Credit Conditions and Consumption, House Prices and Debt: What Makes Canada Different?

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Abstract

We propose new ways to think about the set of equations governing the household sector in general equilibrium models by incorporating real-financial linkages between consumption, household balance sheets and credit markets. Canada shares similarities with the United States and several other countries in that a long-term easing in credit conditions has increased consumption relative to income and expanded household portfolios. Canada differs in that higher house prices relative to income have an overall negative impact on consumption relative to income, except when offset by an easing in credit conditions. Much of the rise in house prices and debt since the late-1990s can be explained by cheaper and easier access to mortgage credit.

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1. Introduction and motivation

Integrating linkages between real and financial variables, particularly in the household sector, into mainstream macroeconomic models is a major and ongoing challenge in the wake of the global financial crisis (Kohn, 2009, Poloz 2015). General equilibrium (GE) models in levels such as the Fed's FRB-US model (Braydon and Tinsley, 1996) and the Bank of Canada's Large Empirical and Semi-structural model (LENS, Gervais and Gosselin, 2014) offer a promising platform for these efforts. This paper proposes new ways to think about the set of equations governing the household sector within such a structure. We test several theoretical propositions on how real-financial linkages may operate and compare our empirical results for Canada with findings for other countries.

Consumption is the largest expenditure sector for most developed countries and is therefore a central element of GE models. Conventional approaches assume that there are two types of representative households: a fixed proportion of "liquidity-constrained" households whose consumption is limited by current income; and "unconstrained" households who respond to changes in permanent income and optimise consumption subject to adjustment costs. The cornerstone model for the latter is the "solved-out" life-cycle/permanent income hypothesis (LCH-PIH) model, where the level of consumption is a function of the discounted present value of expected future "non-property" income (labour income plus net transfers) and net assets.

The two-agent approach contrasts with the heterogeneous agent view of households (Deaton 1991; Carroll 1992, 2001; Kaplan et al. 2016). In this view, households face differing degrees of uninsurable income uncertainty and liquidity (and credit) constraints arising from the composition and distribution of household portfolios. For example, wealthy households (but with mostly illiquid assets) may exhibit "hand-to-mouth" behaviour, or high sensitivity to transitory income changes and low sensitivity to interest rates except through portfolio rebalancing (Kaplan et al. 2014). These features lead to more complex responses for *average* consumption to changes in income, interest rates or wealth.

The reliance on a single net worth term, where all assets have an identical and constant impact on consumption as in Ando and Modigliani (1963) and popular policy models since, is problematic. Cooper and Dynan (2016) review extensive literature on this point. They conclude that the marginal propensity to consume (MPC) out of assets may vary across: asset type, depending on liquidity[1] and access to credit; age, income and wealth distributions; countries, depending on institutional and distributional features; and time, depending on changes in any of the aforementioned characteristics. They also question the assumption that debt has the same impact on consumption as a negative asset. Our paper is consistent with the foregoing ideas and in doing so presents a model of consumption that can incorporate real-financial interactions and structural changes therein.

We focus on the determinants of three key variables: consumption; house prices; and mortgage debt (Chart 9.1). The system could be extended with equations for liquid and illiquid financial assets, and acquisition of housing, thereby comprising the household "block" of a GE model. The starting point for our analysis is the multi-equation approach developed by Muellbauer and Williams (2011) for Australia, Duca and Muellbauer (2013) for the United States, and Aron and Muellbauer (2013) for South Africa. Our approach yields small, tractable, partial equilibrium models that can be estimated with standard techniques to test economic relationships and parameter stability. This approach also makes it possible to trace the credit and monetary policy transmission channels operating on consumption via house prices and debt, since these explicitly modelled within the system.

1. For example, it is highly implausible that cash should have the same MPC as pension wealth.

"Classical" LCH-PIH theory extended to include a separate role for housing wealth (but not mortgage credit constraints) suggests that higher house prices may have a small negative or negligible impact on aggregate consumption (Buiter 2010, Aron et al. 2012). This is because the predicted impact comprises a negative income effect for all households (i.e. higher rent and imputed rent costs), offset by a "pure" wealth effect received by homeowners only. The net of these two effects would then depend on tenure structure (the proportion of renters versus owner-occupiers), which differs across countries. This is an important reason to reject net worth as the relevant measure of wealth for modelling consumption.

When mortgage credit is added to the model, the potential effects of house prices on consumption become more complex and their impact becomes an empirical question. First, house purchases typically involve a mortgage down-payment. This is a major savings event in the household life cycle. The size of the down-payment is determined by house prices and the cost and availability of mortgage credit, none of which are constant. Second, in some countries home-owners have gained access to home equity loans (residentially-secured, revolving lines of credit). This relatively new form of lending enables home-owners to leverage housing capital gains and use the borrowed funds for consumption or investment. This housing collateral effect hinges on country- and time-specific mortgage arrangements. Third, the MPC out of debt may be larger than (minus) that on other assets given its greater liquidity (Otsuka 2004) or due to households' uncertainty about future access to credit (Cooper and Dynan, 2016).
The dichotomy between "liquidity-constrained" (current income) and "unconstrained" (permanent income) consumers is therefore a false one. We propose and test a more generalised and credit-augmented treatment of the household sector.

The paper is organized as follows. Section 2 explains the Latent Interactive Variable Equation System (LIVES) methodology. Section 3 outlines the theoretical foundation for the three equations (consumption, house prices and mortgage debt). Section 4 provides the institutional backdrop, summarizing the major structural changes affecting household borrowing conditions in Canada over recent decades. Section 5 summarizes the data and discusses the approaches used for the observed and unobserved variables. Section 6 outlines the empirical results, including a comparison in Section 6.5 between our LIVES approach and results from a cointegrated VAR approach. Section 7 concludes.

2. Methodology used to estimate latent credit conditions effects

Macroeconomic time series are often subject to structural breaks that cause mean-shifts or location shifts that could be common to several variables (Stock and Watson 1988, Hendry 1997). Hendry and Mizon (2014a and 2014b) argue that models that ascertain the likely persistence of structural breaks ex-post offer the best prospects for mitigating systematic forecast failure. We expand on this idea.

How do we assess the impact of structural changes in household credit arrangements on consumption, house prices and debt? Unlike other countries, Canada does not possess direct time-series data on non-price changes in credit arrangements.[2] An indirect approach is necessary. We propose using a Latent Interactive Variable Equation System (LIVES) approach to control for the impact of a common unobserved structural change in consumption, house prices and mortgage debt. Because we control for a fairly exhaustive set of controls for *other* economic and demographic factors (including interest rates and expectations about future income), thereby ruling out many alternative interpretations, we suggest that the latent variable could be interpreted as non-price mortgage credit conditions (Chart 9.2).

We also explore a control for shifts in access to home equity finance that is not latent but based on debt

2. By contrast, the US Federal Reserve's *Senior Loan Officer Opinion Survey of Bank Lending Practices* provides useful information on unsecured credit conditions since 1966 and on the mortgage market since 1990. The Bank of Japan has surveyed lenders about household credit conditions since 2000, the European Central Bank has done so since 2003 and the Bank of England has done so since 2007.

data (Chart 9.4). The LIVES method yields comparable results with an approach where credit conditions can be directly measured using lender-survey data (Duca and Muellbauer, 2013).[3]

We estimate a system of three equilibrium correction models for consumption, house prices and the mortgage stock using full information maximum likelihood (FIML). We hypothesize that the long-run time series in each equation are approximately cointegrated through their possession of an unobserved stochastic trend. This latent influence is common to all equations and causes an evolving, structural mean-shift. There may also be interaction effects between selected regressors and the latent variable or the control for access to home equity loans.

Each equation contains the latent variable defined as a spline function composed of a linear combination of smoothed step dummies (discussed below). The spline function has a scaled effect in each equation but its slope coefficients are common and jointly estimated. The function captures time variation in the intercept and, via interaction effects, the parameters of key explanatory variables in each equation. This latent variable approach is somewhat different from the literature on time-varying cointegration (Bierens and Martins 2010; Park and Hahn 1999) in that the time variation in our study has an explicit economic interpretation in each equation. This permits the use of institutional knowledge about developments in Canadian housing finance systems, consumption theory and parameter estimates from comparable international studies in order to impose priors on key parameters in the system (including the slope and scale coefficients of the spline function).[4]

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3. See also Muellbauer and Williams 2011, Aron and Muellbauer 2013, Chauvin and Muellbauer 2013, Duca et al. 2016 and Geiger et al. 2016.

4. The LIVES approach, in which sub-systems of equations containing common latent variables are part of a larger GE model, also has potential for dealing with and interpreting other sources of structural change, e.g. in the production sector, a major cause of forecast failure, Hendry (1997) and Hendry and Mizon (2014a).

The generalized LIVES approach is

$$\Delta y_{it} = \varphi_i(\alpha_{i0} + \kappa_i CCI_t + \sum_j \acute{\alpha}_{ijt} z^*_{ijt-1} + \sum_p \alpha_{ip} z_{ipt-1} - y_{it-1}) + \sum_q \acute{\beta}_{iqt} \Delta x^*_{iqt} + \sum_v \beta_{iv} \Delta x_{ivt} + \varepsilon_{it} \qquad (2.1)$$

where $i \in [c, h, m]$,

$\kappa_i \equiv 1$ for equation $i = h$,

$\acute{\alpha}_{ijt} z^*_{ijt-1} = \alpha_{ij} z^*_{ijt-1} + \alpha^*_{ij} CCI_t \left(z^*_{ijt-1} - \overline{z_{ij}}\right)$, and

$\acute{\beta}_{iqt} \Delta x^*_{iq} = \beta_{iq} \Delta x^*_{iqt} + \beta^*_{iq} CCI_t (\Delta x^*_{iqt} - \overline{\Delta x_{iq}})$.

Here, y_i is the dependent variable for equation i, φ_i is the corresponding speed of adjustment to equilibrium and α_{i0} is the intercept or autonomous level of the dependent variable. c, h, m denote the three equations in the system. A spline function representing the unobserved variable (CCI) causes a series of shifts in the intercept scaled by κ_i in each equation. There are two types of long-run explanatory variables: z^*_{ij} are long-run explanatory variables that are tested for an interaction with CCI (that is, whether their coefficients, $\acute{\alpha}_{ijt}$, are time-varying with credit conditions, CCI); z_{ip} are long-run explanatory variables that are not interacted with CCI (that is, their coefficients α_{ip} are time-invariant). There are also two types of short-run variables. Δx^*_{iqt} are short-run explanatory variables that are tested for an interaction with CCI (that is, whether their coefficients $\acute{\beta}_{iqt}$, are time-varying with credit conditions). Δx_{ivt} are short-run explanatory variables that are not interacted with CCI (that is, their coefficients β_{iv} are time-invariant). All parameters should, in principle, be uniquely identified[5] with the exception of κ_i, which we set to 1 in the house price equation.

The history of Canadian credit markets (Section 4) suggests rolling changes in institutional arrangements. Our general-to-specific approach allows for this possibility. The spline function is defined

5. In practice, where data have insufficient independent variation, estimates of interaction effects can be quite imprecise and then all but the most economically relevant parameters can be omitted.

as a linear combination of smoothed step dummies (SSD).[6] We first estimate a very general spline function in the system and then simplify it by omitting insignificant dummies. CCI equals the linear combination of the (significant) dummies, where the jointly estimated slope coefficient on each dummy is τ_s. We opt initially to include a dummy every two years (1982Q1, 1984Q1 and so on) plus an extra dummy around the GFC. We find this strategy is sufficiently general to detect the major latent variable effects on consumption, house prices and debt (see Section 6.4).

$$CCI = \sum_{s=1} \tau_s SSD_s = \tau_1 SSD1982q1 + \tau_2 SSD1984q1 + \dots + \tau_{17} SSD2012q1 \quad (2.2)$$

We therefore hypothesize that the I(1) long-run variables y_i, z_{ij}^*, z_{ip} are cointegrated subject to a mean-shift for y_i and z_{ij}^* captured by CCI.The strong significance of the long-run parameters and of the speeds of adjustment implies, and is implied by, cointegration. As a robustness check, we estimate the parsimonious models as VAR models and apply standard cointegration tests (in Section 6.5).

3. Theoretical models of consumption, house prices and debt

3.1.Consumption

The cornerstone of our consumption model is the Ando-Modigliani-Brumberg-Friedman “solved-out” consumption function.[7] The classical solved-out function, linearized in logs, is

$$\ln c_t/y_t \approx \alpha_0 + \gamma A_{t-1}/y_t + \psi E_t \ln(y_t^p/y_t) \quad (3.1)$$

6. Each SSD takes the values 0.05, 0.15, 0.3, 0.5, 0.7, 0.85, 0.95 and 1 over eight quarters, then remains at 1. SSD thus has an ‘S’ or ‘ogive’ shape.
7. A more detailed derivation is provided in Muellbauer, St-Amant and Williams (2015), i.e. the working paper version of this document.

where the log aggregate consumption to non-property income ratio ($\ln c_t/y_t$) is composed of three elements.[8] First, there is (time-invariant) autonomous consumption share, α_0. Second, there is the ratio of net household wealth (at end of period t-1) to income, A_{t-1}/y_t, scaled by the marginal propensity to consume out of assets (γ). Third, there is the expected log ratio of permanent to current non-property income, $E_t \ln(y^p/y_t)$, scaled by (ψ) to allow for a generalization of the strict permanent income hypothesis. All variables are real per capita levels. $\ln(y_t^p/y_t)$ can be approximated by a weighted moving average of forward-looking income growth rates (Campbell 1987), where k is the time horizon and δ is the discount factor applied to future income

$$\ln(y_t^p/y_t) = \left(\sum_{s=1}^{k} \delta^{s-1} \ln y_{t+s}\right)/\sum_{s=1}^{k} \delta^{s-1}) - \ln y_t. \qquad (3.2)$$

The discount factor δ is assumed to be well below $1/(1+r)$ where r is a risk-free real rate of interest. Together with $\psi < 1$, this reflects the theories, under income uncertainty, of liquidity-constrained consumption of Deaton (1991) and of precautionary saving of Carroll (1992, 2001). Both authors argue that many households discount future income more heavily than in the simple textbook model, and Kaplan and Violante (2014) and Kaplan et al. (2014, 2016) argue that this also applies even for fairly wealthy households when their assets are subject to lumpy transactions costs. Though micro-heterogeneity is central to this literature and there is no “representative agent” in the classical sense, there are profound implications for aggregate data. For example, Anderson et al. (2012) and Kaplan and Violante (2014) confirm strong heterogeneity at the micro level to fiscal policy shocks and the latter

8. Non-property income (y) refers to labour income plus net transfers. See Blinder and Deaton (1985).

paper finds a strong aggregate US consumer response to the 2001 tax rebate. For aggregate data, it therefore makes sense to assume substantial discounting of future income on average.[9]

The dynamic version of Equation (3.1) is

$$\Delta \ln c_t \approx \varphi(\alpha_0 + \gamma A_{t-1}/y_t + \psi E_t \ln(y_t^p/y_t) + \ln(y_t/c_{t-1})) + \lambda \Delta \ln y_t \tag{3.3}$$

where φ is the speed of adjustment to equilibrium and $\Delta \ln y_t$ is the change in log current income. The latter term allows for some time-invariant proportion of households (λ) that are constrained by myopia, liquidity constraints, inattentiveness or use of rules-of-thumb (Campbell and Mankiw 1989, 1991; Gali, Lopez-Salido and Valles 2004; Reis 2006). Equation (3.3) assumes no shifts in the age structure of the population or other factors; constant real interest rates; homogenous net assets ($\gamma = \bar{\gamma}$); and that credit constraints are time-invariant and only affect consumption by determining the proportion of Keynesian consumers, λ. Equations (3.1) and (3.3) are homogenous of degree one such that doubling income and assets doubles consumption.

We now generalize the model and introduce controls for time-varying credit access and allowing marginal propensities to consume of assets to vary by asset type, we augment Equation (3.3):

$$\begin{aligned}\Delta \ln c_t \approx \varphi(\alpha_0 + \kappa_c CCI_t + \acute{\alpha}_{1t} r_{t-1} + \gamma_1 NLA_{t-1}/y_t + \gamma_2 IFA_{t-1}/y_t + \acute{\gamma}_{3t} HA_{t-1}/y_t + \acute{\alpha}_{2t} \ln(hp_{t-1}/y_{t-1}) \\ + \acute{\psi}_t E_t \ln(y_t^p/y_t) + \ln y_t/c_{t-1}) + \acute{\lambda}_t \Delta \ln y_t + \acute{\beta}_{1t} \Delta ue_t \ .\end{aligned} \tag{3.4}$$

The autonomous consumption share now includes the intercept (α_0) and a credit conditions index (CCI) with coefficient κ_c; r is the real household borrowing rate; NLA_{t-1}/y_t is the ratio of net liquid assets

9. We assume that households on average look ahead to incomes over the next 10 years (k = 40) with a quarterly discount factor of δ = 0.95. We rely on a forecasting model that embodies the notion that the deviation of log average permanent income to average current income is explained by deviation of log current income around a trend, augmented by some forward-looking economic variables. See Muellbauer et al. 2015.

(cash and cash-like assets less household debt) to non-property income; IFA_{t-1}/y_t is the ratio of illiquid financial assets (financial assets less liquid assets) to income; HA_{t-1}/y_t is gross housing assets to income; $\ln(hp_{t-1}/y_{t-1})$ is the log ratio of real house prices to income; $E_t\ln(y_y^p/y_t)$ is the log ratio of permanent to current income ; and Δue is the change in the unemployment rate as a proxy for income uncertainty (Kimball 1990). Time subscripts on key parameters denote the form $\acute{\alpha}_{jt}z = \alpha_j z + \alpha_j^* CCI_t(z - \overline{z})$, where $\acute{\alpha}_{jt} = \alpha_j + \alpha_j^* CCI_t$ is the total coefficient on (de-meaned) explanatory variable z subject to potential non-linear parameter shifts due to CCI. If there are no parameter time shifts ($\alpha_j^* = 0$), then $\acute{\alpha}_{jt} = \alpha_j$.

We now explain Equation (3.4) in more detail. The intercept comprises $\alpha_0 + \kappa_c CCI_t$, where the latter term controls for fluctuations in long-run $\ln c/y$ due to (unobserved) shifts in household credit access, particularly mortgage credit access. Due to imperfect information, real interest rates do not perform a Walrasian auctioneer's role in clearing the credit market. Rather, lenders rely on non-price constraints to mitigate default risk. As credit access evolves, households without collateral may need to forego more or less current consumption to save for a deposit. This affects long-run $\ln c/y$ and its obverse, the household savings rate. Movements in log real house prices to income ($\ln hp/y$) could be partially offsetting because, notwithstanding changes in credit availability, higher house prices could still necessitate larger deposits by younger households hoping to enter the housing market.

Whereas equations (3.1) and (3.3) assume homogenous assets, Equation (3.4) adopts a three-part disaggregation of household net worth whereby the marginal propensity to consume out of assets depends on liquidity and credit access. This disaggregation is possible by virtue of working with asset levels relative to income rather than log levels. Net liquid assets can be used to buffer against unanticipated income fluctuations and therefore should have a higher marginal propensity to consume (MPC) compared with other asset types (Otsuka 2004). Subject to the state of mortgage market liberalization, housing collateral could also be used as a buffer stock (Miles 1992; Parkinson et al. 2009). Muellbauer and Williams (2011) and Aron et al. (2012) find that countries with deep mortgage markets and easier access to home equity loans, such as Australia, the United Kingdom and the United States, have a positive housing-wealth MPC that drifts upward as credit conditions become easier. Households with existing housing wealth can spend some of their capital gains through refinancing or drawing against home equity loan products that became popular through the latter half of the 1990s.

Life-cycle/permanent income hypothesis theory predicts that, in the absence of credit effects, price-induced housing-wealth effects on aggregate consumption are small or possibly negative (Aron et al. 2012).[10] Empirical evidence suggests that countries with shallow mortgage markets (conservative down-payment requirements) and no home equity loans, such as Japan (Aron et al. 2012), France (Chauvin and Muellbauer 2013) and Germany (Geiger, Muellbauer and Rupprecht 2016) exhibit negative housing-wealth effects. The foregoing discussion motivates a three-fold disaggregation of household net worth: net liquid assets (NLA, defined as cash and short-term deposits less household debt); illiquid financial assets (IFA, defined as financial assets less liquid assets and therefore mostly directly held pension assets, life insurance and securities); and gross housing assets (HA), where the latter is potentially subject to a parameter shift with access to home equity loans.

A further consideration is that easier credit access could facilitate greater inter-temporal consumption smoothing. In Equation (3.4), this would imply a positive drift in the coefficient on expected future income growth ($\acute{\psi}_t$), a negative drift on the coefficient for (time-varying) real interest rates ($\acute{\alpha}_{1t}$), a downward drift toward zero on the current income growth coefficient as the proportion of non-life-

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10. Housing services are part of the consumption bundle so when house prices rise, the positive consumption effect from wealth gains to home-owners is offset by the negative income effect of higher housing costs (affecting *all* households). The balance of these two oppositely-signed effects depends on the tenure structure.

cycle hypothesis (LCH) consumers declines ($\acute{\lambda}_t$), and perhaps an upward drift toward zero on the change in the unemployment rate coefficient ($\acute{\beta}_{1t}$) for the same reason. Equation (3.4) includes the current income growth term, but the proportion of non-LCH consumers ($\acute{\lambda}$) is now more attributable to "rules of thumb," myopia or lack of sophistication since liquidity constraints are already embodied in CCI.

All of the above hypotheses are statistically testable. Under six testable restrictions, Equation (3.4) reduces to the classical/conventional model in Equation (3.3): (i) $\kappa_c = 0$ so that $\alpha_0 + \kappa_c CCI_t = \alpha_0$; (ii) $\gamma_1 = \gamma_2 = \acute{\gamma}_{3t} = \bar{\gamma}$; (iii) $\acute{\alpha}_{1t} = \alpha_1$; (iv) $\acute{\psi}_t = \psi$; (v) $\acute{\lambda}_t = \lambda$; and (vi) $\acute{\beta}_{1t} = \beta_1$. Using Canadian data, we can reject hypotheses (i) and (ii) but not (iii), (iv), (v) (in fact, $\hat{\lambda}$ is statistically zero in CCI-inclusive estimations) and (vi). In contrast, Gervais and Gosselin (2014) impose these constraints in the long-run solution for non-durable consumption and services, though they allow a slightly more general adjustment process.

3.2. House prices

We follow a supply-and-demand approach to determining house prices. In this approach, the housing demand function is

$$\ln hs = \alpha_0 + \alpha_1 \ln y - \alpha_2 \ln rent + \sum \alpha_f D \qquad (3.5)$$

where hs is housing services (measured by the real per capita net dwelling stock); y is household real per capita non-property income; $rent$ is the price of housing services;[11] and D is a vector of other factors that shift the demand curve. The income elasticity is α_1 and the own-price elasticity is α_2. An efficient housing market where households optimize the rent versus own decision, in the absence of credit

11. $rent$ is unobserved, since we can only impute the price of housing services that owners charge themselves. Market rent data are not a good proxy, given that the characteristics of owned and rented homes are different.

constraints, implies that $rent = hp.ucc$, where hp is the real house price and ucc is the real housing user cost of capital. We can therefore invert the housing demand function as

$$\ln hp = (\alpha_0 + \alpha_1 \ln y - \ln hs + \sum \alpha_f D)/\alpha_2 - \ln ucc. \tag{3.6}$$

The real housing user cost (ucc) can be represented as

$$ucc = r^* + t + c + dep + risk - E_t(\dot{hp}/hp) \tag{3.7}$$

where the real after-tax opportunity cost is r^*; property tax costs are t; transaction costs amortized over the ownership period are c; depreciation costs are dep; the risk premium required to compensate home ownership is $risk$; and, as an offset to these costs, expected housing capital gains are $E_t(\dot{hp}/hp)$.

The inverse housing demand function in Equation (3.6) has useful characteristics. First, it has a clear consumption theory foundation. Since the stock of established housing is large relative to the flow of net new supply, supply can be assumed to be fixed except in the long run. This implies that changes in house prices can be attributed to disequilibrium between demand and supply, caused predominately by shifts in demand. Second, our approach has recourse to international literature that can provide plausible ranges of values for key parameters (e.g. Meen 1990, 1996, 2000, 2002). Such literature suggests that the income elasticity of housing demand (α_1) should be around 1 and the own-price elasticity (α_2) between about (minus) 0.5 and 1. This implies that the elasticity of house prices to income given the stock, α_1/α_2, should lie between 1 and 2.

We augment Equation (3.6) for potential credit conditions effects and for a nominal interest rate effect[12]

12. Standard mortgage amortization contracts fix periodic repayments in nominal terms; while, due to asymmetric information, lenders impose repayment/income-based borrowing tests. The interaction of these two institutional features may give rise to nominal interest rate effects beyond the real rate effects occurring through

$$\ln hp_t \approx h_0 + \kappa_h CCI_t + \; h_2(h_1 \ln y_{t-1} - \ln hs_{t-1}) + h_3 \ln ucc_{t-1} + h_4 \ln i_{t-1} + \sum h_f \, D_{t-1} \quad (3.8)$$

where the intercept is h_0; CCI is a credit conditions index controlling for latent shifts in borrowing conditions; y is real per capita non-property income; hs is real per capita net dwelling stock; i is the nominal mortgage borrowing rate; D is a vector of other long-run demand-side variables; and ucc is the real housing user cost. The parameter $\kappa_h \equiv 1$ is the direct long-run effect of CCI on housing demand normalized to 1 to uniquely identify the components of CCI (see Section 2); the income elasticity of housing demand is $h_1 = \alpha_1$; the inverse own-price elasticity of housing demand is $h_2 = 1/\alpha_2$; h_3 is the elasticity of real house prices with respect to the real user cost and is freely estimated to allow for the possibility that, even in the long run, households do not fully optimize the rent versus own decision (i.e., h_3 does not equal -1 as implied by Equation 3.6); h_4 is the elasticity of real house prices with respect to the nominal mortgage rate; $h_f = \alpha_f/\alpha_2$. In contrast to this model, Gervais and Gosselin (2014) assume that, in the long-run, real house prices in Canada are given by a deterministic trend.

Additional explanatory variables in D could include measures of demographic change, investor demand and housing liquidity. Key gaps in the Canadian data are the lack of data on investor housing purchases (particularly about foreign capital flows into Canadian real estate)[13], and investor mortgage demand. These are important omissions because domestic buy-to-let homeowners could be more sensitive to portfolio considerations, such as net rental cash flow, vacancy risks and prospective relative asset class

the user cost. All else being equal, a fall in inflation and nominal rates raises the maximum qualifying loan size for a given repayment/income ratio. See Kearl (1979), Debelle (2004) and Ellis (2005).

13. Ley (2015), Moos and Skaburkis (2010) and Ley and Tutchener (2001) explore the role of Vancouver and Toronto as “gateway cities” for immigration and housing-related international capital flows over recent decades. They find that property markets have become less tied to local labour markets due to structurally higher housing demand from immigrants, mostly from Hong Kong and Taiwan (in the 1980s and 1990s) and Mainland China (since 2000), with high wealth and unreported income that they continue to earn outside Canada. Canada’s Business Immigration Program appears to have facilitated these trends (CIC 2014).

total returns. Conversely, foreign investors might be less sensitive to domestic economic conditions and policy settings. We explore some of these aspects in Section 6.

3.3. Mortgage debt

The mortgage debt equation is difficult to ground because consumption theory has little to say about debt levels. As noted in the introduction, the LCH-PIH approach relies on a single, homogenous, net asset term. In practice, households' motivations for holding debt may include: human capital investment in education or training; buffering against unexpected income fluctuations; borrowing against anticipated higher future incomes; financing the purchase of lumpy consumer durables, house purchases or home renovations; offsetting suboptimal compulsory pension saving arrangements; and portfolio investment in assets for which the risk-adjusted expected returns are greater than the cost of borrowing (Brueckner 1994). Credit conditions could be an important determinant of debt levels, given asymmetric information (Miles 1992; Brueckner 1994; Leece 1995; Meen 1990, 1996). Income, wealth and mortgage rates may also be important.

We settled on a relatively simple long-run mortgage debt specification

$$\ln ms_t \approx m_0 + \kappa_m CCI_t + m_1 \ln y_{t-1} + m_2 \ln i_{t-1} + m_3 \ln(HA_{t-1}/y_t) \qquad (3.9)$$

where ms is real per capita mortgage stock (excluding residentially secured personal lines of credit); y is real per capita non-property income; i is the nominal mortgage rate; and $\ln(HA_{t-1}/y_t)$ is the log ratio of housing assets (at end-period t-1) to income. The intercept is m_0; κ_m is the coefficient on CCI; m_1 is the income elasticity; m_2 is the elasticity with respect to nominal interest rates; and m_3 is the elasticity of real per capita mortgage debt with respect to the ratio of housing assets to income. In contrast to this model, Gervais and Gosselin (2014) specify the total household debt to income ratio, in the long run, as a function of house prices and a real effective interest rate.

4. Institutional background to the estimations

Canada's housing finance system is considered conservative (Klyuev 2008) or prudent (Schembri 2014) by international standards, and it functioned well during the GFC (Crawford, Meh and Zhou 2013).[14] The Canadian household credit market is dominated by large, domestic banks. Governments have influenced non-price credit conditions by changing the terms and conditions for access to mortgage credit insurance, by changing the regulation of lending institutions, or through fiscal incentives. Credit conditions have also been influenced by financial innovations and by developments, such as the GFC, affecting risk perception.

Schembri (2014) and Crawford, Meh and Zhou (2013) explain that Canada's rigorous regulatory and supervisory framework, coupled with targeted government guarantees of mortgage insurance and securitization products, has a significant impact on mortgage underwriting standards and on the types of mortgage products available. Moreover, certain institutional features may incline Canadian households to be cautious in their use of debt. First, most mortgages are subject to full legal recourse. In contrast, a number of U.S. states (Arizona, California, Oregon, etc.) have non-recourse laws. Second, mortgage interest on a personal residence is not tax deductible in Canada, whereas it is in the United States. Third, about three-quarters of Canadian mortgages are fixed for a period of 5 years or less and most of the rest are at floating rates, exposing borrowers to greater interest rate risk than in the United States where mortgages are typically fixed for 30 years.

14. A detailed institutional history is provided an appendix in Muellbauer, St-Amant and Williams (2015).

We see five distinct periods in the evolution of non-price credit conditions since the early 1980s.

1. *Easing from 1982 to 1988*: A 1980 amendment to the *Bank Act* allowed banks to have mortgage loan subsidiaries (Freedman 1998), stimulating the supply of mortgage loans. Non-price credit conditions also became easier as the effects of the 1981–82 recession on risk aversion subsided. In addition, there was a marked reduction of yearly inflation from above 12 per cent to around 4 per cent, which eased access to credit (Debelle 2004).
2. *Tightening between 1989 and 1992*: This was reflected in tighter bank regulation, including the implementation of the provisions of Basel I and reduced maximum leverage ratios, and heightened risk aversion due to the early 1990s recession.
3. *An underdetermined period from 1992 to about 1999*: There were some developments consistent with easier credit conditions (e.g., inflation falling to around 2 per cent, recovery from the early 1990s recession), but others were consistent with tighter credit conditions (e.g., reduced competition in the credit market due to banks buying trusts, Asian and Russian crises).
4. *Easing from 2000 to 2007*: Access to government-backed mortgage insurance was greatly eased, markets for mortgage-backed securities developed rapidly, bank regulation became looser (lower capital ratios) and new participants entered the household credit market (increased credit supply).
5. *Tightening from 2008 to now*: This was caused by the GFC, by more stringent bank regulations, and by a series of moves to tighten access to mortgage insurance. These tightening developments were partially offset, during the worst of the GFC, by Bank of Canada and government interventions (liquidity injections and fiscal stimuli) to reduce market tensions.

To preview our findings in Section 6.4, our estimated CCI is broadly consistent with this prior information (Chart 9.5) and is well-correlated with growth in household credit aggregates (Charts 9.6).[15]

5. Data, variables and specifications

5.1 Data for the observed and unobserved variables

This section provides an overview of the variables and data used to test the specifications for consumption (Equation 3.4), house prices (Equation 3.8) and mortgage debt (Equation 3.9) as outlined in Section 3. The sample contains 133 quarterly observations from 1982(1) to 2015(1). A detailed data appendix is provided in Muellbauer, St-Amant and Williams (2015).

There are two types of data, observed and unobserved. For observed data, the source is predominately national accounts (CSNA 2012), flow of funds money and credit aggregates, and interest rates from Statistics Canada and the Bank of Canada. We generate a measure of non-property income, essentially net labour income plus net transfer components of household gross disposable income excluding net property income (Blinder and Deaton 1985). We estimate in levels because the long-run relationships between variables-in-levels are of most interest to policy-makers. Because we estimate in levels, a multi-decade quarterly sample is necessary to identify multiple inflection points in trending levels. The trade-off is that, for some time series, quarterly data are available only from the 1990s and must be spliced with earlier data of lower quality or lower frequency.

For unobserved variables, we generate proxies. The four unobserved variables are the credit conditions index (CCI) for mortgage credit access in all equations;[16] an index of housing asset liquidity (HLI) for

15. For period (5), our results suggest only a slight tightening in households' access to mortgage credit after 2010.

access to home equity loans; and the real housing user cost and investor demand in the house price equation. CCI is a jointly estimated latent variable (see Section 2 and Chart 9.2). For HLI (Chart 9.4), we use a four-quarter moving average of the ratio of home equity lines of credit (HELOCs) to mortgage credit. The ratio is set to 0 before 1996Q1 and rebased by subtracting the 1996Q1 value of the series thereafter, since we assume HELOCs were not available in Canada before this time. The real housing user cost is real opportunity costs facing households, plus other ownership costs, less a weighted average of lagged house price growth terms.

Finally, for the house price equation, we explore two variables that could proxy the role of housing investor demand: a four-quarter moving average of the ratio of immigration from Asia to total immigration; and HLI. The former variable appears to capture the uneven flow of Asian immigrants and soon-to-be house seekers into gateway cities where housing supply is tighter, thereby pushing up national house prices. It may also proxy for immigration-linked flows of foreign capital into Canadian real estate (Ley 2015; CIC 2014; Moos and Skaburskis 2010; Ley and Tutchener 2001). The second investor proxy is HLI, incorporating the notion that domestic buy-to-let investors have benefited from easier credit access since HELOCs became widely available after about 1996. HLI is also added to the mortgage debt equation on the grounds that HELOC debt and mortgage debt may be complementary.

16. To some extent, the CCI may capture the latent effects of changes in consumer credit access also, since we have not separately controlled for the latter in the consumption equation. We assume mortgage credit access is the dominant latent impact being captured through CCI in the system.

6. Results[17]

6.1. Consumption equation results

Estimation results for the long-run consumption equations are set out in Table 10.1. In Column (1), we estimate Equation (3.3), the "classic" solved-out consumption model. Note that the speed of adjustment is very slow at 11 per cent per quarter. While the estimates for net worth and income growth expectations are significant, the marginal propensity to consume (MPC) out of net worth is low at about 0.02 and the coefficient on income growth expectations is 1.7. The fact that the latter is well above 1 (i.e., households are consuming in excess of the discounted value of expected future income growth) suggests an important omitted variable such as access to credit.

Column (2) relaxes the previous model by allowing for time-varying interest rates; a three-fold disaggregation of net worth; income uncertainty (by adding the change in the unemployment rate to the dynamics); and lagged consumption growth (also in the dynamics, to account for durable goods stock-building). The speed of adjustment is still slow at 14 per cent per quarter. Real interest rates are not significant. The coefficient on income growth expectations is above 1. The point estimates for the wealth MPCs are 0 for net liquid wealth and gross housing wealth (neither are significant) and 0.04 for illiquid financial wealth.[18] In the dynamics (not shown), the estimated proportion of so-called "liquidity-constrained" households ($\hat{\lambda}$) is 0.09 (compared to 0.13 for the Column (1) model). The results for Column (1) and (2) therefore strongly suggest an omitted structural influence on long-run consumption.

17. Appendix 1 provides a concise summary of the empirical specifications.
18. A recurring feature of the Canadian results is the lack of precision around the coefficients for housing wealth and net liquid assets, since both of these variables are trending variables with few inflection points. We had no difficulty pinning down the coefficient on illiquid financial assets to income at around 0.02–0.04.

Column (3) and (4) provide the credit-channel inclusive estimations. There is a striking improvement in the robustness of the long-run solution and the plausibility of the parameters therein. In Column (3), we estimate Equation (3.4) that incorporates controls for shifts in household borrowing conditions. That is, compared with Column (2), we have added the CCI and log real house prices to income. CCI can be interpreted as the jointly estimated long-run impact of the relaxation of mortgage down-payment constraints (and possibly also consumer credit constraints) on consumption. The effect of CCI is partially offset by the extent to which higher house prices relative to income raise the deposit required to enter the housing market; and by the fall in net liquid assets relative to income, resulting from the rise in debt made possible by more relaxed credit conditions. With these two additional controls, Column (3) shows a marked improvement compared with Column (2). The speed of adjustment more than doubles. The equation standard error is lower and the adjusted R^2 is 0.61. The coefficient on expected income growth is now more plausible at 0.8 (t-stat = 5.8). In the dynamics (not shown), current income growth is now only marginally important ($\hat{\lambda}$ drops to 0.08, t-stat = 1.8) consistent with the fact that we have now controlled for time-varying effects of credit access elsewhere in the specification.

Column (4) adds the long-term change in the ratio of children-to-adult population and an interaction effect between the CCI and the (de-meaned) log real house price to income ratio. The equation standard error is lower at 0.0040 and the adjusted R squared is 0.65. The speed of adjustment is 43 per cent per quarter (t-stat = 6.4). This estimate implies that it takes roughly five quarters for 90 per cent of a shock to the long-run consumption-to-income ratio to fade, and is therefore consistent with the generally-accepted time frame of monetary policy effects on real variables. It is also similar to estimates for Australia (Muellbauer and Williams 2011) and for the U.K. and the U.S. (Aron et al. 2012).

Consumption is positively affected by easier credit (CCI) but negatively affected by higher house prices relative to incomes, since the former decreases the need to save for a mortgage down payment while the latter increases it. The estimated coefficient on CCI ($\hat{\kappa}_c$) is normalized by the house price equation and is well-determined (t-stat = 3.6): all else being equal, if CCI has a +1 per cent long-run impact on real house prices, it has a +0.19 per cent impact on the long-run consumption-to-income ratio. The coefficient on log real house prices to income is -0.15 (t-stat = 1.8). We allow for time variation in this parameter in Column (4), whereas in Column (3), we hold it time-invariant.[19] The coefficient on log house prices to income therefore shifts from -0.15 when CCI = 0 to about -0.02 at peak CCI. In other words, roughly 85 per cent of the negative impact of high house prices (given incomes) on consumption is attenuated by easier non-price credit access.[20]

Real household borrowing rates have modestly negative effects as expected.[21] The coefficient on expected income growth is 0.7, consistent with the view that, if some households are myopic or face credit constraints, current income as well as permanent income should matter for consumption. We tested but found no evidence of interaction effects between CCI and interest rates or expected income growth. Demographic change is proxied by the long-term change in the child-to-adult population, with a modestly significant positive coefficient.[22]

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19. In Column (4), we allow for some attenuation of the negative log house price-to-income effect by interacting that term (de-meaned) with the CCI and defining the coefficient as some fraction of the non-interacted coefficient. Its total coefficient is therefore $\alpha_{2t} = \alpha_2 - \alpha_2^* \alpha_2 CCI_t$. We estimate α_2^* to be 1.4 (t-stat = 2.7).
20. We cannot be too precise about the degree of attenuation. For example, we could also accept a model with an imposed restriction that the log house price to income coefficient is fully offset at peak CCI.
[21] We tested but did not find a significant positive coefficient on deposit interest rates, unlike in Germany (Geiger et al. 2016). This is not surprising since Canadian households are net borrowers.
22. The intuition is that, in comparing steady states, an economy with a lower child/adult ratio will have a broadly similar aggregate household consumption-to-income ratio. But, in the transition to a lower ratio, falls in child-related expenditure will temporarily lower the consumption to income ratio.

We maintain a restriction of 0.07 for the coefficient on the trend-like net assets to income term, aligned with the freely estimated coefficient in Column (3).[23] The marginal propensity to consume (MPC) out of illiquid financial assets is around 0.02 (t stat = 2.9). The housing-wealth coefficient is statistically zero. We tested but found no significant interaction effect between housing wealth and HLI. The evidence thus suggests there is no standard housing-wealth effect or housing collateral channel for Canada.

Charts 9.5a and 9.5b decompose the long-run consumption equation for Column (4) in Table 10.1. The charts depict the estimated coefficient multiplied by each long-run regressor. The charts help to visualize through time the partial-equilibrium contributions of each long-run regressor to the level of log consumption to income. In the 1980s, easier credit access contributes positively, partly offsetting the negative impact of rising down-payment requirements caused by higher log house prices to income. The 1990s rise is mainly explained by higher income expectations, rising financial wealth, lower real interest rates and positive demographic effects. From the late 1990s, there is a heavy drag from indebtedness, weaker income expectations and a rising down-payment requirement caused by rising house prices. These factors are partly offset by easier credit conditions and falling real interest rates.

Chart 9.6 shows the residual of the cointegrating relationship for consumption, which is strongly stationary for models 3 and 4. This reflects our finding that consumption quickly adjusts to long-run fundamentals within about 1½ years. About two-thirds of the quarterly variation is explained by our model, most of which is attributable to the equilibrium correction dynamics. By contrast, the CCI-

23. A likelihood ratio test confirms this is a valid restriction. For Column (3), we also tested the validity of the common coefficient restriction on net liquid assets to income by decomposing the variable into the ratio of liquid assets to income and household debt to income. The coefficient on the former is 0.09 (t-stat = 2.4) and the latter is -0.05 (t-stat = -1.2). A likelihood ratio test confirms no statistically significant difference in system log likelihood by imposing a common coefficient on net liquid assets to income.

exclusive models 1 and 2 show wide and persistent deviation between actual consumption and the fitted long-run solution because of a much longer adjustment horizon of around 4 to 5 years.

The corollary is that the exclusion of controls for shifts in credit access results in poorly-specified long-run consumption models. We explicitly tested six restrictions under which the credit-inclusive model in Equation (3.4), shown in Columns (3) and (4), would reduce to the credit-exclusive model in Equation (3.3), shown in Columns (1) and (2). These are: (i) $\kappa_c = 0$; (ii) $\gamma_1 = \gamma_2 = \acute{\gamma}_{3t} = \bar{\gamma}$; (iii) $\acute{\alpha}_{1t} = \alpha_1$; (iv) $\acute{\psi}_t = \psi$; (v) $\acute{\lambda}_t = \lambda$; and (vi) $\acute{\beta}_{1t} = \beta_1$. We can reject hypotheses (i) and (ii) but not (iii), (iv), (v) (and in fact $\hat{\lambda}$ is statistically zero in CCI-inclusive estimations) and (vi). Likelihood ratio tests confirm that Column (4) is superior to Column (3), and that both models are superior to Columns (1) and (2).

6.2. House price equation results

We find that Canadian house prices have risen along with the evolution of the long-run equilibrium, in particular, the significant decline in mortgage interest rates and the increased availability of mortgage credit. There is little evidence of persistent, unexplained misalignment. Actual house prices in 2015Q1 are close to the sum of the fitted contributions of the long-run explanatory variables (Chart 9.7).

Estimation results are set out in Table 10.2. Columns denoted (2), (3) and (4) correspond to the same system shown for the consumption equation in Table 10.1 (we omit Column [1] for brevity). We restrict the inverse housing demand elasticity (h_2) with respect to price to 1.8 and the income elasticity (h_1) to 1 (discussed below). Column (2) shows the long-run house price equation without controls for credit access. The speed of adjustment is very slow: 3.8 per cent per quarter (t-stat = 1.9). This implies, implausibly, that it takes about 15 years for house prices to adjust to long-run movements in incomes, interest rates and so on. Coefficients on these long-run variables are also poorly determined.

Column (3) shows the CCI-inclusive house price equation where the results are much improved. The speed of adjustment is strongly significant at 10 per cent per quarter (t-stat = 4.4), implying that it takes about 5 years for 90 per cent of a shock to house prices to disappear. This 5-year time frame is much more consistent with Canada's housing finance arrangements. The equation standard error is lower at 0.015 and the adjusted R^2 is now 0.61.[24]

Column (4) shows the CCI-inclusive model plus controls for housing liquidity and demographics. The speed of adjustment and model fit are unchanged relative to Column (3). The combined coefficient on the two interest rate terms is also unchanged at around -0.5. The elasticity with respect to the user cost is around -0.3 (t-stat = -2.2), much less than the strict theoretical prediction of -1 in Equation (3.6) (i.e. an efficient market model where households optimize the rent-own decision in the absence of credit constraints). The elasticity on the nominal borrowing rate is around -0.2 (t-stat = -1.9). This means that a 100 basis point rise in nominal mortgage rates from, say, 2.5 to 3.5 per cent would reduce long-run house prices by about 7 per cent over the subsequent 5 years. If the real user cost increased as well (assuming unchanged inflation, constants and real capital gains expectations at 2.5 per cent), the combined impact would be around -12 per cent over the subsequent 5 years.[25]

24. In Table 10.2, we show the restrictions on h_2 and h_1 being maintained, which reduces the number of free parameters in the system and thereby helps pin down estimates for the consumption equation. Freely estimated, h_2 is 1.6 (t-stat=2.5) and inverting this coefficient gives (minus) the elasticity of housing demand with respect to price of 0.6. These estimates are broadly similar to estimates for other countries. Time-series estimates include -0.5 to -0.6 for the United Kingdom (Meen 1996; Meen and Andrew 1998; Cameron, Muellbauer and Murphy 2006); -0.56 for Australia (Muellbauer and Williams 2011); -0.72 for Germany (Geiger, Muellbauer and Rupprecht 2016); -0.65 for France (Chauvin and Muellbauer 2013). The freely estimated income elasticity (h_1) is 1. The restrictions on h_2 and h_1 are therefore acceptable and in line with the international evidence discussed in Section 3.2. The estimates roughly imply that doubling real incomes, given the housing stock, raises real house prices by a multiple of a little over three.

25. It is difficult to be overly precise about these estimates because we have to make certain assumptions about how a change in mortgage rates translates to a change in the user cost.

Column (4) includes the housing liquidity index as a proxy for the non-price relaxation of credit access for buy-to-let investors. The interpretation is that HELOCs make it easier for investors to find the required deposit to obtain mortgages for their investment properties by withdrawing equity from their existing home(s). Since HELOCs are more expensive than mortgages, they seem to be more of a complement than a substitute for mortgages. Column (4) also includes a small negative impact from the working-age (15–64 years) to total population ratio (the obverse dependency ratio). The interpretation is that households with children or retirees have modestly higher space requirements per person, raising demand for housing. Many retirees live in single-person households and are unable to benefit from scale economies available to couples, especially couples without children. Another long-run variable is the annual average ratio of immigration from Asia to total immigration (lagged two quarters). We find this variable positively and structurally affects house prices, consistent with Ley (2015), Moos and Skaburskis (2010) and Ley and Tutchener (2001).

In the short run (not shown in table), we found significance on two autoregressive terms, outlier and pre-1988 seasonal dummies. The change in the unemployment rate is not significant. We tested housing supply dynamics by adding the lagged four-quarter change in log real net dwelling stock per capita ($\Delta_4 \ln hs_{t-1}$) to the dynamic specification. The coefficient is not significant, indicating that Canadian housing supply dynamics have no significant impact on national house prices in the short run.

Charts 9.8a and 9.8b decompose the long-run solution. There is a drag on real house prices before about 1997 due to declining real income per house. Easier mortgage credit conditions during the 1980s and a rising composition of immigration from Asia contribute positively to house prices. House prices decline and level off after 1990 due to higher user cost and nominal mortgage rates, ongoing declines in incomes per house and tighter credit conditions. Demographics (measured by the obverse dependency ratio) are a slight drag on house prices from about 1994 to 2010. The rise in house prices from 1997 is explained by the cessation of the drag from income per house; lower user cost; lower nominal mortgage rates; growing use of home equity loans (HLI); and further easing in mortgage credit access around 2002 and 2006 that was interrupted by the GFC.

The corollary is that credit-exclusive models do a bad job of explaining the long-run evolution of house prices. It might be tempting to interpret their large and persistent errors (Chart 9.7, Models 1 and 2) as evidence of long-term irrationality in Canadian housing markets. Rather, we show these models are misspecified because they assume that housing demand is not affected by shifts in access to mortgage credit. When controls for these shifts are added, we find well-specified long-run models. House prices mostly adjust to their long-run drivers within about 5 years as mortgage contracts are rolled over. High and rising Canadian house prices since the late-1990s are mostly explained by mortgage credit becoming cheaper and more available.

6.3. Mortgage debt equation results

Results for the long-run mortgage stock equation are set out in Table 10.3. Columns denoted (2), (3) and (4) correspond to the same systems shown for tables 10.1 and 10.2 (for brevity, we omit Column [1]). For Column (2), the absence of any role for interest rates seems counterintuitive and suggestive of misspecification. Column (3) shows the CCI-inclusive mortgage stock equation.[26] The estimated speed of adjustment doubles to 0.12 (t-stat = 6.2), implying that it takes just under 5 years for 90 per cent of a shock to real per capita mortgage debt to disappear. The fact that both house prices and debt adjust to their long-run drivers within about 5 years is consistent with Canada's housing finance arrangements.

26. The freely estimated income elasticity is close to 1, so we impose a unity restriction to reduce the number of free parameters in the system.

Column (4) adds the HLI and performs similarly to Column (3). There is little difference in model fit. The coefficient on the CCI is 0.5 and is highly significant. All else being equal, if CCI has a +1 per cent long-run impact on real house prices, it has approximately half that impact on the long-run mortgage-debt-to-income ratio, given the housing-wealth-to-income ratio. The elasticity with respect to nominal interest rates is -0.11. All else being equal, a 100 basis point (40 per cent) rise in nominal interest rates from current levels (2.5 per cent) would reduce real mortgage debt per capita by about 4 to 5 per cent over the subsequent 5 years. Finally, we add HLI as an additional control for the relaxation of credit access for buy-to-let investors. HLI has a coefficient of 0.9 (t-stat = 1.8). The positive coefficient suggests HELOCs are more of a complement than a substitute for mortgages.[27]

The results in Columns (3) and (4), which include CCI and HLI, are clearly superior to the CCI-exclusive results. Chart 9.9 decomposes the long-run mortgage solution for Column (4). Aside from the early 1980s and early 2000s, mortgage debt/income in Canada has risen steadily over three decades. The rise is explained by significantly lower nominal mortgage rates, especially since the mid-1990s; increased access to mortgage credit (CCI) and HELOC credit (HLI); and rising housing wealth.[28]

6.4. The estimated CCI results

The latent variable (a spline function defined as a linear combination of smoothed step dummies, see Equation 2.2) has a scaled effect in each equation but its slope coefficients are restricted to be identical. The latent variable captures the common long-run variation in consumption, house prices and mortgage

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27. In the short run, two lags of the autoregressive term are significant when the CCI is omitted but are otherwise not significant. They were omitted from the models estimated in columns (3) and (4). Although we retained the delta unemployment rate as a proxy for income uncertainty, its coefficient is not significant.

28. The response of mortgage debt to credit conditions and interest rates, given the housing-wealth-to-income ratio, understate the full response operating indirectly via housing wealth.

debt that is not explained by the economic and demographic regressors.[29] Results are detailed in Table 10.4 and plotted in Chart 9.2. Columns (3b) and (4b) show the more generalized CCI.[30]

The jointly-estimated slope of the latent variable matches the priors well (see Section 4) and is well-correlated with growth in household credit aggregates (Chart 9.3). The latent variable rises strongly during the 1980s, coincident with banks' establishing mortgage loan subsidiaries and disinflation, before retrenching and flattening out coinciding with the early 1990s recession and the implementation of Basel I. The latent variable, together with HLI (which is based on HELOC debt data and is not latent), rises substantially during the 2000s with the strongest increase occurring around the 2006 relaxation of mortgage insurance rules. There is another retrenchment around the 2007–09 crisis, partly offset by counter-crisis policy measures in 2009. There is only a slight retrenchment in the post-crisis years, suggesting that changes to mortgage insurance rules and other regulations have made credit only marginally less easy to obtain.

There are good reasons to interpret the latent variable as a measure of unobserved structural changes in credit access. First, the counterfactual approach, to omit controls for credit, is tantamount to assuming full information shared by lenders and borrowers and unchanging credit markets cleared by interest rates. Yet it is widely accepted that asymmetric information about default risk causes lenders to impose non-price lending constraints (Flemming 1973; Jaffee and Russell 1976; Stiglitz and Weiss 1981, 1992; Stiglitz 1999; Williamson 1986; Miles 1992; Brueckner 1994; Leece 1995; and Meen 1990, 1996).

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29. For comparison purposes, we maintain the same parsimonious specification for columns (3) and (4) and omit the most insignificant dummies. We retained all post-GFC dummies, since these may be of interest to policy-makers. Given clear priors about the GFC, we added an extra dummy for 2007Q3 and lagged the 2008 dummy.
30. Log likelihood tests comparing columns (3) and (3a), and columns (4) and (4a), respectively, confirm that the five restrictions on the parsimonious CCI used in columns (3) and (4) are acceptable.

Second, CCI enters all three equations in a manner predicted by economic theory as set out in Section 3. The slope of CCI is corroborated by the institutional evidence for Canada in Section 4. Our results can be compared with evidence from other countries that use direct (e.g., loan officer survey data) and indirect (e.g., latent variable) methods to control for credit conditions effects.[31] As survey-based measures of household credit access are not available for Canada, we have developed a plausible proxy using a LIVES approach that enables us to estimate statistically-robust long-run models for the dependent variables.

Third, we have shown that omitting CCI and HLI sees a dramatic deterioration in the long-run models. Arguably, instead of CCI, one could try to control for credit access shifts by adding a few dummies to the long-run solutions in individual equations. However, we see this approach as ad hoc and less intuitive than our general-to-specific LIVES method, which exploits the *common* impact of shifting credit conditions on the three dependent variables (in levels). Choosing the "right" dummies would be arbitrary and difficult because regulations and the pace of innovation are continuously evolving.

Fourth, as discussed in the introduction, conventional consumption models rely on a current log income growth term, the coefficient on which represents the (fixed) proportion of "liquidity-constrained" consumers. This simplistic approach is contradicted by the buffer-stock saving models of Deaton (1991) and Caroll (1992, 2001). Moreover, credit constraints are not solely about limits on consumption smoothing: since mortgages dominate household debt, different types of constraints, which are also time-varying, then arise. We have shown that inclusion of CCI and HLI addresses shifts in these constraints and makes the current income growth term in our models empirically redundant.

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31. See Aron et al. (2012) for the United States, the United Kingdom and Japan; Muellbauer and Williams (2011) for Australia; Chauvin and Muellbauer (2013) for France; Geiger, Muellbauer and Rupprecht (2016) for Germany; and Aron and Muellbauer (2013) for South Africa.

Finally, we acknowledge the possibility that the latent variable could be picking up other omitted influences. But this would need to be a common influence on consumption, house prices and debt that is not explained by any of the economic or demographic regressors, and unrelated to identifiable changes in credit market technologies or regulations. Potential examples could be the effect of shifts in unsecured consumer credit access, changes in the distribution of income or wealth, or persistent mis-measurement of one of the included controls. Could CCI be picking up "irrational exuberance"? Possibly, but the user cost term in the house price equation already incorporates extrapolative expectations of recent price appreciation. Moreover, the latent variable does not actually rise very much after the 1990s when house prices are rising strongly (the HLI is rising but that variable is based on HELOC debt data and is not latent). The finding of statistically-robust long-run models suggests the above issues may be minor ones, but could be avenues for future research.

6.5. Robustness checks

We conducted robustness checks for parameter stability, systems estimation methodology, sensitivity to income measurement, and cointegration. Results are discussed in Appendix 2 and available on request. The cointegrated VAR results for the debt equation indicate that housing wealth to income, the mortgage rate and CCI are all endogenous. This seems obvious for housing wealth to income but it is plausible that the mortgage rate and CCI react to household indebtedness also. We find that CCI tends to decline after periods of rising debt/equity, debt/income and debt-servicing ratios. This suggests that while there are long-term exogenous drivers of CCI, such as financial innovations and regulatory changes unrelated to the economic cycle, there is also a countercyclical policy feedback. In other words Canadian regulators tend to tighten (ease) mortgage regulations in response to higher (lower) leverage and debt servicing ratios. The corollary is that while Canada has high-quality lending practices and conservative screening of individual loans, there has been a trend toward easier mortgage credit access over recent decades offset by a policy feedback rule that mitigates excesses from time to time.

7. Conclusion

We propose richer ways to think about some of the key equations governing the household sector "block" within conventional GE models. We tested several theoretical propositions on how real-financial linkages may operate and compare our empirical results for Canada with findings for other countries. We tested models with and without a latent variable. Two models that include the latent variable clearly outperform two models that exclude them. We also tested a non-latent control for the increased availability of home equity loans. We compared the LIVES model results to results from a cointegrating VAR. The estimated latent variable has a profile consistent with prior information about specific changes in Canadian credit market architecture and is well-correlated with growth in credit aggregates. The robust significance of these controls and their role in improving the estimates for the other long-run coefficients offers good evidence of a common unobserved structural influence on Canadian house prices, mortgage debt and consumption.

We interpret our results as indicating that Canada shares similarities with the United States, the United Kingdom, Australia, France, Germany and South Africa in that shifts in credit conditions have real economy effects on consumption as well as on household portfolios. Canada's institutional arrangements appear to have encouraged home equity lending for investment (e.g. rental property purchases and home renovations), contributing to increases in house prices and mortgage debt post 1990s, but not to increases in consumption. Canada differs from the United States, the United Kingdom, Australia and South Africa in that, after controlling for the impact of credit conditions, higher house prices to income appear to have an overall negative impact on consumption. France, Germany, and

Japan share this trait. There is some evidence that Canada's negative house-price-to-income effect was attenuated with easier credit access since the 2000s. The reluctance of Canadian households to borrow to consume housing capital gains could be due to institutional features that encourage households to build housing equity: conservative individual loan screening; full recourse mortgages in most provinces; borrowers being exposed to interest rate risk due to short mortgage renewal periods; and non-deductibility of mortgage interest on personal residences. These are avenues for future research.

Our results also highlight an important feature of Canada's monetary policy transmission mechanism. A decrease in borrowing rates initially raises consumption; however there are also negative, indirect effects on consumption via household balance sheets that take longer to materialize. Consumption adjusts to changes in interest rates within around 1½ to 2 years whereas house prices and debt levels, which in turn dampen consumption, take closer to 5 years. Finally, high and rising real house prices and debt levels in Canada since the late 1990s can be mostly explained by the evolution of the long-run determinants in our model of house prices, particularly cheaper and easier access to mortgage finance. This suggests that the outlook for house prices mostly depends on the future paths of these variables.

8. References

Anderson, E. , A. Inoue and B. Rossi. 2012. "Heterogeneous Consumers and Fiscal Policy Shocks,"2012 Meeting Papers 261, Society for Economic Dynamics.

Ando, A., and F. Modigliani. 1963. "The 'life-cycle' hypothesis of saving: aggregate

implications and tests, " *American Economic Review*, 53(1), 55–84.

Aron, J., J. Duca, J. Muellbauer, K. Murata and A. Murphy. 2012. "Credit, housing collateral and consumption: Evidence from Japan, the U.K. and the U.S." *Review of Income and Wealth* 58(3): 398–423.

Aron, J., and J. Muellbauer. 2013. "Wealth, credit conditions, and consumption: Evidence from South Africa." *Review of Income and Wealth* 59 (S1): S161–S196.

Blinder, A., and A. Deaton. 1985. "The time series consumption function revisited." *Brookings Papers on Economic Activity* (2): 465–621.

Bierens, H., and L. Martins. 2010. “Time-varying cointegration.” *Econometric Theory* 26: 1453–90.

Brayton, F. and P. Tinsley, 1996. “A Guide to FRB/US: A Macroeconomic Model of the United States.” Finance and Economics Discussion Series 1996-42, Federal Reserve Board.

Brueckner, J. 1994. “The demand for mortgage debt: Some basic results.” *Journal of Housing Economics* 3(4): 251–62.

Cameron, G., J. Muellbauer and A. Murphy. 2006. “Was there a British house price bubble? Evidence from a regional panel.” CEPR Discussion Paper No. 5619, London.

Campbell, J. 1987. “Does saving anticipate declining labor income? An alternative test of permanent income hypothesis.” *Econometrica* 55(6): 1249–73.

Campbell, J., and N. G. Mankiw. 1989. “Consumption, income and interest rates: Reinterpreting the time series evidence.” In *NBER Macroeconomics Annual 1989*, 185–246. Edited by O. Blanchard and S. Fischer. Cambridge, MA: MIT Press.

———. 1991. “The response of consumption to income: A cross-country investigation.” *European Economic Review* 35: 723–56.

Carroll, C. D. 1992. "The Buffer-Stock Theory of Saving: Some Macroeconomic Evidence," *Brookings Papers on Economic Activity*, 23(2): 61-156.

Carroll, C. 2001. “A theory of the consumption function, with and without liquidity constraints.” *Journal of Economic Perspectives* 15(3): 23–45.

Chauvin, V., and J. Muellbauer. 2013. “Consumption, household portfolios and the housing market: A flow of funds approach for France.” Paper presented at Bank of France, November.

Canada Mortgage and Housing Corporation (CMHC). 2006. *Canadian Housing Observer*.

Citizenship and Immigration Canada (CIC). 2014. *Evaluation of the Federal Business Immigration Program*. June.

Crawford, A., C. Graham and É. Bordeleau. 2009. “Regulatory constraints on leverage: The Canadian experience.” *Bank of Canada Financial System Review*, Bank of Canada (June): 45–50.

Cooper, D. and K. Dynan. 2016. “Wealth effects and macroeconomic dynamics.” *Journal of Economic Surveys* 30(1): 34-55.

Crawford, A., C. Meh and J. Zhou. 2013. “The residential mortgage market in Canada: A primer.” *Bank of Canada Financial System Review* (December): 53–63.

Deaton, A. 1991. “Saving and liquidity constraints.” *Econometrica* 59 (5): 1221–48.

Debelle, G. 2004. “Household debt and the macroeconomy.” *BIS Quarterly Review*, Bank for International Settlements (March).

Duca, J., and J. Muellbauer. 2013. “Tobin LIVES: Integrating evolving credit market architecture into flow of funds based macro-models.” In *A Flow-of-Funds Perspective on the Financial Crisis*, Vol. 2. Edited by B. Winkler, A. van Riet and P. Bull. Palgrave-Macmillan.

Duca, J., J. Muellbauer and A. Murphy. 2011. "House prices and credit constraints: Making sense of the US experience." *Economic Journal* 121(552): 533–551.

———. 2016. "How financial innovations and accelerators drive booms and busts in U.S. consumption." Forthcoming working paper, Federal Reserve Bank of Dallas.

Ellis, L. 2005. "Disinflation and the dynamics of mortgage debt." BIS Papers 22, Bank for International Settlements.

Flemming, J. 1973. "The consumption function when capital markets are imperfect: The permanent income hypothesis revisited." *Oxford Economic Papers* 25(2): 160–72.

Freedman, C. 1998. "The Canadian banking system." Bank of Canada Technical Report No. 81.

Gali, J., J. Lopez-Salido and J. Valles. 2004. "Rule-of-thumb consumers and the design of interest rate rules." *Journal of Money, Credit and Banking* 36(4): 739–63.

Geiger, F., J. Muellbauer and M. Rupprecht. 2016. "The housing market, household portfolios and the German consumer." Revision of paper for Bundesbank, DFG and IMF conference, "Housing markets and the macroeconomy: Challenges for monetary policy and financial stability," Eltville, June 5–6, 2014.

Gervais, O. and M.A. Gosselin. 2014. Analysing and forecasting the Canadian economy through the LENS model." Bank of Canada Technical Report 102.

Hendry, D. 1997. "The econometrics of macroeconomic forecasting." *Economic Journal* 107(44): 1330–57.

Hendry, D. and G. Mizon. 2014a, "Unpredictability in economic analysis, econometric modelling and forecasting," Journal of Econometrics 182: 186–195.

Hendry, D. and G. Mizon. 2014b., "Why DSGEs fail during crashes." VOX: CEPR's policy portal, 18 June. http://voxeu.org/article/why-standard-macro-models-fail-crises

Jaffee, D.. and T. Russell. 1976. "Imperfect information, uncertainty and credit rationing." *Quarterly Journal of Economics* 90(4): 651–66.

Kaplan, G., G. Violante, and J. Weidner. 2014. "The Wealthy Hand-To-Mouth." *Brookings Papers on Economic Activity*.

Kaplan, G., and G. Violante.2014. "A Model of the Consumption Response to Fiscal Stimulus Payments". *Econometrica* 82(4): 1199-1239.

Kaplan, G, B Moll and G L Violante. 2016. "Monetary Policy According to HANK", Working Papers No. 1602, Council on Economic Policies.

Kearl, J. R. 1979. "Inflation, mortgages, and housing." *Journal of Political Economy* 87: 1115–38.

Kimball, M. 1990. "Precautionary saving in the small and in the large." *Econometrica* 58(1): 53–73.

Klyuev, V. 2008. "Show me the money: Access to finance for small borrowers in Canada." IMF Working Paper 08/22.

Kohn, D. 2009. "Monetary policy and asset prices revisited." *Cato Journal* 29(1): 31–44.

Leece, D. 1995. “Rationing, mortgage demand and the impact of financial deregulation.” *Oxford Bulletin of Economics and Statistics* 57(1): 43–61.

Ley, D. 2015. “Global China and the making of Vancouver’s residential property market.” International Journal of Housing Policy, 1-20.

Ley, D., and J. Tutchener. 2001. “Immigration, globalisation and house prices in Canada’s gateway cities.” *Housing Studies* 16(2): 199–223.

Meen, G. 1990. “The removal of mortgage market constraints and the implications for econometric modelling of UK house prices.” *Oxford Bulletin of Economics and Statistics* 52(1): 1–23.

———. 1996. “Ten propositions in UK housing macroeconomics: An overview of the 1980s and early 1990s.” *Urban Studies* 33(3): 425–44.

———. 2000. “Housing cycles and efficiency.” *Scottish Journal of Political Economy* 47(2): 114–40.

———. 2002. “The time-series behavior of house prices: A transatlantic divide?” *Journal of Housing Economics* 11(1): 1–23.

Meen, G., and M. Andrew. 1998. “On the aggregate housing market implications of labour market change.” *Scottish Journal of Political Economy* 45(4): 393–419.

Mian, A. and A. Sufi. 2014. *House of Debt,* University of Chicago Press: Chicago.

Miles, D. 1992. “Housing markets, consumption and financial liberalisation in the major economies.” *European Economic Review* 36: 1093–1127.

Moos, M. and A. Skaburskis. 2010. “The globalization of urban housing markets: Immigration and changing housing demand in Vancouver.” *Urban Geography* 31(6), 724-749.

Muellbauer, J. 2010. “Household decisions, credit markets and the macroeconomy: Implications for the design of central bank models.” BIS Working Paper 306.

———. 2008. “Housing, credit and consumer expenditure.” In *Proceedings, Economic Policy Symposium, Jackson Hole*, Federal Reserve Bank of Kansas City, 267–334.

Muellbauer, J., and D. Williams. 2011. “Credit conditions and the real economy: The elephant in the room.” Centre for Economic Policy Research, CEPR Working Paper 8386, London.

Otsuka, M. 2004. “Essays on household portfolio and current account dynamics.” PhD. Thesis, John Hopkins University.

Park, J., and S. Hahn. 1999. “Cointegrating regressions with time-varying coefficients.” *Econometric Theory* 15(5): 664–703.

Parkinson, S., B. Searle, S. Smith, A. Stoakes and G. Wood. 2009. “Mortgage equity withdrawal in Australia and Britain: Towards a wealth-fare state?” *European Journal of Housing Policy* 9(4): 365–89.

Poloz, S. 2015. “Integrating financial stability into monetary policy.” Speech to the National Association for Business Economics, Washington D.C., 12 October.

Reis, R. 2006. "Inattentive consumers." *Journal of Monetary Economics* 53(8): 1761–1800.

Schembri, L. 2014. “Housing finance in Canada: Looking back to move forward.” *National Institute Economic Review* 230: R45–57.

Stiglitz, J. 1999. “Interest rates, risk and imperfect markets.” *Oxford Review of Economic Policy* 15(2): 59–76.

Stiglitz, J., and A. Weiss. 1981. “Credit rationing in markets with imperfect information.” *American Economic Review* 71(3): 393–410.

——— . 1992. “Asymmetric information in credit markets and its implications for macro-economies.” *Oxford Economic Papers* 44(4): 694–724.

Stock, J., and M. Watson. 1988. “Testing for common trends.” *Journal of the American Statistical Association* 83(404): 1097–1107.

Williamson, S. 1986. “Costly monitoring, financial intermediation and equilibrium credit rationing.” J*ournal of Monetary Economics* 18(2): 159–79.

9. Charts

Chart 9.1: Log levels of the dependent variables

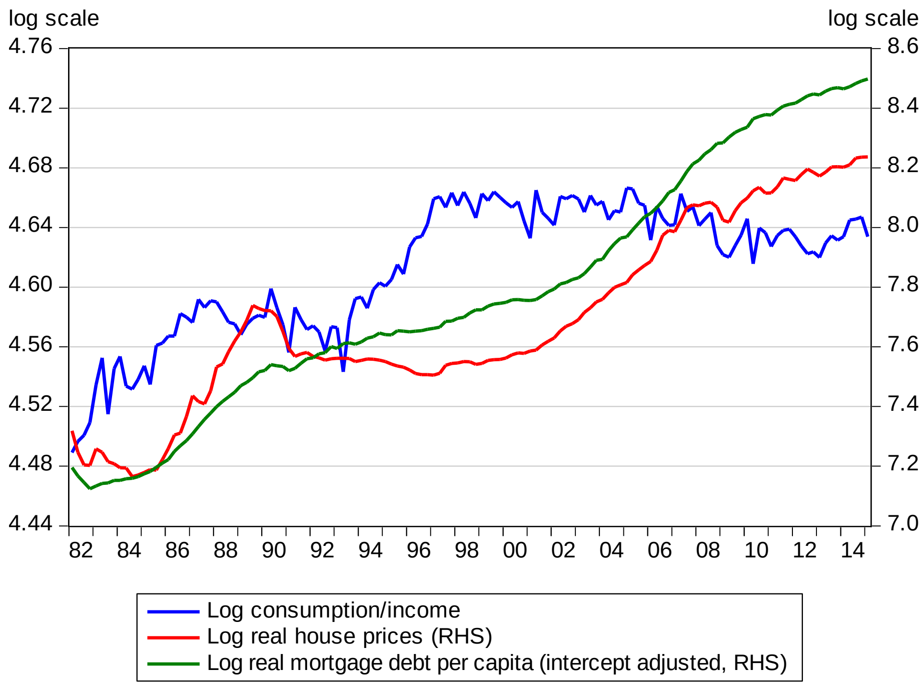


Chart 9.2: Credit conditions effects

(long-run contribution to house price and mortgage debt levels in a partial equilibrium)

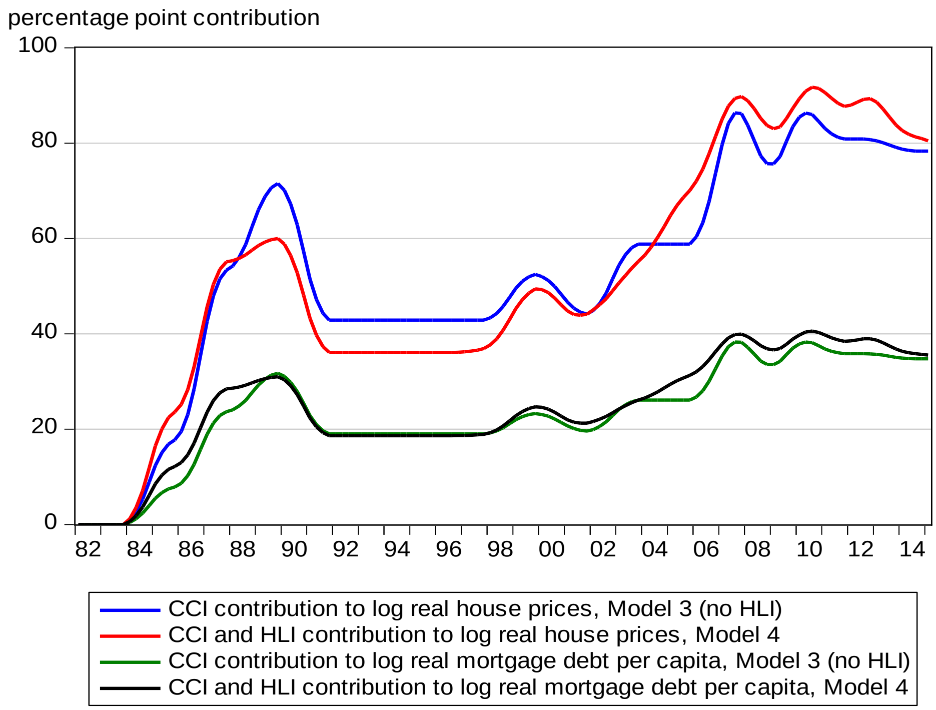


Chart 9.3: The estimated CCI is well-correlated with growth in household credit aggregates

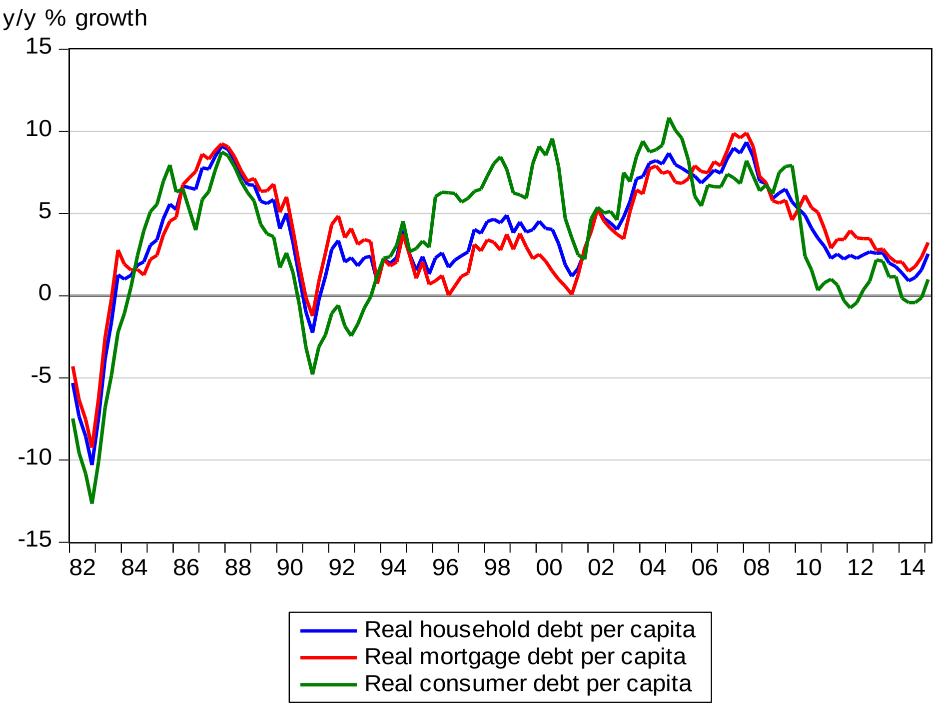
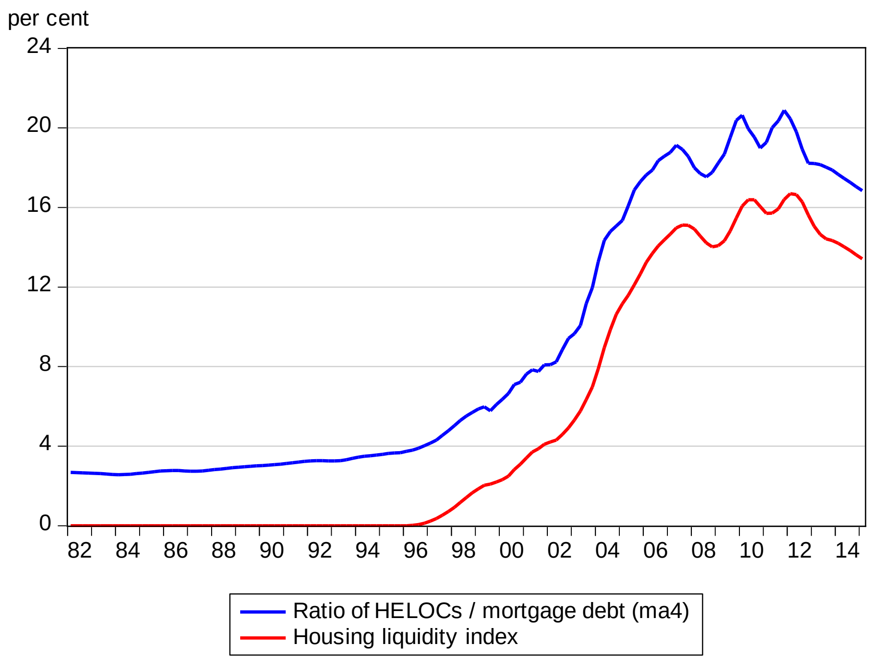
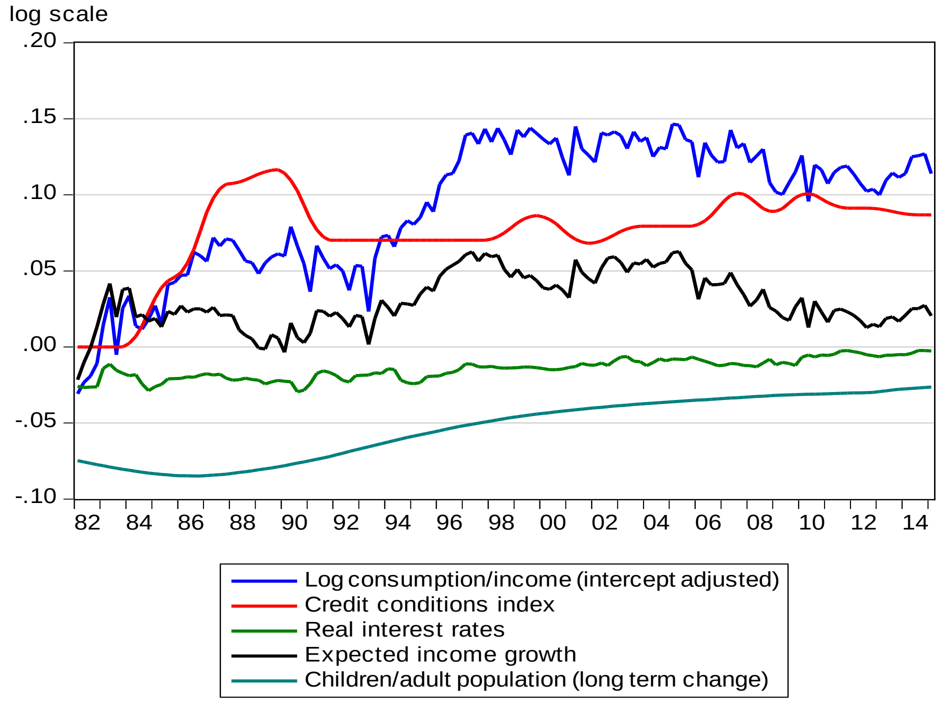


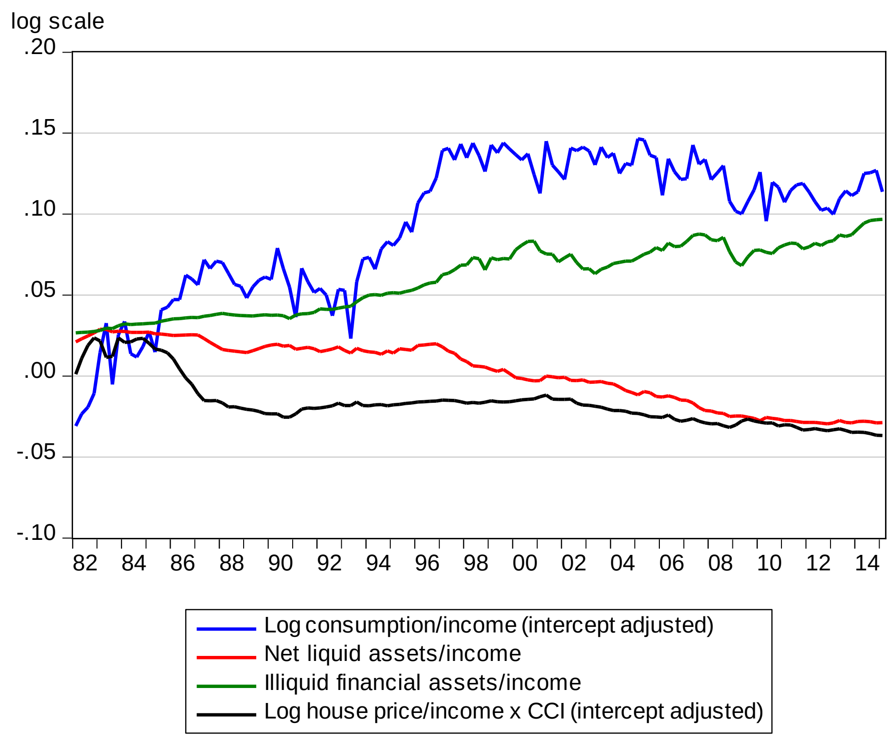
Chart 9.4: Housing liquidity index (HLI)



Note: The housing liquidity index is defined as the four-quarter moving average of the ratio of home-equity lines of credit (HELOC) debt to mortgage debt, set to 0 before 1996Q1 and rebased by subtracting the 1996Q1 value from subsequent values.

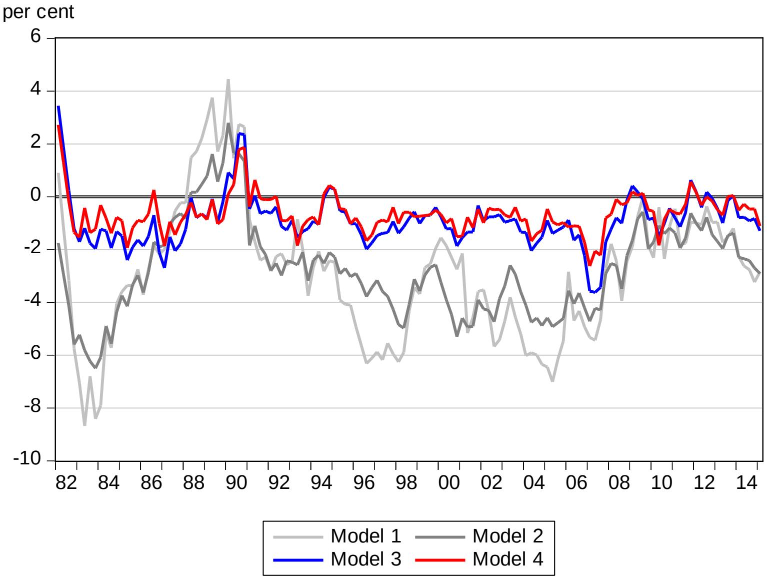
Charts 9.5a and 9.5b: Decomposing the long-run solution for log consumption/income





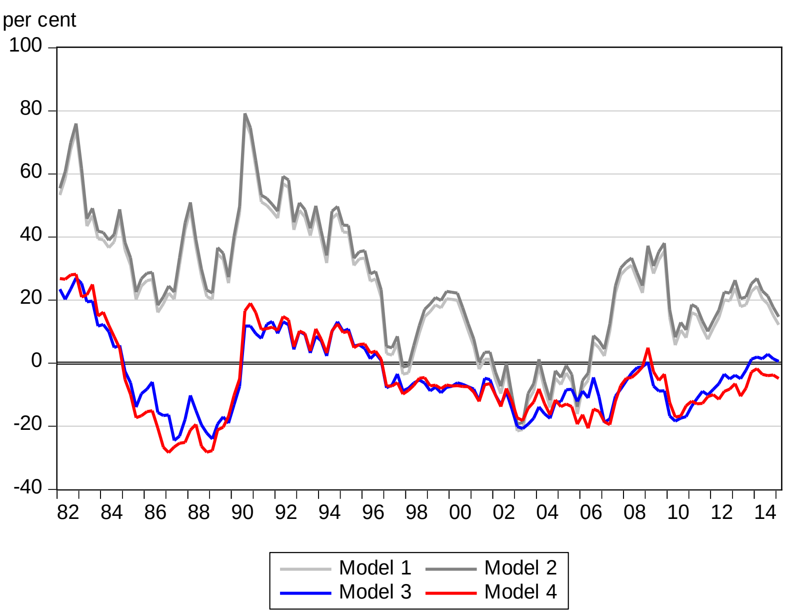
Note: The charts above show the explanatory variables multiplied by their estimated long-run coefficient (Column 4, Table 10.1). Between two points in time, all else being equal, multiplying the change in the y-axis value for a variable by 100 gives its approximate percentage point long-run contribution to the change in the dependent variable in a partial equilibrium.

Chart 9.6: Residual of the long-run solution for log real consumption per capita



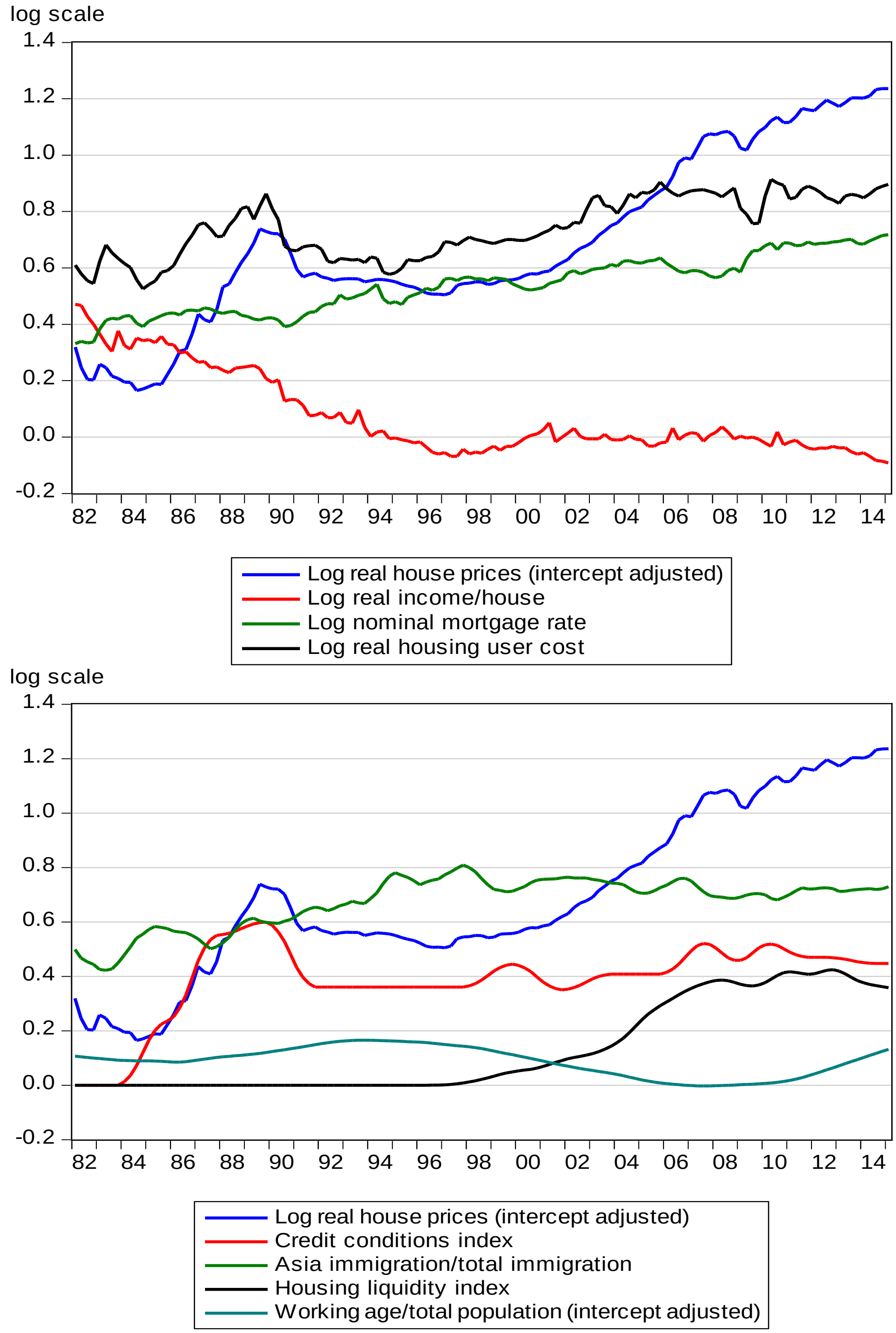
Note: The cointegration residual is defined here as actual log real consumption per capita (at time t) minus the fitted long-run equation set out in columns (1) to (4) in tables 10.1 to 10.4.

Chart 9.7: Residual of the long-run solution for log real house prices



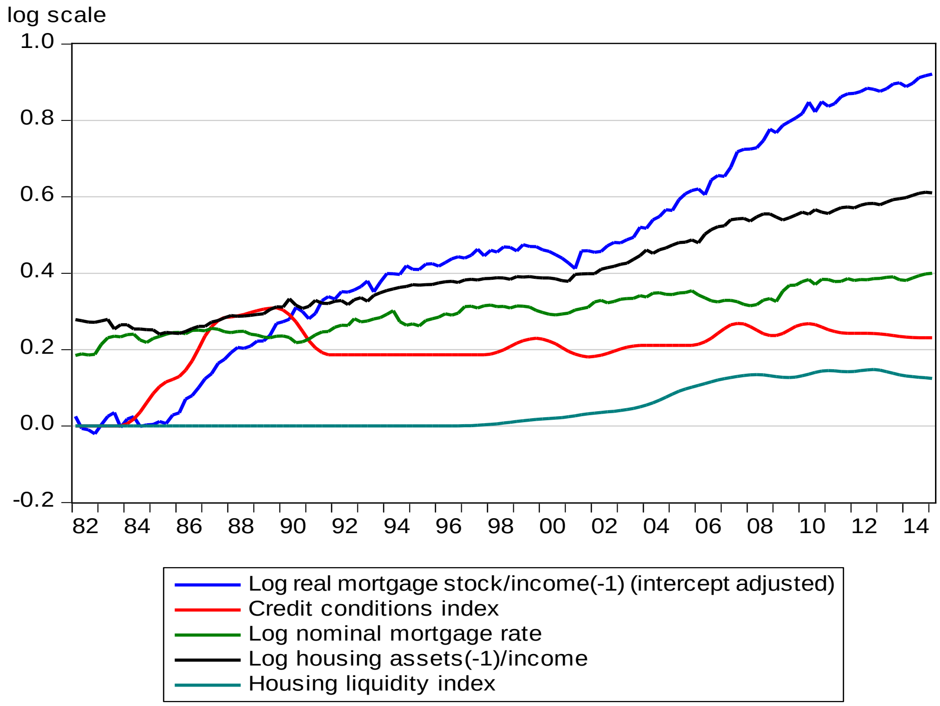
Note: The cointegration residual is defined as actual log real house prices (at time t) minus the fitted long-run equation set out in columns (1) to (4) in tables 10.1 to 10.4. Models (3) and (4), which include controls for credit, suggest house prices in 2015Q1 are around 0 to 5 per cent below their model-fitted value. For models (1) and (2), which exclude credit controls, house prices in 2015Q1 are around 12 to 15 per cent above their model-fitted value.

Charts 9.8a and 9.8b: Decomposing the long-run solution for log real house prices



Note: The charts above show the house price equation explanatory variables multiplied by their estimated long-run coefficient (Column 4, Table 10.2). Between two points in time, all else being equal, multiplying the change in the y-axis value for a variable by 100 gives its approximate percentage point long-run contribution to the change in the dependent variable in a partial equilibrium.

Chart 9.9: Decomposing the long-run solution for log real mortgage stock per capita



Note: The charts above show the mortgage stock equation explanatory variables multiplied by their estimated long-run coefficient (Column 4, Table 10.3). The dependent variable is re-parameterized to incorporate the unity constraint on lagged log income. Between two points in time, all else being equal, multiplying the change in the y-axis value for a variable by 100 gives its approximate percentage point long-run contribution to the change in the dependent variable in a partial equilibrium.

10. Tables

Table 10.1: Estimates of the long-run solution for Canadian consumption

Dependent variable = $\Delta \ln c_t$ Sample: 1982Q1–2015Q1 Sample size: 133 quarterly obs		(1) No credit access effects	(2) No credit access effects	(3) Simple credit access effects	(4) Credit access, HLI & demography
Long-run coefficients for $\ln(c/y)_t$	*Symbol*				
Speed of adjustment	φ_c	0.111 (3.5)	0.136 (4.2)	0.338 (5.7)	0.431 (6.4)
Constant	α_0	4.471 (197.1)	4.551 (52.0)	6.502 (7.8)	7.095 (5.0)
Credit conditions index, CCI_t	κ_c	-	-	0.156 (3.4)	0.194 (3.6)
Real interest rate, r_{t-1}	α_1	-	0.087 (0.2)	-0.248 (-1.4)	-0.286 (-1.9)
Forecast future income growth, $E_t \ln(y_t^p/y_t)$	ψ	1.699 (4.8)	1.187 (3.6)	0.845 (5.8)	0.695 (6.5)
Net total household wealth to income, A_{t-1}/y_t	$\gamma_1 = \gamma_2 = \gamma_3$	0.019 (3.3)	-	-	-
Net liquid wealth to income, NLA_{t-1}/y_t	γ_1	-	-0.038 (-0.5)	0.068 (2.1)	0.07
Illiquid financial wealth to income, IFA_{t-1}/y_t	γ_2	-	0.039 (2.4)	0.035 (3.7)	0.024 (2.9)
Housing wealth to income, HA_{t-1}/y_t	γ_3	-	-0.026 (-1.0)	-	-
Log real house prices to income, $\ln(hp/y)_{t-1}$	α_2	-	-	-0.118 (-2.5)	-0.147 (-1.8)
CCI x log real house prices to income, $CCI_t\,(\ln(hp/y)_{t-1} - mean(.))$	α_2^*	-	-	-	1.409 (2.7)
Long-term change in child/adult population, $child/adpop - ma120(.)$	α_3	-	-	-	0.457 (2.1)
Diagnostics					
Equation standard error		0.0053	0.0049	0.0042	0.0040
Adjusted R^2		0.356	0.451	0.606	0.646
DW		1.68	1.94	2.07	2.11
System log likelihood		1367.46	1381.12	1444.56	1453.13

The table shows coefficients (t-ratios). The estimated generalized credit-augmented consumption function is

$$\Delta \ln c_t = \varphi_c(\alpha_0 + \kappa_c CCI_t + \alpha_1 r_{t-1} + \psi E_t \ln(y_t^p/y_t) + \gamma_1 NLA_{t-1}/y_t + \gamma_2 IFA_{t-1}/y_t + \gamma_3\, HA_{t-1}/y_t + \gamma_3^* HLI_{t-1}(HA_{t-1}/y_t - mean(.))$$
$$+ \alpha_2 \ln(hp/y)_{t-1} - \alpha_2^* \alpha_2 CCI_t(\ln(hp/y)_{t-1} - mean\,(.)) + \ln y_t/c_{t-1}) + \lambda \Delta \ln y_t + \beta_1 \Delta \ln ue + \beta_2 \Delta \ln c_{t-1} + \beta_3(I1990q4 - I1991q1) + \varepsilon_t$$

In Column (1), $A_{t-1}/y_t \equiv NLA_{t-1}/y_t + IFA_{t-1}/y_t + HA_{t-1}/y_t$, which imposes the restriction $\gamma = \gamma_1 = \gamma_2 = \gamma_3$. In columns (3) and (4), γ_3 is insignificant and is set to zero. Column (4) adds demography. For parsimony, in Column (4), the insignificant parameter λ is set to zero. For all equations, outliers are captured by $I1990Q4$ and $I1991Q1$, which are impulse dummies defined as 1 in that quarter and 0 otherwise. Note that Column (4), with λ relaxed, simplifies to Column (1) under the testable restrictions: $\gamma = \gamma_1 = \gamma_2 = \gamma_3$; $\kappa_c = \gamma_3^* = \alpha_2 = \alpha_2^* = \beta_1 = \beta_2 = 0$.

Table 10.2: Estimates of the long-run solution for Canadian house prices

Dependent variable = $\Delta \ln hp_t$ Sample: 1982Q1–2015Q1 Sample size: 133 quarterly obs		(2) No credit access effects	(3) Simple credit access effects	(4) Credit access, HLI & demography
Long-run coefficients for $\ln hp_t$	*Symbol*			
Speed of adjustment	φ_h	0.038 (1.9)	0.104 (4.4)	0.105 (3.1)
Constant	h_0	15.225 (22.3)	17.327 (73.6)	23.296 (4.3)
Credit conditions index, CCI_t	κ_h	-	1.0	1.0
Inverse own-price elasticity, $\ln hs_{t-1}$	h_2	1.8	1.8	1.8
Log real income per capita, $\ln y_{t-1}$	h_1	1.0	1.0	1.0
Log real user cost, $\ln ucc_{t-1}$	h_3	-0.703 (-1.9)	-0.220 (-1.7)	-0.314 (-2.2)
Log nominal mortgage rate, $\ln i_{t-1}$	h_4	-0.397 (-1.6)	-0.282 (-3.0)	-0.200 (-1.9)
Share of immigration from Asia, $aimmig(ma4)_{t-2}$	h_5	2.540 (2.3)	0.948 (3.0)	1.213 (3.6)
Housing liquidity index, HLI_{t-1}	h_6	-	-	2.572 (2.3)
Working-age population (15–64yrs) / total population	h_7	-	-	-8.928 (-1.1)
Diagnostics				
Equation standard error		0.0155	0.0145	0.0146
Adjusted R^2		0.555	0.610	0.607
DW		1.83	2.01	2.02
System log likelihood		1381.12	1444.56	1453.13

The table shows coefficients (t-ratios). The generalized credit-channel augmented house price equation estimated in Column (4) is

$$\Delta \ln hp_t = \varphi_h(h_0 + \kappa_h CCI_t + h_2(h_1 \ln y_{t-1} - \ln hs_{t-1}) + h_3 \ln ucc_{t-1} + h_4 \ln i_{t-1} + h_5 aimmig(ma4)_{t-2} + h_6 HLI_{t-1} - \ln hp_{t-1})$$
$$+ h_7 \Delta \ln hp_{t-1} + h_8 \Delta \ln hp_{t-2} + h_9 \Delta_2 \ln ue_{t-1} + h_{10}(I1983q1 - I1982q4) + h_{11}(I1987q2 - I1987q3) + h_{12} Q4(pre1988) + \mu_t.$$

Outliers are captured by $I1982Q1$, $I1983Q1$, $I1987Q2$, and $I1983Q1$, which are impulse dummies defined as 1 in that quarter and 0 otherwise. $Q4(pre1988)$ is a seasonal dummy that equals 1 in quarter 4 in 1982 through 1987, and 0 in quarter 4 from 1988 onward.

Table 10.3: Estimates of the long-run solution for Canadian mortgage debt

Dependent variable = $\Delta \ln nms_t$ Sample: 1982Q1–2015Q1 Sample size: 133 quarterly obs		(2) No credit access effects	(3) Simple credit access effects	(4) Credit access, HLI & demography
Long-run coefficients for $\ln ms_t$	*Symbol*			
Speed of adjustment	φ_m	0.059 (4.0)	0.122 (6.2)	0.104 (6.5)
Constant	m_0	0.393 (2.9)	0.178 (2.8)	0.245 (2.6)
Credit conditions index, CCI_t	κ_m	-	0.443 (6.9)	0.516 (4.6)
Log real income per capita, $\ln y_{t-1}$	m_1	1.0	1.0	1.0
Log nominal mortgage rate, $\ln i_{t-1}$	m_2	-0.027 (-0.3)	-0.154 (-5.5)	-0.111 (-2.5)
Log housing wealth to income, $\ln HA_{t-1}/y_t$	m_3	0.890 (5.5)	0.4	0.462 (2.5)
Housing liquidity index, HLI_{t-1}	m_4	-	-	0.894 (1.8)
Diagnostics				
Equation standard error		0.0074	0.0063	0.0063
Adjusted R^2		0.377	0.549	0.556
DW		1.80	1.97	2.12
System log likelihood		1381.12	1444.56	1453.13

The table shows coefficients (t-ratios). The generalized credit-augmented mortgage stock equation estimated in Column (4) is

$$\Delta \ln ms_t = \varphi_m(m_0 + \kappa_m CCI_t + m_1 \ln y_{t-1} + m_2 \ln i_{t-1} + m_3 \ln(HA_{t-1}/y_t) + m_4 HLI_{t-1} - \ln ms_{t-1})$$
$$+ m_5 \Delta \ln p_t + m_6 \Delta \ln ms_{t-1} + m_7 \Delta \ln ms_{t-2} + m_7 \Delta \ln ue + v_t.$$

Note that m_5 is not statistically different from -1 so we re-parameterize the dependent variable as delta log nominal mortgage debt per capita: $\Delta \ln nms_t = \Delta \ln ms_t + 1.\Delta \ln p_t$. For columns (3) and (4), m_6 and m_7 are not significant and are set to zero.

Table 10.4: Estimates of the long-run parameters for the credit conditions index (CCI)

Smoothed step dummy (SSD) *Sample: 1982Q1 – 2015Q1*	*Sign prior (see Section 4)*	(3) Simple credit access effects	(3a) Unrestricted CCI	(5) Credit access, HLI & demography	(5a) Unrestricted CCI
SSD 1982Q1	Easing (+)	-	0.052 (0.4)	-	0.096 (0.7)
SSD 1984Q1	Easing (+)	0.177 (3.0)	0.142 (2.5)	0.236 (2.9)	0.226 (2.7)
SSD 1986Q1	Easing (+)	0.356 (4.6)	0.362 (5.1)	0.315 (3.8)	0.297 (3.7)
SSD 1988Q1	Easing (+)	0.182 (2.6)	0.183 (2.3)	0.050 (0.8)	0.021 (0.3)
SSD 1990Q1	Tightening (-)	-0.287 (-3.6)	-0.262 (-3.2)	-0.239 (-3.3)	-0.229 (-3.3)
SSD 1992Q1	Tightening (-)	-	0.003 (0.1)	-	-0.034 (-0.7)
SSD 1994Q1	No prior	-	0.005 (0.1)	-	0.005 (0.1)
SSD 1996Q1	No prior	-	-0.005 (-0.1)	-	0.022 (0.6)
SSD 1998Q1	No prior	0.096 (1.6)	0.086 (1.4)	0.084 (1.5)	0.064 (1.2)
SSD 2000Q1	Easing (+)	-0.083 (-1.5)	-0.071 (-1.4)	-0.095 (-2.0)	-0.098 (-2.1)
SSD 2002Q1	Easing (+)	0.147 (3.1)	0.102 (1.9)	0.058 (1.7)	0.050 (1.2)
SSD 2004Q1	Easing (+)	-	0.090 (1.1)	-	0.003 (0.1)
SSD 2006Q1	Easing (+)	0.298 (4.4)	0.274 (3.5)	0.123 (2.3)	0.122 (2.3)
SSD 2007Q3	Tightening (-)	-0.161 (-2.0)	-0.182 (-2.2)	-0.090 (-1.6)	-0.110 (-2.1)
SSD 2008Q4	Tightening (-)	0.157 (1.5)	0.161 (1.4)	0.092 (1.7)	0.098 (1.7)
SSD 2010Q1	Tightening (-)	-0.073 (-0.8)	-0.065 (-0.7)	-0.063 (-1.3)	-0.066 (-1.3)
SSD 2012Q4	Tightening (-)	-0.025 (-0.5)	-0.002 (-0.04)	-0.023 (-0.7)	-0.022 (-0.6)
Consumption eq. CCI coefficient	κ_c	0.156 (3.4)	0.179 (3.3)	0.194 (3.6)	0.190 (3.6)
Mortgage stock eq. CCI coefficient	κ_m	0.443 (6.9)	0.435 (7.0)	0.516 (4.6)	0.526 (4.4)
System log likelihood		1444.56	1447.68	1453.13	1455.29

The table shows coefficients (t-ratios). $CCI = \sum_{s=1} \tau_s SSD_s = \tau_1 SSD1982q1 + \tau_2 SSD1984q1 + \ldots + \tau_{17} SSD2012q4$. *SSD* is an ogive dummy taking the value 0 up until quarter t-1, then 0.05, 0.15, 0.3, 0.5, 0.7, 0.85, 0.95 and remaining at 1 thereafter.