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Are Canadian House Prices More Closely Linked to Fundamentals than U.S. House Prices?

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I. Introduction

Canadian house prices increased significantly in the period 2000 to 2006, as did U.S. prices but Canadian prices did not experience the dramatic collapse seen in the U.S. in 2007 to 2008 (See Exhibit 1). In this paper, we investigate the hypothesis that Canadian house prices fared better because they continued to be driven by underlying fundamental factors such as income and population. Mikhed and Zemcik (2009) provide evidence that U.S. prices diverged from fundamentals during the period of rapid house price increases during 2000 to 2006. We prepare house price models for the period 2000 – 2009 for both Canada and the U.S., using variables as similar as possible between the two countries, including population, personal income, wages, rents, inflation, stock market returns, mortgage interest rate and gross domestic product as explanatory variables.

![Figure A](source: CUER for Canada, FHFA for U.S.)

We use multiple indices of Canadian house prices including the recently (2009) released Teranet - National Bank of Canada House Price Index. This new index employs the repeat-sales methodology incorporated in the U.S. S&P/Case-Shiller Home Price Index. As part of our analysis, we investigate the qualities of the new index in comparison to alternative Canadian indices and to the S&P/Case-Shiller Home Price Index. We also evaluate whether this new Canadian index performs better than the other Canadian indices in producing a Canadian house price model. For example, the Royal-Lepage Survey of Canadian House Prices was used in the Bank of Canada working paper by Allen, Amano, Byrne and Gregory (2006) despite the fact that this house price data is produced based on telephone surveys of real estate agents.
Our contribution is twofold. Our comparison of the degree to which U.S. and Canada house prices are driven by fundamentals will offer insights about the international differences across house price markets and may offer insights that will help to identify whether house prices are likely to decline after a period of substantial price increases. As a secondary contribution, we will evaluate a new source of Canadian house price data and its usefulness for other researchers.

II. Literature Review

There is a large and well-developed literature on housing demand. Olsen (1987) and Whitehead (1998) provide broad reviews of this empirical literature, but as Zabel (2004) points out, housing demand can mean a variety of different things, from demand for housing services, to demand for individual attributes, to tenure choice, or the spatial allocation of households. Much initial research focused on estimating demand elasticities (see Mayo 1981 for a survey of this literature, or more recently Harmon 1988 and Ermisch 1996). For the most part this literature assumes that parameters are stable over time, and ignores how households adjust to change their demand for housing and non-housing goods with changes in house prices, incomes, and constraints on accessing home-ownership. Bajari, et. al. (2010) present a dynamic structural model where households solve a life cycle consumption problem where housing is an investment good that delivers consumption benefits. In their model, changes in housing market conditions have only a limited effect on housing consumption, operating instead through changes in housing equity or non-housing consumption.

The relationship of our work to housing demand analysis draws mainly on interpreting the contributors to demand, incomes and owner cost variables, as the explanatory fundamentals, along with structure input prices, that move house prices. Wheaton and Nechayev (2008) report that wide consensus exists to support the relationship between house prices and employment, population and income. Drawing on the owner-cost work best associated with Hendershot (1980) and Poterba (1984) interest rates and other owner cost variables are also used in many of these studies, including, for example, Coleman, LaCour-Little and Vandell (2008). Models of this type are linked to the urban price model developed by Capozza and Helsley (1989) in which the price of urban land is decomposed into four elements: agricultural land rent, the cost of conversion, the value of accessibility and the value of expected future rent increases. Abraham and Hendershott (1996) used house price indices in their empirical analysis, reporting that in 1992 the Northeast region of the United States was priced at 30 percent above fundamentals.

Our work does not focus on identifying demand parameters or testing any type of structural model, but on the stability of the relationship between these housing market fundamentals and house prices over time. We assume that supply parameters other than structure inputs remain constant over time. So that even though there is significant variation in supply elasticities across markets, as demonstrated most recently by Saiz (forthcoming) among many others, this does not affect the temporal variation in the relationship between out fundamental variables and housing prices. In comparing differences in the response of housing market outcomes such as prices, construction, and vacancies, to fundamentals across metropolitan areas, Hwang and Quigley (2006) demonstrate that regulation explains an important part of the difference in size and pattern of responses to shocks.
The methodology we present here follows a large number of recent studies that investigate the housing boom and bust cycle of past decade by trying to determine if prices deviate from fundamental values, whether a housing price bubble existed in the house price boom and bust in the US between 2000 and 2009. The term “bubble” was defined in a housing context (perhaps first) by Stiglitz (1990) with later elaboration provided by Case and Shiller (2003). Essentially, bubbles are thought to exist when rapid house price increases result from expectations of future price increases rather than from changes in fundamentals. Bubble hunters include Goodman and Thibodeau (2008), Wheaton and Nechayev (2008) and Mikhed and Zemcik (2009) in the U.S.; Fraser, Hoesli and McAlevey (2008) in New Zealand; Fernandez-Kranz and Hon (2006) in Spain; and Hui and Yue (2006) in China. These studies rely on different but related sets of variables used as fundamental drivers and employ a variety of empirical techniques. Campbell, et.al. (2009) pursue the same question, but do so by studying the fundamental determinants of housing’s rent to price ratio using a dynamic Gordon growth model treatment. They find that variation in the rent to price ratio reflects expected future risk premia for housing more than movements in the real interest rate. By using the rent price ratio, they are able to abstract from fundamentals that operate on house prices through current rents.

Our current study most resembles that used by Mikhed and Zemcik (2009). In their study, they employ unit root and cointegration tests, while we implement a rolling cointegration innovation to the technique that allows us to look more closely at the longitudinal issues. The included explanatory variables, serving as the fundamental drivers, also reflect the choices of Mikhed and Zemcik (2009). As do they, we use rent, a consumer price index, construction material costs, personal income, population, mortgage rates, a stock index and wage data. We extend the work beyond that of Mikhed and Zemcik in that we examine both U.S. and Canadian house price movements. Very little work has been done in the past with Canadian data, with perhaps the sole exception that of Allen, Amano, Byrne and Gregory (2009). Their work uses survey-based house price data from Royal-LePage and Statistics Canada’s house price index for new homes to identify which fundamental house price drivers are significant for particular Canadian cities.

One potential complication in our work involves the time series properties of house prices and fundamental variables. The co-integrating framework used extensively to model the relationship between fundamentals and house prices, and especially in an error-correction framework with incomes (Abraham and Hendershott 1996, Malpezzi 1999, Capozza et al. 2002, and Meen 2002) flows from the stock-flow model of housing used in Poterba (1984), Topel and Rosen (1988), and DiPasquale and Wheaton (1994) . However, Gallin (2006) challenges this approach, arguing that using more powerful panel tests with metropolitan area data, the hypothesis that house prices and incomes are not co-integrated cannot be rejected.
III. Methodology

To examine any core links between housing prices and likely key drivers, we adopt fractional cointegration tools, which generalize the classical I(0)/I(1) cointegration approaches, along with a “rolling forward windows” framework to ascertain whether there is evidence to suggest that relationships have changed over time. Until recently, analysts have restricted attention to ascertaining whether their data are either stationary I(0) (perhaps, trend-stationary) or non-stationary I(1) (a so-called unit-root process). Stochastic shocks affect stationary, I(0), processes but they are quickly “forgotten,” ensuring that the series is “mean-reverting”, perhaps around a deterministic trend. An I(0) series exhibits short or temporary memory with an autocorrelation function that dampens exponentially quickly. In contrast, shocks to a unit-root, I(1), process are not discounted over time, in the sense that all shocks have a permanent effect on the future dynamics of the series. This “permanent memory” feature leads the series to meander widely, exhibiting a stochastic trend, with the sample autocorrelation function having large values at low lags with statistically significant values at very long lags, depending on the generation process; see, Hassler (1994). Such a “knife-edged” dichotomy for a data generating process does not permit the possibility of a middle ground, in which the series possesses long but not permanent memory with high-order autocorrelations that are too large to have arisen from a traditional stationary model. A fractionally integrated process, suggested by Granger and Joyeux (1980) and Hosking (1981), exhibits such a feature, with the degree of persistence dependent on the integration parameter \( d \). Dependent on \( d \), the family allows for nonstationarity with and without mean-reversion; see, e.g., Robinson (2003) for a survey of this process.

Specifically, suppose a series follows an autoregressive moving average (ARMA) process given by:

\[
(1)
\]

where \( B \) is the backshift operator such that \( B^d \) represents a stationary autoregressive process, \( B^{-d} \) represents and invertible moving-average process, and \( d \), the differencing parameters, indicates the integration order of \( X_t \). The series is mean reverting and covariance stationary when \( d=0 \), whereas it is a unit-root process characterized by mean and covariance nonstationarity when \( d=1 \). The series is fractionally integrated for non-integer \( d \) values. If \( d>-1/2 \), the series is invertible and if \( d<1/2 \) it is covariance stationary, but it is covariance nonstationary when \( d\geq1/2 \). For the covariance stationary case the autocorrelations decline at a hyperbolic rate, slower than the geometric decline for an I(0) process. For \( d>0 \) the series is said to have long memory, with \( d \) indicating the extent of the “memory”.

To estimate \( d \) for each series (we consider both levels and first differences), we use Robinson’s (1995) semi-parametric estimator that is based on a log-periodogram regression.\(^1\) For comparison purposes, we also undertake a traditional unit root test and stationarity test: the DF-GLS and KPSS tests. The DF-GLS unit root test, proposed by Elliott et al. (1996), is a modification of the standard augmented Dickey-Fuller (1979) test by de-trending the data prior to estimating the test regression. Given the nature of our levels data, we model with a constant and a linear time trend, but

\(^1\) We follow Robinson in setting the power to be 0.9, which determines the number of ordinates entering the log-periodogram regression.
we only allow for a constant with the first differenced data. Despite using quarterly data, we found no evidence to suggest that any seasonality needed to be accommodated. The DF-GLS null hypothesis is for a unit root with the alternative being that the series is trend stationary for the levels data and level stationary for the first differenced data. Following the recommendations of Ng and Perron (2001), we use their modified Akaike Information Criterion to select the number of augmentation terms. MacKinnon’s (1996) critical value calculations are compared with our chosen 10% significance level to decide the test outcome. The KPSS (Kwiatkowski et al., 1992) test examines the null hypothesis that the series is I(0), with a unit root alternative hypothesis. We use a Bartlett kernel to estimate the frequency zero spectrum with the Newey-West (1994) automatic variable bandwidth selection. The null is specified as trend-stationary for the levels series and level-stationary for the first differenced series. Critical values for the 10% significance level are from Sephton (1995).

Having examined integration orders, we turn to ascertaining whether housing prices are linked with likely fundamental drivers. Let \( \mathbf{x} \) be a K-dimensional vector consisting of a relevant housing price index along with associated drivers, with \( i^{th} \) element \( x_i \), \( i=1, \ldots, K \). Following Robinson and Marinucci (2003), there is fractional cointegration if there exists a \((K\times1)\) vector \( \alpha \neq 0 \) such that

\[
\begin{equation}
\begin{aligned}
\dot{\mathbf{x}} = \mathbf{A} \mathbf{x} + \mathbf{B} \mathbf{u}
\end{aligned}
\end{equation}
\]

where \( \mathbf{A} \) and \( \mathbf{B} \) are matrices of appropriate dimensions and \( \mathbf{u} \) is a vector of deterministic terms. As noted therein, this notion of fractional cointegration requires \( \alpha' \mathbf{A} > 0 \) for some \( \alpha \neq 0 \). The reduction in memory indicates the strength of the long-run relationship linking the variables. Of interest is whether \( \mathbf{A} \) is invertible, as then the errors are stationary, although with more persistence than in the I(0) case, implying a slower rate of convergence to the long-run equilibrium. However, this framework also allows for \( \mathbf{A} \) to be noninvertible, sometimes called “weak fractional cointegration” (see Hualde and Robinson, 2007). In this case, the cointegrating vector does not fully account for the nonstationarity in the variables such that there is still some random wandering around the cointegrating regression but there is still co-movement in the sense of Engle and Granger (1987), as the impulse responses of the residuals decay to zero over time. Clearly, then, the cointegrating relation cannot be given the usual “equilibrium” interpretation.

Traditional I(1)/I(0) cointegration of Engle and Granger (1987) is a special example of this framework, as is also the I(2)/I(1)/I(0) polynomial cointegration of, for instance, Haldrup (1994) that allows for stochastic cointegration at different levels. Specifically, the I(2) variables cointegrate into an I(1) relation that may or may not further cointegrate with other I(1) variables. Clearly, it is also feasible that cointegration at different levels may occur for non-integer \( d \) values. In addition, multicointegration is possible (e.g., Granger and Lee, 1989; Engsted et al., 1997) whereby, for instance, linear combinations of I(1) stochastic processes form an I(0) process (the first level of cointegration) and the cumulated cointegrated residuals, I(1) by definition, cointegrate with the I(1) original variables (a deeper, second, level of cointegration). This form of multicointegration is also extendable to I(2) and fractionally integrated processes.

To examine for some of these possibilities, we consider the regression

\[
\begin{equation}
\begin{aligned}
\ddot{y} = \alpha' \mathbf{x} + \mathbf{d} + \mathbf{e}
\end{aligned}
\end{equation}
\]

where \( \ddot{y} \) is a housing price index, \( \mathbf{x} \) is a vector of fundamentals, \( \mathbf{d} \) is a vector of deterministic terms (we include a constant and linear trend) and \( \mathbf{e} \) is an error term. For the standard I(1)/I(0) cointegration paradigm, Stock (1987) shows that the least squares estimator of the parameters of the
cointegrating vector converge in probability to their true values, albeit at a different rate of convergence than under classical assumptions. Haldrup (1994), who considers the asymptotic properties of the least squares estimator when (2) contains both I(2) and I(1) processes, shows that the order of least squares differs for the I(1) and I(2) variables, but consistency does result when \( r = 2 \) or 0; i.e., when the I(2) variables cointegrate without further cointegration in the system, or when there are more levels of cointegration resulting in an I(0) error. Properties of least squares under fractional integration and cointegration are detailed by, among many others, Cheung and Lai (1993), Robinson and Marinucci (2003) and Chen and Hurvich (2003). Accordingly, we estimate such models by OLS, then explore the degree of integration of the residuals using Robinson’s (1995) semiparametric estimator; a similar approach is undertaken by Gil-Alana (2003). As stability of the cointegrating regression is of interest, we provide plots of the coefficient estimates from the OLS rolling cointegrating regressions.

We also report outcomes from Engle and Granger’s ADF test, which assumes I(1) variables and tests for a unit root in the cointegrating residuals. A 10% significance level was used, with critical values obtained from MacKinnon’s (1996) simulated response surfaces.

When cointegrating vectors are estimated with I(1) variables, Phillips and Hansen (1990) suggest a Fully-Modified OLS (FM-OLS) estimator of the cointegrating regression that accounts for the second-order bias terms from serial correlation and endogeneity that impact the OLS estimator. Kim and Phillips (2001) extend this to fractional cointegration models. As a comparison to using the OLS estimator, we also obtain residuals from using Phillips and Hansen’s FM-OLS estimator, exploring the integration of these residuals in addition to the OLS residuals. For the FM-OLS regressions, we also report outcomes from Hansen’s (1992) test of the null hypothesis of cointegration against the alternative of no cointegration.

IV. Data

To ascertain the sensitivity of our findings to information set and sample size, we examine two cases. Case 1 examines evidence for cointegration between housing prices and six macroeconomic variables (population, personal income, rents, consumer prices, stock market index and mortgage interest rate) and Case 2 adds additional variables likely linked to the housing market (wages and steel, lumber and concrete construction material prices). For Case 1, we have 117 quarterly observations (1980Q4 through 1990Q4) for the U.S. and 102 quarterly observations (1984Q1 through 2009Q2) for Canada. We begin our rolling forward windows with 40 observations, adding one quarter each time to yield 78 windows for the U.S. and 63 windows for Canada. For Case 2, our available sample is shorter ... complete.

For the U.S. we only provide outcomes using one housing price index, while for Canada we consider three indices, described below. The first two Canadian housing price indices are examined for Case 1 with all three considered for Case 2.

The explanatory variables (likely fundamental drivers and housing price indices) used in the housing models were chosen based on those used by Mikhed and Zemcik (2009). They include mortgage rates, wages, stock market returns, inflation, personal income, population, rent and the prices of steel, lumber and concrete. Data was collected for Canada and the U.S. and the specific data sources
are listed in the table below. All series were converted to quarterly periodicity to conform to the housing price series.

<table>
<thead>
<tr>
<th>Explanatory variable</th>
<th>Canadian data source</th>
<th>American data source</th>
</tr>
</thead>
<tbody>
<tr>
<td>Mortgage rates</td>
<td>CMHC Cansim 027-0015, Mortgage rate for 5-year term and fixed rate, available monthly</td>
<td>Freddie Mac Primary Mortgage Market Survey, Mortgage rate for conforming 30-year fixed rate, available monthly</td>
</tr>
<tr>
<td>Steel, lumber and concrete</td>
<td>Statistics Canada Table 329-0039, Industry price index by major commodity, available monthly</td>
<td>Bureau of Labour Statistics, Producer Price Index Industry Data, available monthly</td>
</tr>
<tr>
<td>Stock market</td>
<td>S&amp;P/TSX Composite index, monthly close,</td>
<td>S&amp;P 500 Composite index, monthly close,</td>
</tr>
<tr>
<td>Personal income</td>
<td>Statistics Canada, Cansim Table V498165, available quarterly</td>
<td>Bureau of Economic Analysis, Table 2.1 Personal Income, available quarterly</td>
</tr>
<tr>
<td>Population</td>
<td>Statistics Canada, Table 051-0001, available annually</td>
<td>U.S. Census, available annually</td>
</tr>
</tbody>
</table>

The mortgage rate for the most common form of mortgage was selected, which is the 5-year fixed-rate in Canada and the 30-year fixed-rate in the U.S. These values were available monthly and were converted to a quarterly basis by selecting the rate in effect for the months of March, June, September and December. For Canada this is the 5 year posted rate.

The wage measure selected was the hourly wage for employees in processing and manufacturing in Canada and for unskilled employees in the U.S. The Canadian wages were available monthly with the values at the end of the quarters selected. The U.S. wages were available annually, and straight-line interpolation was used to determine quarterly values. The source for the U.S. wages is MeasuringWorth.com, a data site created by Lawrence Officer, Professor of Economics at University of Illinois, Chicago.

Three indices for construction material prices were extracted for each country. Steel, lumber and concrete were chosen as the materials since they are commonly used in the industry and since prices are widely available. Multiple indices were available for each material, and we selected the one based on the least amount of processing. Our intention was to track the prices of the raw materials.

* * *
These price indices were available monthly, and the values at the end of March, June, September and December were used to determine quarterly values.

The value of the Standard and Poor’s Composite stock market index was extracted at the end of March, June, September and December for the Toronto Stock Exchange and for the New York Stock Exchange. The Canadian index is commonly referred to as the S&P/TSX and the U.S. index is the S&P 500.

Consumer price indices are tracked by the Canadian and U.S. government agencies and are widely available. The rent index is the rental accommodation component of the consumer price index.

Personal income is also widely available from government data sources and is the total amount of income earned by all persons in the country from wages, government transfers and other sources.

Population is measured at the national level. A Canadian census is conducted every 5 years with government estimates of annual values also available. Straight-line interpolation was used to determine quarterly values. The U.S. census is conducted every 10 years but the U.S. Census Bureau reports monthly estimates of population. These monthly estimates were used to construct a quarterly series.

Direct comparison of house price indexes across countries is problematic as there are no indexes used in both countries with similar scope and methodology. For the US, the two most common sources are the Federal Housing Finance Agency (FHFA) indexes using a repeat sales like methodology created from Freddie Mac and Fannie Mae mortgage transaction data and the S&P/Case-Shiller repeat sales indices created for 10 cities (20 cities since 2000) from transaction records.\(^2\) The former covers a national geography, but only includes mortgages that are securitized by Freddie and Fannie. This can result in a biased series as higher priced houses that are more likely to be purchased with non-conforming loans and the share of mortgages securitized by Freddie and Fannie fell and then rose over the sub-prime boom and bust. The latter covers a limited number of cities, making it an imprecise measure of “national” prices. There is as well some concern that the large number of foreclosure sales in the last few years results in a downward bias in the estimated house price level. In an effort to be as “national” as possible, we use the FHFA price index. We sue the “transactions” sample, so that the index is based on reported transaction prices and excludes appraised prices from refinancings.

The Canadian house price index most similar to the S&P/Case-Shiller house price indices is the relatively recent series issued by a partnership between Teranet and National Bank.\(^3\) However, this


\(^3\) [http://www.housepriceindex.ca/](http://www.housepriceindex.ca/)
index only covers six of the largest Canadian cities, and is limited to three prior to 1998. It is the only transaction-based index available in Canada and is based on all transactions in a metropolitan area. Statistics Canada as part of the CPI has a house price index for new homes that is available at national, provincial, and metropolitan area aggregates. The index is based on the prices reported in a survey of builders of new single family detached homes, where there is an attempt to get repeat observations from the same house type in the same development, but this cannot be assured for any city or period. For some cities, the Statistics Canada new house price index does not appear to reflect actual price movements. A third choice for a Canadian house price index is complied by the UBC Centre for Urban Economics and Real Estate (CUER) from numbers reported by Royal LePage in their Survey of Canadian House Prices. Royal LePage surveys its members, asking for the estimated sales price of consistent house types in multiple communities within a single metropolitan area, and then across all major Canadian cities and in all provinces. Although it is survey rather than transaction data, it performs quite well when compared to transaction based repeat sales measures. The compilation price index is the average of the indices for each of nine Canadian metropolitan areas, reflecting a mix of national geography and size (the nine cities include nearly 50 percent of the Canadian population, weighted by population. In this study we use the Statistics Canada national new house price index (hpi) and the weighted average of major cities index using the Royal LePage survey data (LePage1) to measure the temporal variation in Canadian house prices.

Case 1: Orders of Integration

Using the levels of the series, Tables 1 and 2 summarize the outcomes for the DF-GLS and KPSS tests, along with the minimum and maximum fractional integration parameters. The DF-GLS test suggests that each series is at least I(1), with rejection of the unit root hypothesis occurring for few windows. Typically, the KPSS test supports the DF-GLS results, although there are many windows for which the two tests provide conflicting results, a feature often observed with I(1)/I(0) tests. The range of estimated fractional integration parameters does not support stationarity for any series, but rather that the series are nonstationary long memory processes, possibly reasonably modelled as I(1). Robinson (1995) also provides a test of equivalence of the fractional integration parameters, which only rejects equivalence (at least at the nominal 1% level) for five of the 78 windows for the U.S. and for nine of the 63 windows for Canada.

<table>
<thead>
<tr>
<th>Series</th>
<th>DF-GLS</th>
<th>KPSS</th>
<th>d range</th>
</tr>
</thead>
<tbody>
<tr>
<td>H₀: I(1)</td>
<td></td>
<td>H₀: TS</td>
<td>min, max</td>
</tr>
</tbody>
</table>

Table 1. US: Outcomes for order of integration using levels series

---

4 The six are Calgary, Halifax, Montreal, Ottawa, Toronto, and Vancouver. The three: Halifax, Montreal, and Vancouver.

5 Over the 1983-93 trough to peak for Vancouver prices, a real repeat sales index of single family house transaction prices increased by 114 percent. This compares with a decrease of 4 percent for the Statistics Canada series. The correlation with the real repeat sales index is 0.16.


7 108 percent for the Royal LePage series and are 0.95 and
We repeated this part of the analysis for the first differenced data, examining whether there is evidence of double unit roots or greater degrees of nonstationarity than suggested by a unit root; Mikhed and Zemčík (2007) find that some of the U.S. variables were better modelled as I(2) rather than I(1). Results are reported in Tables 3 and 4.
Table 3. US: Outcomes for order of integration using first differenced series

<table>
<thead>
<tr>
<th>Series</th>
<th>DF-GLS</th>
<th>KPSS</th>
<th>d range</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\Delta$(mtgrates)</td>
<td>support $H_0$</td>
<td>support $H_0$</td>
<td>-0.072, 0.187</td>
</tr>
<tr>
<td>$\Delta$(cpi)</td>
<td>support $H_0$</td>
<td>support $H_0$</td>
<td>-0.164, 0.371</td>
</tr>
<tr>
<td>$\Delta$(rent)</td>
<td>support $H_0$ except for $S_{end} = 114-117$</td>
<td>support $H_0$ except for $S_{end} = 45-80, 104-116$</td>
<td>-0.023, 0.426</td>
</tr>
<tr>
<td>$\Delta$(hpi)</td>
<td>support $H_0$ except for $S_{end} = 82-112$</td>
<td>support $H_0$</td>
<td>0.216, 0.778</td>
</tr>
<tr>
<td>$\Delta$(pers. inc.)</td>
<td>reject $H_0$ except for $S_{end} = 44, 66-92, 94, 98, 100-112$</td>
<td>support $H_0$ except for $S_{end} = 63-115$</td>
<td>-0.175, 0.482</td>
</tr>
<tr>
<td>$\Delta$(s&amp;p)</td>
<td>reject $H_0$ except for $S_{end} = 65-72, 75, 77, 81, 82, 84, 71, 78-80$</td>
<td>support $H_0$ except for $S_{end} = 68, 70, 71, 78-80$</td>
<td>-0.262, 0.221</td>
</tr>
<tr>
<td>$\Delta$(pop)</td>
<td>support $H_0$</td>
<td>reject $H_0$</td>
<td>0.269, 0.758</td>
</tr>
</tbody>
</table>

Notes: LS = level stationary; d is the integration parameter; $S_{end}$ = end observation # for window; tests were undertaken at the 10% significance level; observation 1 corresponds to 1980Q4.

More often than not, we see lack of consistency between the DF-GLS and KPSS tests, except for U.S. and Canadian personal income, U.S. and Canadian population, U.S. stock prices, and a significant number of windows for U.S. hpi. The d estimates provide some guidance in reconciling these differences. Negative d values suggest a process that is nonstationary but less so than for a unit root process, whereas d values less than 0.5 suggest a process that is stationary but with more persistence than an I(0) process. The results suggest that it is likely reasonable to approximately model mortgage rates, personal income, consumer prices, stock price indices and rents as I(1) processes but that the population and housing price series may have a greater degree of nonstationarity than implied by a unit root, although this finding is window dependent. Interestingly, Canada’s Royal LePage housing price series exhibits less persistence than Statistic’s Canada ??????. Not surprisingly, there is no evidence to suggest that the U.S. or Canadian first differenced processes have the same degree of integration. However, for both the U.S. and Canada, Robinson’s (1995) F-test supports that $\Delta$(pop) and $\Delta$(hpi) have equivalent fractional integration parameters for all windows, whereas the test only supports the null that Canada’s $\Delta$(pop) and $\Delta$(LePage1) have the same d for approximately one-third of the windows. This latter finding is not surprising given the d estimates reported in Table 4.

Table 4 Canada: Outcomes for order of integration using first differenced series

<table>
<thead>
<tr>
<th>Series</th>
<th>DF-GLS</th>
<th>KPSS</th>
<th>d range</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\Delta$(mtgrates)</td>
<td>support $H_0$</td>
<td>support $H_0$</td>
<td>-0.072, 0.187</td>
</tr>
<tr>
<td>$\Delta$(cpi)</td>
<td>support $H_0$</td>
<td>support $H_0$</td>
<td>-0.164, 0.371</td>
</tr>
<tr>
<td>$\Delta$(rent)</td>
<td>support $H_0$ except for $S_{end} = 114-117$</td>
<td>support $H_0$ except for $S_{end} = 45-80, 104-116$</td>
<td>-0.023, 0.426</td>
</tr>
<tr>
<td>$\Delta$(hpi)</td>
<td>support $H_0$ except for $S_{end} = 82-112$</td>
<td>support $H_0$</td>
<td>0.216, 0.778</td>
</tr>
<tr>
<td>$\Delta$(pers. inc.)</td>
<td>reject $H_0$ except for $S_{end} = 44, 66-92, 94, 98, 100-112$</td>
<td>support $H_0$ except for $S_{end} = 63-115$</td>
<td>-0.175, 0.482</td>
</tr>
<tr>
<td>$\Delta$(s&amp;p)</td>
<td>reject $H_0$ except for $S_{end} = 65-72, 75, 77, 81, 82, 84, 71, 78-80$</td>
<td>support $H_0$ except for $S_{end} = 68, 70, 71, 78-80$</td>
<td>-0.262, 0.221</td>
</tr>
<tr>
<td>$\Delta$(pop)</td>
<td>support $H_0$</td>
<td>reject $H_0$</td>
<td>0.269, 0.758</td>
</tr>
</tbody>
</table>
Table of Results

<table>
<thead>
<tr>
<th>Variable</th>
<th>Test Statistic</th>
<th>Notes</th>
</tr>
</thead>
<tbody>
<tr>
<td>Δ(mtgrates)</td>
<td>support H₀</td>
<td>support H₀, -0.143, 0.104</td>
</tr>
<tr>
<td>Δ(cpi)</td>
<td>support H₀</td>
<td>support H₀, -0.139, 0.325</td>
</tr>
<tr>
<td>Δ(rent)</td>
<td>support H₀</td>
<td>except for Sₜₚₜ = 63-69, 71, 78, 80, 88, 90-92, 94-97, 101</td>
</tr>
<tr>
<td>Δ(hpi)</td>
<td>reject H₀ except for Sₜₚₜ = 40-43</td>
<td>reject H₀ except for Sₜₚₜ = 40, 43, 0.535, 0.817</td>
</tr>
<tr>
<td>Δ(LePage1)</td>
<td>support H₀</td>
<td>except for Sₜₚₜ = 67-75, 77, 80, 88, 97-101</td>
</tr>
<tr>
<td>Δ(pers. inc.)</td>
<td>support H₀</td>
<td>except for Sₜₚₜ = 70, 71, 73-88, 99-102</td>
</tr>
<tr>
<td>Δ(tsx)</td>
<td>support H₀</td>
<td>except for Sₜₚₜ = 80-87, 101, 102</td>
</tr>
<tr>
<td>Δ(pop)</td>
<td>support H₀</td>
<td>support H₀, 0.688, 0.898</td>
</tr>
</tbody>
</table>

Notes: LS = level stationary; d is the integration parameter; Sₜₚₜ = end observation # for window; tests were undertaken at the 10% significance level; observation 1 corresponds to 1984Q3.

The implications from this first stage for fractional cointegration with respect to Robinson and Marinucci’s (2003) condition that for some are:

For the U.S., it is likely that hpi and pop have greater, but similar, degrees of nonstationarity than implied by a unit root, and the other variables may be reasonably approximated by a unit root process;

For Canada, using hpi as the housing price index, the results suggest that hpi and pop have greater, but similar, degrees of nonstationarity than implied by a unit root, and the other variables may be reasonably approximated by a unit root process. For Canada, using LePage1 as the housing price series, LePage1 has more nonstationarity than a unit root process, but perhaps less than for pop, whereas an I(1) assumption seems reasonable for the other variables. Such features may imply, for many windows, that LePage1 and pop do not have the same d value.

Turning to estimation of the cointegrating regressions, Figures 1 to 3 provide plots of the OLS coefficient estimates, excluding for the deterministic terms, over the windows, with the dates that appear on the horizontal axis corresponding to the end quarter used in each of the rolling forward regressions. The coefficients vary widely, relatively so for mortgage rates, rent and population, indicating the instability of the link between housing prices and these drivers. Considering traditional
I(1)/I(0) cointegration tests, irrespective of window, the EG-ADF and Hansen tests suggest no evidence of cointegration for the Canadian case, whether we use hpi or LePage1 as the dependent variable. The U.S. results are just as overwhelmingly against cointegration, with the EG-ADF test always supporting the null of noncointegration. In this case, the Hansen test results do suggest cointegration for 18 of the 78 windows, but not so for the other windows.\(^8\)

To ascertain whether there is evidence of fractional cointegration, we estimated the long-memory parameter \(d_u\) of the residuals. Figures 4 and 5 provide graphical evidence for this feature for our models. In each figure, we plot the estimated long-memory parameter obtained for the FM-OLS and OLS residuals\(^9\), along with the maximum \(d\) value for the variables included in the regressions. Examining Figure 4, which provides the U.S. outcomes, we see that the residuals dominantly have a lower degree of integration, typically in the nonstationary region. Aside from windows ending in the mid-1990s, the long-memory estimates for the residuals are reasonably stable, suggesting that U.S. housing prices are fundamentally linked with these drivers but that the equilibrium errors exhibit long-memory such that shocks persist. The evidence presented here does not suggest that the links between housing prices and its fundament drivers substantially altered during the 2000 to 2006 price run up, in contrast to the claims of Mikhed and Zemčík (2007).

The Canadian outcomes (Figures 4 and 5) are less stable than observed for the U.S., with evidence of possibly stationary \((d<0.5)\) residuals for windows that include observations up to the end of the 1990s. However, fundamental shifts occurred thereafter, with little evidence of fractional cointegration for windows ending with observations in the early 2000 to end of 2005, the period over which Canadian housing prices substantially increased. These features suggest that Canadian housing prices may have diverged from fundamentals over this period. Some evidence of nonstationary fractional cointegration reappears when the last three to four years of observations are included. Although some characteristics differ, these results are similar whether hpi or LePage1 are used as the dependent variable in the cointegrating regressions. All in all, the findings suggest that the links between housing prices and the drivers we examined (mortgage rates, cpi, personal income, rent, stock prices and population) are less stable for Canada than for the U.S.

**Figure 1. US: Coefficient estimates from hpi OLS cointegrating regressions**

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\(^8\) The explicit results for the EG-ADF and Hansen tests are available on request.

\(^9\) We recognize that our estimates of \(d\) for the errors may be upwardly biased due to omitting short run dynamics. We leave an exploration of this for future work.
Figure 2. Canada: Coefficient estimates from hpi OLS cointegrating regressions
Figure 3. Canada: Coefficient estimates from LePage OLS cointegrating regressions
Figure 4. US: long-memory estimates from cointegrating regressions
Figure 5. Canada: long-memory estimates from cointegrating regressions (HPI equations)
Figure 6. Canada: long-memory estimates from cointegrating regressions (LePage1 equations)

To be completed....
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