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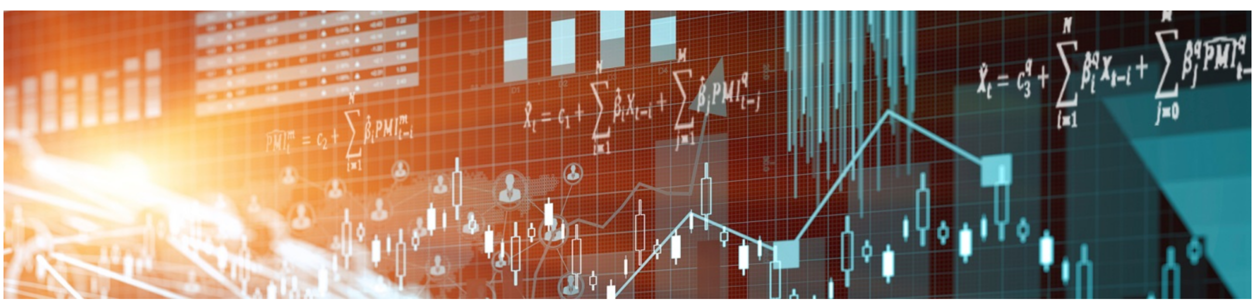
Exchange Rates, Retailers, and Importing: Theory and Firm-Level Evidence



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Exchange Rates, Retailers, and Importing: Theory and Firm-Level Evidence

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Abstract

We develop a model with firm heterogeneity in importing and cross-border shopping among consumers. Exchange-rate appreciations lower the cost of imported goods, but also lead to more cross-border shopping; hence, the net impact on aggregate retail prices and sales is ambiguous. Using Canadian firm-level data from 2002 to 2012, we find empirical support for several predictions of the model. We then estimate the model-implied exchange-rate elasticities of aggregate retail prices and sales. Our benchmark results indicate a deflationary effect of appreciations on retail prices and a small positive effect on sales. From 2002 to 2012, the Canadian exchange rate appreciated by 57%, which, according to our model, led to a 6.5% reduction in the retail price index. We also find that the estimated elasticities of aggregate retail prices and sales grew over this period, driven by import growth from China. This suggests that the transmission of exchange-rate movements to prices has grown since the early 2000s, which has consequences for the role of Canada's flexible exchange rate regime in supporting inflation stability.

Bank topics: Exchange rates, International topics, Service sector
JEL codes: F10, F14, L81

Résumé

Nous élaborons un modèle qui simule les importations d'un groupe hétérogène d'entreprises et les achats transfrontaliers de particuliers. Comme les hausses du taux de change font à la fois baisser le coût des biens importés et augmenter les achats transfrontaliers, leur incidence nette sur les prix et les ventes de détail agrégés est ambiguë. À partir de données sur les entreprises canadiennes couvrant la période de 2002 à 2012, nous obtenons des résultats qui corroborent de manière empirique plusieurs prévisions de notre modèle. Nous nous fondons ensuite sur celui-ci pour estimer l'élasticité des prix et des ventes de détail agrégés par rapport au taux de change. Nos résultats principaux indiquent que les hausses du taux de change ont un effet déflationniste sur les prix et un léger effet positif sur les ventes. De 2002 à 2012, le dollar canadien s'est apprécié de 57 %, ce qui a fait reculer l'indice des prix de détail de 6,5 %, selon notre modèle. Nous constatons aussi que les élasticités estimées des prix et des ventes de détail agrégés vont en augmentant sur cette période, sous l'effet de la croissance des importations en provenance de la Chine. La transmission des variations du taux de change aux prix pourrait donc s'être accrue depuis le début des années 2000, ce qui a des conséquences pour le rôle du régime de changes flottants du Canada dans le maintien de la stabilité du taux d'inflation.

Sujets : Taux de change, Questions internationales, Secteur des services
Codes JEL : F10, F14, L81

Non-Technical Summary

Understanding the effect of exchange-rate movements on prices and output is of central importance for monetary policy. In this paper, we study the impact of exchange-rate movements on aggregate retail sector prices and sales in a small open economy and apply our analysis to the case of Canada.

We develop a model with retail firms that differ in terms of size and productivity. Retail firms have the opportunity to sell both domestically produced goods and imports, which are imported either directly or indirectly through a domestic wholesale sector. Goods are sold to domestic consumers, who also buy foreign retail goods through cross-border shopping. Exchange-rate appreciations have two effects on the domestic retail sector. On the supply side, appreciations lower the cost of imported goods, which lowers retail prices and yields higher sales. On the demand side, appreciations lower the cost of foreign retail goods, which entices consumers to purchase from the foreign retail sector at the expense of the domestic sector. These two forces work in opposite directions, and therefore the net impact of exchange-rate appreciations on retail sector sales and prices is ambiguous.

Using Canadian firm-level data from the Annual Retail Trade Survey, the Canadian Import Register and other administrative data for the years 2002–2012, we test several predictions from our model. We find that a real exchange-rate appreciation increases the sales of direct-importing firms relative to other retailers, and that importing firms do not adjust their margins in response to exchange-rate movements. These findings are consistent with the theoretical predictions of the model. We also find, as other studies have, that retailers in close proximity to the Canada-U.S. border experience lower sales due to exchange-rate appreciations. This result is also consistent with the model, where exchange-rate appreciations lead to greater spending on cross-border retail goods, which lowers sales for domestic retail firms. Overall, our firm-level empirical results provide support for our theoretical framework.

Next, we quantify our model-derived equations for how the retail price index and aggregate retail sales are affected by exchange rates for the years 2002–2012. Our results indicate that appreciations have a deflationary effect on retail prices. From 2002 to 2012, the Canadian exchange rate appreciated by 57%, which, according to our model, led to a 6.5% reduction in the retail price index. With regard to aggregate retail sales, we find that appreciations have a small positive effect, as the negative effect of increased cross-border shopping is more than offset by the positive impact of lower-cost imported inputs.

We also find that the retail price index and aggregate retail sales became more sensitive to exchange-rate movements over this period, largely due to growth in Canadian retail imports from China.

Our analysis contributes to a large literature on exchange-rate pass-through to consumer prices. Ours is among the first papers in this literature that explores the role of direct importing by retailers in determining pass-though using firm-level data. Our results are consistent with existing research that finds that exchange-rate pass-through at the retailer level is high. We contribute to this literature by examining the distinction between direct-importing retail firms and other retailers and providing both theoretical and empirical evidence that direct importers experience higher sales in response to exchange-rate appreciations, relative to other retailers.

Our research also contributes to the literature on international trade and the retail sector. Several of the studies in this literature find that Canadian retail sales are negatively affected by appreciations, based on data from an earlier period that lacked information on firm-level importing. In contrast, our results indicate that the appreciation of the Canadian dollar from 2002 to 2012 had a positive effect on aggregate retail sales.

1 Introduction

Understanding the effect of exchange-rate movements on prices and output is a central aim in international economics. How far these movements influence prices can affect consumer price inflation, and is therefore important for the conduct of monetary policy. More generally, how fluctuations in the exchange rate affect the economy has consequences for the relationship between economic openness and macroeconomic outcomes. This paper studies the impact of exchange-rate movements on aggregate retail sector prices and sales in a small open economy.[1] We focus on quantifying the role of retail imports in transmitting exchange-rate changes to aggregate outcomes in the sector.

To guide our empirical analysis, we develop a model with firm heterogeneity in productivity and cross-border shopping among consumers. In the model, retail firms can import either directly or indirectly by purchasing imported goods from a domestic wholesale sector. Exchange-rate appreciations lower the cost of imported goods, but also lead to more cross-border shopping. Our model yields a closed-form expression for the exchange-rate elasticity of the retail price index, which is ambiguous in sign. Exchange-rate appreciations are deflationary due to pass-through of lower-cost imported goods to retail prices. But they are also inflationary, since greater cross-border shopping lowers the measure of domestic retail firms that operate. This in turn raises the price index through a love-of-variety effect in the model. We also derive an expression for the exchange-rate elasticity of aggregate retail sales. The sign of this elasticity is also ambiguous, capturing both the positive supply-side and negative demand-side effects of an appreciation.

Using Canadian firm-level data from the Annual Retail Trade Survey, the Canadian Import Register and other administrative data for the years 2002–2012, we test several predictions from our model. We find that a real exchange-rate appreciation increases the sales of direct-importing firms relative to other retailers. We also find that the difference in the sales-cost ratio between direct-importing and other retailers is unaffected by exchange-rate movements. These findings are consistent with the theoretical predictions of the model. Direct-importing firms experience cost shocks when the exchange rate appreciates. These shocks are fully passed through to lower retail prices, leading to higher sales relative to other

[1] In many developed countries, the retail sector contributes to gross domestic product (GDP) a share that is comparable with the manufacturing sector, and the retail sector is highly dependent on imports of goods for resale. In 2012, the Canadian employment share of the retail sector was 11.8% as compared with 10% for the manufacturing sector (Statistics Canada Labour Force Survey, Statistics Canada Table 14-10-0355). Despite these considerations, the retail sector has received relatively little attention in this literature compared with manufacturing.

firms.[2] We also find, as other studies have, that retailers in close proximity to the Canada-U.S. border experience lower sales due to exchange-rate appreciations. This result is also consistent with mechanisms from the model, where exchange-rate appreciations lead to greater spending on cross-border retail goods, which lowers sales for domestic retail firms. Overall, our firm-level reduced-form results provide empirical support for the theoretical framework we use in our subsequent model-based analysis.

Next, we quantify our model-derived equations for the exchange-rate elasticities of the retail price index and aggregate retail sales for the years 2002–2012. We estimate several of the parameters needed to calculate these elasticities using a combination of industry-level and firm-level data. Our results indicate that appreciations have a deflationary effect on retail prices. From 2002 to 2012, the Canadian dollar appreciated by 57%[3], which, according to our model, led to a 6.5% reduction in the retail price index. With regard to aggregate retail sales, we find that appreciations have a small positive effect, as the negative effect of increased cross-border shopping is more than offset by the positive impact of lower-cost imported inputs. We also find that the exchange-rate elasticities of the retail price index and aggregate sales increased over this period, largely due to growth in Canadian retail imports from China.

Our findings contribute to a large literature on exchange-rate pass-through to consumer prices.[4] Ours is among the first papers in this literature that explore the role of direct importing by retailers in determining pass-though using firm-level data. Our model and reduced-form firm-level results are consistent with research by Goldberg and Hellerstein (2013), Gopinath et al. (2011), and Nakamura and Zerom (2010), all of which highlights that retailers largely adjust their prices to reflect changes in costs, and hence pass-through at the retailer level is high.[5] Much of the existing literature focuses on documenting and explaining incomplete pass-through of exchange-rate movements in the manufacturing and/or wholesale sector. In contrast, we emphasize that direct importing by retail firms leads to higher pass-through of

[2] Devereux et al. (2017) report, based on data from the 2002–2006 period, that 92% of overall Canadian imported goods (in value terms) are invoiced in U.S. currency. This suggests that a Canadian dollar (CAD) appreciation should lower the cost of nearly all goods imported into Canada, relative to domestic goods.

[3] We use the U.S. dollar as a reference currency throughout the paper.

[4] Examples of recent contributions to this literature include Atkeson and Burstein (2008), Auer and Schoenle (2016), Berman et al. (2012), Corsetti et al. (2008), Devereux et al. (2017), Goldberg and Campa (2010), Goldberg and Hellerstein (2013), Burstein and Gopinath (2014), Gopinath et al. (2011), Hellerstein (2008), Nakamura and Zerom (2010), and Rodriguez-Lopez (2011).

[5] Goldberg and Hellerstein (2013) find that cost changes are largely passed through to prices based on retail-firm estimates for beer sold in Chicago. Gopinath et al. (2011) report similar findings based on estimates from a retail chain that operates in both the U.S. and Canada, as do Nakamura and Zerom (2010) based on estimates for the coffee industry. Gopinath and Itskhoki (2011) also highlight evidence that pass-though at the retail level is high, and consider the role of this and other empirical regularities in determining aggregate price adjustments.

exchange-rate movements to consumer prices. A recent study by Auer et al. (2018) has a similar focus to our own. These authors examine the impact of the 2015 Swiss franc appreciation on import prices, retail prices, and consumer spending based on transaction-level data. Our analysis is different from theirs in that we examine the distinction between direct-importing retail firms and other retailers, and provide both theoretical and empirical evidence that direct importers experience higher sales in response to exchange-rate appreciations, relative to other retailers.

Our research also contributes to the literature on international trade and the retail sector.[6] Our paper is related to the strand of this literature that studies the relationship between exchange-rate movements, cross-border shopping, and retail sector performance. Research in this literature has largely focused on cross-border shopping between Canada and the U.S., and our paper is closest to the recent papers by Baggs et al. (2016) and Baggs et al. (2018).[7] Baggs et al. (2016) find that appreciations in the CAD had negative effects on the sales of small retail firms during the period 1986–1997. Baggs et al. (2018) find that the magnitude of this negative effect declined in the mid-2000s, and argue this can be partially explained by more restrictive border controls following 9/11, leading to a reduction in responsiveness of cross-border travel to exchange-rate appreciations. Our paper differs from Baggs et al. (2016) and Baggs et al. (2018) in that they study the average response of small retailers (i.e., those with 20 or fewer employees), whereas we focus on aggregate retail sales and price responses to exchange-rate movements. We provide the first estimates of the effects of CAD appreciations on the aggregate sales and prices of Canadian retailers using a theoretical framework that incorporates cross-border shopping and importing by firms. Our paper is further differentiated in that it is the first in this literature to use retail firm-level importing data to quantify the aggregate and firm-level effects of an appreciation. While Baggs et al.'s (2018) search-theoretic model does consider importing, their empirical analysis does not incorporate firm-level import data. Our results highlight the importance of the import channel. We find that the appreciation of the Canadian dollar from 2002 to 2012 had a positive effect on aggregate retail sales, and that the magnitude of the exchange-rate elasticity increased over time as a result of growth in retail

[6] Recent research in this literature includes: Bernard et al. (2009), Basker and Van (2010), Bernard et al. (2010), Raff and Schmitt (2012), Bernard et al. (2015), Raff and Schmitt (2016), Cole and Eckel (2018), Emlinger and Poncet (2018), Meinen and Raff (2018), and Smith (2019).

[7] Other studies in this literature include Campbell and Lapham (2004) and Chen et al. (2017), who also find evidence of a significant effect of exchange-rate shocks on retail firm performance as a result of cross-border shopping. Chandra et al. (2014) do not examine retail firm performance, but do find that exchange-rate movements change the propensity of Canadians and Americans to cross-border shop.

imports from China.

The rest of the paper proceeds as follows. In section 2, we present the model. In section 3, we describe our data and test several theoretical predictions of the model using a reduced-form analysis. In section 4, we estimate the model-derived exchange-rate elasticities of aggregate retail sales and prices. In section 5, we conclude.

2 Model

In this section, we develop a small open-economy model featuring retail firms that are heterogeneous in productivity and importing, as well as cross-border shopping among consumers. We model retail firms in the spirit of Melitz (2003), in which exporters from the original framework are replaced with importing retailers.[8] Following the approach of Ahn et al. (2011), we include a wholesale sector in the model to account for the fact that a large fraction of retail goods is imported indirectly by intermediaries.

To model cross-border shopping, we include a cross-border good as an additional (exogenously priced) product in the constant elasticity of substitution (CES) consumption bundle of households. We therefore abstract from the details of cross-border shopping that have been emphasized elsewhere in the literature. The simplicity of our model in this respect allows us to derive closed-form analytical expressions for the exchange-rate elasticities of aggregate retail sales and the retail price index.

2.1 Households

We consider an economy with a unit measure of households, each with 1 unit of labour and the following utility function:

$$U = (\sum_{n=0}^{N} Q_n^{(\sigma-1)/\sigma} + Q_B^{(\sigma-1)/\sigma})^{\sigma/(\sigma-1)},$$

[8]Kasahara and Lapham (2013) also develop a model with firm heterogeneity in productivity and importing. Their model considers the dual decision to import and export among firms, where imported inputs are combined with domestic inputs in a CES production function that exhibits productivity gains through "love of variety." By comparison, our model entails fewer significant departures from the original Melitz (2003) model. This allows us to draw upon the well-established theoretical results from Melitz (2003) in the model-based empirical analysis in this paper.

where Q_B is the quantity of the cross-border good consumed by households, and Q_n is the CES aggregate of varieties from country n, adopting the convention that the domestic country is denoted as $n = 0$.[9] The imported varieties from country $n \in [1, N]$ include directly and indirectly imported varieties. The parameter σ is the elasticity of substitution across varieties, and we make the standard assumption that $\sigma > 1$.

There is a single foreign firm that sells the cross-border good, Q_B. Consumers cross-border shop by purchasing this good at an exogenous price, τ_B. This price, τ_B, includes the foreign price of the good converted into domestic currency and trade costs associated with cross-border shopping.

2.2 Retailers

All retailers are assumed to sell a domestic variety, which is produced using labour as the sole variable input. The cost of producing the domestic variety also entails a fixed operating cost f.[10] Retailers also have the option of selling imported varieties from up to N countries. Retailers produce imported varieties with a technology similar to the production function used for the domestic variety. However, imported inputs are used instead of labour as the variable input. Retailers purchase imported inputs by one of two means: direct importing or indirect importing. Direct importing entails the payment of a source-country-specific variable cost, τ_n, and fixed cost, $\eta_n f$. Alternatively, retailers can purchase imported inputs from a perfectly competitive domestic wholesale sector at a source-country-specific price $\tau_n^w = (1 + a)\tau_n$, where $a > 0$. Indirect importing via wholesalers also entails a country-specific fixed cost, $\eta_n^w f < \eta_n f$. Thus, as in Ahn et al. (2011), indirect importing entails a higher variable cost and lower fixed cost relative to direct importing.

In producing retail goods (domestic or imported), retailers are heterogeneous in that output is proportional to their idiosyncratic productivity, φ. We define d_n^w as an indicator variable that is 1 if a retailer indirectly imports from the wholesale sector, and 0 otherwise. The retail cost function for any source-

[9] We assume that the elasticity of substitution across varieties within each of the CES aggregates is also equal to σ, $Q_n = \left(\int_{\Omega_n} q_n(\omega)^{(\sigma-1)/\sigma} \partial\omega \right)^{\sigma/(\sigma-1)}$.

[10] Labour is the numeraire, and the nominal wage is normalized to unity.

country-specific variety can then be fully summarized as follows:

$$c_n(\varphi, d_n^w) = \frac{(1 + d_n^w a)\tau_n q_n(\varphi, d_n^w)}{\varphi} + (d_n^w \eta_n^w + (1 - d_n^w)\eta_n)f,$$

where $q_n(\varphi, d_n^w)$ is the retailer's source-country-specific output function. For the domestic variety, the indirect import indicator variable is 0 by assumption, $d_0^w = 0$, as it is produced with domestic labour. Therefore the variable cost of the domestic variety is equal to the nominal wage, which is normalized to 1, $\tau_0 = 1$. We also normalize the fixed cost of producing the domestic variety to 1, $\eta_0 = 1$. The productivity parameter φ is assumed to be drawn from a common continuous cumulative distribution function (CDF), $G(\varphi)$, with probability density function $g(\varphi)$.

We assume that the retail market is characterized by monopolistic competition. Facing CES demand for their differentiated goods, retailers use the following pricing function to maximize profits: $p_n(\varphi, d_n^w) = (1 + d_n^w a)\tau_n/(\rho\varphi)$, where $\rho \equiv (\sigma - 1)/\sigma$.[11]

We now describe the endogenous firm-level productivity cut-offs that determine retailers' operating and import status. The operating cut-off, φ_0, is defined by the condition that a retailer with this productivity level makes zero profits from selling its domestic variety, $\pi_0(\varphi_0, 0) = 0$. The indirect-importing cut-off, φ_n^w, is defined such that a retailer of this type makes zero profits from indirectly importing from country n, $\pi_n(\varphi_n^w, 1) = 0$. Finally, the direct-importing cut-off, φ_n, is defined as the productivity level at which a retailer's profits are equal across the two modes of importing from country n, $\pi_n(\varphi_n, 1) = \pi_n(\varphi_n, 0)$. The assumption that variable trade costs are lower for direct relative to indirect importing, $\tau_n < \tau_n^w$, implies that the most productive firms will have higher profits from importing directly (i.e., firms with $\varphi > \varphi_n$ import directly).[12] As in Melitz (2003), the trade participation cut-offs, φ_n^w and φ_n, can be expressed relative to the operating cut-off:

$$\varphi_n^w = \varphi_0 \tau_n^w {\eta_n^w}^{\frac{1}{\sigma-1}}; \quad \varphi_n = \varphi_0 \left(\frac{\eta_n - \eta_n^w}{\tau_n^{1-\sigma} - {\tau_n^w}^{1-\sigma}} \right)^{\frac{1}{\sigma-1}}. \tag{1}$$

[11] Our assumption that retail firm markups are constant is consistent with existing evidence from the literature (see Gopinath and Itskhoki (2011)). It is also consistent with results from our empirical analysis in section 3, where we find that importing retail firms do not adjust their sales-cost ratio, relative to other firms, in response to exchange-rate movements.

[12] Our assumptions regarding the trade cost parameters are sufficiently weak to allow for a variety of different outcomes with regard to indirect importing. For example, if the fixed cost of indirect importing is sufficiently low, then all operating firms will import either via wholesalers or through direct importing. Conversely, if the fixed cost of indirect importing is exorbitantly higher, then wholesalers may be bypassed completely, with all imported inputs being imported directly.

Combining the cut-off equations in (1) with the CES demand and pricing functions, we can write the profit functions as follows:

$$\pi_n(\varphi, d_n^w) = f\left(\left(\frac{\varphi}{\varphi_0}\right)^{\sigma-1}((1+d_n^w a)\tau_n)^{1-\sigma} - d_n^w\eta_n^w - (1-d_n^w)\eta_n\right).$$

Next, we solve for mean retail profits, $\tilde{\pi}$:[13]

$$\tilde{\pi} = \int_{\varphi_0}^{\infty}\pi_n(\varphi,0)\frac{g(\varphi)}{1-G(\varphi_0)}\partial\varphi + \sum_{n=1}^{N}\left(\int_{\varphi_n^w}^{\varphi_n}\pi_n(\varphi,1)\frac{g(\varphi)}{1-G(\varphi_0)}\partial\varphi + \int_{\varphi_n}^{\infty}\pi_n(\varphi,0)\frac{g(\varphi)}{1-G(\varphi_0)}\partial\varphi\right)$$

$$= fk(\varphi_0) + \sum_{n=1}^{N}\left(\frac{1-G(\varphi_n^w)}{1-G(\varphi_0)}\eta_n^w fk(\varphi_n^w) + \frac{1-G(\varphi_n)}{1-G(\varphi_0)}(\eta_n - \eta_n^w)fk(\varphi_n)\right), \tag{2}$$

where $k(\varphi) \equiv \left(\frac{\tilde{\varphi}(\varphi)}{\varphi}\right)^{\sigma-1} - 1$, and $\tilde{\varphi}(\varphi) = \left(\frac{1}{1-G(\varphi)}\int_{\varphi}^{\infty}\xi^{\sigma-1}g(\xi)\partial\xi\right)^{1/(\sigma-1)}$.

We assume that retail firms must pay a fixed cost to enter the industry, f_e, which implies the free-entry condition: $\tilde{\pi} = f_e\left(1-G(\varphi_0)\right)^{-1}$. The unique equilibrium value of φ_0 is determined by combining this free-entry condition with equation (2).

2.3 Aggregate Retail Sales and Retail Price Index

In this section we derive analytical expressions for aggregate retail sales and the retail price index, which are the key equations used in the model-based empirical analysis of this paper.

The household budget constraint requires that total income equals aggregate expenditures on the sum of retail and cross-border goods, $1 = PQ + P_BQ_B$. Note that the structure of CES demand is such that the ratio of expenditures on the cross-border good relative to those on any other variety can be written as:

$$\frac{P_BQ_B}{p_n(\varphi, d_n^w)q_n(\varphi, d_n^w)} = \left(\frac{\tau_B}{p_n(\varphi, d_n^w)}\right)^{1-\sigma}.$$

In particular, the ratio of expenditures on the cross-border good relative to expenditures on a domestic

[13] Detailed derivations of the trade participation cut-offs, the firm-level profit function, and the mean retail profits are provided in sections 2.1, 2.2, and 2.3 of the online appendix, respectively.

variety produced by a retailer with cut-off productivity, φ_0, is

$$\frac{P_B Q_B}{p_0(\varphi,0)q_0(\varphi,0)} = (\tau_B \rho \varphi_0)^{1-\sigma} \quad \Longrightarrow \quad P_B Q_B = (\tau_B \rho \varphi_0)^{1-\sigma} p_0(\varphi,0) q_0(\varphi,0). \tag{3}$$

Recall that the operating cut-off, φ_0, is defined by the zero-profit condition $\pi_0(\varphi_0, 0) = 0$. This zero-profit condition implies the following sales equation for a retailer with cut-off productivity: $p_0(\varphi, 0)q_0(\varphi, 0)$ $= \sigma f$. Combining this expression with equation (3) and the household budget constraint yields:

$$PQ = 1 - P_B Q_B = 1 - (\rho \tau_B \varphi_0)^{1-\sigma} \sigma f, \tag{4}$$

which is our analytical expression for aggregate retail sales.

We define $\bar{P}$ as the aggregate price index, $\bar{P} = \left(P^{1-\sigma} + \tau_B^{1-\sigma}\right)^{1/(1-\sigma)}$. That is, the aggregate price index is a CES aggregate of the retail price index, P, and the price of the cross-border good, τ_B. Rearranging the aggregate price index yields the following equation for the retail price index:

$$P = \left(\bar{P}^{1-\sigma} - \tau_B^{1-\sigma}\right)^{1/(1-\sigma)}. \tag{5}$$

In the next section, we use equations (4) and (5) to derive analytical expressions for the exchange-rate elasticities of aggregate retail sales and the price index.

2.4 Exchange-Rate Elasticities of Aggregate Retail Sales and the Retail Price Index

We denote the nominal exchange rate as E, where an increase in E indicates an appreciation in the domestic currency. We assume that imports from every country are denoted in a common foreign currency, and E denotes the units of this foreign currency that can be exchanged for one unit of the domestic currency. We assume that the domestic country, as a small open economy, takes E as exogenous. An appreciation is assumed to have the effect of decreasing the variable cost of importing from any foreign country, $\partial \tau_n / \partial E < 0$, for all $n \in [1, N]$[14] and decreasing the cost of cross-border shop-

[14]Our assumption that the wholesale sector is perfectly competitive implies that the wholesale sector, like the retail sector, fully passes through any changes in its costs to its prices. While existing research supports our assumption of complete pass-through for retailers, several studies have found evidence of incomplete cost pass-through in the wholesale sector. We acknowledge that this is a strong assumption, but it provides us with a sufficiently tractable framework to derive and estimate an exchange-rate elasticity of the retail price index from our model. In section 4.4 we show that our estimates of this elasticity are similar to estimates from the exchange-rate pass-through literature that uses aggregated consumer price index (CPI) data.

ping, $\partial \tau_B / \partial E < 0$.[15] In section 4.1 we empirically validate our assumptions that variable import trade costs and the cost of the cross-border good are decreasing with respect to the exchange rate. Throughout this section, we define all elasticities using the following notation: $\varepsilon_x^y = \partial ln(y)/\partial ln(x)$. Starting from equation (4), the elasticity of aggregate retail sales with respect to the exchange rate is:

$$\varepsilon_E^{PQ} = (\sigma - 1)\frac{P_B Q_B}{PQ}(\varepsilon_E^{\varphi_0} - |\varepsilon_E^{\tau_B}|). \tag{6}$$

We next focus on deriving a representation for the exchange-rate elasticity of the operating cut-off, $\varepsilon_E^{\varphi_0}$, which can be estimated using firm- and product-level data. It is relatively straightforward to estimate the remaining variables in equation (6), as is discussed in detail in section 4.1.

To derive an analytical expression for $\varepsilon_E^{\varphi_0}$, we begin by noting that the free-entry condition can be written $\tilde{\pi}(1 - G(\varphi_0)) = f_e$. Equating this representation of the free-entry to equation (2), multiplied through by $1 - G(\varphi_0)$, yields

$$fe = f(1 - G(\varphi_0))k(\varphi_0) + \sum_{n=1}^{N} (\eta_n^w f(1 - G(\varphi_n^w))k(\varphi_n^w) + (\eta_n - \eta_n^w)f(1 - G(\varphi_n))k(\varphi_n)). \tag{7}$$

Next we differentiate both sides of equation (7) with respect to the exchange rate, E, and rearrange to derive the following equation for the elasticity $\varepsilon_E^{\varphi_0}$:[16]

$$\varepsilon_E^{\varphi_0} = \frac{\sum_{n=1}^{N} \left(\eta_n^w (k(\varphi_n^w) + 1)(1 - G(\varphi_n^w))|\varepsilon_E^{\tau_n}| + (\eta_n - \eta_n^w)(k(\varphi_n) + 1)(1 - G(\varphi_n))|\varepsilon_E^{\tau_n}|\right)}{(k(\varphi_0) + 1)(1 - G(\varphi_0)) + \sum_{n=1}^{N} (\eta_n^w (k(\varphi_n^w) + 1)(1 - G(\varphi_n^w)) + (\eta_n - \eta_n^w)(k(\varphi_n) + 1)(1 - G(\varphi_n)))}$$

$$= \sum_{n=1}^{N} |\varepsilon_E^{\tau_n}| \frac{P_n Q_n}{PQ}. \tag{8}$$

That is, the elasticity $\varepsilon_E^{\varphi_0}$ is equal to the weighted import trade cost elasticity, where the weights are given by the import share of retail sales for each respective source country. Note that the numerator

This provides some external validation for our approach; however, we note that an extension of our framework that accounts for incomplete pass-through in the wholesale sector is an interesting topic for future research.

[15] In our setting, the τ_B represents the price of the cross-border good from the perspective of domestic consumers. When the domestic currency appreciates, this price declines, since foreign retail goods are priced in foreign currency.

[16] A detailed derivation of the elasticity of the operating cut-off with respect to the exchange rate is provided in section 2.4 of the online appendix.

of the source-country-specific import shares, P_nQ_n, is inclusive of sales generated from directly and indirectly imported varieties.[17] Substituting equation (8) into equation (6) yields

$$\varepsilon_E^{PQ} = (\sigma - 1)\frac{P_B Q_B}{PQ}\left(\sum_{n=1}^{N}\left(|\varepsilon_E^{\tau_n}|\frac{P_n Q_n}{PQ}\right) - |\varepsilon_E^{\tau_B}|\right). \tag{9}$$

The tension between the supply- and demand-side effects of an appreciation are captured in the elasticity in equation (9). The benefits of an appreciation for retail-sector sales result from a decline in import trade costs, $\varepsilon_E^{\tau_n} < 0$. This supply-side effect is larger if import trade costs are highly sensitive to changes in the exchange rate (i.e., if the magnitude of $\varepsilon_E^{\tau_n}$ is large), and if the import share of sales is high. The negative effect of an appreciation for retail-sector sales results from a decline in the cross-border shopping cost, $\varepsilon_E^{\tau_B} < 0$. This demand-side effect is larger if the cost of cross-border shopping is highly sensitive to changes in the exchange rate (i.e., if the magnitude of $\varepsilon_E^{\tau_B}$ is large).

We now derive an analytical expression for the exchange-rate elasticity of the retail price index. Starting from equation (5), the elasticity of the retail price index is:

$$\varepsilon_E^P = -\left(\frac{\bar{P}}{P}\right)^{1-\sigma}|\varepsilon_E^{\varphi_0}| + \left(\frac{\tau_B}{P}\right)^{1-\sigma}|\varepsilon_E^{\tau_B}|, \tag{10}$$

where we make use of the result that $\varepsilon_E^{\bar{P}} = -\varepsilon_E^{\varphi_0}$, which follows from the fact that $\bar{P} = (\sigma f)^{1/(\sigma-1)}/(\rho\varphi_0)$, as derived in appendix 6.1. Next, making use of the definition of P, and the representation of $\varepsilon_E^{\varphi_0}$ from equation (8), we can rewrite equation (10) as

$$\varepsilon_E^P = -\frac{1}{1-(\bar{P}/\tau_B)^{\sigma-1}}\sum_{n=1}^{N}\left(|\varepsilon_E^{\tau_n}|\frac{P_n Q_n}{PQ}\right) + \frac{(\bar{P}/\tau_B)^{\sigma-1}}{1-(\bar{P}/\tau_B)^{\sigma-1}}|\varepsilon_E^{\tau_B}|. \tag{11}$$

Finally, we note that, as a consequence of CES demand, $P_B Q_B = \bar{P}\bar{Q}\left(\bar{P}/\tau_B\right)^{\sigma-1}$. Making use of this expression and the fact that $\bar{P}\bar{Q} = PQ + P_B Q_B$, we rewrite equation (11) as

$$\varepsilon_E^P = -\frac{\bar{P}\bar{Q}}{PQ}\sum_{n=1}^{N}\left(|\varepsilon_E^{\tau_n}|\frac{P_n Q_n}{PQ}\right) + \frac{P_B Q_B}{PQ}|\varepsilon_E^{\tau_B}|, \tag{12}$$

where $\bar{P}\bar{Q}/(PQ)$ is the ratio of total expenditures (inclusive of retail and cross-border expenditures)

[17] By country-specific "import shares," we specifically mean the ratio of sales of (directly and indirectly) imported goods from country n relative to total retail sales.

relative to retail expenditures. Equation (12) captures the effect of an appreciation on the retail price index in an intuitive way. An appreciation reduces import trade costs, $\varepsilon_E^{\tau_n} < 0$, which lowers the price of imported varieties, and this puts downward pressure on the retail price index. This supply-side effect is larger when the exchange-rate elasticity of variable import trade costs is high, when the import share of sales is high, and when the share of domestic retail in total expenditures is high. In contrast, an appreciation lowers the cost of cross-border shopping, $\varepsilon_E^{\tau_B} < 0$, and this puts upward pressure on the retail price index. The decline in the cost of cross-border shopping increases demand for the cross-border good, and this has a negative effect on the total measure of retail varieties.[18] Under CES demand, a reduction in the measure of varieties results in an increase in the retail price index. This inflationary demand-side effect is larger when the exchange-rate elasticity of cross-border costs is high, and when the share of cross-border spending in total retail spending is high.

To summarize, changes in the exchange rate may either decrease or increase aggregate retail sales and the retail price index depending on the relative magnitudes of the supply- and demand-side effects respectively. This ambiguity motivates our model-based analysis in section 4. In that section we discuss how each of the variables in the elasticity equations for aggregate retail sales and the retail price index, equations (9) and (12) respectively, can be estimated from a combination of firm-level and product-level data.

2.5 Theoretical Firm-Level Predictions

Our theoretical framework has implications for how a retail firm's sales and prices respond to exchange-rate movements. In this section we derive two theoretical predictions, which we subsequently test with firm-level data in section 3.

Our first theoretical prediction relates to the impact of an exchange-rate movement on retail firm-level sales. Formally, we make the following proposition:

Proposition 1: Consider a retailer that sells imported varieties from $N' \leq N$ countries. The exchange-rate elasticity of the retailer's sales will be higher if it imports directly, rather than indirectly,

[18]Interestingly, changes in cross-border trade costs affect the retail price index only through influencing the total measure of varieties. The operating and trade participation cut-offs are unaffected by changes in τ_B, since these equilibrium cut-offs depend only on the retail supply-side parameters of the model. This can be seen from the trade participation cut-off equations in (1), and from the equations that determine the operating cut-off (equation (7) and the free-entry condition).

from all $N^{'}$ countries.

Proof: The logarithm of retail firm-level total sales, including sales of domestic and imported varieties, is given by the equation

$$log\left(p(\varphi)q(\varphi)\right) = log\left(\sigma f\left(\frac{\varphi}{\varphi_0}\right)^{\sigma-1}\left(1+\sum_{n=1}^{N^{'}}(1+d_n^w a)^{1-\sigma}\tau_n^{1-\sigma}\right)\right). \tag{13}$$

Differentiating equation (13) with respect to the logarithm of the exchange rate yields the following retail firm-level sales elasticity equation:

$$\varepsilon_E^{pq} = (1-\sigma)\varepsilon_E^{\varphi_0} + (\sigma-1)\frac{\sum_{n=1}^{N^{'}}(1+d_n^w a)^{1-\sigma}\tau_n^{-\sigma}|\varepsilon_E^{\tau_n}|}{1+\sum_{n=1}^{N^{'}}(1+d_n^w a)^{1-\sigma}\tau_n^{1-\sigma}}. \tag{14}$$

We define $\varepsilon_E^{pq}|_{d^w=0}$ and $\varepsilon_E^{pq}|_{d^w=1}$ as the retail firm-level sales elasticity for a direct and indirect importer, respectively. Proposition 1 implies that $\varepsilon_E^{pq}|_{d^w=0} > \varepsilon_E^{pq}|_{d^w=1}$, the proof of which follows directly from equation (14):

$$\varepsilon_E^{pq}|_{d^w=0} > \varepsilon_E^{pq}|_{d^w=1} \iff \frac{\sum_{n=1}^{N^{'}}\tau_n^{-\sigma}|\varepsilon_E^{\tau_n}|}{1+\sum_{n=1}^{N^{'}}\tau_n^{1-\sigma}} > \frac{(1+a)^{1-\sigma}\sum_{n=1}^{N^{'}}\tau_n^{-\sigma}|\varepsilon_E^{\tau_n}|}{1+\sum_{n=1}^{N^{'}}(1+a)^{1-\sigma}\tau_n^{1-\sigma}}$$

$$\iff 1+\sum_{n=1}^{N^{'}}(1+a)^{1-\sigma}\tau_n^{1-\sigma} > (1+a)^{1-\sigma}+\sum_{n=1}^{N^{'}}(1+a)^{1-\sigma}\tau_n^{1-\sigma} \iff 1 > (1+a)^{1-\sigma},$$

where the final inequality holds under the assumptions that $a > 1$ and $\sigma > 1$. □

The intuition for this result is as follows. Direct importers pay a lower variable cost for imports relative to indirect importers. As a result, imports make up a larger fraction of the sales of direct importers, and therefore they have a higher elasticity of sales with respect to the exchange rate compared with indirect importers. In section 3 we test the theoretical prediction in Proposition 1 by estimating the exchange-rate elasticity of sales for direct importers relative to other firms.

Our second proposition pertains to the response of firm-level prices to exchange-rate variation. Ideally, we would state our model's prediction of the exchange-rate elasticity of firm-level prices, and test this hypothesis with firm-level data. Unfortunately, unit prices are not available in our data, so we instead frame our second proposition in terms of the sales-cost ratio. As we establish below, the CES demand structure and monopolistically competitive market structure assumptions in our model have implications for the sales-cost ratio. This allows us to use firm-level data on the sales-cost ratio in section 3 to test the hypothesis that follows from our assumptions regarding the market structure in our model.

Prior to stating our second proposition, it is important to formalize the concept of the sales-cost ratio from a theoretical perspective. We define the sales-cost ratio as:

$$\nu(\varphi, d_n^w) = \frac{\sum_{n=0}^{N} p(\varphi, d_n^w) q(\varphi, d_n^w)}{\sum_{n=0}^{N} (1 + d_n^w a) \tau_n q(\varphi, d_n^w)} = \frac{\sum_{n=0}^{N} p(\varphi, d_n^w) q(\varphi, d_n^w)}{\sum_{n=0}^{N} cogs(\varphi, d_n^w)}. \tag{15}$$

That is, the sales-cost ratio is the ratio of total sales to the total cost of good sold (cogs). We conceptualize cogs as costs that are paid by the retailer to purchase the goods that it sells. Importantly, our definition of cogs does not incorporate the firm's productivity parameter, which we think of as affecting the firm's output after the goods enter the firm's possession. For example, firm-level productivity in our retail setting may be thought of as the idiosyncratic quality of the shopping experience that is provided by the retailer.

Having defined the sales-cost ratio, we state our second proposition as follows:

Proposition 2: A firm's sales-cost ratio depends only on its firm-level productivity and the CES markup, and therefore is unaffected by exchange-rate movements or its import status.

Proof: The result follows immediately by multiplying the numerator and denominator of equation (15) by $\varphi\rho$, and from the definition of the firm's monopolistically competitive price, which together imply $\nu(\varphi, d_n^w) = \nu(\varphi) = 1/\varphi\rho$. □

Intuitively, the higher a firm's CES markup, $1/\rho = \sigma/(\sigma - 1)$, the higher its price and its sales-cost ratio, whereas the higher a firm's productivity, φ, the lower its price and its sales-cost ratio. Proposition 2 implies that the exchange-rate elasticity of the firm-level sales-cost ratio is zero, and that the sales-cost ratio is unaffected by import status after controlling for firm-level productivity.

To summarize, the model developed in this section augments the canonical Melitz (2003) model by integrating retail-firm heterogeneity, importing, and cross-border shopping among consumers in a simple and concise way. From this framework, we derive closed-form analytical expressions for the exchange-rate elasticities of aggregate retail sales and the price index (equations (9) and (12), respectively). Both of these equations intuitively depend on the share of imported goods in domestic retail sales: as this share increases, exchange-rate movements have a more positive (negative) effect on aggregate retail sales (the retail price index). In section 4, we derive estimates of the share of imported goods in domestic retail sales from product-level data, and show that import growth from China has pushed up these shares throughout the 2002–2012 period. We also derive all other parameters required to compute equations (9) and (12), which permits us to quantify both elasticities for Canada throughout 2002–2012.

3 Data and Reduced-Form Analysis

In this section we introduce the firm-level data that are used throughout our empirical analysis. We also use reduced-form regression analysis to provide empirical support for the two theoretical firm-level predictions that were derived in section 2.5.

Our empirical analysis uses microdata from the Canadian Annual Retail Trade Survey (ARTS) merged with firm-level import data derived from Statistics Canada's Import Register (IR). Data from the ARTS are used for the two outcome variables of interest in our reduced-form analysis, firm-level sales and the sales-cost ratio, denoted $p_{i,t}q_{i,t}$ and $\nu_{i,t}$ respectively.[19] We compute the sales-cost ratio by dividing sales by cogs, $\nu_{i,t} = pq_{i,t}/cogs_{i,t}$, where $cogs_{i,t}$ is the cost of goods sold by firm i in year t. $cogs_{i,t}$ is defined implicitly in the ARTS by the following inventory accounting formula: $\Delta inventories_{i,t} = purchases_{i,t} - cogs_{i,t}$, where $purchases_{i,t}$ is the cost of new purchases and $\Delta inventories_{i,t}$ is the difference between closing and opening inventories.[20] The sales-cost ratio corresponds closely to the "gross margin," which is a commonly used measure of firm profitability.[21]

Our objective is to test our theoretical predictions regarding the effect of changes in the Canada-U.S. real exchange rate on firm-level sales and the sales-cost ratio. We define the real exchange rate as $RER_{j,t} = E_{USD/CAD,t}P_{j,CAN,t}/P_{j,US,t}$, such that an increase in $RER_{j,t}$ reflects a real appreciation

[19] In the case where a firm operates in multiple retail NAICS four-digit industries, the ARTS requires that the firm complete separate surveys for each of its respective industries. For the sales variable, the data are even further disaggregated, as firms are required to report their sales at each of their store locations. To match the level of aggregation in our import data, we aggregate the ARTS data to the firm level. In aggregating variables to the firm level, we define a firm's four-digit NAICS classification according to the sector in which the firm has the largest fraction of its sales. All nominal firm-level values in this paper are converted to real values using NAICS four-digit CPI deflators. Throughout this paper, our use of the term "firm" corresponds with Statistics Canada's definition of an enterprise.

[20] We censor our data by excluding all non-positive values for sales and cost of goods sold. For our $\nu_{i,t}$ regressions, we also exclude outlier observations for the sales-cost ratio, defined as cases where the ratio is greater than 5. This exclusion results in our dropping less than 1% of the firm-level observations in our sample.

[21] Specifically, the "gross margin" is commonly defined as $gross\ margin_{i,t} = (sales_{i,t} - cogs_{i,t})/sales_{i,t}$; hence, our measure is related such that $\nu_{i,t} = 1/(1 - gross\ margin_{i,t})$.

in Canadian retail sector j. In calculating the real exchange rate, we use data from Statistics Canada for the nominal exchange rate, $E_{USD/CAD,t}$, and the Canada-U.S. North American Industry Classification System (NAICS) four-specific relative CPI, $P_{j,CAN,t}/P_{j,US,t}$. We define an indicator variable for import status, $1(Importer_{i,t})$, which takes a value of one if a firm had a positive value of imports in the IR. We acknowledge many retailers do not directly import, yet nevertheless sell products that are imported by an intermediary. Thus our import indicator variable distinguishes direct-importing retailers from other firms that may or may not import indirectly.

Figure 2 reports how the Canada-U.S. exchange rate varied between 2002 and 2012. There are two important takeaways from this figure. The first is that annual fluctuations in the exchange rate were strongly correlated with oil price fluctuations throughout our period of study, and hence can be interpreted as exogenous from the perspective of retail firms in a small open economy like Canada. The second is that the Canadian dollar appreciated significantly over the course of 2002–2012, which likely had an impact on both firm and consumer decisions regarding the purchase of goods priced in U.S. dollars.

For our reduced-form empirical analysis, we adapt the standard trade premium regression framework that is commonly used in the empirical trade literature, as follows:

$$y_{i,j,t} = \beta_0 + \beta_1 ln(RER_{j,t}) + (\beta_2 + \beta_3 ln(RER_{j,t}))\, 1(Import_{i,t}) + (\beta_4 + \beta_5 ln(RER_{j,t}))\, ln(dist_{i,t})$$

$$+\boldsymbol{Z}_{j,t}\gamma + \delta_j + u_{i,j,t}, \tag{16}$$

where the dependent variable, $y_{i,j,t}$, is either the logarithm of sales or the logarithm of the sales-cost ratio. The key variables of interest are the logarithm of the real exchange rate, $ln(RER_{j,t})$, the import indicator, $1(Import_{i,t})$, and the interaction between these two variables.

Regression equation (16) also includes a distance variable, $ln(dist_{i,t})$, that accounts for proximity to the U.S. border.[22] We centre the $ln(dist_{i,t})$ variable to have a mean zero, so that the regression coefficients can be interpreted as marginal effects for a firm that is located at the average log distance from the U.S. border. We also include an interaction of this distance variable with the real exchange

[22] This variable is calculated in three steps. First, we calculate the Euclidean distance to the U.S. border from each retail store in ARTS. We approximate the location of each store using the centroid of its forward sorting area (the geographic area defined by the first 3 digits of the store's postal code – there are 1,620 FSAs in Canada). Second, we calculate each store's share of its parent firm's total sales. Using the store-specific distances from step one, and the shares from step two as weights, in the third step we calculate the weighted average distance to the U.S. border for each firm-level observation in the sample.

rate variable, $ln(RER_{j,t})$. The inclusion of the distance variable is motivated by research by Baggs et al. (2016), who show that Canadian retailer sales are negatively affected by real exchange-rate appreciations, and that this effect is attenuated with distance to the U.S. border. Baggs et al. (2016) attribute this result to cross-border shopping to the U.S. by Canadian consumers, which increases when the Canadian dollar appreciates, reducing demand at Canadian retail stores.

The matrix $\boldsymbol{Z}_{j,t}$ contains a set of country-level and industry-level control variables that might also affect retail sales. These control variables include: a measure of annual household consumption in each NAICS four-digit industry j, $ln(HHCons_{j,t})$, the Goods and Services Tax rate, GST_t, the logarithm of real median household income, $ln(Inc_t)$, the real interest rate RIR_t, and a variable t to allow for a trend in the dependent variable reflecting other dynamic factors that are not controlled for by the other regressors.[23] Finally, we include a dummy variable for each NAICS four-digit industry to control for any time-invariant sector-specific factors that may affect our dependent variables at the industry level.

Table 1: Summary Statistics

	Mean	Std. Dev.	Obs.
$pq_{i,t}$	44,717,000	322,536,645	89,968
$\nu_{i,t}$	1.541	0.526	89,172
$1(Imp_{i,t})$	0.284	-	89,968
$RER_{j,t}$	0.870	0.204	89,968
$dist_{i,t}$	239.704	297.385	89,968
GST_t	0.060	0.009	89,968
Inc_t	56,307.524	1,360.666	89,968
RIR_t	0.0173	0.0148	89,968
$HHcons_{j,t}$	831.151	1,913.843	89,968

All dollar values are in real 2002 Canadian dollars. The variables $pq_{i,t}$ and $\nu_{i,t}$ are firm-level sales and the sales-cost ratio, respectively, both sourced from the microdata in ARTS. The variable $\nu_{i,t}$ is censored to exclude firms with a sales-cost ratio greater than 5. $RER_{j,t}$ is the NAICS four-digit specific real exchange rate sourced from Statistics Canada. $1(Imp_{i,t})$ is a variable that takes a value of 1 if a firm has a positive value of imports in the IR, and 0 otherwise. The variable $dist_{i,t}$ is the firm-specific weighted average distance from the U.S. border (for details on the calculation of this variable, see footnote 22). GST_t is the level of the Canadian Goods and Services Tax, sourced from the Canadian Revenue Agency. The variable Inc_t is the level of annual real median income, sourced from Statistics Canada. RIR_t is the Canadian real interest rate, sourced from the World Bank. The variable $HHcons_{j,t}$ is the average value of household expenditures in each NAICS four-digit industry, sourced from Statistics Canada.

Summary statistics for the variables included in our regression analysis are reported in Table 1. On average, annual firm-level retail sales are almost 45 million CAD in our data; standard deviation in retail

[23] Data for these variables are sourced from the Canadian Revenue Agency, Statistics Canada, and the World Bank. All control variables are annual.

sales is more than nine times this figure, which indicates significant heterogeneity in retail firm size and motivates our theoretical approach in section 2. The average sales-cost ratio among firms in our data is roughly 1.5, and the dispersion of this ratio is as low as the coefficient of variation, 34%. The average share of direct importers among retail firms in our data is roughly 0.28. This share rose substantially over our period of study, from 0.13 in 2002 to 0.39 in 2012.

3.1 Reduced-Form Results: OLS, FE, and PSM

We estimate regression equation (16) by ordinary least squares (OLS), fixed effects (FE), and propensity score matching (PSM) estimation, with standard errors clustered at the firm level. Our goal is to estimate the impact of exchange-rate movements on retail firm sales and sales-cost ratios. Results are reported in Table 2. The OLS regressions in columns 1 and 4 control for various observable characteristics that might influence firm sales. The FE regressions in columns 2 and 5 additionally control for any time-invariant unobserved heterogeneity at the firm level. The PSM regressions in columns 3 and 6 attempt to identify the causal effect by matching each importer in our sample with a control firm that doesn't import directly but has a statistically similar propensity to import based on observable characteristics.[24] The two dependent variables, $ln(pq_{i,t})$ and $\nu_{i,t}$, correspond to the logarithm of sales and the sales-cost ratio measure.

We first discuss results related to non-direct-importing firms, since these results apply most directly to the existing literature. In column 1, which reports results with log sales as the dependent variable, the coefficient on the real exchange rate variable is negative but statistically insignificant; in column 2, this coefficient is negative and highly significant; and in column 3, this coefficient is, like in column 1, negative but statistically insignificant. These results suggest that the sales of a non-direct-importing retailer at average log distance from the U.S. border are either neutral or negatively affected by a real appreciation. According to columns 1 and 2, any negative effect of appreciations on sales is attenuated with distance from the U.S. border, as the interaction between the distance variable and the real exchange rate is positive and highly significant. These results are consistent with evidence from Baggs et al. (2016), who argue that Canadian retail firms face a loss in demand due to cross-border shopping during

[24]The PSM technique applied here follows a similar approach taken by Meinen and Raff (2018) and others. Details of how we implement this technique are described in the appendix.

Table 2: Descriptive Regression Analysis of Retailers and Real Exchange Rates

	$ln(pq_{i,t})$	$ln(pq_{i,t})$	$ln(pq_{i,t})$	$\nu_{i,t}$	$\nu_{i,t}$	$\nu_{i,t}$
VARIABLES	(OLS)	(FE)	(PSM)	(OLS)	(FE)	(PSM)
$ln(RER_{j,t})$	-0.0649	-0.223***	-0.239	-0.0499**	0.00878	-0.00258
	(0.0651)	(0.0420)	(0.217)	(0.0220)	(0.0205)	(0.110)
$1(Imp_{i,t})$	0.500***	0.0943***	0.210***	0.0833***	-0.00238	-0.00625
	(0.0258)	(0.0102)	(0.0417)	(0.00766)	(0.00481)	(0.0174)
$ln(RER_{j,t})1(Imp_{i,t})$	0.208**	0.234***	0.495***	-0.0229	-0.00679	-0.0609
	(0.0858)	(0.0406)	(0.141)	(0.0268)	(0.0210)	(0.0818)
$ln(dist_{i,t})$	-0.238***	0.0650**	-0.146***	-1.86e-05	-0.00721	-0.00875
	(0.00933)	(0.0303)	(0.0179)	(0.00336)	(0.0109)	(0.00716)
$ln(RER_{j,t})ln(dist_{i,t})$	0.153***	0.102***	0.0417	0.00452	-0.0220*	-0.0119
	(0.0327)	(0.0239)	(0.0704)	(0.0119)	(0.0122)	(0.0349)
$ln(HHcons_{j,t})$	1.552***	0.922***	1.572***	0.0994***	0.116***	0.346***
	(0.0812)	(0.0564)	(0.244)	(0.0253)	(0.0244)	(0.106)
GST_t	-0.960	-1.327*	8.924***	0.662	-0.386	0.654
	(1.165)	(0.712)	(3.351)	(0.444)	(0.349)	(1.673)
$ln(Inc_t)$	3.388***	5.118***	-3.486	-0.788***	0.331*	-0.822
	(0.603)	(0.361)	(2.809)	(0.229)	(0.177)	(1.367)
RIR_t	-2.056***	-2.683***	2.598*	0.262*	-0.143	-0.163
	(0.365)	(0.233)	(1.401)	(0.138)	(0.104)	(0.750)
t	-0.0363***	-0.0449***	0.0331	0.00783***	0.00195	0.00538
	(0.00509)	(0.00311)	(0.0213)	(0.00187)	(0.00147)	(0.00979)
Observations	89,968	89,968	18,339	89,172	89,172	18,181
R-squared	0.271	0.034	0.225	0.298	0.009	0.337
Number of firms	30,562	30,562	8,106	30,345	30,345	8,047

Standard errors are clustered at the firm level. All specifications include a full set of NAICS four-digit dummy variables. *** p<0.01, ** p<0.05, * p<0.1. The reduction in the sample size between the OLS/FE and PSM specifications reflects the fact that the PSM estimator uses only the matched sample of direct-importing firms. For each estimator (OLS, FE, and PSM), the sample size for the sales-cost ratio, $\nu_{i,t}$, is slightly smaller than the log sales outcome variable, $ln(pq_{i,t})$, as the variable $\nu_{i,t}$ is censored to exclude firms with a sales-cost ratio greater than 5.

periods of appreciation in the CAD. However, the coefficient estimate for this parameter under the PSM specification in column 3 is statistically insignificant.

In columns 4–6, which report results with the sales-cost ratio as the dependent variable, the coefficient on the real exchange rate variable is negative and statistically significant in the OLS specification, but not statistically different from 0 in the FE or PSM specifications. Thus, our FE and PSM results are consistent with our second theoretical proposition, which indicates that non-direct-importing retail firms did not adjust their sales-cost ratios when the exchange rate appreciated over the 2002–2012 period.

Next, we discuss results related to direct-importing firms. In column 1, the estimated coefficient for the import indicator variable, $1(Importer_{i,t})$, is positive and highly significant under OLS, which may be attributed either to large/growing firms self-selecting into importing, or to a causal effect of importing on sales. Both of these channels are found to be significant in previous studies that examine importing firms.[25] In column 2, which reports results with firm-level fixed effects, the estimated coefficient for the import indicator variable is also positive and significant, although it is less than one-fifth the magnitude of the coefficient in column 1. This difference suggests that the OLS estimate is influenced by selection bias due to unobserved time-invariant firm-level attributes, which are controlled for with firm-level fixed effects. In column 3, the estimated coefficient for the import indicator variable is positive and highly significant. This indicates that direct importing by retailers leads to higher sales.

A more novel finding in columns 1–3 is the positive and significant coefficients on the interaction between the import indicator and the real exchange rate. This result is consistent with our first theoretical proposition, as it suggests that Canadian retailers that are direct importers fare better than other firms during an appreciation. The estimated coefficients on the interaction term in the OLS and FE specifications are very similar, suggesting that the bias associated with unobserved firm-level attributes is not a large issue for this coefficient estimate. In fact, our PSM results in column 3, which can be most closely aligned with causality, suggest a larger positive impact of exchange-rate appreciations on importer sales relative to the results in columns 1 and 2. Together, these results indicate that the exchange-rate appreciations of the CAD during the period 2002–2012 led to higher sales for direct-importing firms relative to other retailers. These results are intuitive, since a real appreciation acts as a supply shock that lowers costs for importing retailers, which can then be passed on to lower retail prices, induce expenditure-switching

[25]For examples of studies that provide evidence of both these channels, see Amiti and Konings (2007), Halpern et al. (2015), Kasahara and Lapham (2013), and Kasahara and Rodrigue (2008).

among consumers, and yield higher sales relative to non-importers. Of course, such a transmission might not take place if retail firms raise their sales-cost ratios and offset the fall in import prices, a possibility we examine in columns 4–6.

In column 4, the import indicator variable has a positive and statistically significant relationship with the sales-cost ratio, whereas the corresponding coefficients in columns 5 and 6 are insignificant. The former OLS estimate could reflect self-selection of more profitable firms into importing. The latter result indicates that, once self-selection into importing is accounted for, there is no evidence that firms that enter into importing raise their sales-cost ratios. The lack of a statistically significant effect on the sales-cost ratio for the FE and PSM specifications is consistent with Proposition 2.

For the interaction between the import indicator and the real exchange rate, the coefficient is statistically insignificant in columns 4–6. Hence, we find no evidence that direct-importing retail firms adjust their sales-cost ratios, relative to other firms, in response to exchange-rate appreciations. This is also consistent with Proposition 2, and provides support for the exchange-rate transmission mechanism in our model. That is, an exchange-rate appreciation lowers the intermediate input costs for direct-importing retailers. These retailers do not adjust their sales-cost ratios, and hence pass though these lower costs to lower prices, inducing greater sales relative to other retailers.

The remaining coefficient estimates in Table 2 are broadly consistent with economic intuition and previous research. We omit a detailed discussion of these results, as our primary motivation is to test the theoretical predictions and assumptions of our model by focusing on the relationships between the outcomes and key variables of interest in our regression results.

All together, our results in Table 2 provide evidence of several key relationships that are relied upon in our model-based analysis.[26] We find evidence that sales of firms that import directly are more sensitive to exchange-rate movements than sales of firms that do not, *ceteris paribus*, which is consistent with Proposition 1 from the model. This result is supported by evidence from our OLS specification that controls for numerous firm and economic-wide variables, our FE specification that controls for time-invariant unobserved firm-level characteristics, and our PSM specification, which conditions on import propensity according to observed firm characteristics. Consistent with Proposition 2, we also find no

[26]As a robustness check, we also considered a version of the specification in (16) that includes a continuous measure of imports instead of the importing indicator variable, where the imports are defined at the firm level as the total value of imports. Results from this specification are qualitatively similar to those reported in Table 2.

robust evidence that direct-importing retail firms adjust their sales-cost ratios, relative to other firms, in response to exchange-rate movements. This result is supported by evidence from our OLS, FE, and PSM specifications. Together, these two results support the central point of our analysis, which is that the import channel is crucial for the transmission of exchange-rate movements to retail firm sales and prices (see equations (9) and (12)). Finally, we also find evidence that firms in close proximity to the Canada-U.S. border experience losses in sales, relative to firms further away from the border, due to exchange-rate appreciations. This result is supported by evidence from our OLS and FE specifications, although results from the PSM specification are not statistically significant. Our OLS and FE results are consistent with the cross-border shopping mechanism that is built into our model, where exchange-rate appreciations lower the cost of cross-border goods, thus attracting domestic consumers to foreign retailers at the expense of domestic retailers (see equation (9)).

4 Model-Based Analysis

In this section we estimate the exchange-rate elasticities of aggregate retail sales and the price index using the model-derived equations (9) and (12), respectively. We begin by discussing the model parameter estimates that are used in estimating both of these elasticity equations.

4.1 Model Parameter Estimates

Our elasticity equations require estimates of the following model parameters: the source-country-specific exchange-rate elasticity of variable importing trade costs, $\varepsilon_E^{\tau_n}$; the exchange-rate elasticity of the cost of cross-border shopping, $\varepsilon_E^{\tau_B}$; the import share, P_nQ_n/PQ, for each country n; and the CES elasticity of substitution, σ. In what follows, we estimate each of these model parameters using a combination of firm-level and product-level data. We also require an estimate of the ratio of cross-border shopping expenditures relative to Canadian retail expenditures, P_BQ_B/PQ. As discussed below, for this ratio we make use of estimates from previous research. Throughout our model-based empirical analysis, we consider three import source countries/regions: the U.S., China, and the rest of the world (ROW).

4.1.1 Exchange-Rate Elasticity of Import Trade Costs ($\varepsilon_E^{\tau_n}$)

We use the theoretical model to derive two estimating equations that identify exchange-rate elasticities of import trade costs, $\varepsilon_E^{\tau_n}$, which are specific to each source country. We derive the estimating equations as a function of the variable import trade cost, τ_n. Then we substitute the following reduced-form equation for variable trade costs into the estimating equation:

$$ln(\tau_{n,t}) = \alpha_0 + \alpha_1 ln(E_{USD/CAD,t}) + \alpha_2 t, \tag{17}$$

where we have added a time subscript to τ_n to reflect the time variation in import trade costs; the variable t is a time trend variable; and $E_{USD/CAD,t}$ is the USD-CAD nominal exchange rate.[27] The coefficient $\alpha_1 = \varepsilon_E^{\tau_n}$ is the country-specific exchange-rate elasticity of variable import costs, which is a key parameter needed for our model-based empirical analysis. As noted below, our estimation framework involves running country-specific regressions, and therefore any time-invariant country-specific factors affecting trade costs (e.g., distance) are absorbed in the coefficient α_0.

The dependent variable in our first estimating equation is the logarithm of country-specific firm-level purchases (i.e., imports) divided by firm-level domestic p urchases.[28] The Melitz (2003) structure of our model implies that this ratio is equal to $1 - \sigma$ multiplied by the logarithm of country-specific variable import costs, τ_n. Substituting in equation (17) for this term yields our first estimating equation:

$$ln\left(\frac{purchases_{i,n,t}}{purchases_{i,0,t}}\right) = (1-\sigma)\alpha_0 + (1-\sigma)\alpha_1 ln(E_{USD/CAD,t}) + (1-\sigma)\alpha_2 t + u_{i,n,t}, \tag{18}$$

where $purchases_{i,n,t}$ is the value of retail firm i's imports from country n, and $purchases_{i,0,t}$ is the value of the retailer's domestic purchases. We approximate domestic purchases by subtracting the value of the

[27] We use the USD-CAD nominal exchange rate rather than other bilateral exchange rates or the nominal effective exchange rate for Canada to accord with the dominance of USD invoicing in Canadian imports. Our data do not include information on currency of invoice, but Devereux et al. (2017) report that roughly 92% of Canadian imported goods (in value terms) were invoiced in USD during the period 2002–2008, which overlaps with our period of study. This suggests that, for nearly all importing firms in Canada, changes in the USD-CAD nominal exchange rate are more relevant than changes in other bilateral Canadian exchange rates.

[28] In our model-based analysis we use purchases rather than cogs as our measure for the intermediate input variable. Recall from section 3 that cogs is defined: $cogs_{i,n,t} = purchases_{i,n,t} - \Delta inventories_{i,n,t}$. In theory cogs is a more accurate measure as it accounts for inventories, but unfortunately we do not observe country-specific changes in inventories. However, we note that at the firm level the total change in inventories is typically very small relative to purchases. Theoretically, we do not make a distinction between purchases and cogs.

firm's total imports (from all countries) from its total purchases of goods for resale. For each importing source country, equation (18) is estimated by a panel fixed effects regression.[29] In our regressions we include NAICS four-digit industry dummies and cluster our standard errors at the firm level. Given an estimate of σ, we can calculate our key parameter of interest $\alpha_1 = \varepsilon_E^{\tau_n}$ from the regression coefficient on the exchange rate variable $ln(E_{USD/CAD,t})$. As discussed in section 4.1.4, we use an estimate of $\sigma = 4.02$ in our model-based analysis.

The dependent variable for our second estimating equation is the logarithm of retailer i's imports from country n, $ln(purchases_{i,n,t})$, which from our model implies the following estimating equation:

$$ln(purchases_{i,n,t}) = ln((\sigma - 1)f) + (1 - \sigma)ln(\tau_{n,t}) + (1 - \sigma)ln(\varphi_{0,t}) + \sigma ln(\varphi_{i,t}), \qquad (19)$$

where we have added a time subscript to the operating cut-off, $\varphi_{0,t}$, and a time and firm subscript to the firm's productivity parameter, $\varphi_{i,t}$, to reflect the inter-temporal and cross-sectional variation in these parameters. These parameters did not appear in equation (18) because they are in both the numerator and denominator of the import-domestic purchases ratio. Relative to the first estimating equation, the presence of cut-off and firm-specific productivity parameters makes identification more challenging. However, an advantage of the second estimating equation is that it is directly comparable to the approach we will use to estimate the exchange-rate elasticity of the cost of cross-border shopping, $\varepsilon_E^{\tau_B}$. Using a similar methodology to estimate $\varepsilon_E^{\tau_n}$ and $\varepsilon_E^{\tau_B}$ is important, since the exchange-rate elasticities of aggregate retail sales and prices depend critically on the relative magnitudes of these parameters.

We control for the firm-specific productivity parameter in equation (19), estimating a panel fixed effects regression. To control for time variation in the operating cut-off, we make use of the free-entry condition: $\tilde{\pi}_t = f_e\left(1 - G(\varphi_{0,t})\right)^{-1}$. Assuming that the distribution function G is invertible, we rearrange this condition to express the logarithm of the operating cut-off as a non-linear function of average retail profits, $\tilde{\pi}_t$, and the fixed cost of entry, f_e. We approximate this non-linear function using a fourth-order polynomial in average profits, which we denote $h(\tilde{\pi}_t)$.[30] As in our first estimating equation, we

[29] We note that according to our theory, the OLS estimator is unbiased, since the firm-level heterogeneity parameter, φ, cancels in the import-domestic cost ratio in equation (18). However, we estimate the fixed effects specification to control for other time-invariant dimensions of heterogeneity.

[30] We calculate average profits by subtracting retailers' mean annual rent and leasing expenses (our measure of fixed costs) from mean variable profits. As defined by the model, mean variable profits are expressed as mean sales divided by the elasticity of substitution, σ.

substitute equation (17) into equation (19), which allows us to rewrite our second estimating equation as follows:

$$ln(purchases_{i,n,t}) = \zeta_0 + (1-\sigma)\alpha_1 ln(E_{USD/CAD,t}) + (1-\sigma)\alpha_2 t + h(\tilde{\pi}_t) + u_{i,n,t}. \tag{20}$$

The right-hand side of our second estimating equation is the same as (18), except for the inclusion of the polynomial profit function and the redefined constant term, $\zeta_0 = ln((\sigma-1)f)+(1-\sigma)\alpha_0$. As in the first estimating equation, we include NAICS four-digit industry dummies and cluster our standard errors at the firm level. For each importing country, equation (20) is estimated by panel fixed effects regression. For the profit and import cost variables, we convert from nominal to real profits using the expenditure-based CPI for Canada. The key parameter of interest is again $\alpha_1 = \varepsilon_E^{\tau_n}$, which is calculated post-estimation from the coefficient on the exchange rate variable $ln(E_{USD/CAD,t})$, and using our estimate of σ = 4.02.

The results from estimating regression equations (18) and (20) are reported in Table 3, columns 1–3 and 4–6, respectively. For equation (20), our preferred specification, the estimates of the coefficient on the nominal exchange rate, $ln(E_{USD/CAD,t})$, are positive and statistically significant at the 5% level across all three source countries. This is consistent with the logic that a CAD appreciation lowers the cost of imported goods (priced in USD) relative to goods that are sourced domestically (priced in CAD), which leads to a rise in the import share. For both specifications, our estimate of this coefficient is largest for imports from China, slightly lower for imports from the U.S., and lower still for imports from ROW.

To derive values for $\varepsilon_E^{\tau_n}$, we simply divide our estimated coefficients on the nominal exchange rate by $(1-\sigma)$ as reflected in equations (18) and (20). Using the value of $\sigma = 4.02$, our country-specific exchange-rate elasticities of import costs (in absolute value) are $|\varepsilon_E^{\tau_{US}}| = 0.238$, $|\varepsilon_E^{\tau_{China}}| = 0.323$, and $|\varepsilon_E^{\tau_{ROW}}| = 0.201$. The relatively large elasticity estimate for China, combined with the rapid growth in Chinese imports during our study period, implies that there were considerable cost savings to retailers that imported from China. This, according to our model, contributes to deflationary pressure and Canadian retail sales growth in response to an appreciation of the CAD.

We also estimate p-values for the null hypothesis that these exchange-rate coefficient estimates are equal across all three source countries, which we report in the row below the R-squared values in Table 3. The p-values are high enough that we fail to reject the null hypothesis at any conventional level of

Table 3: Estimates of the Elasticity of Import Trade Costs with Respect to the Exchange Rate

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable	$ln\left(purchases_{i,n,t}/purchases_{i,0,t}\right)$			$ln(purchases_{i,n,t})$		
Import Country (n)	U.S.	China	ROW	U.S.	China	ROW
$ln(E_{USD/CAD,t})$	0.572**	0.878**	0.325	0.719***	0.973***	0.606**
	(0.266)	(0.344)	(0.287)	(0.249)	(0.323)	(0.257)
$time_t$	-0.0250**	0.102***	0.0142	-0.104***	0.0636**	-0.0363
	(0.0118)	(0.0156)	(0.0132)	(0.0212)	(0.0268)	(0.0229)
Polynomial profit function	No	No	No	Yes	Yes	Yes
NAICS 4-digit dummy variables	Yes	Yes	Yes	Yes	Yes	Yes
Firm-level fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	21,057	12,839	16,109	22,248	13,967	17,301
Number of firms	7,480	4,588	5,790	7,660	4,760	5,967
R-squared	0.004	0.060	0.004	0.009	0.099	0.012
pval: $ln(E_{USD/CAD,t})$ equal $\forall n$	0.466	0.466	0.466	0.670	0.670	0.670
pval: $time_t$ for all countries equal $\forall n$	5.91e-10	5.91e-10	5.91e-10	6.22e-06	6.22e-06	6.22e-06
$\lvert\epsilon_E^{\tau_n}\rvert$	0.190	0.291	0.108	0.238	0.323	0.201

Standard errors are clustered at the firm level. *** p<0.01, ** p<0.05, * p<0.1. The estimates of $|\epsilon_E^{\tau_n}|$ are calculated using the coefficient on the exchange rate variable, $(1-\sigma)\varepsilon_E^{\tau_n}$, and our estimate of σ=4.02 from Table 5.

statistical significance. Nevertheless, the economic significance of the differences in the estimates is large, as the trade cost elasticity estimate for China is more than one and a half times larger than for ROW. We therefore use the source-country-specific elasticities in columns 4–6 as our baseline estimates in our model-based analysis.

4.1.2 Exchange-Rate Elasticity of Cross-Border Shopping Costs ($\varepsilon_E^{\tau_B}$)

Our approach to estimating the exchange-rate elasticity of cross-border shopping, $\varepsilon_E^{\tau_B}$, closely follows the methodology used in (20) to estimate $\varepsilon_E^{\tau_n}$. From equation (4), the logarithm of cross-border expenditures can be written as:

$$ln(P_{B,t}Q_{B,t}) = ln(\sigma f \rho^{1-\sigma}) + (1-\sigma)ln(\tau_{B,t}) + (1-\sigma)ln(\varphi_{0,t}). \tag{21}$$

As in (20), we control for the operating cut-off using a fourth-order polynomial in average profits, defined as $h(\tilde{\pi}_t)$.[31] For cross-border trade costs, τ_B, we again make use of the reduced-form specification given by equation (17). We therefore rewrite our cross-border shopping estimating equation (21) as follows:

$$ln(P_{B,t}Q_{B,t}) = ln(\sigma f \rho^{1-\sigma}) + (1-\sigma)\alpha_0 + (1-\sigma)\alpha_1 ln(E_{USD/CAD,t}) + (1-\sigma)\alpha_2 t + h(\tilde{\pi}_t) + u_t, \tag{22}$$

where $P_{B,t}Q_{B,t}$ is quarterly Canadian aggregate expenditures on cross-border shopping, sourced from Statistics Canada Table 387-005.[32] While these cross-border expenditures data are available from quarter 1 of 1986 to quarter 2 of 2012, the data used for our average profit variable are only available from 2000 to 2012. We therefore estimate equation (22) using observations from quarter 1 of 2000 to quarter 2 of

[31] As equation (21) is estimated using quarterly data, we require quarterly data on average profits. Following the specification of equation (20), we calculate average profits by subtracting retailers' mean rent and leasing expenses from mean variable profits. We do not observe quarterly data on rent and leasing expenses, so we calculate this variable by dividing annual mean rent and leasing expenses by 4. As in the specification of equation (21), mean variable profits are defined as mean quarterly sales divided by the elasticity of substitution, σ. We calculate average quarterly Canadian retail sales using Statistics Canada Table 080-0020 (seasonally adjusted aggregate retail sales from the Monthly Retail Trade Survey, which excludes the sales of non-store retailers), and Statistics Canada Table 527-0013 (quarterly estimates for the number of Canadian firms for the retail sector).

[32] Statistics Canada Table 387-005 reports seasonally adjusted quarterly data from Canada's international travel account, including figures for total spending abroad by Canadians. Total spending abroad includes purchases of both goods and services. Some of these expenses are incurred during trips that are unrelated to cross-border shopping, and hence these figures provide an overestimate of cross-border shopping expenses. Nevertheless, this series provides the best measure of Canadian cross-border shopping expenses in the data available for our analysis.

2012.[33]

We report estimation results for specifications with and without the fourth order polynomial profit function, and in some specifications we also include a post-9/11 dummy variable (beginning in quarter 4 of 2001). For the cross-border expenditure and import cost variables, we convert from nominal to real profits using the annual average of the monthly aggregate CPI for Canada reported by Statistics Canada. We estimate equation (22) by OLS using heteroskedastic-robust standard errors.

The results from estimating equation (22) are reported in Table 4. Across all columns, the estimate of the coefficient on the nominal exchange rate, $ln(E_{USD/CAD,t})$, is positive and statistically significant at the 5% level. This is consistent with our conjecture that a CAD appreciation lowers the cost of cross-border shopping to the U.S. relative to goods that are sold in Canada, which encourages cross-border shopping expenditures. Columns 3 and 4 correspond to estimates from (22), which include the polynomial profit function. Results from our preferred specification, which includes the polynomial profit function and controls for the drop in cross-border shopping after 9/11, are reported in column 4.

Table 4: Estimates of the Elasticity of Cross-Border Shopping Costs with Respect to the Exchange Rate

	(1)	(2)	(3)	(4)
$ln(E_{USD/CAD,t})$	0.295***	0.225**	0.294***	0.241**
	(0.104)	(0.109)	(0.0929)	(0.0945)
$time_t$	0.00651***	0.00935***	0.00252	0.00428**
	(0.00127)	(0.00114)	(0.00208)	(0.00204)
$9/11_t$		-0.145***		-0.0969***
		(0.0202)		(0.0275)
Polynomial profit function	No	No	Yes	Yes
Observations	50	50	50	50
R-squared	0.868	0.939	0.942	0.951
$\lvert\epsilon_E^{\tau_B}\rvert$	0.0978	0.0746	0.0975	0.0801

Robust standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1. The estimates of $\lvert\epsilon_E^{\tau_B}\rvert$ are calculated using the coefficient on the exchange rate variable, $(1-\sigma)\varepsilon_E^{\tau_n}$, and our estimate of σ=4.02 from Table 5.

To derive a value for $\varepsilon_E^{\tau_B}$, we divide our estimated coefficient on the nominal exchange rate by $(1-\sigma)$ as reflected in equation (22), using a value of $\sigma = 4.02$. The exchange-rate elasticity of cross-border shopping costs is $\left|\varepsilon_E^{\tau_B}\right| = 0.0801$. Note that the magnitude of this estimate is considerably

[33] Our results are robust to excluding observations from 2000 and 2001. We have included these additional two years of data in order to include specifications that include a post-9/11 dummy variable.

smaller than our estimates of the import trade cost elasticities, $|\varepsilon_E^{\tau_{US}}| = 0.238$, $|\varepsilon_E^{\tau_{China}}| = 0.323$, and $|\varepsilon_E^{\tau_{ROW}}| = 0.201$. This suggests that variable import costs are more responsive than the cost of cross-border shopping to changes in the exchange rate.[34]

4.1.3 Import Share (P_nQ_n/PQ)

For our model-based analysis we require estimates of the share of source-country-specific imports in total retail sales, P_nQ_n/PQ. We calculate these import shares using total consumer good purchases from the input-output (IO) tables made publicly available by Statistics Canada.[35]

Columns 1–3 in Table 5 report the import share of retail sales from the U.S., China, and ROW, respectively, as calculated from the Statistics Canada IO tables. Taking the sum across columns 1–3, the overall import share in total retail sales rose from 0.4664 in 2002 to 0.5091 in 2012. The import growth during this period was largely driven by imports from China, which accounted for 7.5% of total imports in 2002 and 15.3% in 2012. In contrast, imports from the U.S. and ROW accounted for 55.4% and 37.2% respectively in 2002, and fell to 48.4% and 36.4% respectively by 2012.

4.1.4 Elasticity of Substitution (σ)

We cannot identify the elasticity of substitution, σ, without imposing parametric assumptions on the ex-ante productivity distribution, $G(\varphi)$. To obtain a baseline value of σ for our analysis, we assume that all firms are homogenous in productivity and normalize aggregate productivity to 1. In this case the CES markup is equal to the aggregate sales-cost ratio, $\sigma/(\sigma-1) = PQ/COGS$. Using this formula, we report the aggregate sales-cost ratio in column 4 and the corresponding value of σ in column 5 of Table 5 using Statistics Canada data.[36] We are reassured that σ is very stable over our observation period, and by the fact that its average value, $\sigma = 4.02$, is very close to 4, which is the value used in Melitz and

[34]We also considered several alternative specifications that are not reported in Table 4. As an alternative to the nominal exchange rate, we considered the Canadian-Dollar Effective Exchange Rate Index constructed by the Bank of Canada. This yields similar results to those yielded with the nominal exchange rate. As additional regressors, we considered an interaction term between the nominal exchange rate and the 9/11 indicator variable, and a "great trade collapse" indicator variable that equals 1 from the fourth quarter of 2008 to the fourth quarter of 2009 (inclusive), and 0 in all other quarters. Both of these variables yielded insignificant coefficients, and their inclusion had little effect on the estimate of the coefficient on the nominal exchange rate, $ln(E_{USD/CAD,t})$.

[35]In section 3 of the online appendix, we provide a detailed description of the method we use to estimate the import shares reported in columns 1–3. These estimates cover a broad range of core consumer goods, including final motor vehicles, gasoline, consumer food products, clothing, household goods, and personal care goods.

[36]Columns 4 and 5 use Statistics Canada Tables 080-0011, 080-0012, 080-0023, 080-0030, and 080-0028.

Table 5: Estimates of Source-Country-Specific Import Share of Retail Sales and CES Parameter

	(1)	(2)	(3)	(4)	(5)
Year	U.S. $\frac{P_nQ_n}{PQ}$	China $\frac{P_nQ_n}{PQ}$	ROW $\frac{P_nQ_n}{PQ}$	$PQ/COGS$	σ
2002	0.2583	0.0348	0.1733	1.328	4.050
2003	0.2435	0.0365	0.1691	1.321	4.120
2004	0.2386	0.0433	0.1694	1.317	4.158
2005	0.2406	0.0524	0.1704	1.310	4.222
2006	0.2382	0.0595	0.1755	1.327	4.057
2007	0.2447	0.0630	0.1754	1.333	3.999
2008	0.2419	0.0689	0.1845	1.331	4.020
2009	0.2233	0.0726	0.1844	1.343	3.913
2010	0.2328	0.0731	0.1909	1.345	3.899
2011	0.2355	0.0754	0.1936	1.338	3.958
2012	0.2463	0.0776	0.1852	1.345	3.899

Columns 1–3 provide country-specific measures of the import share calculated from Statistics Canada IO Tables 36-10-0403-01, 36-10-0424-01, 36-10-0418-01, and 36-10-0417-01, and UN Comtrade data. Column 4 is the ratio of aggregate retail sales relative to the costs of goods sold, and calculated from Statistics Canada Tables 080-0011, 080-0012, 080-0023, 080-0030, and 080-0028. Column 5 is the value of the CES elasticity of substitution σ that is implied by the aggregate sales-cost ratio for a benchmark model without firm heterogeneity, $\sigma/(\sigma-1) = PQ/COGS$.

Redding (2013) and several other papers in the empirical trade literature.

It is worth noting that our results are not particularly sensitive to the value of σ that is used in estimating our model-derived aggregate exchange-rate elasticity equations. In fact, our estimate of the exchange-rate elasticity of aggregate retail sales, equation (9), is unaffected by the value of σ that is used.[37] Furthermore, in estimating the exchange-rate elasticity of retail prices, equation (10), the parameter σ affects the magnitude but not the sign of our estimate.[38] In this respect, our qualitative results on the effect of exchange-rate variation on aggregate retail sales and prices are very robust to alternative specifications of the parameter σ.

4.1.5 Cross-Border Shopping Share of Canadian Retail Expenditures (P_BQ_B/PQ)

We rely on existing literature for our estimate of the ratio of cross-border shopping expenditures relative to Canadian retail expenditures, P_BQ_B/PQ. Corbi (2014) estimates this ratio using data from

[37] This is because elasticity equation (9) requires only an estimate of $(\sigma-1)\varepsilon_E^{\tau_n}$ and $(\sigma-1)\varepsilon_E^{\tau_B}$, which are regression coefficients on the exchange rate variable that are estimated in sections 4.1.1 and 4.1.2, respectively.

[38] Although σ does not appear directly in equation (10), we need an estimate of this parameter to obtain estimates of $\varepsilon_E^{\tau_n}$ and $\varepsilon_E^{\tau_B}$ from the regression results in sections 4.1.1 and 4.1.2, respectively.

Statistics Canada's Canadian System of Macroeconomic Accounts (CSMA) for the period 2006–2012. Included in Corbi's estimate of cross-border shopping expenditures are the outlays that Canadian households make on goods that are purchased in the U.S. and then brought back to Canada, goods that are delivered to Canada by post and courier, and motor vehicle imports purchased directly by consumers. These figures are derived for the CSMA from several different sources.[39]

Corbi (2014) provides a range of estimates for the total value of cross-border shopping expenditures based on different assumptions regarding expenses incurred during trips to the U.S. that are not brought back into Canada (e.g., meals abroad, accommodation). The estimates are based on both annual and quarterly seasonally adjusted series. For the purposes of our analysis, we use the values from the "high spending" scenario in Corbi (2014), since this is the scenario that is most closely related to how we model cross-border shopping.[40]

In Figure 3, we reproduce the ratio of cross-border shopping expenditures relative to Canadian retail expenditures. From the first quarter of 2006 to the fourth quarter of 2012, the cross-border shopping share fluctuated in the range from 1.6% to 2.4%. This share fluctuates positively with the exchange rate, which is consistent with a negative value for $\varepsilon_E^{\tau_B}$ as estimated in section 4.1.2. As we do not have estimates of cross-border shopping expenditures for the duration of our observation period, we set $P_B Q_B/(PQ)$ at a constant value equal to 1.96%, which is the average value of the "high spending" scenario ratio reported by Corbi (2014) over the period 2006–2012.

4.2 Exchange-Rate Elasticity of Retail Sales

In this section we present our estimates of the exchange-rate elasticity of aggregate retail sales. Our estimates use our model-derived elasticity equation and the model parameter estimates from section 4.1.

[39] These sources include: the International Travel Survey, which collects reported spending from respondent Canadian residents on overnight and same-day trips to the U.S.; customs data, combined with Canada post data, which permit calculation of the estimated value of goods delivered from abroad into Canada by post and courier; and data mainly from the Register of Imported Vehicles, which are used to derive an estimate for the value of motor vehicles brought into Canada from the U.S. by Canadian households. For the denominator, retail trade sales derived from data collected in the Monthly Retail Trade Survey are used to estimate total retail sales in Canada.

[40] In particular, we choose the "high spending" scenario because in our model, τ_B includes all travel costs that are associated with cross-border shopping.

Our elasticity equation, also reported as equation (9) in section 2, is the following:

$$\varepsilon_E^{PQ} = (\sigma - 1)\frac{P_B Q_B}{PQ}\left(\sum_{n=1}^{N}\left(|\varepsilon_E^{\tau_n}|\frac{P_n Q_n}{PQ}\right) - |\varepsilon_E^{\tau_B}|\right).$$

This equation allows us to distinguish the positive contribution of lower-priced imports, captured by the weighted import trade cost elasticity, $\sum_{n=1}^{N}\left(|\varepsilon_E^{\tau_n}|P_n Q_n/(PQ)\right)$, and the negative contribution of lower-priced cross-border goods, captured by $|\varepsilon_E^{\tau_B}|$, for aggregate retail sales.

Column 1 of Table 6 reports our estimates of the weighted import trade cost elasticity from 2002 to 2012, using the IO-based source-country-specific import shares, $P_n Q_n/(PQ)$, as reported in Table 5.[41] This weighted elasticity grew by roughly 12% from 2002 to 2012, and this growth is almost entirely due to the change in China's contribution. China's contribution to the weighted elasticity increased both because Canadian retail imports from China grew rapidly over this period, and because our estimated Chinese trade cost elasticity, $|\varepsilon_E^{\tau_{China}}|$, is larger than the trade cost elasticity for the U.S. and the ROW.

In column 2 we present our estimates of the exchange-rate elasticity of aggregate sales, ε_E^{PQ}. The aggregate sales elasticity is positive in all years, as the weighted import trade cost elasticity is greater than our estimate of the exchange-rate elasticity of cross-border shopping, $\varepsilon_E^{\tau_B} = 0.0801$. The aggregate sales elasticity increased by 48% over our sample period, and this growth is again due to the growth of Chinese imports. The magnitude of the retail sales elasticity is small in all years. For example, taking the estimate from 2012, a 10% appreciation in the exchange rate would result in a 0.0241% increase in retail sales. Mechanically, the small magnitude of our sales elasticity estimates is due to the fact that cross-border shopping accounts for only 2% of household expenditures. As households already spend 98% of their budget on domestic retail goods, the added sales from consumers switching away from the cross-border good results in a very small percentage increase in domestic sales.

Column 3 reports the counterfactual elasticity estimates for the years 2002–2012 under the case where retail import shares are set to 0 for all source countries, $P_n Q_n/(PQ) = 0$ for all n. Under this counterfactual scenario with no importing, the effect of an exchange-rate appreciation on sales is negative, as the exchange-rate elasticity of aggregate retail sales is -0.0047. Comparing this counterfactual case to our baseline estimates in column 2 shows that the mitigating effects of importing on retail sales

[41] All specifications in Table 6 use the following parameter estimates, which are assumed to be constant over time: $\sigma = 4.02$, $P_B Q_B/(PQ) = 0.0196$, $|\varepsilon_E^{\tau_B}| = 0.0801$, $|\varepsilon_E^{\tau_{US}}| = 0.238$, $|\varepsilon_E^{\tau_{China}}| = 0.323$, and $|\varepsilon_E^{\tau_{ROW}}| = 0.201$.

Table 6: Elasticity of Retail Sales with Respect to the Exchange Rate

	(1)	(2)	(3)
Year	$\frac{\sum_{n=1}^{N} \lvert\varepsilon_E^{\tau_n}\rvert P_n Q_n}{PQ}$	ε_E^{PQ}	No Importing ε_E^{PQ}
2002	0.1076	0.00163	-0.00473
2003	0.1038	0.00140	-0.00473
2004	0.1049	0.00146	-0.00473
2005	0.1085	0.00168	-0.00473
2006	0.1113	0.00184	-0.00473
2007	0.1139	0.00200	-0.00473
2008	0.1170	0.00218	-0.00473
2009	0.1137	0.00199	-0.00473
2010	0.1175	0.00221	-0.00473
2011	0.1194	0.00232	-0.00473
2012	0.1210	0.00241	-0.00473

All specifications use the following parameter estimates, which are assumed constant over time: $\sigma = 4.02$, $P_B Q_B/(PQ) = 0.0196$, $|\varepsilon_E^{\tau_B}| = 0.0801$, $|\epsilon_E^{\tau_{US}}| = 0.238$, $|\epsilon_E^{\tau_{China}}| = 0.323$, and $|\epsilon_E^{\tau_{ROW}}| = 0.201$. Estimates of $P_n Q_n/(PQ)$ for each respective country in each year are from Table 5. Column 3 reports the elasticity when imports are set to 0.

more than offset the negative effects of cross-border shopping, such that the net effect of an appreciation on retail sales is positive in all years in our observation period.

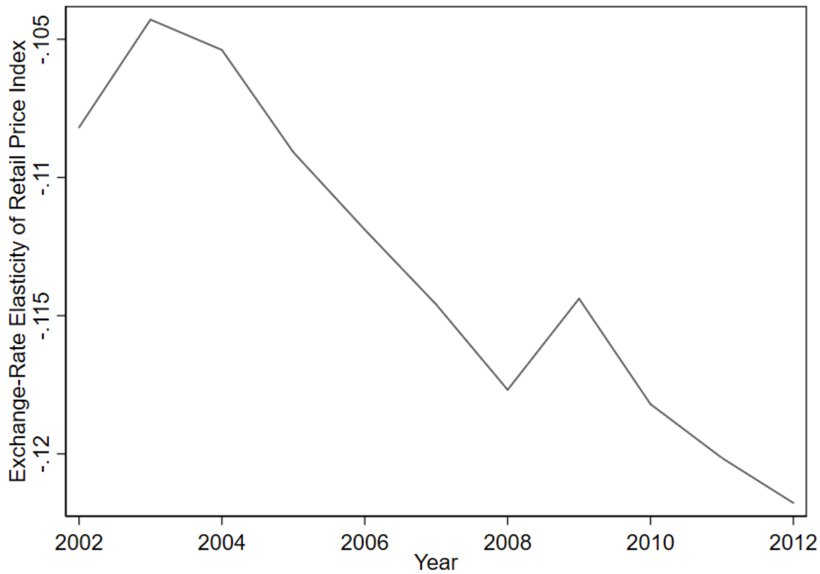
4.3 Exchange-Rate Elasticity of the Retail Price Index

In this section we present our estimates of the exchange-rate elasticity of the retail price index. Our elasticity equation, also reported in equation (10) in section 2.4, is the following:

$$\varepsilon_E^P = -\frac{\bar{P}\bar{Q}}{PQ} \sum_{n=1}^{N} \left(|\varepsilon_E^{\tau_n}| \frac{P_n Q_n}{PQ} \right) + \frac{P_B Q_B}{PQ} |\varepsilon_E^{\tau_B}|. \tag{23}$$

Intuitively, the weighted import trade cost elasticity, $\sum_{n=1}^{N} \left(|\varepsilon_E^{\tau_n}| P_n Q_n/(PQ) \right)$, has a deflationary effect, as cheaper imports lower retail prices. This term is weighted by the term $\bar{P}\bar{Q}/(PQ)$, which is the ratio of total expenditures (inclusive of retail and cross-border sales) to retail expenditures. Meanwhile, the exchange-rate elasticity of cross-border shopping costs, $\varepsilon_E^{\tau_B}$, has an inflationary effect on retail prices, since greater cross-border shopping leads to a decline in the measure of domestic retail firms. This reduction in the measure of domestic retailers raises the retail price index as a result of the love-of-variety

Figure 1: Exchange-Rate Elasticity of Retail Price Index



The following parameter estimates are used in calculating the elasticities, which are assumed constant over time: $\sigma = 4.02$, $P_BQ_B/(PQ) = 0.0196$, $|\varepsilon_E^{\tau B}| = 0.0801$, $|\epsilon_E^{\tau US}| = 0.238$, $|\epsilon_E^{\tau China}| = 0.323$, and $|\epsilon_E^{\tau ROW}| = 0.201$. Estimates of $P_nQ_n/(PQ)$ for each respective country in each year are from Table 5.

property of CES demand. However, this inflationary effect is very small, as the cross-border elasticity is weighted by the ratio of cross-border shopping expenditures relative to Canadian retail expenditures, $P_BQ_B/(PQ)$. Using Corbi's (2014) estimate of $P_BQ_B/(PQ) = 0.0196$ and our estimate of $\varepsilon_E^{\tau B} = 0.0801$, we find that the inflationary effect of an exchange-rate appreciation on retail prices due to cross-border shopping is 0.00157. We show below that this effect is dominated by the deflationary effects from the reduction in import costs in every year from 2002–2012.

Figure 1 presents our estimates of the exchange-rate elasticity of the retail price index, ε_E^P, using the parameter estimates from section 4.1.[42]

The elasticity of the aggregate retail price index with respect to the exchange rate has a mean value of -0.1132 over the period 2002–2012. The magnitude of this elasticity increases over our period of study by 13%, which mainly reflects growth in Canadian retail imports from China during this period. Overall, results from Figure 1 imply that exchange-rate appreciations contributed significantly to deflation in Canada's retail price index over the 2002–2012 period.

[42]The ratio $\bar{P}\bar{Q}/(PQ)$ is calculated using our estimate of $P_BQ_B/(PQ) = 0.0196$ from Corbi (2014), $\bar{P}\bar{Q}/(PQ) = (1 + P_BQ_B/(PQ)) = 1.0196$.

4.4 Discussion

Our results in sections 4.2 and 4.3 provide evidence that the import cost channel was an important contributor to the exchange-rate elasticities of aggregate retail sales and the retail price index for Canada throughout 2002–2012.

Between 2002 and 2012, the Canadian exchange rate appreciated in value by 57%, based on annual estimates derived from Bank of Canada data. According to our model, this appreciation led to a 0.1% increase in aggregate Canadian retail sales, and a 6.5% reduction in the retail price index.[43]

These results revise several narratives that come out of the existing literature for Canada. Numerous existing studies have found that exchange-rate appreciations have a negative impact on Canadian retail sales, and cite the role of cross-border shopping in generating this result. Baggs et al. (2018) find that more restrictive border controls following 9/11 reduced cross-border travel and the responsiveness of cross-border travel to exchange-rate appreciations. They show that the disruption in cross-border travel resulting from 9/11 considerably reduced the negative effects of an appreciation on the sales of small Canadian retailers. The fact that our study period focuses exclusively on the post-9/11 era is likely an important factor in explaining the qualitatively different conclusions in our paper relative to the earlier literature. Another important difference is that previous authors largely rely on data from the 1990s or earlier, which is an era when retail imports were less important for Canada. Our findings suggest that Canada experienced sizable growth in retail imports throughout 2002–2012, mostly originating from China. This growth put upward pressure on the exchange-rate elasticity of Canadian retail sales, contributing significantly to our finding that exchange-rate appreciations had a positive effect on aggregate retail sales over this period.[44]

Regarding the degree of exchange-rate pass-through to consumer prices, our estimates fall in a range similar to those that come out of the existing literature.[45] Meanwhile, several existing studies have found

[43] These estimates are calculated by using the average value of the retail sales and price elasticity estimates over the 2002–2012 period, and multiplying these average elasticities by the percentage change in the Canada-U.S. nominal exchange rate between 2002 and 2012, 57%. The magnitude of both the sales and price elasticities grew over the period, so the estimated effects of the CAD appreciation are smaller if we use the 2002 elasticities and larger if we use the 2012 elasticities.

[44] As discussed in section 4.2, our findings indicate that the exchange-rate elasticity of retail sales would have been negative throughout 2002–2012 if the import channel had been shut down. In periods prior to the 2000s, and especially prior to the 1990s, Canadian retail imports might well have been below the threshold where this elasticity turns positive, and hence the aggregate impact of exchange-rate appreciations on Canadian retail sales would have been negative.

[45] For example, Goldberg and Campa (2010) provide estimates for the exchange-rate elasticity of the CPI across a set of 21 industrialized countries, and report an average of -0.15. For Canada, Savoie-Chabot and Khan (2015) estimate the long-run elasticity of the Canadian CPI with respect to the nominal Canada-U.S. exchange rate to be roughly -0.06 using quarterly data

that this elasticity declined over time for industrialized countries, and for Canada in particular, based on evidence up to the late 1990s.[46] Our analysis updates this narrative, as our results show that throughout the 2002–2012 period, this elasticity rose by 13% for Canada, due to the rising importance of consumer goods imports from China.

5 Conclusion

This paper studies the impact of exchange-rate movements on retail sector sales and prices in a small open economy. We develop a model that includes retail firm heterogeneity in productivity, importing, and cross-border shopping among consumers. We use the model to derive expressions for the exchange-rate elasticities of aggregate retail sales and the retail price index. The sign of both elasticities is ambiguous, indicating that exchange-rate appreciations could have a positive or negative impact on aggregate retail sales and the retail price index, depending on several model parameters.

We then examine firm-level empirical evidence on the impact of exchange-rate movements on the sales and sales-cost ratios of Canadian retailers over the 2002–2012 period. Our findings suggest that retail firms that import directly experience higher sales, relative to other firms, due to exchange-rate appreciations. We find no evidence that direct-importing retail firms adjust their sales-cost ratios relative to other firms in response to exchange-rate movements, which suggests that importers fully pass through the lower cost of imported goods to retail prices. We also find that firms that are close to the Canada-U.S. border, and hence more exposed to pressures from cross-border shopping, experience losses in sales relative to the typical Canadian firm, due to exchange-rate appreciations. This evidence is altogether consistent with the mechanisms from the model.

Finally, we quantify the effect of exchange-rate appreciations on Canadian aggregate retail sales and the retail price index, using model parameter estimates from Canadian microdata and other sources. Our results indicate that the negative impact of increased cross-border shopping is more than offset by the positive impact of lower-cost imports for the retail sector; hence, the aggregate impact of exchange-rate appreciations on retail sales is positive. Meanwhile, we find that appreciations had a deflationary effect

for the period 1995 to 2013. In a somewhat older study, Bailliu and Bouakez (2004) note, “It has been traditionally estimated that about 20% of a persistent change in the Canadian dollar is reflected in the core CPI” (pg. 23).

[46] For evidence related to industrialized countries, see Bailliu and Bouakez (2004) and Goldberg and Campa (2010). For evidence related to Canada, see Bouakez and Rebei (2008).

on retail prices. From 2002 to 2012, the Canadian exchange rate appreciated by 57%, which, according to our model, led to a 6.5% reduction in the retail price index.

In the context of the literature, our results are largely in line with the findings of several other papers. Several studies have found that, while exchange-rate pass-through among wholesalers and manufacturers may be incomplete, pass-through at the retailer level is high.[47] Our findings corroborate these results, and furthermore bring emphasis to the potential role for direct importing by retail firms in bypassing other firms and, in turn, enhancing exchange-rate pass-through. Meanwhile, several studies have found that exchange-rate appreciations had a negative impact on Canadian retail sales, and that exchange-rate pass-through to retail prices has fallen over time in Canada, based on evidence from before the 2000s (Baggs et al. (2016), Bouakez and Rebei (2008)). Our findings from throughout the 2002–2012 period indicate that, supported by the strength of the Canadian dollar and declining costs of imports from China, the share of imports in Canadian retail goods was both large and increasing over time. This led to a positive impact of exchange-rate appreciations on Canadian retail sales, and growth in the degree of exchange-rate pass-through to Canadian retail prices over this period.

Our findings have important implications for policy-makers. From the perspective of monetary policy, our results suggest that the transmission of exchange-rate movements to Canadian prices has grown since the early 2000s, which has consequences for the role of Canada's flexible exchange rate regime in supporting inflation stability. More generally, our results point to the benefits of importing for inducing lower prices and higher sales in the Canadian retail sector when the exchange rate appreciates.

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6 Model Appendix

6.1 Solving the Model

In this section we solve for the remaining variables that determine the equilibrium of the model. In particular, we derive analytical expressions for the equilibrium mass of varieties and the aggregate price index.

The mass of varieties is defined as $\bar{M} = M_0 + \sum_{n=1}^{N}(M_n^w + M_n)$, where M_0 is the mass of domestic varieties, M_n^w is the mass of retailers indirectly importing from country n, and M_n is the mass of retailers directly importing from country n. Following Melitz (2003), we define the weighted average productivity, $\bar{\tilde{\varphi}}$, as follows:

$$\bar{\tilde{\varphi}} = \left\{ \frac{1}{\bar{M}} \left[M_0 \tilde{\varphi}(\varphi_0)^{\sigma-1} + \sum_{n=1}^{N} \left((M_n^w + M_n) \tau_n^{w\,1-\sigma} \tilde{\varphi}(\varphi_n^w)^{\sigma-1} + M_n(\tau_n^{1-\sigma} - \tau_n^{w\,1-\sigma}) \tilde{\varphi}(\varphi_n)^{\sigma-1} \right) \right] \right\}^{\frac{1}{\sigma-1}}. \tag{24}$$

Using these definitions of $\bar{M}$ and $\bar{\tilde{\varphi}}$, we derive an equation for aggregate retail sales as follows:[48]

$$PQ = \int_{\varphi_0}^{\infty} \left(\frac{\varphi}{\varphi_0} \right)^{\sigma-1} \sigma f \frac{g(\varphi)}{1 - G(\varphi_0)} M_0 \partial\varphi$$

$$+ \sum_{n=1}^{N} \left(\int_{\varphi_n^w}^{\varphi_n} \left(\frac{\varphi}{\varphi_0} \right)^{\sigma-1} \tau_n^{w\,1-\sigma} \sigma f \frac{g(\varphi)}{1 - G(\varphi_0)} M_0 \partial\varphi + \int_{\varphi_n}^{\infty} \left(\frac{\varphi}{\varphi_0} \right)^{\sigma-1} \tau_n^{1-\sigma} \sigma f \frac{g(\varphi)}{1 - G(\varphi_0)} M_0 \partial\varphi \right)$$

$$= \bar{M}(\bar{\tilde{\varphi}}/\varphi_0)^{\sigma-1} \sigma f. \tag{25}$$

[48] A detailed derivation of this representation of aggregate retail sales is provided in section 2.5 of the online appendix.

Next we derive households' aggregate expenditures on retail goods as a function of $\bar{\tilde{\varphi}}$, $\bar{M}$, and the aggregate price index, $\bar{P}$. Note that the aggregate price index, $\bar{P}$, is distinct from the retail price index, P, since the aggregate index incorporates retail prices *and* the price of the cross-border good.[49] We derive an expression for aggregate expenditures on retail goods as follows:[50]

$$PQ = \int_{\varphi_0}^{\infty} p_0(\varphi, 0)^{1-\sigma} \bar{P}^{\sigma-1} \frac{g(\varphi)}{1 - G(\varphi_0)} M_0 \partial\varphi$$

$$+ \sum_{n=1}^{N} \left(\int_{\varphi_n^w}^{\varphi_n} p_n(\varphi, 1)^{1-\sigma} \bar{P}^{\sigma-1} \frac{g(\varphi)}{1 - G(\varphi_0)} M_0 \partial\varphi + \int_{\varphi_n}^{\infty} p_n(\varphi, 0)^{1-\sigma} \bar{P}^{\sigma-1} \frac{g(\varphi)}{1 - G(\varphi_0)} M_0 \partial\varphi \right)$$

$$= \bar{M} \left(\bar{P} \bar{\tilde{\varphi}} \rho \right)^{\sigma-1}. \tag{26}$$

Finally, we use equations (25) and (26) to solve for the aggregate price index, $\bar{P}$, and the mass of varieties, $\bar{M}$:[51]

$$\bar{P} = (\sigma f)^{\frac{1}{\sigma-1}} \frac{1}{\rho\varphi_0}, \qquad \bar{M} = \frac{\bar{P}^{1-\sigma} - \tau_B^{1-\sigma}}{(\rho\bar{\tilde{\varphi}})^{\sigma-1}}.$$

Note that $\bar{\tilde{\varphi}}$ is *not* a function of the mass of varieties, despite the fact that M_0, M_n^w, M_n, and $\bar{M}$ appear in the definition of $\bar{\tilde{\varphi}}$ in equation (24). To see why this is true, note that $(M_n^w + M_n)$ and M_n can be written as a function of the mass of domestic varieties, M_0, the cut-off, φ_0, and the parameters of the model. M_0 can then be factored so that $\bar{\tilde{\varphi}}$ is a function of the ratio $M_0/\bar{M}$, which also depends only on φ_0, and the parameters of the model. Therefore, φ_0 is the only endogenous variable that determines $\bar{\tilde{\varphi}}$, $\bar{P}$, and $\bar{M}$. As discussed in section 2.2, the equilibrium value of φ_0 is determined by the free-entry condition, $\tilde{\pi} = f_e \left(1 - G(\varphi_0)\right)^{-1}$, and the expression for mean retailer profits in equation (2).

[49] That is, the aggregate price index is $\bar{P} = \left(P^{1-\sigma} + \tau_B^{1-\sigma}\right)^{1/(1-\sigma)}$, and the retail price index is $P = \left(P_0^{1-\sigma} + \sum_{n=1}^{N} \left(P_n^{w\,1-\sigma} + P_n{}^{1-\sigma}\right)\right)^{1/(1-\sigma)}$, where P_0 is the CES price index for domestic retail varieties, and P_n^w and P_n are, respectively, the CES price indexes for indirect and direct imported varieties from country n.

[50] A detailed derivation of this representation of aggregate expenditures is provided in section 2.6 of the online appendix.

[51] A detailed derivation of the aggregate price index and the mass of varieties is provided in section 2.7 of the online appendix.

7 Empirical Appendix

7.1 Propensity Score Matching Technique

The results derived from OLS and FE specifications provide evidence that direct-importing retailers perform significantly better, in terms of sales, than other retailers during exchange-rate appreciations. In principle, this result could be driven by either a causal effect of exchange-rate appreciations on direct-importing-firm sales (relative to other firms), or by self-selection of high-performing retail firms into importing in response to appreciations. Both of these effects are consistent with the theoretical model derived in section 2. While firm-level fixed effects control for any time-invariant components of self-selection bias, changes over time at the firm level could lead to self-selection into importing during exchange-rate appreciations and, therefore, stand in the way of causal interpretation for results from the OLS and FE regressions.

To identify the causal effect of exchange-rate changes on direct-importing-firm performance, we apply a PSM technique, taking motivation from recent work by Meinen and Raff (2018) and others. Our approach uses PSM to create matches between a treatment group of direct-importing firms and a control group of non-direct-importing retailers that have a similar propensity to import as the control group. We then estimate equation (16) using OLS on our matched treatment and control firms only, with control firms weighted equally to their matched treatment firms.

The first step in this technique involves estimating import propensity based on observables. We follow Meinen and Raff (2018) and employ the following probit model:

$$P\left(1(Importer_{i,t})\right) = \phi\left\{X_{i,t-1}\right\}, \tag{27}$$

where $\phi(\cdot)$ denotes the normal CDF and $X_{i,t-1}$ contains a set of firm-level variables, lagged by one year, that are significantly correlated with direct importing. In our application, the vector $X_{i,t}$ includes the following variables: the logarithm of labour productivity (value-added per employee), the logarithm of number of employees, wage share in total sales, and total sales growth. Labour productivity and employment are included in $X_{i,t}$ since both are positively associated with firm-level direct importing according to our model and other evidence. Wage share in total sales is included to capture the firm's cyclical position, and sales growth is included as an additional measure of productivity. Our probit regressions also include industry and year dummies. All variables included in the right-hand side of (27) are also included by Meinen and Raff (2018) in their analysis.[52] This exercise produces estimates for propensity scores that vary across firm-year observations in our sample.

To estimate (27), we require data for $labour\ productivity_{i,t}$, $number\ of\ employees_{i,t}$, $wage\ share_{i,t}$, and $sales\ growth_{i,t}$. We define $labour\ productivity_{i,t}$ = $value\text{-}added_{i,t}$/($number\ of\ employees_{i,t}$), where $number\ of\ employees_{i,t}$ = ($wages\ and\ salaries_{i,t}$) / $wage_{i,t}$; $wages\ and\ salaries_{i,t}$ is derived from the income statement of the firm, as reported to tax authorities, and reported in the National Accounts Longitudinal Microdata File (NALMF); and $wage_{i,t}$ = ($total\ firm\ payroll_{i,t}$) / ($average\ number\ of\ employees_{i,t}$), where both $total\ firm\ payroll_{i,t}$ and $average\ number\ of\ employees_{i,t}$ are derived from Payroll Deductions and Remittances data, as reported to tax authorities, and reported in NALMF. We define $wage\ share_{i,t}$ = ($wages\ and\ salaries_{i,t}$)/($total\ sales_{i,t}$). Finally, $sales\ growth_{i,t}$ is defined as annual change in $total\ sales_{i,t}$. Values of $total\ sales_{i,t}$, $value\text{-}added_{i,t}$, $wages\ and\ salaries_{i,t}$, and $wage_{i,t}$ are deflated using industry-level CPIs. We drop all values that are less than 0 and more than five standard deviations above or below the median.

We perform a nearest neighbour matching (with replacement) based on the estimated propensity scores to generate probability weights, to be used for equal weighting of treatment and control groups.[53]

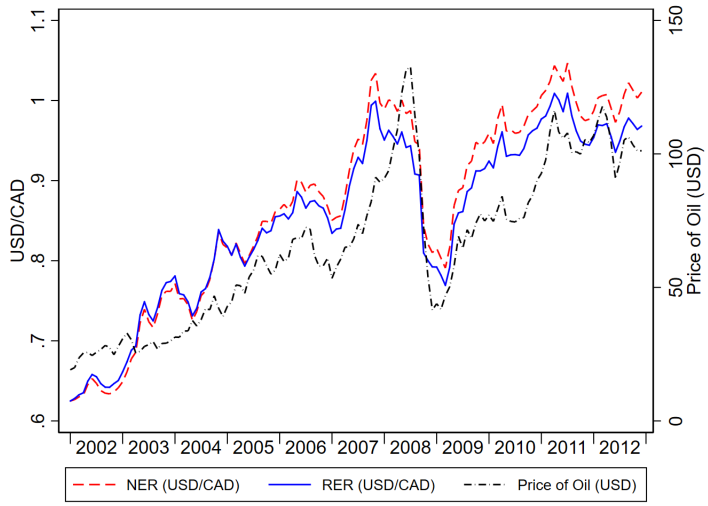
Finally, we estimate regression equation (16) by OLS, with standard errors clustered at the firm level, weighted by our estimated probability weights.

[52]Meinen and Raff (2018) also include the logarithm of labour productivity squared. We exclude this variable since it was not found to be statistically significant in the estimation.

[53]We modify propensity scores to have non-overlapping coverage for each year, to ensure that any two observations with the same propensity score fall in a common year. To ensure common support, we delete all observations where propensity scores for the treatment group (importers) fall outside of the estimated range found for the control group (non-importers).

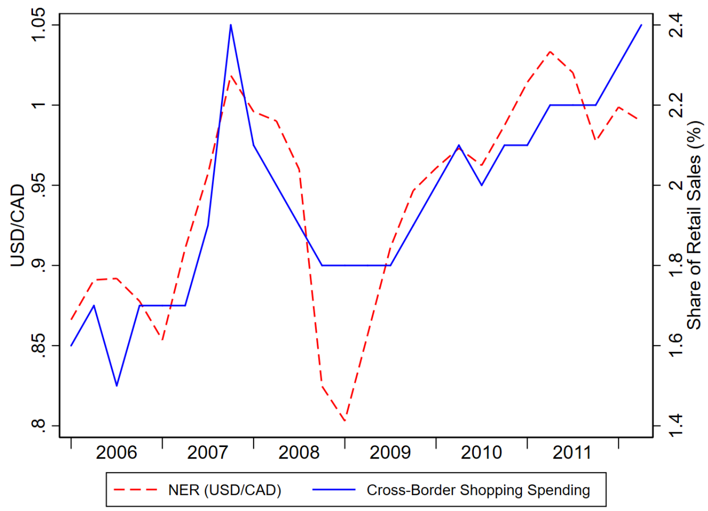
7.2 Figures

Figure 2: Canada-U.S. Exchange-Rate Dynamics, 2002–2012



Values for the Canada-U.S. nominal exchange rate are derived from Bank of Canada data for the monthly average spot rate. Values for the Canada-U.S. real exchange rate are produced from the nominal exchange rate deflated by the ratio of the Canada CPI to U.S. CPI, produced by Statistics Canada and the Bureau of Labor Statistics, respectively. Values for the price of oil are derived from monthly averages for the price of Brent crude taken from the International Monetary Fund's International Financial Statistics.

Figure 3: Canada-U.S. Exchange Rate and Cross-Border Shopping, 2006 Q1–2012 Q4



Values for the ratio of cross-border shopping expenditures to Canadian retail expenditures are reproduced from the "high spending" scenario in Corbi (2014). Values for the nominal Canada-U.S. exchange rate are derived from Bank of Canada data for the monthly average spot rate.