higher.

B. *Robustness to Specification Changes and Falsification Tests*

In table 3 we present results from a number of different specifications and variable definitions. We use the set of controls corresponding to columns 3 and 6 of table 2. The first two columns of table 3 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, using information from Google on differences in average driving speeds relevant for different subdivisions. We add 26 minutes to the driving time to account (very roughly) for border wait times.[27] In columns 5 and 6, we use our secondary measure of distance (detailed in the online appendix). Relative to the primary measure used in table 2, it has the advantage of taking into account not just the nearest port but the five ports that residents of the CD use most frequently. It has the disadvantage of using the geographic center of the CD as the origin point, which exaggerates distances severely for some large divisions.

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 day trips and is insignificant for overnight trips. Travel costs, whether measured in terms of distance or time, have a negative and strongly significant effect on the probability of crossing the border—much more so for day trips than overnight ones.

There are a number of other robustness checks that we conducted, the results of which are contained in table 7 in the online appendix. We included a quadratic term for distance, but it was not statistically significant and did not 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). We examined whether commuters—residents of Canada who work in the United States—affect our results, since these travelers cross the border daily regardless of the exchange rate

[27] There is no source for nationwide wait time data. Twenty-six minutes is the median wait for travelers entering the United States during from 7:00 a.m. to noon at the two largest ports in British Columbia, using daily data from 2006 to 2010. Whatcom Council of Governments.

TABLE 3.—ALTERNATIVE SPECIFICATIONS OF REGRESSION AND TRAVEL COSTS

Method:	Log Odds (OLS)		Fractional Probit			
Stay:	Day Trip	Overnight	Day Trip	Overnight	Day Trip	Overnight
θ_0: constant	25.40***	−2.28	5.07***	−5.22***	10.33***	−4.61***
	(3.14)	(1.87)	(1.82)	(1.01)	(2.47)	(1.08)
θ_1: $\ln e_t$	−1.55***	−2.00***	−0.65***	−0.75***	−0.65***	−0.75***
	(0.25)	(0.15)	(0.13)	(0.12)	(0.13)	(0.12)
θ_2: $(\ln e_t)^2$	3.73***	0.20	0.93***	−0.15	1.03***	−0.16
	(0.73)	(0.63)	(0.33)	(0.24)	(0.32)	(0.24)
θ_3: ln distance or time	−1.14***	−0.28***	−0.89***	−0.19***	−0.56***	−0.14***
	(0.07)	(0.04)	(0.06)	(0.02)	(0.05)	(0.02)
$\ln P(g)_{ct}$	−0.15	−0.42***	−0.05	−0.13***	−0.03	−0.13***
	(0.13)	(0.08)	(0.05)	(0.02)	(0.05)	(0.02)
$\ln W_{ct}$	−2.41***	−0.19	−0.64***	0.26***	−0.94***	0.25**
	(0.30)	(0.17)	(0.18)	(0.10)	(0.24)	(0.10)
Post-9/11	−0.25***	−0.18***	−0.13***	−0.14***	−0.12***	−0.14***
	(0.06)	(0.04)	(0.03)	(0.03)	(0.03)	(0.03)
Observations	24,232	33,771	63,000	63,000	63,000	63,000
R^2	0.51	0.28	0.57	0.08	0.51	0.08

Standard errors clustered by census division except columns 1 and 2, where SEs also clustered by month-year. Regressions include month, province FEs. Significant at *10%, **5%, ***1%. Driving time in columns 3 and 4; port-use weighted average distances in columns 5 and 6.

and therefore are not the type of travelers whom the model considers.[28] Dropping the census divisions where commuters made up 10% or more of travelers leads to very similar results.

While all regressions in table 2 control for the average seasonal pattern in travel using month dummies, we also estimated regressions (table 9 in the online appendix) that directly include monthly weather data, as measured for the principal city in each province. In the absence of month effects, higher mean temperatures raise crossing propensities. When month effects are included, weather does not have a significant effect on day trips, indicating that the regular pattern of the seasons explains same-day travel behavior but idiosyncratic weather deviations do not. Overnight trips do respond to unusually bad weather, with sharp dips in months with high snowfall. 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 real exchange rate and distance terms enter the crossing equation (12) 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 does not improve the R^2 relative to the equation implied by our model, as shown in column 3 of table 7 in the online appendix.

A second way to validate the model draws on the interpretation of the seasonality captured in the month effects. If seasons matter because they increase marginal travel costs, then they should have a greater impact on residents living farther from the border. Furthermore they should have no interaction with the shopping benefits of crossing captured in the real exchange rate. To test these predictions, we interact the month dummies with $\ln D_c$ and, separately, with $\ln e_t$. We find strong distance-month interactions for same-day travel: the marginal impact of distance falls from −0.55 in January to −0.48 in July. The elasticity of travel with respect to the RER, however, does not have any significant seasonal pattern.[29] These results suggest a seasonal pattern to travel costs and add support for the model's implication of independence between shopping benefits and travel costs.

C. 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. Reexpressing the net benefits of crossing, $v_X - v_S$ in equation (3), using the parametric forms for $B(\ln e)$, and $\ln \tau(D)$, and setting $\zeta = 0$ we obtain

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

Totally differentiating by e and D and rearranging yields

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

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.

[28] Although commuters constitute under 6% of travelers, they make up a disproportionate share in certain census divisions, such as Essex (35% commuters), just across the border from Detroit.

[29] The p-value on the restriction that all month-RER interactions are zero is 0.7.

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 2, and -1.02 for time, based on column 3 of table 3. 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 day-trippers 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 US\$51, the 2010 median expenditure in the ITS of day-trippers who spent a positive amount.

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,[30] and 2010 Canadian median hourly wages of US\$23.34 per hour.[31] Using means instead of medians for D (56 miles) and X (\$152) leads to travel cost estimates of \$1.66 per mile and \$68.34 per hour. These travel cost estimates are at the high end of the range reported in the literature on shopping within national markets.[32]

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. As shown table 10 in the online appendix, the evaluation of $-\hat{\theta}_3/(\hat{\theta}_1 + 2\hat{\theta}_2 \ln e)$ in 2010 ranges from -0.60 to -0.62, with the normal distribution in the middle. The monetary travel costs differ by only a few cents per mile, demonstrating robustness to specific distributional assumptions.

D. *Quantification: Crossing Elasticities and Crosser Gains*

In this section we consider three counterfactual exercises: (a) a 10% appreciation of the Canadian dollar in any given year, (b) a doubling of wait times at the border, and (c) a replay 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 4.—Counterfactual Effects on Same-Day Travel Probabilities

	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

Columns 1 and 2 of table 4 show the impact, in two different years, on the number of cross-border trips from 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 3 in table 2. 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 U.S. dollar, in the past fifty years.

Table 4 reveals differences in the implied exchange rate elasticities across locations and time. In a comparison of the three Ontario census divisions, elasticities are larger for communities located farther 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 D.2 in the online appendix shows, Alberta, Saskatchewan, and Manitoba (provinces with high elasticities) have relatively few inhabitants located at or very close to the border. This finding is consistent with our discussion at the end of section III and in the online appendix. 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 increases 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 responsiveness of consumers to changes in relative prices. When the Canadian dollar is at its weakest (2002), the Canada-wide elasticity of 0.80 (first row of table 4) is almost the same as the average elasticity of 0.81 of the Blonigen and Wilson (1999) estimate for Canada-U.S. trade in goods. At the strongest levels of the RER, elasticities for travel are three

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

[31] See CANSIM Table 2820070.

[32] Chiou and Muehlegger (2008) estimate that consumers would be willing to travel to a location 2.7 miles farther 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.

times as large as those observed for goods. One reason that 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 upfront investments in marketing, distribution, and logistics.

Columns 3 and 4 of table 4 show the effect of increasing wait times at the border. We use the specification from column 3 of table 3, which had assumed a wait time of 26 minutes at the border. In our counterfactual experiment, we double this to 52 minutes.[33] 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, which 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, our model predicts a very large decrease in trips for a given increase in wait times for a province such as Ontario with a large population very close to the border.[34] 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. For a community with mass N_c of potential monthly crossers, aggregate surplus is the integral over individuals for whom $\zeta > \zeta_c^*$:

$$\begin{aligned} G_c &= N_c \int_{\zeta_c^*}^{\infty} [B - \ln \tau_c + \zeta] dF(\zeta) \\ &= \underbrace{(B - \ln \tau_c + \mathbb{E}[\zeta \mid \zeta > -B + \ln \tau_c])}_{\text{Average crosser's gain}} \underbrace{F[B - \ln \tau_c] N_c}_{\text{Number of crossers}}. \end{aligned} \tag{14}$$

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 $N(\mu, \sigma^2)$, we can compute the average crosser's gain as

$$\begin{aligned} G_c/n_c &= (B - \ln \tau_c) + \mu + \sigma \frac{\phi[(\mu + B - \ln \tau_c)/\sigma]}{\Phi[(\mu + B - \ln \tau_c)/\sigma]} \\ &= \sigma \left(\mathbf{Z}\hat{\boldsymbol{\theta}}_c + \frac{\phi[\mathbf{Z}\hat{\boldsymbol{\theta}}_c]}{\Phi[\mathbf{Z}\hat{\boldsymbol{\theta}}_c]} \right), \end{aligned} \tag{15}$$

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 in equation (15), 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.[35] 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%.[36] The gains to the average crosser are approximately three times as high when the appreciation starts from an already strong Canadian dollar.[37] 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%.[38]

Our final counterfactual 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 RER, gas prices, and incomes. The average crosser incurs a 3.4% reduction in welfare.

E. Reconciliation with Reduced-Form Estimates

We now return to a key result obtained in the stylized facts section: the elasticity of crossings with respect to the RER is 25% lower for Americans than for Canadians in the first two columns of table 1. An asymmetry when the Canadian dollar is strong is a prediction that follows from equation (5) since low e would reduce $\tilde{z}$, raising (in absolute value) the Canadian elasticity while lowering the U.S. one. However, the quartile specifications show that the asymmetry is found even when the RER is close to 1. Here we investigate whether differences in population distributions may be partially responsible for the observed lower crossing elasticities of U.S. residents.

We do not have data on the geographic distribution of U.S. crossers. However, we can use our estimates to simulate cross-border travel by Canadian residents in the event that

[33] Note that this increase in wait times needs to occur for exogenous reasons, such as reduced staffing at the border or an increase in the time taken to process each vehicle.

[34] The online appendix contains a figure displaying the different geographical distributions of population across Canadian provinces.

[35] This 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.

[36] The remainder, 0.31%, is attributable to the weighted product of the changes.

[37] As mentioned in section III, the benefits from crossing could be limited by car size constraints that prevent the crosser from taking full advantage of lower prices.

[38] See tables 11 and 12 in the online appendix for detailed results.

FIGURE 4.—POPULATION AND DISTANCE TO THE BORDER

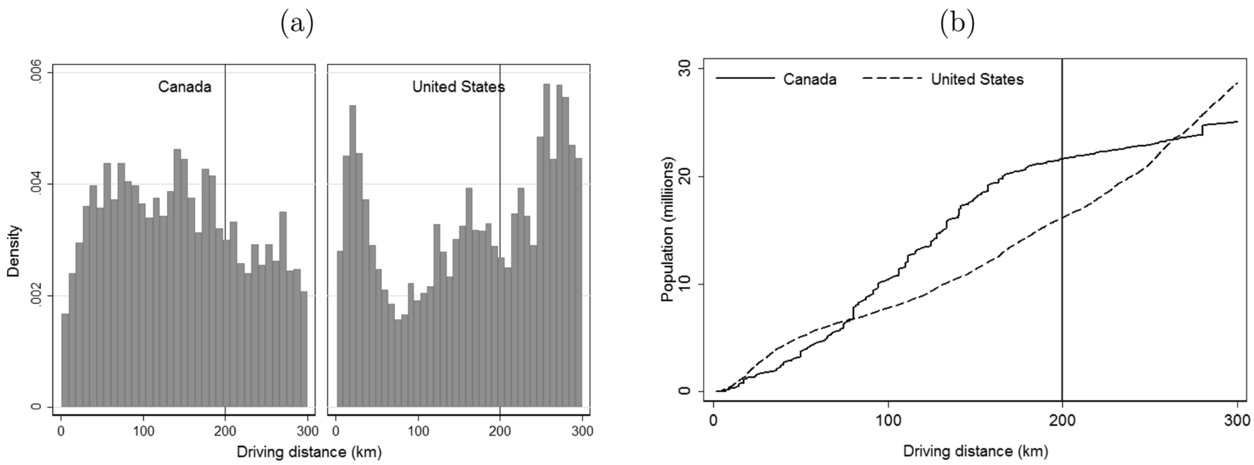


TABLE 5.—COUNTERFACTUAL TRAVEL ELASTICITIES, WITH SIMULATED U.S. DATA

	2002 ($e = 1.3$)		2005 ($e = 1.01$)		2010 ($e = 0.88$)	
Table 3 Specification	Canada	United States	Canada	United States	Canada	United States
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 as in table 4. U.S. elasticities simulated using census tract populations, with estimated coefficients from table 2.

their geographic distribution resembled that of the U.S. population most likely to make cross-border shopping trips. For this exercise, we use U.S. population and driving distances at the census tract level. We impose a cutoff distance of U.S. census tracts to the border of 200 km in order for the set of included census tracts to generally resemble the Canadian census divisions that are likely to have same-day crossers; this distance bound contains about 97.5% of Canadian same-day crossers. For each U.S. census tract we compute the predicted crossing probability, corresponding to estimating equation (12). We then conduct a counterfactual exercise similar to section IVD by increasing the exchange rate by 10% in order to calculate elasticities.

Figure 4 shows the differences between the United States and Canada in terms of population density and distance to the border.[39] 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 United States (within about 200 km of the border) is generally less densely populated than a similar distance cutoff in Canada.[40] These different distributions and population densities can affect crossing elasticities, as explained in the online appendix.

The comparison of Canadian and (simulated) U.S. elasticities is shown in table 5. We calculate these elasticities for 2002 and 2010, in order to correspond to table 4, as well as for 2005, which had a value of the RER close to 1. We present elasticities corresponding to each of the three specifications from table 2. Note that the elasticities for Canada in 2002 and 2010 using the column 3 specification are the same as those reported in table 4. The elasticities for the United States use the distribution of population across U.S. census tracts but applied to Canadian data on incomes and gas prices and using the coefficients estimated on the Canadian population in table 2. Although 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 U.S. census tracts to Canadian provinces in a somewhat arbitrary fashion.[41]

The results of table 5 suggest that changing the distribution of population in Canada to more closely reflect that of the northern United States would lower the elasticity of crossings with respect to the RER. In each year and given any of the three specifications of table 2, the elasticity using U.S. population data is lower than using Canadian data. In the most conservative estimate—that of column 3—the simulated elasticities are about 10% lower using the U.S. population distribution. In table 1 levels specification,

[39] The figures were constructed by calculating the driving distance from each census tract to the closest land border. Details are provided in the online appendix.

[40] The exception is the region within about 70 km of the border, containing the large U.S. cities of Buffalo and Detroit but no similar-sized Canadian cities.

[41] We assigned each U.S. census tract to the Canadian province that is across the border from the closest port to that census tract.

Americans have 25% lower elasticities. In other words, using the U.S. population distribution explains 40% of the difference in elasticities between Canadians and Americans. There are a variety of potential explanations for the remainder of the difference, but they lie outside our model. The U.S. side of the border might have a greater density of retail networks or offer a greater variety of goods than what is available on the Canadian side of the border. In addition, the Canadian media seem to accord more attention to the level of the currency, which may prime Canadians to travel in response to favorable shifts.

V. Conclusion

On average, each person living within a three-hour drive of the Canada-U.S. border makes more than one cross-border car trip per year.[42] In this paper, we develop and estimate a model of cross-border travel. In line with the shopping motive in our model, U.S.-Canada border crossings are heavily influenced by exchange rate changes. Furthermore, the elasticity of same-day crossings with respect to the exchange rate 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 travel costs. Our estimates range between \$30 and \$68 per travel 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 Canadians reside more than 81 miles away. We use our 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 U.S. and Canadian travelers to the exchange rate.

[42] Twenty-two million Canadians and 24 million Americans reside in this region.

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