

The Law of One Price in Unfamiliar Places: The Case of International Real Estate*

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October 27, 2011

Abstract

International commodity market arbitrage and its corollary, the law of one price, are generally based on moving the commodity between markets to exploit price differences, making allowance for prevailing exchange rates. This form of arbitrage is clearly impossible for services and immobile objects such as real estate. However, there is the possibility of another form of arbitrage when the buyer can move to the object instead of the object moving to the buyer. Such a situation is possible for the market for international real estate such as recreational properties in exclusive mountain ski resorts and oceanfront estates. By constructing repeat sales indices for several such markets, we find that the price dynamics of international properties differs from local properties in terms of the impact of exchange rates. We show there is a significant long-term equilibrium relationship between the levels of relative prices and exchange rates. Rising values of a country's currency dampen local currency prices of internationally-traded properties relative to domestic properties, and vice versa. Lastly, we provide direct evidence in support of the law of one price for the case of international real estate. (R32, F31)

Keywords: exchange rates; international real estate; law of one price; one-way arbitrage

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1 Introduction

The English language emphasizes the inflation compensation quality of *real* estate, referring to its relative fixity of value in terms of a broad basket of goods and services. In the French and German languages it is the fixity of location that is emphasized, that is, the quality of its immobility, the French and German equivalents for real estate being respectively *immobilier* and *Immobilie*. It is the fixity of location that makes each piece of real estate unique, and which prevents arbitrage in its common form of buying where an item is cheap and selling where it is expensive and thereby bringing disparate prices closer together. However, it is not necessary to buy, move and sell an object to bring prices together. Rather, choosing between alternative suppliers of the same or similar objects can lead prices to comparable levels in an adjustment process that might be referred to as ‘one-way arbitrage’, perhaps relying on the process of price discovery that Leon Walras called *tatonnement*.¹

Just as one kilo of Granny Smith apples will be priced more or less the same as another kilo on nearby stalls in the same market, so will similar housing lots on the same street or even in different but otherwise comparable neighborhoods. Nevertheless, as the distance between alternative properties increases, price disparities grow. Even one-way arbitrage will have limited effect driving housing lot prices closer together when these lots are in distant cities. Real estate prices in different countries would seem to be even more independent, with no force driving anything close to the law of one price whereby relative local currency prices are reflective of exchange rates.

However, this paper argues that there are markets where real estate prices are connected to exchange rates, these being the international markets for luxury properties in exclusive mountain ski resorts and oceanfront estates, and housing that straddles some international borders. The prices of such properties should differ from local, segmented properties in terms of the impact of exchange rates. Rising values of a country’s currency, by making properties more expensive to potential foreign buyers, should dampen demand and local currency prices of internationally traded properties relative to purely domestic properties. The adjustment of local prices is necessary to induce local buyers to pick up the slack left by the decline in the interest of foreign buyers. Similarly, declining values of a country’s currency raises demand and local currency prices of international relative to domestic properties. This induces local buyers to free up some supply. Exchange rate changes should also affect local currency prices in international properties in one country

¹For an explanation and application of one-way arbitrage in the context of covered interest arbitrage see Deardorff (1979).

versus another, rising in the country with the depreciated currency and falling in the country with an appreciated currency, or at least a change in relative prices. The more international the participation in a particular property market the larger is likely the price effect of exchange rates.

To investigate our hypothesis, we conduct a number of empirical tests on a collection of real estate transaction data mainly from areas in the state of Washington on the U.S. side of the border and the neighboring province of British Columbia on the Canadian side of the border. The list of areas includes famous ski resorts, oceanfront properties, and a number of local control markets. We also examine the very unusual situation of a property market spanning the U.S. - Canadian border where the only access to U.S. homes is via Canada. We can judge the relative appeal of different areas in part by their recent proportion of foreign ownership, allowing us to classify those areas with a non-negligible share of foreign owners as international markets.

We construct so-called relative price indices (RPI), which are the ratios of real estate price indices of two distinct geographical areas. The two areas of comparison can be in the same country, e.g. an area with foreign ownership such as an ocean-front community relative to a local benchmark with little or no foreign participation such as a close-by city. Alternatively, we can directly compare two similar areas of interest on different sides of the border with each other (e.g. a Canadian and a U.S. ski resort).

We apply two econometric frameworks to the data that emphasize different aspects of the question as to whether real estate prices in certain areas are influenced by movements in the exchange rate between the local currency and the currency of potential foreign buyers in agreement with the law of one price.² First, we investigate if *year-over-year changes* in relative prices can be in part explained by recent changes in the exchange rate while controlling for other factors. This approach imposes little structure, but has the advantage that it will reliably apply even to shorter time series. We find strong evidence that changes in relative prices of international versus local areas are strongly affected by recent changes in the exchange rate in the hypothesized way on each side of the border, while they are not in the case of control areas. The evidence is weaker, but still present when comparing international areas across borders directly.

Second, we check within a vector error correction model (VECM) if there is a significant long-term equilibrium relationship between *the levels* of relative prices and exchange rates. In other words, we consider whether there is evidence of cointegration. We find that

²Not surprisingly, foreign buyers on the Canadian side are overwhelmingly from the U.S., while most buyers on the U.S. side are from Canada. This makes the exchange rate between the Canadian dollar and the U.S. dollar a natural candidate to focus on.

prices of some international markets relative to local benchmarks exhibit a long-term relationship with exchange rates, while none of the control (non-international) markets do. Using relative prices of international markets on each side of the border, the results are even more striking, with all international market comparisons showing evidence of cointegration, while none of the comparisons of non-international markets do. Lastly, within the VECM framework we can directly test the law of one price (LOP) for the case of internationally traded properties and find that the hypothesis that LOP holds cannot be rejected for several international cross-border markets.

The paper is structured as follows. Section 2 provides a theoretical context of the impact of exchange rates on prices under market clearing. Section 3 discusses our contribution within the existing literature. Section 4 details the sources of all data involved and the cleaning process of the individual sales data. Section 5 describes the construction of the price indices and the empirical frameworks used, while Section 6 presents the empirical analysis. Lastly, Section 7 concludes.

2 Motivation

All over North America, housing prices have experienced major ups and downs since the 1990s. Fueled by low interest rates and economic expansion, the trend of the 1990s and early 2000s came to a crashing halt when massive sub-prime mortgage defaults triggered a global financial crisis. Yet the ups and downs have not occurred homogeneously across all markets. For example in recent years, prices in the prime ski resort of Whistler, B.C. have gone against overall price trends of nearby Vancouver properties.

What has caused this divergence? Given Whistler's prominence in the Winter Olympics of 2010, one might actually suspect that Whistler would have performed better than the nearby Vancouver market throughout the price cycle, not frequently worse. However, Whistler represents a relatively international market with around 20 percent of properties belonging to foreign owners, the overwhelming majority of whom are from the U.S.

Actual and prospective U.S. owners care about the value of Canadian properties in terms of the U.S. dollar. In the last few years the Canadian dollar has shown enormous strength against other currencies, especially the U.S. dollar. From a low of 62 U.S. cents at the beginning of 2002, it has traded above parity with the U.S. dollar on several occasions between 2007 and 2011. What does this mean for prospective buyers/current owners? From a Canadian perspective, property values in Whistler have risen less than other domestic markets, while from a U.S. perspective, Whistler prices still gained substantially.

The exchange rate can account for the difference.

2.1 The demand for international properties

It seems reasonable to assume that the demand for properties in an international real estate market is partially determined by the exchange rate that translates the price of the property into the buyer's home currency. Imagine a two-country world in which properties in a particular market M_I in the subject, or "home country" appeal to potential buyers from both countries. Assume that foreign demand for properties in the subject market M_I follows a certain demand schedule vis à vis the price translated into the foreign currency. At the same time the domestic demand schedule is in terms of the domestic currency.

If the home currency appreciates against the foreign currency, any property in the home subject market will become more expensive in terms of foreign currency. As a result foreign demand will fall along the foreign demand schedule, while domestic demand remains unchanged, at least initially. The decrease in total demand should cause a decline in home currency price until the sum of a lower quantity of foreign demand and an increasing quantity of domestic demand equals the existing housing stock.

The new equilibrium price after a change in the exchange rate will depend on the domestic and foreign shares of market demand as well as the price elasticities of demand of both countries' buyers. In order to see how, assume a fixed supply $\bar{Q}$ equal to domestic and foreign demand, Q and Q^* , which depend on the price of property in the buyer's currency. The foreign price is the domestic currency price p translated by the exchange rate π . Demand also depends on domestic and foreign income (or wealth) of prospective buyers, y and y^* . In equilibrium we have

$$Q(p, y) + Q^*(\pi p, y^*) = \bar{Q}.$$

Total differentiation maintaining equilibrium yields

$$\begin{aligned} \frac{\partial Q}{\partial p} dp + \frac{\partial Q}{\partial y} dy + \frac{\partial Q^*}{\partial \pi p} (\pi dp + p d\pi) + \frac{\partial Q^*}{\partial y^*} dy^* &= 0 \\ -Q|\eta|\dot{p} + Q|\omega|\dot{y} - Q^*|\eta^*|\dot{p} - Q^*|\eta^*|\dot{\pi} + Q^*|\omega^*|\dot{y}^* &= 0 \end{aligned}$$

where dots signify rates of change and the Greek characters in absolute values are elas-

ticities of demand with respect to price and wealth, specifically

$$\begin{aligned} |\eta| &= -\frac{p}{Q} \frac{\partial Q}{\partial p} & |\eta^*| &= -\frac{\pi p}{Q^*} \frac{\partial Q^*}{\partial \pi p} \\ |\omega| &= \frac{y}{Q} \frac{\partial Q}{\partial y} & |\omega^*| &= \frac{y^*}{Q^*} \frac{\partial Q^*}{\partial y^*}. \end{aligned}$$

Solving for changes in price yields

$$\dot{p} = -\frac{Q^*|\eta^*|}{Q^*|\eta^*| + Q|\eta|} \dot{\pi} + \frac{Q|\omega|}{Q^*|\eta^*| + Q|\eta|} \dot{y} + \frac{Q^*|\omega^*|}{Q^*|\eta^*| + Q|\eta|} \dot{y}^*. \quad (1)$$

Neglecting the elasticities for a moment, this equation has an intuitive structure. The negative influence of appreciation of the subject country's currency on local property prices depends on the relative share of foreign demand in the market; the larger the foreign share the larger the effects of the exchange rate. The same holds for the influence of changes in domestic and foreign income or wealth. In addition, the elasticities act as multipliers and can strengthen or subdue the influence of domestic versus foreign market share and income on prices. It seems reasonable to assume that foreign demand is generally a lot more elastic than is domestic demand. Foreign individuals that look for property abroad can choose among many areas (e.g. ski resorts) in more than one country. Under these assumptions the relative impact of the exchange rate could be magnified beyond the pure share of foreign ownership.

Similarly, we can broaden the meaning of "foreign demand". Imagine a potential domestic buyer in the home country is interested in acquiring property in an international market. Even though he is not actually a foreigner, in his mind properties in market M_I still compete with other international property markets in other countries. For instance, for a wealthy buyer from New York, a property in Aspen, Colorado may be directly competing with one in Switzerland, assuming that a slightly more expensive flight and a few more hours of travel time are not of major concern. Then, even though the target property is priced in home currency, the exchange rate may still affect his demand via alternatives in other countries. Thus, instead of dividing demand into domestic and foreign, one might divide it into local vs. international demand, where the former describes potential buyers who are bound to the local market because of employment, family and social ties. Buyers in the international market are not so constrained.

To achieve an exact correspondence with the empirical setup later on, we can take an additional step and consider the relative price change between the international market M_I and a strictly local market M_L that lies in close geographical proximity and for which

no foreign or international demand exists. With $Q^* = 0$ in the local market, Equation 1 collapses to

$$\dot{p}_L = \frac{|\omega_L|}{|\eta_L|} \dot{y}_L$$

where the subscript L denotes the respective variables for the local market. Then, for small changes (or in a log specification) it holds that

$$\begin{aligned} \dot{p} - \dot{p}_L &\approx -\frac{Q^*|\eta^*|}{Q^*|\eta^*| + Q|\eta|} \dot{\pi} + \frac{Q|\omega|}{Q^*|\eta^*| + Q|\eta|} \dot{y} - \frac{|\omega_L|}{|\eta_L|} \dot{y}_L + \frac{Q^*|\omega^*|}{Q^*|\eta^*| + Q|\eta|} \dot{y}^* \\ &\approx -\frac{Q^*|\eta^*|}{Q^*|\eta^*| + Q|\eta|} \dot{\pi} + \left[\frac{Q|\omega|}{Q^*|\eta^*| + Q|\eta|} - \frac{|\omega_L|}{|\eta_L|} \right] \dot{y} + \frac{Q^*|\omega^*|}{Q^*|\eta^*| + Q|\eta|} \dot{y}^* \end{aligned} \quad (2)$$

where the second approximation is valid under the reasonable assumption that changes in income for local residents are highly correlated for the two markets. Consequently, our argument on relative share of ownership and elasticities carries over from absolute prices to prices relative to a local benchmark. Clearly, this model leaves out many other potential influences on house prices. However, the relative specification is particularly useful here because the impact of other factors will be weakened or eliminated to the extent that their effects are common to the two related markets.

Lastly, one can also compare prices of similar markets in the home and the foreign country, say M_I and M_I^* . We see that in this case the effects of the exchange rate on the relative price change $\dot{p} - \dot{p}^*$ (each in their home currency) are amplified by going in opposite directions while the effects from wealth or income are weakened.

In what follows this paper will investigate if there is empirical evidence that exchange rates are an important determinant of housing prices for markets in which foreign buyers own a non-negligible share of property. We also investigate whether at the same time exchange rates do not matter for purely local markets for which there is no interest by foreign buyers.

3 Literature

Our study is located at the intersection of two large strands of literature. On the one hand, we contribute to the literature that explains house price dynamics by suggesting a new factor, the exchange rate, that can play a major role for a particular set of real estate markets. On the other hand, we contribute to the expansive literature on purchasing

power parity (PPP) and the law of one price (LOP) by finding that the relative version of LOP can be successfully applied to a commodity that one generally would not associate with the notion of arbitrage, namely what we have called international real estate.

3.1 Determinants of house prices

Determining potential economic determinants of house prices is, not surprisingly, a central topic in the vast and growing real estate literature. For instance, Case and Shiller (1990) provide evidence that per capita income, population growth and construction costs help explain house prices. Peek and Wilcox (1991) find significant effects for real after-tax interest rates, unemployment and demographic factors. Lamont and Stein (1999) find that leverage increases sensitivity of house prices to fundamental shocks. Further, Case and Shiller (1989, 1990) establish that house prices exhibit predictability. Capozza, Hendershott, and Mack (2004) find that this predictability changes with city size, income growth, population growth and construction costs. Gyourko, Mayer, and Sinai (2006) show that the increase in national income inequality and the inelastic supply of land in attractive locations lead to different growth patterns in house prices across the United States.

The literature that specifically investigates the connection between exchange rates and real estate prices is almost non-existent. Bensen et al. (1997, 1999) relate the CAD/USD exchange rate to property prices in two areas just south of the Canadian border in Washington State, which are also part of our sample. Bensen et al. (1997) find that the level of house prices in Point Roberts, WA - a small patch of the United States accessible only through Canada - correlate positively with the exchange rate and with Vancouver, B.C. house prices over a 10 year period from 1984 to 1994. However, the analysis is done in levels only. The regression of non-stationary time series raises econometric issues and makes the interpretation of the results difficult. In addition, their Exhibit 4 shows that the prices of all of Whatcom County, WA homes - of which Point Roberts is a very small part - increase about as much as Point Roberts prices, thus challenging the evidence for the exchange rate having been responsible. Bensen et al. (1999) conduct a similar analysis for the Bellingham, WA housing market over the same time period and find a positive effect, again for the level of the exchange rate with accompanying non-stationarity issues. According to our data, the total foreign ownership share within Bellingham is around 1 percent. Canadian buyers are thus unlikely to have had an effect.

We are careful to account for the non-stationary nature of exchange rates and house prices in both our econometric frameworks, either by explicitly accounting for potential cointegration or by using year-over-year changes rather than price levels. Second, by using

price changes in excess of price movements of nearby local markets or other international markets, we neutralize common trends to help avoid false positives.

3.2 Law of one price in real estate markets

For a historical perspective on the literature on purchasing power parity, PPP, and international dimension of the law of one price, LOP, we refer the reader to the literature reviews by Rogoff (1996) and Taylor (2006). The general tenor of this literature is that, while the notion that traded goods prices should be equal across countries is highly intuitive, evidence for PPP in its absolute version is hard to establish, even in the long run. Transportation costs, duties, taxes, non-traded inputs, especially labor and additional service components like warranties segment markets. In fact, Engel and Rogers (1996) and Parsley and Wei (1996) find that price differences exist not only across country borders, but also between U.S. cities with price discrepancies increasing with distance.

Generally, empirical studies have tested the relative version of PPP in aggregate price levels or LOP in particular goods, i.e. the notion that relative *changes* in prices offset *changes* in the exchange rate. The evidence for relative PPP focusing on rates of change is much stronger than the evidence for exchange rate levels, but depending on the good, it may take considerable time until the adjustment in prices has been made to offset movements in exchange rates.

Even beyond time for adjustment, prices of the same good may change by different degrees between countries for reasons that are independent of the cost of providing the good and exchange rates, such as strategic considerations by firms in competition with others or with power to price discriminate. These phenomena have been referred to as pricing-to-market (Krugman, 1987) leading to incomplete pass-through of exchange rates. Goldberg and Knetter (1997) provide a thorough overview of the literature on LOP and the connection with incomplete pass-through and pricing-to-market.

Previous tests of LOP for specific goods include Goldberg and Verboven (2005) who test the convergence of car prices across Europe. Asplund and Friberg (2001) compare prices payable in either of 2 currencies for the same goods in the same location, Scandinavian duty-free stores, and find that price deviations outside some “band of inaction” are quickly corrected. This result corresponds for instance to the findings of Yeyati et al. (2009) who test LOP for equities that are listed in different countries and find that price deviations are arbitrated away quickly once they exceed narrow non-arbitrage bands.

The parts of our empirical investigation of similar areas in different countries are a direct test of relative LOP for one particular type of good, that of internationally traded

real estate. Unmovable pieces of property are not what one would normally investigate when seeking support for the law of one price. In a financial context, arbitrage involves selling a security at one price while simultaneously buying it at a different price. In an international trade context, we can also think of one-way arbitrage, whereby potential buyers choose to buy a good from the cheapest location (after accounting for exchanges rates, transaction costs etc.) until prices have converged. Real estate is generally thought to be immune from this kind of arbitrage. The old saying that ‘all real estate is local’ expresses the multitude of frictions that keep property prices from say, Detroit even remotely approaching those of New York City. Instead, prices respond to the availability of jobs, incomes, local amenities, infrastructure, climate, energy costs, scarcity of land, taxes and so on.

We suggest that properties that are appealing to buyers of second (or more) properties are not ‘local’, especially when they have certain attributes (e.g. waterfront or world-renown ski hills). Given the relatively low cost of global travel in modern times, wealthy international home buyers can potentially choose from many locations around the world. Foreign demand should be less dependent on local economic conditions or borrowing constraints and more affected by the appeal of local amenities and even the social status that comes with owning a place in a well-known international destination.

4 Description of data

4.1 Canadian house price data

We acquire most of the Canadian property data from Landcor Corp., the local vendor of assessment and sales data for the province of British Columbia. Records of sales begin between 1971 and 1976 depending on the area. Whenever possible, we use sales up until the end of June 2011. The data include information about the type and use of the property, sales date and price, a history of permits issued and some limited property attributes. For the larger control areas, we use sales data that ends in 2005 or 2007 and append public house price index data from the local real estate boards that are based on the implied sales price of a typical home within each area. The most recent ownership data used to judge which markets are international versus local are from July 2011 and contain the address of each property as well as the mailing address of the owner that is used to send out property tax assessments.

4.2 U.S. house price data

On the U.S. side, ownership data are derived from the assessment rolls of Whatcom County (December 2009) - the location of Mount Baker ski resort - and San Juan County (June 2011) in Washington State provided by the respective county assessor offices. We acquire sales data from a local real estate service firm (Real Market Data LLC) that has been recording sales activity in those counties since 1984. In the case of Whatcom County, we also use data on building permits provided by the county assessor office. Sales (1971 - 2011) and ownership (as of 2011) data for the famous Aspen ski resort town were acquired from the Pitkin County Assessor's Office. Repeat-sales house price indices for additional control markets such as Washington State, Seattle (WA) and Denver (CO) are taken from the website of the Federal Housing Finance Agency (FHFA)³.

4.3 Additional data sources

The exchange rate between the Canadian and the United States Dollars (price of 1 USD in CAD) and the Federal Reserve dollar index are taken from DATASTREAM. Average per capita or family income data for various levels of aggregation (state, county, metropolitan statistical area) are downloaded from the websites of Statistics Canada and the St. Louis Federal Reserve. We employ annual data on the top percentiles of wage earners in the U.S. (based on work by Piketty and Saez (2003), updated up to 2008), in Canada and the province of B.C. (based on work by Saez and Veall (2005), and Veall (2010)). In addition, we construct a measure of what we call mortgage debt capacity (MDC), i.e. given income, maximum amortization period and current mortgage rates, what is the amount that can be borrowed so that monthly mortgage payments do not exceed a certain threshold of gross income?⁴ Lastly, we collect monthly unemployment rates for the states of Washington and Colorado, the province of B.C. as well as the national U.S. rate. Annual variables are interpolated by splines to yield quarterly data points, while for monthly data, we simply use the observation from the last month of each quarter. All variables representing dollar amounts are in nominal terms.

4.4 Repeat Sales Indices

We construct Repeat Sales Indices (RSI) following the flexible Fourier approach of McMillen and Dombrow (2001) which is discussed in more detail later. For each market, separate in-

³Available at www.fhfa.gov

⁴McQuinn and O'Reilly (2008) demonstrate that this measure is well suited to explain house prices.

dices are computed for condominiums and/or single family dwellings, whenever sufficient transaction data are available.

It is well-known that all repeat-sales methodologies are subject to sample selection issues (e.g. Gatzlaff and Haurin, 1998). Additional biases can arise due to heterogeneity in times between sales (Goodman and Thibodeau, 1998) and the frequency of resales (Clapp and Giaccotto, 1999). In the spirit of these findings, we filter along a number of dimensions to limit the impact of any of these effects. We filter by transaction type ('improved', i.e. not vacant) where available, and impose a minimum price floor (in real terms of 2005 dollars) by area and type to make the sample more homogenous. Sales pairs that lie less than 6 months apart are discarded (likely flips), as were transactions for which more than one sales price was recorded on the same date (potential data errors).

To increase the probability that the quality of the property remained constant between sales, the following filters are applied. If a building permit was issued between two transactions for a particular property, the transaction pair is discarded from the sample. Because the history of permits is not necessarily complete for some areas, large changes in price from one sale to the next may be due to some unobserved change in property characteristics. We delete pairs from the sample when the price increases (decreases) by more than 100% (50%) in real terms for transactions within one year, and more than 150% (60%) within 2 years. Beyond that, increases (decreases) of more than 200% (67%) in real terms are dropped from the sample. Finally, we delete transactions if more than 10 years has passed since the previous transaction for areas for which we do not have permit data.

4.5 Sample description

[Table 1 about here.]

Table 1 lists the property markets for which we have individual sales data, and with few exceptions, ownership data as well. We divide markets by country and further classify them as either 'local' or 'international', the latter being markets that can be considered to hold some appeal to foreign buyers. In the context of our sample, these are mostly famous ski resorts on both sides of the border and some waterfront communities that are more expensive than local markets and potentially have a large share of foreign ownership. Except for two control markets for which foreign ownership is likely negligible, we have address data of the owners allowing us to estimate the share of foreign and out of province/state owners for each location. These addresses are used by the local assessor

offices to send out property tax bills. We also list the proportion of properties for which the address is a post office box. It is quite common for foreigners to use local intermediaries who in turn use PO boxes. In general, the true foreign ownership share is likely higher than that which is visible. This holds in particular for the international locations in the United States.

Notable in Table 1 is the large cross-sectional dispersion in the percentage share of foreign ownership, reaching from essentially zero percent for some local markets to an astonishing 56 percent for Point Roberts, WA. As can be seen in the third-to-last column, foreign ownership mostly means U.S. owned for Canadian properties and vice versa. The exception is Aspen, Colorado where the already low visible foreign ownership share is dispersed among many foreign countries. However, this does not mean that Aspen should be considered a local market. Rather, aside from the potentially high share of hidden foreign owners (PO boxes) we find an extremely high share of out-of-state owners. As explained in the previous section, an owner from the Eastern United States essentially has the same choices as a foreign owner, i.e. Aspen is not much closer than European resorts. That in turn means that even a potential buyer in the same country as the area of interest will be sensitive to exchange rates when there are alternatives outside of the country.

To help the reader unfamiliar with this sample of property markets, which with one exception are all situated either in the Canadian province of British Columbia or the adjacent state of Washington in the United States, let us give some details on the individual areas and our motivation for including them. The international locations are mostly large and well-known ski resorts (Big White, B.C.; Whistler, B.C.; Mt. Baker, WA; and Aspen, CO) that attract non-local home buyers, as is apparent from Table 1. The local markets are typically cities in the immediate vicinity of these international markets, such as Vancouver (for Whistler) or Bellingham (for Mt. Baker).

We include the remaining areas for their particular location and features. West Vancouver, a separate municipality close to Vancouver, is added because it is a very expensive property market with ocean-front estates that should potentially be interesting to foreigners, even though there is no visible foreign ownership. We use Abbotsford, B.C. in a cross-border comparison with Bellingham, WA because the two cities are of similar size and are a mere 30 miles apart with the U.S.-Canadian border inbetween. Second, the Canadian Gulf Islands and U.S. San Juan Islands are part of the same set of small islands in the Pacific Northwest, often no further apart than a few miles. It just so happens that the U.S.-Canadian border divides them in two. We further split the Gulf Islands

into waterfront properties and those without access to the water to investigate if even within markets there are differences with regards to the impact of the exchange rate; waterfront properties are more likely international. Finally, the oceanfront communities of Tsawwassen, B.C. and Pt. Roberts, WA border on each other with no distance inbetween them at all. These pairs seem ideal to test if prices for similar property markets in close proximity move together (accounting for the exchange rate). We argue in this paper that the mechanism by which this adjustment can occur is foreign demand that views similar markets in different sides of the border as close substitutes. Our hypothesis is that *ceteris paribus*, changes in the CAD/USD exchange rate should impact prices of real estate in those markets with a degree of foreign ownership, but not those where properties are almost entirely held domestically.

5 Methodology

5.1 Repeat Sales Indices

We start by estimating repeat sales indices (RSI) for the housing markets in the sample employing the Fourier approach suggested by McMillen and Dombrow (2001) (henceforth MD). MD motivate their approach by the observation that house prices do not jump abruptly, but move rather smoothly over time in a way that makes them suitable to be approximated by a sequence of Fourier terms. Further, as time becomes a continuous variable the approach uses the information contained in the time of sale more efficiently than traditional approaches which assume that sales on the first of the quarter have more in common with sales at the end of that quarter than with sales on the previous day.

For our purposes the MD approach has the additional benefit that sparse sales data do not lead to missing observations in the price index, a point that is especially important for the small markets in our sample. This allows us to conduct the empirical part at the higher, quarterly frequency than would be possible using the standard repeat sales approach or even a hedonic approach. The downside to this method is that it introduces some serial correlation into the index because the functions used to fit the data are smooth sine and cosine functions. We correct for this problem by using Newey and West (1987) corrected standard errors, where the optimal lag length is determined following Newey and West (1994). The approach of MD leads to a regression setup with the log change in the property price between times of sale $t > s$ on the left and a sequence of Fourier terms

on the right:

$$\begin{aligned} \log(P_{i,t}/P_{i,s}) = & \alpha_1(z_{i,t} - z_{i,s}) + \alpha_2(z_{i,t}^2 - z_{i,s}^2) \\ & + \sum_{q=1}^Q [\lambda_q(\sin(qz_{i,t}) - \sin(qz_{i,s})) + \gamma_q(\cos(qz_{i,t}) - \cos(qz_{i,s}))] + \epsilon_{i,t,s} \end{aligned} \quad (3)$$

where z are transforms of times of sales over the full period $[T_0, T_1]$, i.e. $z_{i,\tau} = 2\pi \frac{\tau - T_0}{T_1 - T_0} \in [0, 2\pi]$. As for the optimal number of Fourier terms Q , MD suggest to use the Schwarz Information Criterion (SIC) for guidance

$$SIC(Q) = \log\left(\frac{1}{N} \sum_{i=1}^N \epsilon_i^2\right) + 2 \frac{\log(N)}{N} (Q + 1)$$

In the context of their data, MD choose very small values for Q based on the SIC. Having markets with up to 40 years of data and over 10,000 repeat sales in our sample, the optimal Q is generally much larger with values above 20. However, for the particularly sparse markets, such as Pt. Roberts and Mt. Baker, the algorithm chooses $Q = 3$ or 4 because higher values of Q are heavily penalized by the SIC given the small number of observations. Visual inspection reveals that the resulting price indices are too smooth and not able to capture the true variation of prices over time. There is no reason to believe that smaller markets are by nature smoother than large markets. Given the trade-off between capturing the fundamental dynamics of all markets in the sample and overfitting low-observation markets, we deviate from MD in the following way. We set a maximum value for $Q_{\max}$ that is used if a market extends over the entire sample period (1971Q1-2011Q2). If a market covers only a certain part of the sample period, it is assigned a proportionately smaller Q , rounded to the nearest integer. That way each repeat sales index has approximately the same degrees of freedom per unit time. Values of $Q_{\max} \in [14, 20]$ resulted in sensible and consistent graphs even for the smaller markets. All reported results are based on $Q_{\max} = 17$.

Because many of the markets we investigate are small, most fundamental factors employed in the existing literature are either not available at the disaggregated local level or they are available in widely spaced intervals, e.g. only in census years. Our key focus is the incremental effect of the exchange rate on local house prices beyond other fundamental factors that affect real estate prices more broadly. We thus choose to form a relative price index (RPI) that describes the relative performance of some particular property market of interest relative to a benchmark. For instance, on the Canadian side we may

evaluate the house prices of international property markets relative to the price index of the nearby City of Vancouver; on the U.S. side, we may evaluate prices of small markets within Washington State relative to the entire state of Washington. Thereby, factors that affect both markets equally are either cancelled out or they are significantly weakened if the effects are of different magnitudes. The first set of RPIs that we consider in our tests are within-country comparisons as the examples above. The second set of RPIs are cross-border comparisons between similar markets in different countries, e.g. between two ski resorts or between two markets that lie in close proximity to each other on different sides of the border. The relative price index (RPI) of market M relative to benchmark B at the end of quarter t is defined as the log ratio of the two price indices:

$$\text{RPI}_t = \log \left(\frac{\text{RSI}_{M,t}}{\text{RSI}_{B,t}} \right) \quad (4)$$

We employ two empirical methods to provide an answer to the question as to whether some property markets do indeed react to changes in the exchange rate via a foreign demand channel. Since our question is closely related to the law of one price (LOP), a natural way to investigate this problem is to look for a *long-run* relationship between prices and the exchange rate. As we know from the extant literature, prices may take considerable time to converge. Thus, a vector error correction model (VECM) where deviations from the long-run equilibrium affect price dynamics until equilibrium is restored seems well suited for the task. In particular, a VECM framework allows for a direct statistical test of LOP.

The downside to VECM is that tests of the long-run equilibrium require long time series and the crucial assumption that the equilibrium does not change over time. With up to 140 quarters of observations we are satisfying the first criterion. The assumption of no structural breaks is harder to maintain, in particular because it is unlikely that properties were bought across borders 30 years ago to the extent they are today. Anecdotally, we find that foreign ownership in the ski resort of Whistler in the year 1991 was approximately 5 percent, much lower than today. Therefore the hypothesized long-run relationship may have been weak in the early years of the sample. For this reason, we test each set of RPIs first in a less structural setting, simply asking if recent *changes* in the exchange rate affect *changes* in relative property prices, followed by the more formal VECM framework.

5.2 Univariate framework

To reduce the degrees of freedom, we opt for year-over-year changes in RPIs and exchange rates instead of quarterly differences. Second, we lag all independent variables by two quarters assuming that it takes some time before exchange rates or other factors can affect property prices via demand from foreign buyers. After all, even for local buyers and sellers, real estate transactions often take months to complete. It is possible that the exact time lag may be different from market to market. Indeed, in unreported results we find that the strongest effect for some cross-border market pairs that are far apart in distance occurs at a lag of one year, while for pairs that border each other directly, the effect seems to be more immediate. Nevertheless, we maintain a lag of two quarters throughout the analysis for consistency.

The dependent variable is (the logarithm of) the rolling 4-quarter change of house prices in one area relative to that in the benchmark area. The key independent variable is the logarithm of the rolling 4-quarter change in the CAD/USD exchange rate π lagged by $l = 2$ quarters.⁵ We also control for a small set of additional effects X (lagged by $k = 2$ quarters) such as income growth or unemployment:

$$\Delta_4 \text{RPI}_t = \alpha + \beta \Delta_4 \log(\pi_{t-l}) + \gamma' X_{t-l} + \epsilon_t \quad (5)$$

Using annual changes with quarterly data makes it necessary to adjust for serial correlation, which we do following Newey and West (1987) and choosing the optimal lag length as in Newey and West (1994) via GMM.

5.3 Error correction framework

We investigate the long-run relationship between relative prices indices (RPI) and the exchange rate in the setting of a vector error correction model (VECM). First, we test for the *existence* of a long-run relationship, asking if the two time series are cointegrated using the trace statistics by Johansen (1991). Assuming that there is such a relationship, we are interested in finding out which of the two series, prices or the exchange rate, react to any disequilibrium and undergo adjustment. In our context, it seems unlikely that a disequilibrium in relative house prices would causally affect the exchange rate. We therefore assume the exchange rate to be exogenous and test if it is prices that react to

⁵For some U.S. markets in our sample, the non-local demand is mostly not Canadian. Thus in these cases we consider it more appropriate to use the trade-weighted effective exchange rate against major U.S. trading partners, also known as the U.S. dollar index. However, results change little.

the disequilibrium, i.e. testing weak exogeneity.

More formally, define $y_t = (\text{RPI}_t, \log(\pi_t))' \in \mathbb{R}^2$, then the $\text{VECMX}(p,q)$ representation, including a set of control variables $X \in \mathbb{R}^k$, has the form

$$\Delta y_t = \sum_{i=1}^{p-1} \gamma_i \Delta y_{t-i} + \sum_{i=1}^{q-1} \delta_i X_{t-i} + A(B'y_{t-1} + \eta) + \epsilon_t \quad (6)$$

where the first row of $B \in \mathbb{R}^{2 \times 2}$ can be normalized to ones. If tests for cointegration are positive, the rank of both A and B is $r = 1$. Thus, 2×2 matrices A and B are reduced to vectors $(\alpha_1, \alpha_2)'$ and $(1, \beta)'$ and vector η collapses to a scalar c , and there is a single long-term equilibrium vector $(1, \beta, c)$ affecting changes in either variable through $\alpha_j < 0$, the rate of convergence back towards equilibrium. A test of weak exogeneity means testing the hypothesis $H_0 : \alpha_1 = 0$.

5.4 Tests of LOP

The VECM framework can be used to take the analysis one step further. Under cointegration the long-term equilibrium is

$$0 = (1, \beta, c)(\text{RPI}, \log(\pi), 1)' \quad (7)$$

Substituting (4), this solves to

$$\exp^{-c} = \frac{\text{RSI}_M \pi^\beta}{\text{RSI}_B} \quad (8)$$

In the case of a cross-border comparison of two similar markets, this equation has a very simple interpretation. If β is close to negative one (assuming that prices in M are in CAD and prices in B are in USD), this equation is the relative version of the law of one price: once prices are adjusted for the exchange rate, their ratio is constant over time. Thus, testing the hypothesis $H_0 : \beta = -1$ is a formal test of LOP for the case of international properties.

6 Results

6.1 Tests relative to domestic benchmarks

[Table 2 about here.]

We begin with the analysis of property markets of interest relative to a domestic, near-by benchmark market. Table 2 presents those results for the Canadian markets relative to the domestic control market Vancouver, B.C. The table is split into international markets, which according to Table 1 have substantial foreign ownership, and local markets, where foreign ownership is muted or non-existent. In this and the next two tables, we control for local and international income-related effects. At the local level we include what we have called mortgage debt capacity; given a fixed proportion of income, maximally allowed amortization period and prevailing mortgage rates, what size of mortgage can the average family afford? On the foreign side, we use the income threshold of the 99th-percentile of U.S. income earners. Year-over-year changes in both variables are added besides year-over-year changes in the exchange rate (all lagged by two quarters) in order to explain year-over-year changes in relative property prices between the two markets. The column labeled 'FX' presents the coefficient estimates on the exchange rate variable. We find that all international markets, but none of the local markets have significantly positive exposure to the exchange rate. This means that an increase in the price of a U.S. dollar in Canadian dollars, i.e. a decline in the Canadian dollar, leads to a significant appreciation in relative prices for all four international markets in local currency, but for none of the local markets, as hypothesized. Remarkable is the difference between the two Gulf Islands markets. While oceanfront properties strongly react to exchange rate changes, properties away from the water do not. This is even more surprising because the islands are quite small and many non-waterfront properties still have some ocean view. The difference, we suggest, lies in the fact that oceanfront properties are a somewhat segmented market attracting interest from outside the region and from abroad, while properties that do not have that 'oceanfront' feature do not. This is reflected in the foreign ownership shares of 7 vs. 2 percent.

[Table 3 about here.]

Table 3 presents the corresponding results for the U.S. markets. Here, the roles of local and foreign income variables are reversed. The first is now the MDC of the average family in the corresponding U.S. state, while foreign top income is based on the 99th-percentile of Canadian wage earners. Because we know from Table 1 that neither Aspen nor the San Juan Islands have any meaningful Canadian ownership, the CAD/USD exchange rate may not fully capture changes in foreign demand. For these two cases, we use the U.S. dollar index instead and we use top U.S. rather than Canadian income because of the relative higher out-of-state ownership.

Note that now the sign of the coefficient on the exchange rate should be negative for international markets. An increase in the U.S. dollar relative to the Canadian dollar should lead to less foreign demand for international properties and thus to a decline in the prices of those markets. With the exception of Aspen, this is what we find.⁶ Neither of the two local markets seem to be affected by the exchange rate.

Notable is that R^2 are quite high, particularly for the international markets. The coefficient on local MDC is highly significant in many markets. The negative sign is reasonable because average local income should be a major driver of the underlying benchmark market. Accounting for the capacity to borrow strengthens the significance because most principal residence purchases are financed, whereas purchases of recreational properties making up of the international markets are less dependent on the financing constraints of average households. The financial constraints of buyers of second or third homes are much more likely tied to things like global stock markets than to average incomes.

Rather puzzling, however, is the negative coefficient on top foreign income. Here we would have expected a positive sign. After all, foreign buyers should increase demand for international properties with rising incomes, leaving local prices unaffected. To make sure our results are robust, we repeat the analysis using local and foreign unemployment rates as alternative proxies for income or wealth effects. The results can be found in the appended Tables A.1 and A.2. We find that the effects of the exchange rate remain unchanged. Further, the sign of the coefficients on unemployment rates are as one would expect: a high level of local unemployment raises the relative price index because it depresses the benchmark market. High foreign unemployment likely limits foreign demand and thus depresses prices of international properties resulting in a negative loading on the RPI.

[Table 4 about here.]

Table 4 shows the results from the estimation of a vector error correction model (VECM) based on the same data used in the previous two tables. First, we conduct tests for cointegration of relative prices and the exchange rate. We consider 2 specifications of the VECM. ‘Case 2’ (based on common textbook treatments) is the specification shown in Equation 6, while ‘Case 3’ replaces the constant term η in the ECM term with a drift in the process.

Test results are based on the Johansen (1991) trace statistics, where the null (alternative) hypothesis is that there are at most (more than) r cointegrating vectors. In order

⁶We do not use the condominium index for Aspen here, because we were unable to find a benchmark condominium index for the U.S. with a sufficiently long history.

to conserve space, we do not show test statistics and critical values. Rather the numbers displayed are the difference between the test statistic and the critical value at the 5 percent significance level. Whenever the value shown is positive (negative) it means that we can (cannot) reject the null hypothesis that the rank of the error correction term is at most 0 or 1, respectively. Taken together in the context of a 2-dimensional test, rejecting $H_0 : r \leq 0$ and not rejecting $H_0 : r \leq 1$ means that the data support cointegration of vector y with one cointegrating vector $(1, \beta)$.

If either case supports cointegration, α and β are shown (with ‘Case 2’ being given preference, in case both show cointegration). α denotes the speed at which relative prices adjust to correct an existing disequilibrium between the levels of RPI and the exchange rate. For ease of interpretation the number displayed is annualized, e.g. 0.10 would mean that about 10 percent of the distance from equilibrium is corrected by changes in the RPI per year.

Table 4 shows that 5 out of 7 within-country pairs with ownership characteristics of an international market show signs of cointegration between relative property prices and the exchange rate; none of the control pairings with local markets do. With one exception, β estimates have the correct sign, i.e. RPIs of within-Canada pairs should have a negative β , whereas within-U.S. pairs a positive β given the definition of the exchange rate.

The second to last column of Table 4 reports p-values from a test of weak exogeneity for the relative price index, i.e. as to whether changes in RPI are affected by the disequilibrium, $H_0 : \alpha = 0$. The fact that p-values are generally quite low means that relative prices do react to disequilibrium. Again, we repeat the same analysis in Table A.3 using unemployment rates as controls and find that 4 out of 7 international RPIs show evidence of cointegration, all of them with the correct sign, while none of the purely local RPIs seem to be cointegrated.

6.2 Cross-border comparisons

[Table 5 about here.]

The preceding results provide evidence that prices in international real estate markets are affected by exchange rate movements, in the short-term via recent changes in the exchange rate and in the long run via reversion to some equilibrium level. We now focus on pairing up similar property markets directly to investigate if their relative prices are also affected by the exchange rate. For example, does a rise in the Canadian dollar, by making properties in Canadian ski resorts more expensive relative to U.S. ski resorts,

cause a decline in the local currency price of the Canadian market and a rise in the U.S. dollar price of the U.S. market? As before, we first investigate if this adjustment occurs in the short-run as a response to recent shocks to the exchange rate, and second if there is a long-run relationship between similar international markets. As controls, we also consider two cross-border pairs of comparable cities with city-wide little foreign ownership that are close to the border and to each other, namely Seattle, WA versus Vancouver, B.C. and Bellingham, WA versus Abbotsford, B.C. Are such pairs of local real estate markets unrelated with respect to the exchange rate, unlike the international markets?

Table 5 shows the univariate regression results from pairing up two comparable international or local markets on either side of the U.S.-Canadian border. In this case we control for year-over-year changes in affordability taking the form of average local income (at the state level) on the U.S. side of the border and the previously described local mortgage debt capacity (MDC) on the Canadian side.⁷

The results are weaker than for within-country pairs of international versus local markets. Nevertheless, for 3 out of 6 international pairs we find some limited effect of recent exchange rate changes on relative prices, while as expected the two control pairs show no signs of significance at all.

The weakness in the results for international pairs could be due to lack of availability of more suitable comparisons. For instance, the pairing of the two Pacific Northwest island markets (row 5) is imperfect because foreign ownership in one of the markets, San Juan Islands, is essentially zero. Any exchange rate effect would be dependent on American buyers deciding on which side of the border to buy. Secondly, in our motivation we assume that by constructing relative price indices we could control the effects of other driving forces of property prices. Unfortunately, this seems to apply less in the case of cross-border comparisons, and the controls are not doing a good enough job of picking up those other effects. In unreported tests, using unemployment rates and other income variables, results are changed little.

Nevertheless, we do find positive results for the ski resort pairs that are closest together in distance, Whistler versus Mt. Baker and Big White versus Mt. Baker. Furthermore, results are particularly strong for the two oceanfront communities that border directly on each other, Tsawwassen, B.C. and Pt. Roberts, WA. With both markets being in the majority owned by Canadians and both being in essentially the same location, they can be considered very close substitutes. As expected exchange rate movements appear to

⁷We cannot use MDC for both sides, because changes in interest rates are highly correlated and lead to near-collinearity between the two affordability measures.

have a strong effect. Our results support those of Engel and Rogers (1996), namely that distance may play a role in the price adjustment process.

[Table 6 about here.]

We now turn to the VECM framework for the cross-border pairs. Table 6 shows the results using the exact same data as in the previous table. We find strong evidence of cointegration in all 6 international pairs and in none of the controls. Only one of the markets shows a counter-intuitive positive sign in the cointegration vector. Tests for weak exogeneity show that for all international pairs, RPIs do react very strongly to deviations from the long-run equilibrium, while that is not the case for the control markets. The fact that we find weaker evidence when using year-over-year changes but very strong evidence in the VECM framework would suggest that the process of price adjustment is much more drawn out across borders and large distances as well as time relative to the within-country comparisons.

Finally, in the last column of Table 6 we report p-values from a test of the hypothesis $H_0 : \beta = -1$. As shown in Equation 8, this can be interpreted as a direct test of the law of one price. P-values above, say, 10 percent give reasonable confidence that the LOP holds in the long term. The table shows this to be the case for 3 international market pairs. Support for the LOP in the context of real estate underlines the view that a kind of one-way arbitrage holds even for immovable commodities where the owner is the one moving to the good to take advantage of price differentials in a foreign country.

7 Conclusion

In this paper we provide evidence for a connection between exchange rates and real estate prices, at least when it comes to international properties such as exclusive mountain ski resorts and oceanfront estates that are attractive to foreigners. First, using ratios of repeat sales indices for a number of international property markets on both sides of the Canadian-U.S. border, we show in a simple, univariate framework that *changes* in the exchange rate affect *changes* in prices of internationally traded properties relative to domestically-held, local property markets. An appreciation of the local currency relative to foreign currency leads to a decrease in the price of the international market relative to the local market, and vice versa.

Second, results from a VECM estimation support the idea that prices for international areas relative to either local benchmarks or comparable areas abroad are often cointegrated

with the exchange rate that translates local currency into the currency of the foreign buyer. That is, in the long run the *levels* of prices of these international markets react to the *level* of the exchange rate, whereas none of the local market pairs do.

Lastly, by comparing the relative price of two comparable international markets on different sides of the Canadian-U.S. border, we conduct a direct test of the relative law of one price (LOP) and find for some areas that the coefficient for the exchange rate is not significantly different from one, thus providing evidence in support of the relative LOP for international property markets.

We believe that there are numerous pockets of international real estate. Canadian snow birds compete with U.S. home buyers in Palm Springs, Palm Desert, Miami Beach and Phoenix. Middle-Eastern buyers are a demand component in upscale neighborhoods in London and Nice. Wealthy Russians have at various times, depending on the ruble, been buyers along the French Riviera. Finally, Monaco is a powerful magnet for wealthy individuals from all over the world. Only lack of appropriate data prevents us from exploring some of these. We trust this paper will encourage others to consider investigating some of these international markets for real estate.

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Table 1: SALES AND OWNERSHIP STATISTICS

Panel A : Canada									
	Sales Data				Ownership distribution				
	# Rep. Sales	Begin	End	Avg. Price (2005)	# Prop.				out of province
						Total	Foreign % U.S.	PO Boxes	
International markets									
Big White, B.C. (Condo)	1,205	1974Q1	2011Q1	\$298,613	1,559	21%	45%	7%	11%
Gulf Is., B.C. (W/F) (SFD)	640	1971Q2	2011Q1	\$609,164	1,396	7%	84%	8%	5%
Whistler, B.C. (Condo)	3,563	1976Q1	2011Q1	\$501,744	2,711	21%	62%	6%	4%
Whistler, B.C. (SFD)	1,401	1972Q2	2011Q1	\$1,368,312	2,560	10%	54%	8%	2%
Local markets									
Abbotsford, B.C. (SFD)	13,826	1975Q1	2005Q2	\$274,048	n/a	< 1%			
Gulf Is., B.C. (non-W/F) (SFD)	2,251	1971Q1	2011Q1	\$342,976	2,974	2%		3%	3%
Tsawwassen, B.C. (SFD)	6,584	1974Q1	2011Q1	\$500,380	3,948	0%		0%	0%
Vancouver, B.C. (Condo)	48,039	1975Q1	2005Q2	\$317,499	23,600	1%		1%	2%
Vancouver, B.C. (SFD)	28,382	1975Q1	2005Q2	\$702,063	n/a	< 1%			
West Vancouver, B.C. (SFD)	10,058	1974Q1	2007Q3	\$1,238,350	11,192	1%		1%	0%
Panel B : United States									
	Sales Data				Ownership distribution				
	# Rep. Sales	Begin	End	Avg. Price (2005)	# Prop.				out of state
						Total	Foreign % Can.	PO Boxes	
International markets									
Aspen, CO (Condo)	2,176	1971Q1	2011Q2	\$1,211,488	2,617	2%	31%	17%	53%
Aspen, CO (SFD)	725	1979Q3	2011Q2	\$3,330,140	1,174	1%	23%	15%	44%
Mt. Baker, WA (SFD)	564	1984Q1	2011Q2	\$162,065	1,255	15%	89%	22%	4%
Pt. Roberts, WA (SFD)	495	1984Q2	2011Q2	\$254,039	1,462	56%	97%	11%	4%
Local markets									
Bellingham, WA (SFD)	10,282	1983Q4	2009Q4	\$289,876	16,929	1%		5%	2%
San Juan Is., WA (SFD)	1,389	1986Q1	2011Q2	\$465,903	8,190	0%		1%	14%

This table shows all property markets for which we have individual sales data. Areas are split by country and by the nature of the area. International markets are those considered to be appealing to foreign buyers such as ski resorts and oceanfront communities, while local markets hold little appeal to foreigners. '# Rep. Sales' shows the number of repeat sales available to construct price indices for the empirical analysis, '# Prop.' the number of properties in the ownership data. The date range of sales and the average property price in the year 2005 are also depicted. The last four columns detail the ownership distribution by individual market. Under 'Foreign', we list the total share of visible foreign ownership as well as the percentage of U.S. (Canadian) owned properties as a share of total foreign ownership, based on address data. In addition, the share of PO Box addresses and out of province (state) addresses is shown.

Table 2: GMM RESULTS FOR CANADIAN MARKETS

Area of Interest	Inter- cept	FX	Local MDC	Foreign Top Inc.	# Obs	Sample Begin End	# NW Lags	adj. R^2
International Markets								
Big White, B.C. (Condo)	0.02 [0.86]	0.58 **[2.08]	-0.38 *[-1.72]	-0.08 [-0.23]	128	1977Q3 2009Q2	4	13.03%
Whistler, B.C. (Condo)	0.05 [1.37]	0.60 **[2.25]	-0.47 ***[-2.94]	-0.15 [-0.30]	128	1977Q3 2009Q2	3	19.31%
Whistler, B.C. (SFD)	0.05 **[2.23]	0.57 ***[3.17]	-0.25 **[-2.45]	-0.44 [-1.46]	128	1977Q3 2009Q2	1	17.06%
Gulf Is., B.C. (W/F) (SFD)	0.08 ***[3.15]	0.47 ***[2.77]	-0.49 ***[-3.10]	-0.74 ***[-2.81]	128	1977Q3 2009Q2	3	32.38%
Local Markets								
Gulf Is., B.C. (nW/F) (SFD)	0.06 ***[3.34]	0.07 [0.51]	-0.35 ***[-2.66]	-0.82 ***[-3.43]	128	1977Q3 2009Q2	4	23.29%
Tsawassen, B.C. (SFD)	0.03 *[1.95]	0.12 [1.39]	-0.08 [-1.11]	-0.47 ***[-2.66]	128	1977Q3 2009Q2	3	12.09%
West Vancouver, B.C. (SFD)	-0.01 [-1.44]	0.02 [0.24]	-0.01 [-0.23]	0.22 [1.60]	128	1977Q3 2009Q2	8	1.68%

This table depicts the results for quarterly univariate time series regressions for select Canadian markets relative to the close-by benchmark market of Vancouver, B.C.:

$$\Delta_4 \text{RPI}_t = \alpha + \beta \Delta_4 \log(\pi_{t-l}) + \gamma' X_{t-l} + \epsilon_t$$

Markets are classified as international markets with high foreign ownership and local markets with negligible foreign ownership. The dependent variable is the year-over-year change (in logs) of property prices in the area of interest relative to the price of the benchmark. All independent variables are lagged by $l = 2$ quarters. π_t is the nominal exchange rate (price of 1 USD in CAD); X_t is a vector of two income-related control variables: 1) year-over-year changes in the mortgage debt capacity (MDC) of the average family in Vancouver, B.C. and 2) year-over-year changes in the income of the 99th-percentile in the United States. 'FX' represents the estimate of coefficient β above. '# Obs' is the number of quarterly observations available and '# NW Lags' is the optimal lag length for the Newey and West (1994) adjustment to correct for serial correlation. Newey-West t-statistics are shown in brackets below each coefficient.

Table 3: GMM RESULTS FOR U.S. MARKETS

Area of Interest	Inter- cept	FX	Local MDC	Foreign Top Inc.	# Obs	Sample Begin End	# NW Lags	adj. R^2
International Markets								
Aspen, CO (SFD)	0.05 **[2.16]	0.02 [0.13]	-0.63 ***[-4.13]	0.41 *[1.73]	115	1980Q4 2009Q2	4	30.07%
Mt. Baker, WA (SFD)	0.06 *[1.73]	-0.77 ***[-2.68]	-0.13 [-0.76]	-1.52 **[-2.29]	93	1985Q2 2008Q2	6	26.59%
Point Roberts, WA (SFD)	0.01 [0.32]	-0.92 ***[-3.28]	-0.30 **[-2.07]	0.40 [0.75]	93	1985Q2 2008Q2	5	40.98%
Local Markets								
Bellingham, WA (SFD)	0.00 [0.21]	-0.16 [-1.40]	0.00 [0.11]	-0.06 [-0.30]	115	1979Q4 2008Q2	10	2.81%
San Juan Is., WA (SFD)	-0.02 [-1.22]	0.02 [0.22]	-0.00 [-0.02]	0.37 [1.62]	89	1987Q2 2009Q2	10	12.19%

This table depicts the results for quarterly univariate time series regressions for select U.S. markets relative to house prices in all of Washington State. The exception is Aspen, CO where the comparison is made relative to the U.S. national house price index:

$$\Delta_4 \text{RPI}_t = \alpha + \beta \Delta_4 \log(\pi_{t-l}) + \gamma' X_{t-l} + \epsilon_t$$

Markets are classified as international markets with high foreign ownership and local markets with negligible foreign ownership. The dependent variable is the year-over-year change (in logs) of property prices in the area of interest relative to the price of the benchmark. All independent variables are lagged by $l = 2$ quarters. π_t is the nominal exchange rate (price of 1 USD in CAD); X_t is a vector of two income-related control variables: 1) year-over-year changes in the mortgage debt capacity (MDC) of the average family in the respective state and 2) year-over-year changes in the income of the 99th-percentile of the population in Canada. However, for the two markets with no significant Canadian ownership (Aspen and San Juan Islands), the second control variable is based on the income of the 99th-percentile in the United States and the exchange rate is the U.S. dollar index. 'FX' represents the estimate of coefficient β above. '# Obs' is the number of quarterly observations available and '# NW Lags' is the optimal lag length for the Newey and West (1994) adjustment to correct for serial correlation. Newey-West t-statistics are shown in brackets below each coefficient.

Table 4: VECM FOR WITHIN-COUNTRY PROPERTY MARKET PAIRS

Area of Interest	Type	# Obs	Johansen Cointegration Tests					Add. Tests				
			Case 2 (constant in ECM)		Case 3 (drift in process)		If $r = 1$					
			$r \leq 0$	$r \leq 1$	$r = 1?$	$r \leq 0$	$r \leq 1$	$r = 1?$	α	β	$H_0 : \alpha = 0$	$H_0 : \beta = \pm 1$
Panel A: High foreign ownership share												
Canada												
Big White, B.C. Whistler, B.C. Whistler, B.C. Gulf Is., B.C. (W/F)	Condo	126	0.55	-3.86	YES	3.28	1.40	NO	-0.09	-1.54	0.0077	0.5013
	Condo	126	3.68	-0.70	YES	3.99	4.16	NO	-0.09	-1.30	0.0836	0.8382
	SFD	126	-0.20	-0.74	NO	2.60	2.72	NO			0.9008	
	SFD	126	-0.58	-5.56	NO	3.42	-0.76	YES	-0.19	0.58	0.0012	0.0030
United States												
Aspen, CO	SFD	115	-1.04	-8.73	NO	-5.18	-3.63	NO			0.0042	
Mt. Baker, WA	SFD	95	20.78	-8.44	YES	23.62	-3.38	YES	-0.36	0.41	0.0000	0.0000
Point Roberts, WA	SFD	95	-0.63	-7.33	NO	3.68	-2.22	YES	-0.22	1.35	0.0001	0.1826
Panel B: Low foreign ownership share												
Canada												
Gulf Is., B.C. (nW/F) Tsawassen, B.C. West Vancouver, B.C.	SFD	126	-5.05	-4.66	NO	-1.00	0.48	NO			0.0179	
	SFD	126	-7.22	-5.39	NO	-3.64	-0.10	NO			0.0439	
	SFD	126	-2.30	-4.36	NO	1.66	0.89	NO			0.1097	
United States												
Bellingham, WA	SFD	115	-7.54	-8.22	NO	-4.86	-3.02	NO			0.0034	
San Juan Is., WA	SFD	91	-1.31	-5.03	NO	-1.88	-1.73	NO			0.0053	

This table reports the results from fitting a 2-dimensional vector error correction model (VECM) to a relative price index (RPI) and the log of the CAD/USD exchange rate π at the quarterly frequency. The relative price index is the log ratio of two price indices comparing areas with a high share of foreign ownership to a local benchmark area with little foreign ownership. Defining $y_t = (RPI_t, \log(\pi_t))'$, the VECMX(p, q) can be written as follows:

$$\Delta y_t = \sum_{i=1}^{p-1} \gamma_i \Delta y_{t-i} + \sum_{i=0}^q X_{t-l-i} + A(B'y_{t-1} + \eta) + \epsilon_t$$

Reported results are based on $p = 2, q = 1$. X_t is a vector of two income-related control variables: year-over-year changes in 1) average local mortgage debt capacity (MDC) and 2) in the 99th-percentile of foreign income (lagged by $l = 2$ quarters). MDC is based on family incomes in Vancouver, B.C. for Canadian markets, and in the respective state for U.S. markets. 'Foreign' is the top U.S. income for Canadian markets, and vice versa. However, for the two markets with no significant Canadian ownership (Aspen and San Juan Islands), the second control variable is based on the income of the 99th-percentile in the United States and the exchange rate is the U.S. dollar index. Cointegration test results are based on the Johansen (1991) trace statistics. In order to conserve space, numbers displayed are the difference between the test statistic and the critical value at 5 percent. A positive number means rejection of the null, a negative number the opposite. In the context of a 2-dimensional test, rejecting $H_0 : r \leq 0$ and not rejecting $H_0 : r \leq 1$ indicates cointegration of y with unique cointegrating vector $(1, \beta)$. In this case, the column ' $r = 1$ ' contains a 'YES', while 'NO' means that cointegration is rejected by the data. 'Case 2' (based on common textbook treatments) is the specification depicted above, while 'Case 3' replaces the constant term η in the ECM term with a drift in the process. The second to last column reports p-values from a test of weak exogeneity for the relative price index, i.e. that changes in RPI are not affected by the disequilibrium, or in short $H_0 : \alpha = 0$. If either case supports cointegration, α (annualized) and β are shown. In addition, the last column contains the p-value from the F-test of the restriction $\beta = -1$ ($\beta = 1$) for Canadian (U.S.) comparisons.

Table 5: GMM RESULTS FOR CROSS-BORDER PAIRS

Can. Area	Comparison of U.S. area	Type	Inter- cept	FX	Income U.S.	MDC Canada	# Obs	Sample Begin End	# NW Lags	adj. R^2
International ski resorts										
Whistler, B.C.	Aspen, CO	Condo	0.06 [1.07]	0.13 [0.51]	-0.55 [-0.62]	-0.45 ***[-2.97]	132	1977Q3 2010Q2	3	5.40%
Whistler, B.C.	Aspen, CO	SFD	-0.07 [-1.20]	0.02 [0.06]	0.35 [0.38]	0.61 ***[3.76]	119	1980Q4 2010Q2	7	11.89%
Whistler, B.C.	Mt. Baker, WA	SFD	0.02 [0.29]	0.72 *[1.81]	0.41 [0.42]	-0.08 [-0.26]	101	1985Q2 2010Q2	10	8.28%
Big White, B.C.	Mt. Baker, WA	C/S	-0.05 [-1.55]	0.46 *[1.94]	1.18 *[1.98]	-0.05 [-0.19]	101	1985Q2 2010Q2	4	9.32%
Ocean-front real estate (close proximity to each other)										
Gulf Is., B.C. (W/F)	San Juan Is., WA	SFD	0.10 ***[4.79]	0.14 [0.95]	-1.78 ***[-4.52]	0.06 [0.45]	93	1987Q2 2010Q2	1	26.00%
Tsawwassen, B.C.	Pt. Roberts, WA	SFD	0.05 [1.56]	0.69 ***[3.01]	-1.10 *[-1.89]	0.07 [0.47]	101	1985Q2 2010Q2	4	29.66%
Control markets (comparable cities close to the border)										
Abbotsford, B.C.	Bellingham, WA	SFD	-0.01 [-0.84]	-0.09 [-0.86]	-0.04 [-0.11]	0.29 ***[4.87]	100	1985Q1 2009Q4	1	13.33%
Vancouver, B.C.	Seattle, WA	SFD	0.03 [0.97]	-0.46 [-1.53]	-0.55 [-0.76]	0.02 [0.22]	132	1977Q3 2010Q2	39	4.51%

This table depicts the results for quarterly univariate time regressions for relative price indices between two similar property markets (e.g. ski resorts) on different sides of the U.S.-Canadian border:

$$\Delta_4 \text{RPI}_t = \alpha + \beta \Delta_4 \log(\pi_{t-l}) + \gamma' X_{t-l} + \epsilon_t$$

Markets are classified by type. The dependent variable is the year-over-year change (in logs) of property prices in the area of interest relative to the price of the benchmark. All independent variables are lagged by $l = 2$ quarters. π_t is the nominal exchange rate (price of 1 USD in CAD). X_t is a vector of control variables, the year-over-year changes in 1) average U.S. per capita income of the respective state and in 2) average mortgage debt capacity (MDC) based on family income in the province of B.C. 'FX' represents the estimate of coefficient β above. '# Obs' is the number of quarterly observations available and '# NW Lags' is the optimal lag length for the Newey and West (1994) adjustment to correct for serial correlation. Newey-West t-statistics are shown in brackets below each coefficient.

Table 6: VECM FOR CROSS-BORDER PROPERTY MARKET PAIRS

Can. Area	Comparison of U.S. area	Type	# Obs	Johansen Cointegration Tests				Add. Tests		
				Case 2 (constant in ECM) $r \leq 0$	Case 3 (drift in process) $r \leq 1$	If $r = 1$ α	β	$H_0 : \alpha = 0$	$H_0 : \beta = -1$	
International ski resorts										
Whistler, B.C.	Aspen, CO	Condo	132	80.32	-3.42	YES	-1.10	YES	-0.36	0.0000
Whistler, B.C.	Aspen, CO	SFD	121	78.41	-2.24	YES	0.02	NO	-0.34	0.0007
Whistler, B.C.	Mt. Baker, WA	SFD	103	10.10	-5.56	YES	-0.81	YES	-1.08	0.8113
Big White, B.C.	Mt. Baker, WA	mixed	101	4.97	-6.23	YES	-1.49	YES	-0.85	0.6260
Ocean-front real estate (close proximity to each other)										
Gulf Is., B.C. (W/F)	San Juan Is., WA	SFD	95	33.23	-6.54	YES	-2.35	YES	0.43	0.0000
Tsawwassen, B.C.	Pt. Roberts, WA	SFD	103	22.27	-4.35	YES	-0.52	YES	-0.82	0.1054
Control markets (comparable cities close to the border)										
Abbotsford, B.C.	Bellingham, WA	SFD	102	-3.26	-1.49	NO	2.49	NO		0.2786
Vancouver, B.C.	Seattle, WA	SFD	134	-5.40	-3.94	NO	1.30	NO		0.1463

This table reports the results from fitting a 2-dimensional vector error correction model (VECM) to a relative price index (RPI) and the CAD/USD exchange rate π at the quarterly frequency. The relative price index is the log ratio of two price indices comparing an area of interest in Canada with a similar area in the United States. Defining $y_t = (RPI_t, \log(\pi_t))'$ and including a vector of (lagged) control variables X_{t-l} , the VECMX(p, q) can be written as follows:

$$\Delta y_t = \sum_{i=1}^{p-1} \gamma_i \Delta y_{t-i} + \sum_{i=0}^q X_{t-l-i} + A(B'y_{t-1} + \eta) + \epsilon_t$$

Reported results are based on $p = 2, q = 1$ and X_t is a vector of two control variables, year-over-year changes in average local per capita income of the corresponding U.S. state and in the mortgage debt capacity of the average family in the province of B.C. (lagged by $l = 2$ quarters). Cointegration test results are based on the Johansen (1991) trace statistics. In order to conserve space, numbers displayed are the difference between the test statistic and the critical value at 5 percent. A positive number means rejection of the null, a negative number the opposite. In the context of a 2-dimensional test, rejecting $H_0 : r \leq 0$ and not rejecting $H_0 : r \leq 1$ indicates cointegration of y with unique cointegrating vector $(1, \beta)$. In this case, the column ' $r = 1$ ' contains a 'YES', while 'NO' means that cointegration is rejected by the data. 'Case 2' (based on common textbook treatments) is the specification depicted above, while 'Case 3' replaces the constant term η in the ECM term with a drift in the process. The second to last column reports p-values from a test of weak exogeneity for the relative price index, i.e. that changes in RPI are not affected by the disequilibrium, or in short $H_0 : \alpha = 0$. If either case supports cointegration, α (annualized) and β are shown. In addition, the last column contains the p-value from the F-test of the restriction $\beta = -1$.

Table A.1: ADDITIONAL GMM RESULTS FOR CANADIAN MARKETS

Area of Interest	Inter- cept	FX	Local U-Rate	Foreign U-Rate	# Obs	Sample Begin End	# NW Lags	adj. R^2
International Markets								
Big White, B.C. (Condo)	0.35 *[1.72]	0.54 **[1.98]	0.03 [0.35]	-0.23 **[-2.58]	138	1976Q3 2010Q4	4	16.46%
Whistler, B.C. (Condo)	0.37 **[2.16]	0.54 **[2.10]	0.08 [1.38]	-0.30 ***[-4.18]	136	1977Q1 2010Q4	4	24.54%
Whistler, B.C. (SFD)	0.35 ***[3.04]	0.56 **[2.57]	0.09 **[2.08]	-0.30 ***[-4.80]	138	1976Q3 2010Q4	9	33.08%
Gulf Is., B.C. (W/F) (SFD)	0.26 *[1.89]	0.59 ***[3.64]	-0.06 [-1.11]	-0.07 [-0.91]	138	1976Q3 2010Q4	8	13.65%
Local Markets								
Gulf Is., B.C. (nW/F) (SFD)	0.17 *[1.83]	0.16 [1.13]	-0.02 [-0.54]	-0.07 [-1.09]	138	1976Q3 2010Q4	2	3.40%
Tsawassen, B.C. (SFD)	0.15 **[2.43]	0.15 [1.53]	-0.01 [-0.48]	-0.07 [-1.61]	138	1976Q3 2010Q4	17	6.69%
West Vancouver, B.C. (SFD)	0.06 [0.98]	-0.07 [-0.76]	0.02 [0.72]	-0.06 **[-2.25]	140	1976Q3 2011Q2	11	7.20%

This table depicts the results for quarterly univariate time series regressions for select Canadian markets relative to the close-by benchmark market of Vancouver, B.C.:

$$\Delta_4 \text{RPI}_t = \alpha + \beta \Delta_4 \log(\pi_{t-l}) + \gamma' X_{t-l} + \epsilon_t$$

Markets are classified as international markets with high foreign ownership and local markets with negligible foreign ownership. The dependent variable is the year-over-year change (in logs) of property prices in the area of interest relative to the price of the benchmark. All independent variables are lagged by $l = 2$ quarters. π_t is the nominal exchange rate (price of 1 USD in CAD); X_t is a vector of two control variables, the local and foreign unemployment rates of the province of B.C. and the United States, respectively. 'FX' represents the estimate of coefficient β above. '# Obs' is the number of quarterly observations available and '# NW Lags' is the optimal lag length for the Newey and West (1994) adjustment to correct for serial correlation. Newey-West t-statistics are shown in brackets below each coefficient.

Table A.2: ADDITIONAL GMM RESULTS FOR U.S. MARKETS

Area of Interest	Inter- cept	FX	Local U-Rate	Foreign U-Rate	# Obs	Sample Begin End	# NW Lags	adj. R^2
International Markets								
Aspen, CO (SFD)	0.17 [1.14]	0.11 [0.65]	0.11 [1.37]	-0.18 [-1.47]	121	1980Q4 2010Q4	2	1.75%
Mt. Baker, WA (SFD)	-0.03 [-0.23]	-0.38 ***[-2.91]	0.10 **[2.06]	-0.08 [-1.42]	103	1985Q2 2010Q4	2	13.82%
Point Roberts, WA (SFD)	0.10 [0.60]	-0.72 ***[-3.39]	-0.03 [-0.21]	-0.03 [-0.50]	103	1985Q2 2010Q4	6	27.91%
Local Markets								
Bellingham, WA (SFD)	-0.03 [-0.62]	-0.08 [-0.89]	0.04 [1.45]	-0.02 [-0.96]	126	1979Q4 2011Q1	11	4.47%
San Juan Is., WA (SFD)	0.08 [1.12]	-0.03 [-0.34]	0.16 **[1.99]	-0.21 ***[-2.86]	95	1987Q2 2010Q4	8	26.25%

This table depicts the results for quarterly univariate time series regressions for select U.S. markets relative to house prices in all of Washington State. The exception is Aspen, CO where the comparison is made relative to the U.S. national house price index:

$$\Delta_4 \text{RPI}_t = \alpha + \beta \Delta_4 \log(\pi_{t-l}) + \gamma' X_{t-l} + \epsilon_t$$

Markets are classified as international markets with high foreign ownership and local markets with negligible foreign ownership. The dependent variable is the year-over-year change (in logs) of property prices in the area of interest relative to the price of the benchmark. All independent variables are lagged by $l = 2$ quarters. π_t is the nominal exchange rate (price of 1 USD in CAD); X_t is a vector of two control variables, the unemployment rates of the local state and the province of B.C., respectively. However, for the two markets with no significant Canadian ownership (Aspen and San Juan Islands), the second control variable is based on the U.S. national unemployment rate and the exchange rate is the U.S. dollar index. 'FX' represents the estimate of coefficient β above. '# Obs' is the number of quarterly observations available and '# NW Lags' is the optimal lag length for the Newey and West (1994) adjustment to correct for serial correlation. Newey-West t-statistics are shown in brackets below each coefficient.

Table A.3: ROBUSTNESS VECM FOR WITHIN-COUNTRY PROPERTY MARKET PAIRS

Area of Interest	Type	# Obs	Johansen Cointegration Tests										Add. Tests	
			Case 2 (constant in ECM)		Case 3 (drift in process)		If $r = 1$		$H_0 : \alpha = 0$	$H_0 : \beta = \pm 1$				
			$r \leq 0$	$r \leq 1$	$r = 1?$	$r \leq 0$	$r \leq 1$	$r = 1?$	α	β				
Panel A: High foreign ownership share														
Canada														
Big White, B.C.	Condo	136	3.70	-2.90	YES	7.27	2.28	NO	-0.11	-1.43	0.0039	0.5780		
Whistler, B.C.	Condo	140	17.49	-0.25	YES	21.18	4.30	NO	-0.18	-1.15	0.0000	0.7937		
Whistler, B.C.	SFD	138	6.71	2.06	NO	7.95	5.59	NO			0.0358			
Gulf Is., B.C. (W/F)	SFD	138	9.80	0.16	NO	11.32	3.86	NO			0.0008			
United States														
Aspen, CO	SFD	125	-4.23	-8.53	NO	0.28	-3.27	YES	-0.08	1.92	0.0021	0.0271		
Mt. Baker, WA	SFD	107	39.62	-4.74	YES	41.92	-0.75	YES	-0.38	0.37	0.0000	0.0000		
Point Roberts, WA	SFD	107	-4.99	-4.38	NO	-6.36	-2.29	NO			0.0218			
Panel B: Low foreign ownership share														
Canada														
Gulf Is., B.C. (nW/F)	SFD	138	-2.41	-2.81	NO	0.05	2.13	NO			0.6187			
Tsawassen, B.C.	SFD	138	-3.03	-3.66	NO	-1.48	0.02	NO			0.2486			
West Vancouver, B.C.	SFD	139	-9.41	-6.98	NO	-5.68	-2.22	NO			0.6076			
United States														
Bellingham, WA	SFD	130	-11.68	-6.74	NO	-7.07	-1.48	NO			0.1680			
San Juan Is., WA	SFD	99	-6.98	-5.45	NO	-5.81	-3.44	NO			0.0076			

This table reports the results from fitting a 2-dimensional vector error correction model (VECM) to a relative price index (RPI) and the log of the CAD/USD exchange rate π at the quarterly frequency. The relative price index is the log ratio of two price indices comparing areas with a high share of foreign ownership to a local benchmark area with little foreign ownership. Defining $y_t = (\text{RPI}_t, \log(\pi_t))'$, the $\text{VECMX}(p, q)$ can be written as follows:

$$\Delta y_t = \sum_{i=1}^{p-1} \gamma_i \Delta y_{t-i} + \sum_{i=0}^q X_{t-l-i} + A(B'y_{t-1} + \eta) + \epsilon_t$$

Reported results are based on $p = 2, q = 1$. X_t is a vector of two control variables: for Canadian markets the unemployment rates of the province of B.C. and the U.S. national average, for U.S. markets the unemployment rates of the local state and the province of B.C. (all lagged by $l = 2$ quarters). However, for the two markets with no significant Canadian ownership (Aspen and San Juan Islands), the second control variable is based on the U.S. national unemployment rate and the exchange rate is the U.S. dollar index. Cointegration test results are based on the Johansen (1991) trace statistics. In order to conserve space, numbers displayed are the difference between the test statistic and the critical value at 5 percent. A positive number means rejection of the null, a negative number the opposite. In the context of a 2-dimensional test, rejecting $H_0 : r \leq 0$ and not rejecting $H_0 : r \leq 1$ indicates cointegration of y with unique cointegrating vector (1, β). In this case, the column ' $r = 1?$ ' contains a 'YES', while 'NO' means that cointegration is rejected by the data. 'Case 2' (based on common textbook treatments) is the specification depicted above, while 'Case 3' replaces the constant term η in the ECM term with a drift in the process. The second to last column reports p-values from a test of weak exogeneity for the relative price index, i.e. that changes in RPI are not affected by the disequilibrium, or in short $H_0 : \alpha = 0$. If either case supports cointegration, α (annualized) and β are shown. In addition, the last column contains the p-value from the F-test of the restriction $\beta = -1$ ($\beta = 1$) for Canadian (U.S.) comparisons.