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Credit conditions and the housing price ratio: Evidence from Ireland's boom and bust

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A B S T R A C T

The Great Recession starting in 2007 has refocused attention on the importance of understanding housing market dynamics as contributors to macroeconomic fluctuations. While the sale-to-rent ratio of housing prices is generally regarded as a fundamental barometer of housing market health, the study of its determinants remains in its infancy. This paper examines the housing price ratio in Ireland since 2000, a period including an extreme housing market cycle. Using new data on first-time buyer loan-to-value ratios, a one-step error correction model of the housing price ratio in Ireland is presented for the first time. It finds clear evidence that, alongside user cost, credit conditions were central in determining equilibrium in the housing market. Throughout, and especially earlier in the sample, there is rapid adjustment of the housing price ratio to its implied equilibrium relation. There is evidence that the housing market regime changed during the period, in 2010 and again in 2014/2015. The preferred specifications imply that a ten percentage point increase in the median first-time buyer loan-to-value was associated with a 9% rise in sale prices, holding other factors – including rental prices and the system wide ratio of credit to deposits – constant. In addition to an understanding of the Irish market, the findings contribute to the evidence base for macroprudential policies that focus on mortgage lending and also hint at how housing market history may differ across rising and falling markets in forming expectations of capital gains.

1. Introduction

The OECD housing boom and bust of the 1990s and 2000s has reminded economists of the importance of housing. It is typically the single most important class of consumption good, making up for example 32% of the U.S. urban CPI basket, and is also the most prevalent investment asset, comprising 54% of US household wealth (Luckett, 2001). Unsurprisingly, there is strong evidence of the link between housing and broader economic outcomes, not just for recent economic history (Davis and Heathcote, 2005; Leamer, 2007) but the entire postwar era and even predating the Industrial Revolution (e.g. Eichholtz et al., 2012; Holly and Jones, 1997).

The housing cycle was particularly acute in Ireland. The period from the mid-1990s to 2007 was one of very strong economic growth in Ireland, initially export-led but in later years fuelled by readily available cheap credit and an unprecedented building boom (see, for example, Devitt et al., 2007). From 2007, the economic downturn was severe. Nominal GNP fell from € 163bn in 2007 to € 128bn in 2011, while government finances deteriorated sharply, with a fiscal deficit of 10% of output by 2010. Unemployment rose from below 5% in 2007 to almost 15% by 2011, while large inward migration flows changed to emigration. Central to the dramatic change in Ireland's economic fortunes was the end of a domestic real estate boom, which had seen nominal house prices rise four-fold in the decade to 2007. By late 2012, prices had fallen by more than half.

This paper presents the first model of the ratio of sale prices to rental prices in the Irish housing market (hereafter, "housing price ratio"). The case of Ireland in the 2000s exemplifies the links between housing and other aspects of the economy, including financial stability, the labour market, government finances, and public service provision. Research examining its housing boom and bust, however, remains scarce. This paper builds on existing research for other economies on the housing price ratio, which stresses the importance of credit conditions, as well as the user cost. To do this, it uses error correction methods and quarterly data for the period 2000–2016, including a new series on the typical loan-to-value for Irish first-time buyers. The paper has three main findings. Firstly, it finds that models with measures of credit conditions strongly outperform those without. Secondly, it finds that throughout the period 2000–2016, but particularly in the earlier part of the period, there is rapid adjustment of the housing price ratio to its implied equilibrium relation. Thirdly, by investigating different samples and windows, it is clear that there were changes in the housing market regime during the period: the data suggest changes in 2010 and again in 2014/2015.

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This paper is related to existing work on the housing price ratio. A starting point is Himmelberg et al. (2005), who show the large increase in sale prices for housing in the U.S. in the decade to 2004 (both in absolute terms and relative to rental prices) can be attributed in large part to dramatic falls in the user cost of housing. The user cost reflects interest rates (less any deductibility), expected capital gain, property taxes and maintenance costs (Poterba, 1984). Research by Duca et al. (2016) shows the limitations of relying solely on user cost to explain housing market outcomes: they find that the inclusion of credit conditions, as measured by the average loan-to-value for first-time buyers, notably improves models of the U.S. housing price ratio for the period 1981–2007 (see also their earlier work, Duca et al. (2011)). Compared to models without credit conditions included, the augmented specification gives better model fits, reasonable speeds of adjustment, and stable long-run relationships with sensible and more precisely estimated income and user cost coefficients. The omission of credit conditions may also affect other studies of the housing price ratio, including Ambrose et al. (2013). Their analysis of Amsterdam sale and rental prices from 1650 to 2005 found "persistent and long-lasting deviations between housing market fundamentals and prices" and that such mispricing occurs mainly through sale prices, rather than rental prices.

Existing work on the Irish housing market includes work in the late 1970s and early 1980s, following publication of official housing prices. Such literature (see, for example, Kenneally and McCarthy, 1982; Thom, 1983) typically tried to include some measure of credit constraints, a feature notably absent in the next phase of research, which dates from the late 1990s and 2000s. At this time, there was a concern among policymakers about a potential bubble in housing, reflected in the ratio of sale prices to income and in the housing price ratio. A number of papers examined Irish housing prices but, without the inclusion of credit conditions, they struggled to generate meaningful results (e.g. IMF, 2003; Stevenson, 2003; Roche, 2004; McQuinn, 2004; Murphy, 2005; Rae and van den Noord, 2006). The closest to a theoretically-grounded analysis of the relationship between credit conditions and housing prices is Murphy (2005), who uses an inverted demand error-correction model, including a dummy variable for financial liberalization.

The absence of credit conditions in analyses of Irish housing prices was partially addressed by Addison-Smyth et al. (2009). They present a two-equation system of average mortgage levels and house prices that builds on McQuinn and O'Reilly (2008), which relies on the mortgage repayment (affordability) and the "funding rate", the ratio of the outstanding level of mortgage lending to total domestic deposits. This is found to have considerable power in explaining average mortgage levels over the period and a plot of fundamental house prices including this factor matches price developments more closely than the more restricted model of mortgage levels.

Research on user costs for Irish housing is similarly scarce. Research by Barham (2004) found that the user cost associated with owning housing in the Irish market was negative for large parts of the period from 1976 on, principally due to the favourable tax treatment afforded owner-occupancy. More recently, Browne et al. (2013) updated this analysis, finding that the user cost is dominated by expected capital gain, where this is measured with the annual gain over the last four years. Neither paper attempts to econometrically link user cost with housing prices. Negative user costs are, of course, somewhat problematic as this implies infinite demand for purchasing housing. True user costs may not be negative, however, if measured costs exclude psychological and other costs of purchasing (and moving, if relevant). Nonetheless, if user costs were negative, this highlights the importance of understanding changes in credit constraints.

The principal contribution of this paper lies in its presentation of the first econometric analysis of the housing price ratio in Ireland. In addition, the principal result – that, with credit conditions included, the ratio appears to adjust very swiftly to changes in its determinants – underscores the finding of Duca et al. (2011) about the importance of including some measure of credit supply, in order to accurately model and thus better understand housing market outcomes. Lastly, more minor contributions include the construction of credit conditions series for Ireland during its boom and bust and a theoretical bridge between the inverted demand and price-rent ratio approaches to modeling housing prices.

The rest of this paper is as follows. Section 2 outlines the basic economic theory involved, including a framework for connecting up inverted demand and housing price ratio approaches to modeling housing markets and Section 3 provides details on the data used in this analysis. Section 4 presents the main empirical analysis and results, while Section 5 concludes.

2. Theory

Theoretically, demand for a good depends on its prices, the income of consumers and other demand shifters. Applied to housing, suppose that in any given period t, the quantity of housing demanded, h_t, can be approximated linearly by:

$$ln(h_t) = -\alpha ln(hp_t) + \beta ln(y_t) + z_t \tag{1}$$

where hp_t refers to the real housing price, y_t to (real) household income and z_t to demand shifters, as discussed below. As the supply of housing is fixed in the short run, the demand function can be inverted, giving:

$$ln(hp_t) = (\beta ln(y_t) - ln(h_t) + z_t)/\alpha \tag{2}$$

Where the income elasticity of demand, β, is one, this simplifies further, with house prices being determined by the log income per house ($\frac{y}{h}$) and other demand shifters, z. This applies to the housing price for both sale and rental properties. Demand shifters unique to sale properties (denoted z_t^S) including user costs (described in more detail below) and credit conditions (as discussed earlier), as well as demographics, which would also affect rental prices. Demand shifters that affect both market and implicit rents can be denoted by z_t^R.

Where income, housing supply and demographics affect both sale and rental prices, this implies that dividing through by rental prices leaves, in addition to a constant, z_t^S, asset factors that affect sale properties, in particular user cost and credit conditions. This logic connects the theory outlined above, which corresponds to the inverted demand approach for modeling housing prices, to the alternative, modeling the housing price ratio, outlined below.

This ratio is related to the concept of financial arbitrage (Poterba, 1984). In an equilibrating market, sale prices will reflect the discounted future stream of rental prices: $hp_t = rent_t/\rho_t$, where ρ_t represents a discount rate. Where discount rates match interest rates r_t, and where housing is subject to costs of depreciation and maintenance (δ_t), costs of transaction and taxation (τ_t), and expected capital gains (κ_t), this means that the housing price ratio in period t (hpr_t) depends on the user cost in that period[1]:

$$hpr_t = 1/(r_t + \delta_t + \tau_t - \kappa_t) \tag{3}$$

In log formulation, and allowing for flexibility in relation to the relative importance with which the various factors affect the ratio:

$$ln(hpr_t) = \beta_0 + \beta_1 r_t + \beta_2 \delta_t + \beta_3 \tau_t + \beta_4 \kappa_t \tag{4}$$

where the expectation is that β_1, β_2 and β_3 are negative and β_4 is positive (greater expected capital gains push up house prices). Tax relief on mortgage interest will also affect the net cost of capital but can be accounted for using the correct measure of r_t.

To this classic specification of the housing price relation can be

[1] This is typically thought of in annual terms. For example, market participants may use a rule such as: "what multiple of annual rent is this property worth?" Equivalently, one could consider the ratio of rents to house prices as being the percentage dividend on housing as an asset.

added at least three further factors that may be relevant. As outlined in Duca et al. (2011) and other related work, credit conditions affect the equilibrium ratio of prices to rents. This is the focus here. In addition, a risk premium term, π_t, should be included in user cost. This is not measured, however, and has thus so far defied inclusion in any empirical setting, a challenge not overcome here. Lastly, Kim (2008) outlines a case where, if a house provides a different level of rental service to an owner-occupier than to a tenant, and houses are rented out reflecting this "rental efficiency", then the ratio of prices to rents will be positively related to rates of home ownership, θ_t. As high-frequency data on changes in tenure are not available, this factor is similarly left for future work. Empirically, where uc_t refers to the user cost term described above, the log form allows the estimation of the long-run relationship between the various factors:

$$ln(hpr)_t = \beta_0 + \beta_1 r_t + \beta_2 \delta_t + \beta_3 \tau_t + \beta_4 \kappa_t + \beta_5 CCI_t(+\beta_6 \theta_t + \beta_7 \pi_t) + \epsilon_t \quad (5)$$

where *CCI* is specified such that an increase reflects an easing of credit conditions; thus the expectation is that β_4, β_5 and β_6 are positive, with all other coefficients (apart from the intercept) negative.

3. Data

This section first presents the housing price series, both sale and rental, and their combination to form the housing price ratio over the period 2000–2016, before presenting the series underpinning the various regressors of interest, in particular user cost and credit conditions.

3.1. Sale prices

Sale price data over time come from two main sources. For the period from 2005Q1 on, the official Central Statistics Office (CSO) residential property prices index (RPPI) is used. For the period prior to 2005Q1, the quasi-official ESRI index, based on mortgages issued by Permanent TSB, is used. Both use hedonic price methods (Duffy, 2004; O'Hanlon, 2011). The figures given are for quarterly averages. To convert housing prices into a level from an index, the national average housing price is taken from the 2016 in Review report by listings website daft.ie.[2] This average is based on 2011 Census weights for all households.

As an alternate source, the daft.ie index, which covers the period from 2006Q1 on, is used. Its index as published in 2016 was based on constant weights for the entire period; the CSO RPPI is based on market transaction weights, which may introduce cyclicality into trends over time, in particular when combined with rental prices. As the publication of the daft.ie series only dates from 2006Q1, quarter-on-quarter changes in the ESRI index are used to extend the series back to 2000. Both the CSO-RPPI and daft.ie indices above are at national level. Given the potential for different market trends by region over the period covered, and indeed the differences in tenure by location, Dublin-specific series are also constructed, using the same two sources. This gives four series for sale prices from 2000 on, covering both national (black) and Dublin (grey), using CSO (solid) and daft.ie (dashed lines).

These are shown in log levels and changes in Figs. 1 and 2. What is striking is the overall similarity in trends, regardless of the source and scope. As can be seen, the four series are highly correlated, even when the pre-2005 period is excluded (all series use ESRI figures for the earlier years). In log-levels, the average correlation 2006–2016 is 97%, with the least correlated pair being the daft.ie National and CSO Dublin series (92%). In log-changes, again the correlation is typically very strong, with a pairwise average of 86%. (The least correlated series again are the daft.ie National and CSO Dublin series (76%), with changes in the two CSO series 95% correlated over the period.) What is obvious from the Dublin figures is that the CSO series is more volatile than its daft.ie counterpart, in terms of quarterly changes. The standard deviation of changes over the period 2005–2016 is 0.047 for the CSO Dublin series, almost one quarter higher than the figure for the daft.ie Dublin series (0.039). While the CSO National series is also more volatile than the daft.ie National series, the difference is smaller (0.038 compared to 0.034).

3.2. Rental prices

Rental price data over time come from the CSOs Consumer Price Index (CPI) sub-component, 'Private Rents'. This is available at quarterly frequency throughout the period and selection of base year has only trivial effects on the implied changes per period. However, the underlying methodology producing the CPI index is unclear. Methodological notes accompanying the CPI state only that data for the sub-component, which makes up 4.4% of the 2011 basket, is from one of 126 "special inquiries conducted", where inquiries are "made by post, telephone, e-mail along with internet price collection".

For that reason, daft.ie is again used as an alternate source. This source uses hedonic methods equivalent to the sale price report, including the use of Census weights to aggregate from regional markets to a national average monthly rent. The full daft.ie rental price index extends back to January 2002, with CPI quarterly changes used for the quarters prior to this. In both cases, the indices were set to levels using the 2016Q4 national average monthly rent, from the daft.ie Rental Report, which is estimated using the same procedure as the average listing price. This choice of weighting means that issues of comparability are not a first-order concern in this dataset. Clearly, there are likely to be some unobserved differences between sale and rental homes but these differences, while they may have a level effect, are not time-varying based on the consistent mix-adjusted and weighted composition of the daft.ie indices.

Again, it may be possible that trends in Dublin are different to those elsewhere in the country. For that reason, a daft.ie series is calculated for Dublin, based on published reports from 2006. For the period from 2000 to 2006, a sample of individual rental listings from daft.ie (which extend back to 2002Q1) and from the Evening Herald newspaper (for the 1999Q4-2002Q1) was compiled. A hedonic regression was run on this dataset, giving a Dublin-specific rental index for the entire period 2000–2016. Figs. 3 and 4 show the log levels and changes in the three rental series (CSO National, daft.ie National and daft.ie Dublin) for the period analysed. Again, as with the sale price series, the three rental series are very highly correlated. The daft.ie Dublin series appears to be the most volatile of the three, with a standard deviation of 0.029, compared to 0.024 for the other two series. It is worth noting, though, that all three standard deviations for rental series are below those for sale series.

3.3. Housing price ratios

The various sale and rental price data series give four series for the ratio of sale prices to rental prices. The first two are national: CSO RPPI relative to CSO CPI rents, and daft.ie listed sale prices to rents. The second pair relate to the Dublin market only, using either CSO or daft.ie sale prices for the capital from the mid-2000s on (and ESRI prices prior to this). In both cases, the rental series used is the Dublin-specific one, based on daft.ie listings to 2002 and extended back by Evening Herald listings.

Figs. 5 and 6 show the raw sale price to rent price ratios, in log-levels (expressed as the log of the sale price to the annual rental). Again, the four series are highly correlated in levels, with an average pairwise correlation of 98%. The volatility of quarterly changes, however, means that the correlations of changes in the sale-rent price ratio are lower but still strongly positive (an average of 65%). Given the consistent

[2] daft.ie is Ireland's largest property listings website and claims to cover more than 95% of both sales and rental listings in the country.

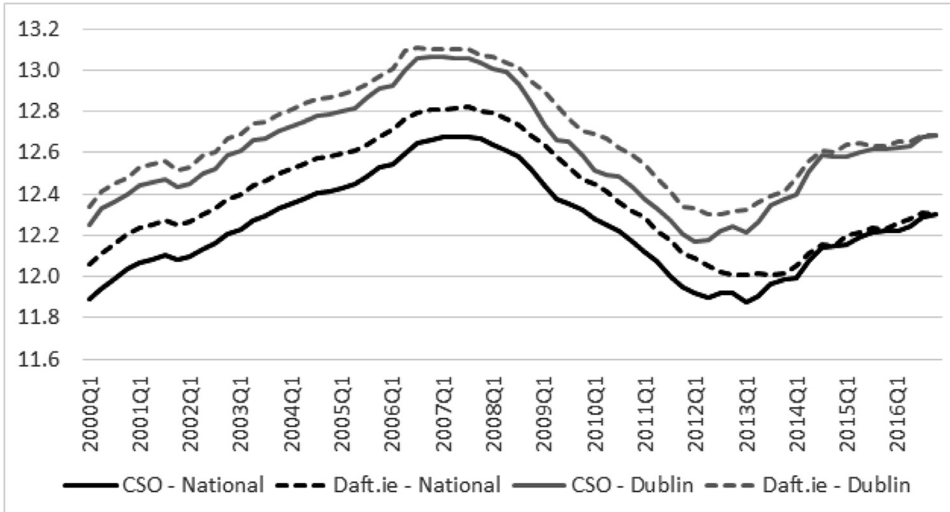

Fig. 1. Sale prices for Irish housing (logs), by source and scope.

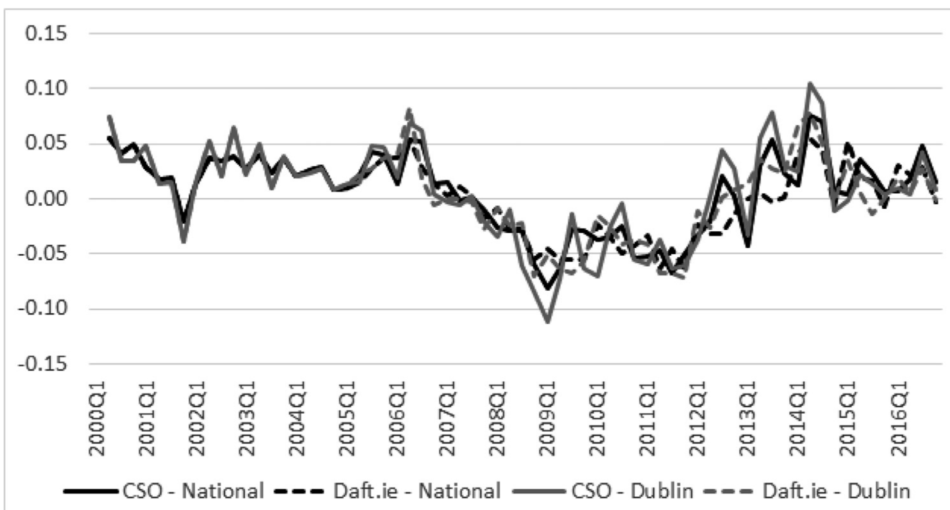

Fig. 2. Change in log sale prices for Irish housing, by source and scope.

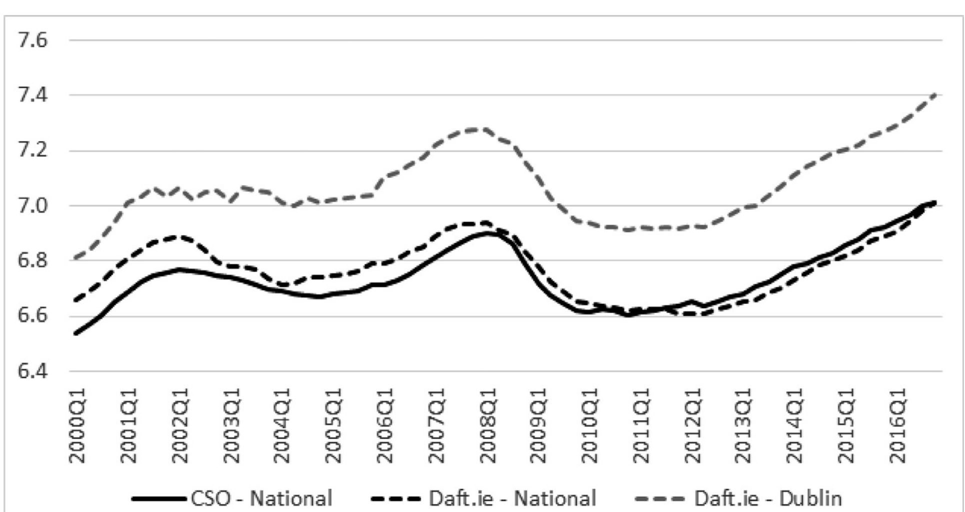

Fig. 3. Rental prices for Irish housing (logs), by source and scope.

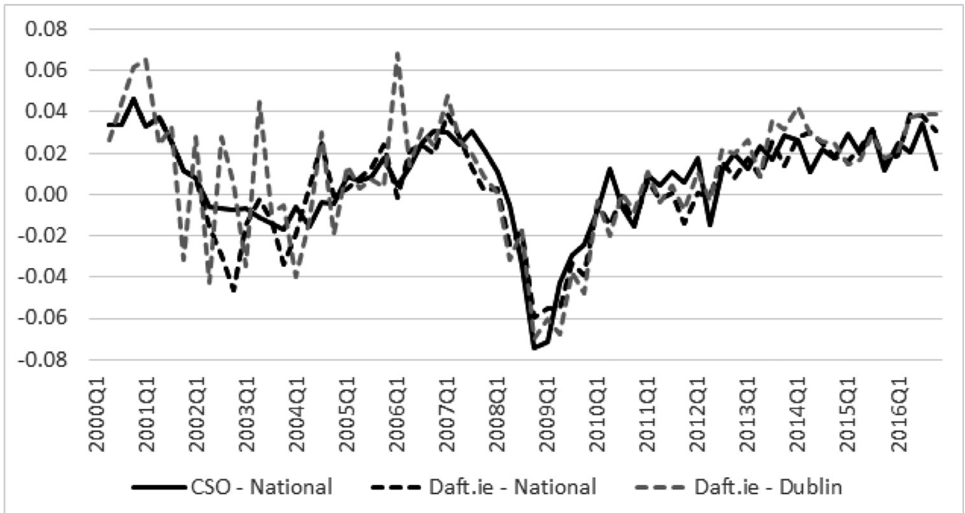

Fig. 4. Change in log rental prices for Irish housing, by source and scope.

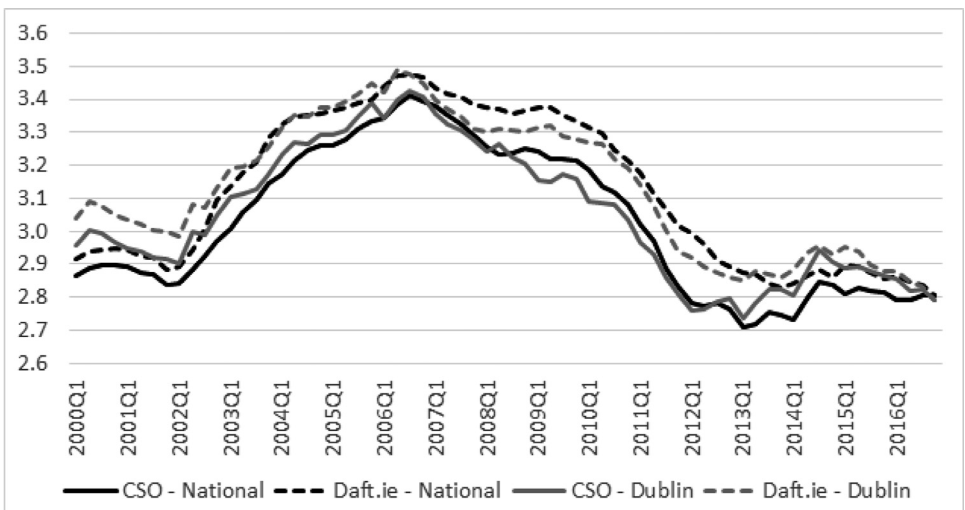

Fig. 5. Sale-to-rent price ratios for Irish housing (logs), by source and scope.

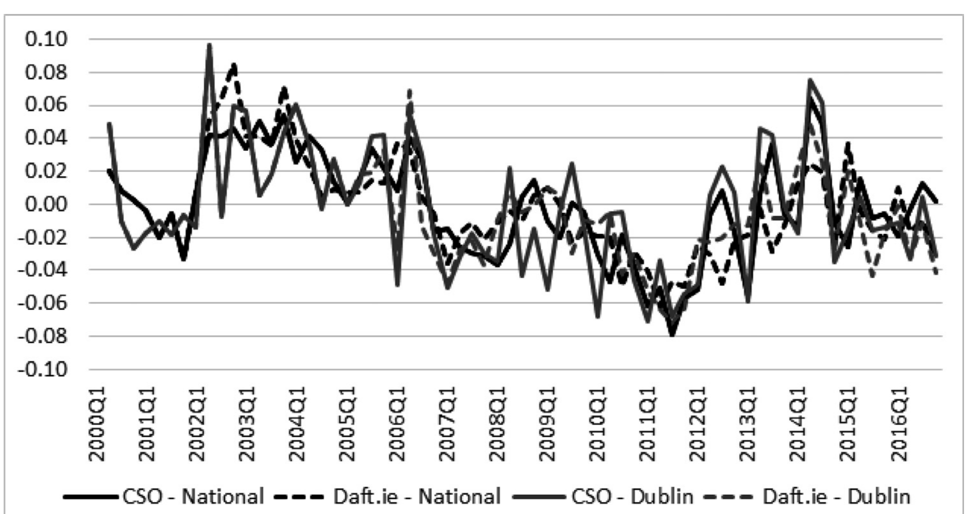

Fig. 6. Change in log sale-to-rent price ratios for Irish housing, by source and scope.

weighting both over time and across market segments, and given that other variables are measured at the national level, the national housing price ratio using daft.ie data is preferred, although this will be verified empirically.

The overall picture presented by these is relatively clear. For the period 2000–2002, the sale price for Irish housing was roughly 19 times the annual rent, before rising steadily to a peak of over 30 in late 2006 and early 2007. This fell back below 20 in 2011 and was close to 17 (roughly 2.85 in logs) for the entire period 2012–2016. There is some evidence that for the final two years of the sample, the housing price ratio fell slightly, with a quarterly average increase in sale prices of 1.8% but an average rental price increase of 2.5%. The first-order changes, however, and thus the focus of this analysis, are the rise and subsequent correction in the housing price ratio in the decade to 2011.

3.4. User cost

The two main time-varying components of the user cost of owner-occupied housing, typically regarded as the principal determinant of the housing price ratio, are mortgage interest rates and expected capital gains. Mortgage interest rates are measured using official data (from the CSO) and are presented in nominal after-tax terms, taking into account any mortgage interest relief available; relief was abolished in 2013. The series varies from a high of 4.7% in mid-2008 to a low of 2.5% in mid-2009 but, with the exception of that one quarter, is within one percentage point of the mean (3.7%) for the full 2000–2016 period throughout.

In contrast with minimal changes in borrowing costs, expected capital are likely to have varied substantially during the period. To measure expected capital gains, the approach here follows the existing literature in thinking of expectations as driven in large part by adaptive backward-looking expectations (see, for example, Duca et al., 2016). As with Muellbauer (2007), a combination of one- and four-year annualised changes in the sale price of housing are used. These are shown, together with the after-tax rate of interest for mortgage credit, in Fig. 7. While the mean in both cases is low and positive, similar to the interest rate, the spread is an order of magnitude larger (standard deviations of 12% and 13% respectively). For roughly half the period observed, the one- and four-year increases in sale prices are greater than 10% or less than −10%.

3.5. Credit conditions

Non-price conditions in the mortgage credit market are measured in two principal ways. One is a ratio of mortgage credit to household deposits, a system-wide measure of leverage. This rose from less than 90% in 2000 and 2001 to a peak of 150% in early 2008, before falling back to less than 90% after 2014. At least some of this later fall is driven by securitization, including a 20 percentage point fall in late 2011, which means this may not accurately capture changes in conditions faced by individual households. For that reason, a series capturing median loan-to-value specifically for first-time buyers is used. The source of this data is the Central Bank of Ireland, who have access to loan-level data from Irish-owned banks. Data for the period prior to 2011 relates to the stock of mortgage debt on bank balance sheets at that point, exploiting information on date of origination. Figures for 2011–2016 come from contemporaneous updates to information provided to the Central Bank. As shown in Fig. 8, the typical loan-to-value rose from two thirds in 2000 and 2001 to over 90% in 2005 and 2006. In other words, the multiple of a deposit loaned out to first-time buyers on average rose from 3 to 12. Since 2009, the typical LTV has been very stable, at close to 85% (a multiple of roughly 7).

Other User Costs. User cost includes property taxes, maintenance and depreciation. Ireland did not have an annual property tax for the period to 2013, relying instead on stamp duties, i.e. transaction taxes. The percentage rate that applied was subject to certain bands, but for a first-time buyer of a house of average value, the rate was 0% throughout most of the 2000s. Thereafter, a rate of 0.18% of the market value in May 2013 was applied on an annual basis. The CSO Household Budget Survey 2010 gives an estimate of the amount spent on maintenance; based on spending and housing prices in 2010, households spend on average about 0.5% of the value of their dwelling on maintenance. However, setting a fixed proportional cost of maintenance means that this does not vary over the period and thus is irrelevant for a dynamic model of changes in house prices over time. A similar point can be made for financial and psychological costs of moving.

4. Empirical analysis

The empirical analysis is presented in three parts. In the first part, using the daft.ie national housing price ratio as the preferred series, a

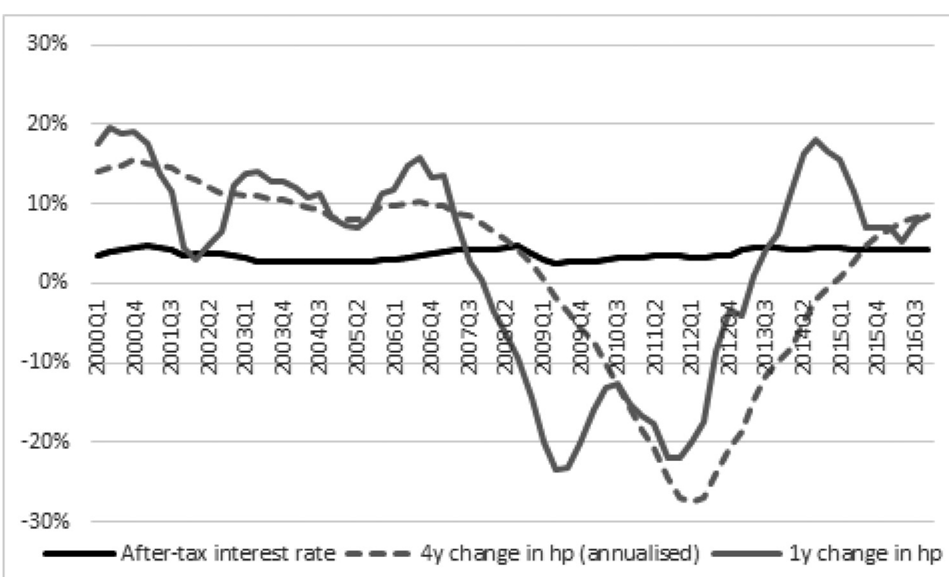

Fig. 7. Measures of user cost of Irish housing.

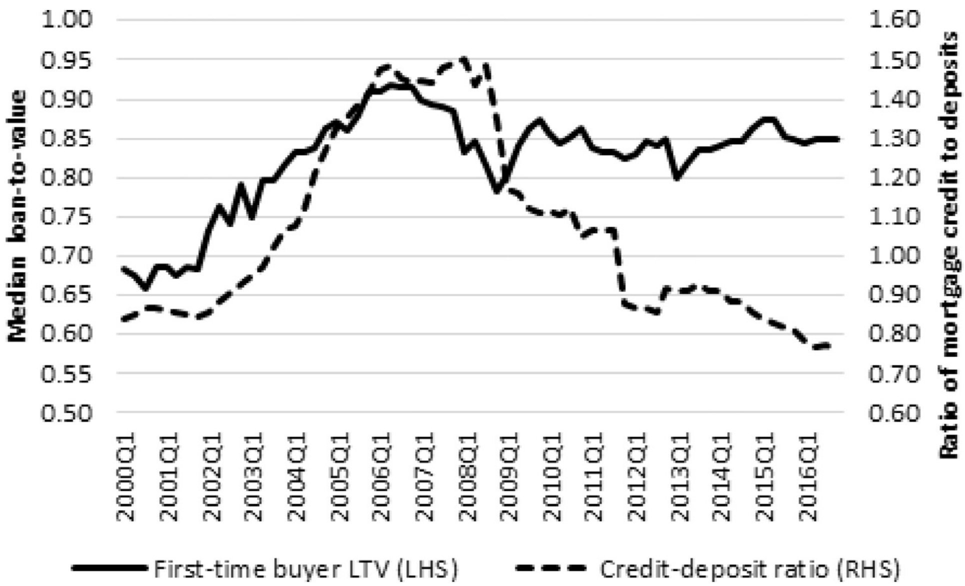

Fig. 8. Measures of credit conditions in the Irish housing market.

baseline empirical specification is established for the period where the housing price ratio undergoes substantial change (2000–2012). Secondly, this is cross-checked using alternative datasets. Lastly, tests for regime stability are undertaken, by extending the dataset and using rolling regressions to explore the stability of key parameters of interest.

Identification and Causality. Throughout, a one-step error-correction framework is employed, combining both long-run fundamental determinants of the housing price ratio (elements of the user cost and measurements of credit conditions) as well as short-run dynamics. Error correction methods rely on the use of lagged levels, together with potential short-run dynamic effects, as the regressors, and the current change in the outcome of interest as the regressor. Identification of the underlying long-run relation comes through this temporal asymmetry: as discussed by Angrist and Pischke (2008), this overcomes the problem of bad controls, as variables measured before the variable of interest was determined cannot be the outcomes in the causal nexus.[3]

4.1. Establishing a baseline

In establishing a baseline model, a varying set of long-run determinants of the housing price ratio are included, alongside a fixed set of dynamic terms. Where credit conditions are not included, two dynamic terms are included: the lagged value of the dependent variable, capturing any momentum; and the contemporaneous change in rents, capturing the extent to which changes in the ratio reflect changes in the denominator. In addition, where credit conditions are included, the contemporaneous change in credit conditions – as measured through either or both the loan-to-value or credit-deposit ratio – is also included.

Standard models of the equilibrium housing price ratio suggest just one fundamental long-run determinant: the user cost. As discussed above, the two main time-varying components of the user cost are after-tax rates of interest and the expected capital gain. Column (1) in 1 presents the results for a model with only the interest rate included, while Column (2) adds expected capital gain. Without capital gains, the interest rate term is insignificant and incorrectly signed. Adding capital gains, adjustment of the ratio to its suggested long-run equilibrium value is slow (3.4% per quarter or a rate of less than 15% a year), although all coefficients have the predicted sign: lower interest rates and higher past sale price inflation are associated with a higher ratio of sale to rental prices. The dynamic terms are important: without them, adjusted R-squared falls from 79% to 51% and the root mean square error (RMSE) rises from 0.0148 to 0.0228 (the mean of the dependent variable is -0.0013 while the standard deviation is 0.03). Columns (3) and (4) exclude iteratively 4-year and 1-year changes in sale prices and, taking Columns (2)-(4) together, it appears to be the shorter history (the most recent year) that has the greater explanatory power.

Columns (5) and (6) add two separate measures of credit conditions: the typical loan-to-value for first-time buyers (LTV) and the mortgage credit to deposit ratio (MCDR). Both noticeably improve the fit of the model, as shown by three key metrics. Firstly, speed-of-adjustment (SOA) improves from 3% per quarter to 12%-16%, depending on the specification. Furthermore, both the adjusted R-squared and in particular RMSE improve, the latter from 0.015 to roughly 0.0125. The additional dynamics terms are statistically significant, while the principal coefficients of interest remain in line with theoretical predictions. (The lack of variation in the after-tax interest rate means that this term is noisy and imprecisely estimated.)

Column (7) in Table 1 presents the full baseline model, which includes both LTV and MCDR measures of credit conditions. This model has the fastest SOA, higher adjusted R-squared and lowest RMSE of all the models presented in this section. The one-step error correction specification allows quarterly changes in the ratio of sale to rental prices in the Irish housing market, 2000–2012, to reveal an underlying long-run equation for that ratio as well as the speed with which the ratio converges to that equilibrium relation. In particular, it indicates that for the period in question, a ten percentage point increase in the median first-time buyer LTV was associated with a 9% rise in sale prices, holding other factors – including rental prices and the system wide ratio of credit to deposits – constant.[4] Similarly, an increase in MCDR of 10 percentage points, as happened between the second and third quarters of 2004, was associated with an increase in the housing price ratio of almost 5%.

There are limits to interpreting these coefficients causally. As outlined above, the nature of the ECM method means that it is not possible for changes in the housing price ratio to affect credit conditions, as these are not entered contemporaneously here. That said, for a full

[3] Nonetheless, a single-equation system cannot capture determinants of other key variables, including credit supply. This issue is discussed in more detail in the final section.

[4] Where the MCDR term is omitted and LTV is the only credit conditions term included, the equivalent sale price increase is 24%.

Table 1
A baseline model of the housing price ratio, 2000–2012.

	Column (1) b/se	Column (2) b/se	Column (3) b/se	Column (4) b/se	Column (5) b/se	Column (6) b/se	Column (7) b/se
Constant	0.069	0.142**	0.135**	0.143**	0.165***	0.461***	0.364**
	(0.049)	(0.045)	(0.043)	(0.049)	(0.045)	(0.096)	(0.103)
HPR (t-1)	–0.023	–0.034**	–0.032**	–0.037**	–0.122***	–0.163***	–0.167***
	(0.013)	(0.012)	(0.011)	(0.013)	(0.026)	(0.035)	(0.035)
Interest rate (t-1)	0.080	–1.070*	–1.027*	–0.784	–0.510	–1.899***	–1.228*
	(0.424)	(0.433)	(0.422)	(0.453)	(0.431)	(0.497)	(0.573)
4y hp change (t-1)		0.021		0.105**	0.126**	0.037	0.083
		(0.041)		(0.030)	(0.045)	(0.035)	(0.047)
1y hp change (t-1)		0.145**	0.164***		0.088	0.105*	0.082
		(0.051)	(0.035)		(0.046)	(0.044)	(0.044)
LTV (t-1)					0.292***		0.151
					(0.072)		(0.091)
MCDR (t-1)						0.109***	0.076*
						(0.028)	(0.033)
Delta HPR (t-1)	0.816***	0.285*	0.285*	0.505***	0.325**	0.356**	0.379**
	(0.078)	(0.131)	(0.130)	(0.113)	(0.118)	(0.113)	(0.115)
Delta rent	–0.239*	–0.693***	–0.727***	–0.360***	–0.665***	–0.704***	–0.660***
	(0.101)	(0.148)	(0.133)	(0.097)	(0.133)	(0.131)	(0.131)
Delta LTV					0.234*		0.186
					(0.102)		(0.100)
Delta MCDR						0.100*	0.068
						(0.043)	(0.047)
R-squared	0.705	0.795	0.798	0.763	0.849	0.853	0.862
RMSE	0.0177	0.0147	0.0146	0.0159	0.0126	0.0125	0.0121
N	51	51	51	51	51	51	51

Note: Regression output, where dependent variable is the change in the log housing price ratio, using daft.ie series for sale and rental prices for the period 2000–2012. Two sets of variables are included in each model: determinants of the long-run relationship (all lagged) and short-run/dynamic terms. Columns show specifications suggested by economic theory. For each variable, coefficients are shown above standard errors in parentheses; asterisks beside coefficients denote statistical significance, where relevant: *, ** and *** signify the 10%, 5% and 1% levels respectively.

Table 2
Summary of results across different data series.

	CSO National	daft.ie National	CSO Dublin	daft.ie Dublin
Interest rate	–7.496	–7.360	–7.613	–10.513
4y change	0.821	0.499	0.645	0.801
1y change	0.243	0.494	0.282	0.233
Loan-to-value	1.100	0.904	0.865	0.663
Credit-deposit	0.380	0.457	0.366	0.355
Adj R-sq	0.821	0.862	0.697	0.678
RMSE	0.0138	0.0121	0.0211	0.0201

Note: Table shows implied coefficients in long-run equilibrium relation for the housing price ratio and summary statistics from underlying error-correction regression.

causal interpretation, a richer system of equations would be needed, to understand the determinants of credit conditions. These may include interest rates, for example: lower interest rates may encourage banks to offer a higher loan-to-value. What is clear, though, is that, based on the evidence of the Irish housing market 2000–2012, looser credit conditions were associated – *ceteris paribus* – with a higher ratio of sale prices to rental prices.

4.2. Consistency across datasets

As outlined in Section 3, four separate series for the housing price ratio were calculated, based on coverage (Ireland as whole, or Dublin alone) and source (CSO and daft.ie). While *a priori* official data may appear more attractive to use, as outlined above, the exact methodology behind the rental series is not stated. In addition, the weighting of segments and markets in both CSO series is unclear. In contrast, both the daft.ie sale and rental series use the same methodology and weighting, making comparisons of average sale and rental prices less prone to issues around bias. In relation to scope, the Dublin market is likely to be more homogeneous. However, the regressors – in particular interest rates and credit conditions – are measured for the national level.

A check of the baseline model described above for each of the four housing price ratio series confirms that the daft.ie national *hpr* series performs strongest.[5] A summary is presented in Table 2. Implied long-run coefficients on the net nominal interest rate and on the credit conditions terms are typically stable across all four series. However, the two Dublin series have significantly lower overall model fit (as measured by adjusted R-squared) and significantly larger RMSE (over 0.02 compared to 0.012-0.014). Of the four series, the daft.ie national series has the lowest RMSE and highest adjusted R-squared. While its speed of adjustment is not the largest, it is no less precisely estimated than the other series (as measured by the *t*-statistic).

4.3. Parameter and regime stability

The results from the two preceding subsections suggest that inclusion of credit conditions, using either or both loan-to-value or the credit-deposit ratio significantly improves the fit, over classical models of the housing price ratio. Given the extreme housing market conditions that prevailed in Ireland during the early 21st century, it is worth investigating whether similar housing market regimes applied in the boom (2000–2007), bust (2008–2013) and post-crash recovery (from 2013 on). Results are presented for each of these three periods, alongside the original period analysed (2000–2012) and an extended period (2000–2016), in Table 3. What is immediately clear is that it is the boom period (2000–2007) that is driving the long-run relation for the period 2000–2012. The relation largely holds when 16 additional

[5] This is not simply an artefact of the process outlined in the previous section and holds true for other parsimonious specifications including credit conditions.

Table 3
Model stability across market periods.

	2000–2012 b/se	2000–2016 b/se	2000–2007 b/se	2008–2013 b/se	2013–2016 b/se
Constant	0.364**	0.474***	0.653**	0.416	4.219
	(0.103)	(0.104)	(0.226)	(0.435)	(2.834)
HPR t-1	–0.167***	–0.179***	–0.329**	–0.115	–0.997
	(0.035)	(0.037)	(0.092)	(0.116)	(0.613)
Interest rate t-1	–1.228*	–1.830**	–0.609	0.830	5.270
	(0.573)	(0.549)	(0.666)	(1.556)	(8.926)
4y hpchange t-1	0.083	0.035	–0.480	0.277	–0.961
	(0.047)	(0.036)	(0.523)	(0.170)	(0.644)
1y hpchange t-1	0.082	0.115*	0.322*	–0.070	0.742
	(0.044)	(0.044)	(0.138)	(0.176)	(0.549)
LTV t-1	0.151	0.033	0.429	–0.039	–0.896
	(0.091)	(0.044)	(0.213)	(0.308)	(1.569)
MCDR t-1	0.076*	0.118***	0.089	–0.048	–1.096
	(0.033)	(0.029)	(0.055)	(0.103)	(0.902)
Delta HPR t-1	0.379**	0.306**	0.407**	0.012	0.076
	(0.115)	(0.108)	(0.135)	(0.309)	(0.662)
Delta rent	–0.660***	–0.739***	–0.540*	–0.492	0.086
	(0.131)	(0.132)	(0.193)	(0.374)	(1.310)
Delta LTV	0.186	0.128	0.511**	-0.151	–1.457
	(0.100)	(0.098)	(0.155)	(0.215)	(1.852)
Delta MCDR	0.068	0.107*	0.249	0.034	–0.164
	(0.047)	(0.048)	(0.125)	(0.067)	(0.715)
R-squared	0.862	0.763	0.863	0.565	–0.284
RMSE	0.0121	0.0146	0.0109	0.0125	0.0232
N	51	67	31	24	16

Note: Regression output, where dependent variable is the change in the log housing price ratio, using daft.ie series for sale and rental prices for different periods. Two sets of variables are included in each model: determinants of the long-run relationship (all lagged) and short-run/dynamic terms. Columns show different market periods. For each variable, coefficients are shown above standard errors in parentheses; asterisks beside coefficients denote statistical significance, where relevant: *, ** and *** signify the 10%, 5% and 1% levels respectively.

quarters, covering the years 2013–2016, are added, although of the two credit conditions terms, only the MCDR term retains statistical significance.

Focusing solely on the period 2008–2013, when the ratio corrected (sale prices fell by significantly more than rental prices), the model weakens dramatically. The coefficient on the interest rate changes sign, while no variable is statistically significant at conventional levels. This may in part due to the small sample size (24 quarters). However, six variables remain statistically significant in the earlier period, which is only slightly longer (31 quarters). The model breaks down completely in the post-2012 period, with a negative adjusted R-squared and no interpretable speed of adjustment and thus long-run relation.[6]

One potentially interesting finding is the set of results for the one- and four-year housing price changes across specifications. In general, theory allowing for backward-looking housing price expectations would suggest these terms have positive coefficients in the long-run relation: the more sale prices have increased recently, the more this will shift out current demand in anticipation of similar increases in the near future. This is the case for both one- and four-year increases in the 2000–2012 sample. Splitting the sample reveals an asymmetric pattern, however: in times of rising sale prices (2000–2007 and 2013–2016), one-year changes play a far more important role, while when prices were falling (2008–2013), it is four-year changes that are in line with economic theory (and closer to conventional statistical significance).

4.4. Parameter stability

A final test of the model's stability is undertaken using both recursive and rolling regressions. For recursive regressions, the baseline model is applied to an ever-increasing window size, starting with a 32-quarter window and extending ultimately to cover the entire period 2000–2016. For rolling regressions, the baseline model is applied to a moving 32-quarter window, with the earliest window starting in 2000Q1 and the latest window starting in 2010Q1. An overview of the speed of adjustment and implied coefficients in the long-run equilibrium relation is presented graphically in Figs. 9, 10, 11, 12, 13, 14 and discussed below.

In relation to speed of adjustment (the coefficient on the lagged housing price ratio), the recursive regression produces a very stable coefficient once the sample extends into the 2010s (typically around –0.2). The coefficient is much larger in absolute value in earlier smaller samples. It is always statistically significant. For the rolling regression, there is no obvious trend, with a coefficient typically around -0.2, and largely statistically significant, although final sample have wider confidence intervals.

For the three user cost variables, the results vary. The coefficient on the interest rate is growing in absolute size over time, from less than -1 to almost -2 as the sample grows. Once the sample extends past 2008, it is almost always statistically significant. In the rolling regression, it is statistically significant and negative for most early samples but increasingly imprecisely estimated thereafter and drifting positive. This is likely due to the small variation in interest rates over the sample. For the one-year lagged change in sale prices, the coefficient is positive and marginally statistically significant throughout in the recursive sample, but there is a step down in magnitude once the 2010s are included. Similarly, in the rolling regression, the coefficient is largest at the start and falling thereafter with no statistical significance after the 2002–2009 window. Lastly, for the four-year change in prices, the coefficient in the recursive regression is negative for early samples and while it turns positive from 2011 end-date it is not statistically significant. In rolling regression windows, it is positive throughout from 2003 to 2010 windows on but statistically significant only in parts (especially in samples starting in 2006/2007).

Lastly, there are the two credit conditions variables in the long-run relation: the first-time buyer loan-to-value and the ratio of mortgage credit to deposits. The inclusion of both variables – suggested by the

[6] This remains the case if MCDR terms, and the four-year change, are omitted, to allow a parsimonious specification, given there are just 16 quarters of data.

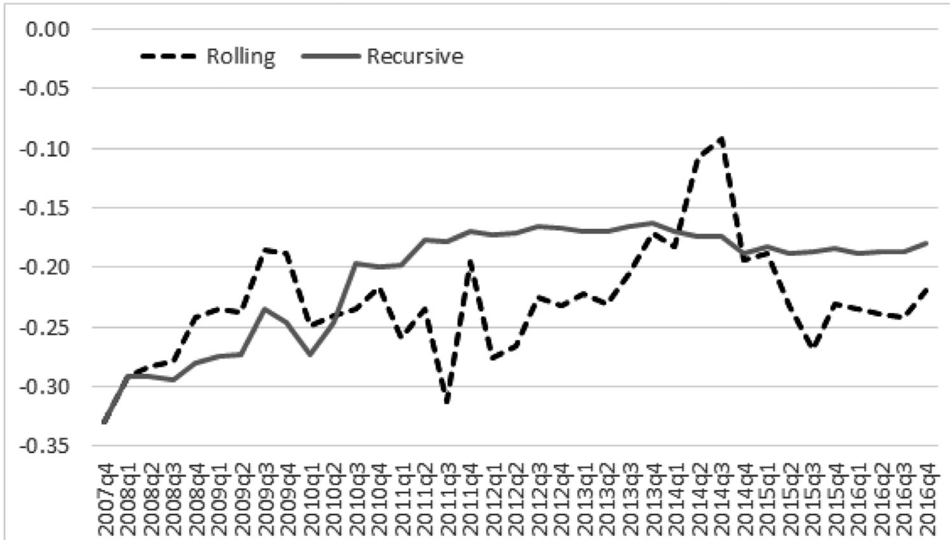

Fig. 9. Speed of adjustment to long-run equilibrium relation, by window end-date.

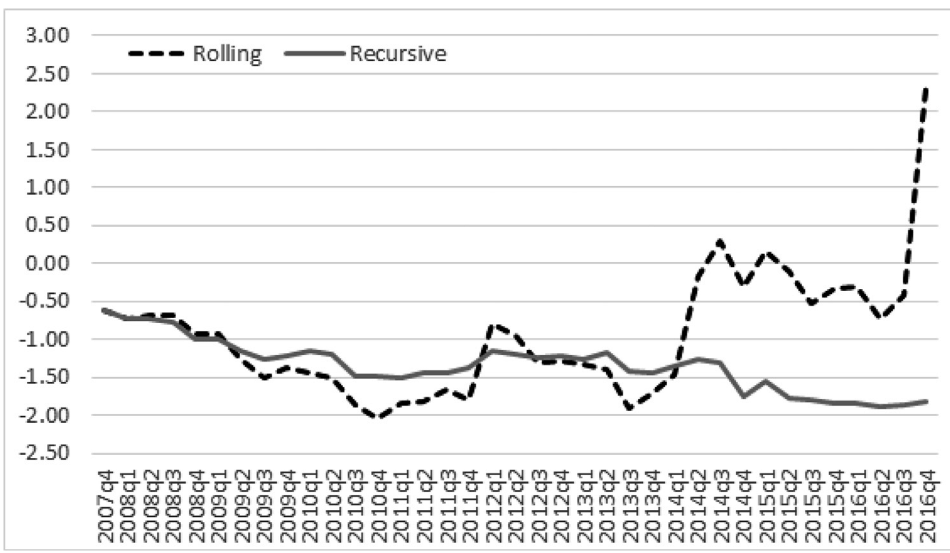

Fig. 10. Implied long-run coefficient on interest rate, by window end-date.

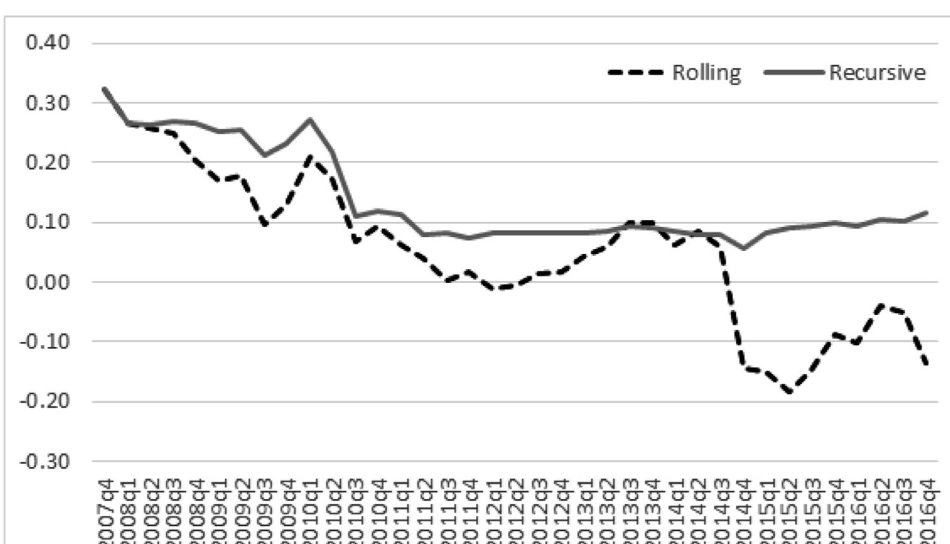

Fig. 11. Implied long-run coefficient on one-year change in housing prices, by window end-date.

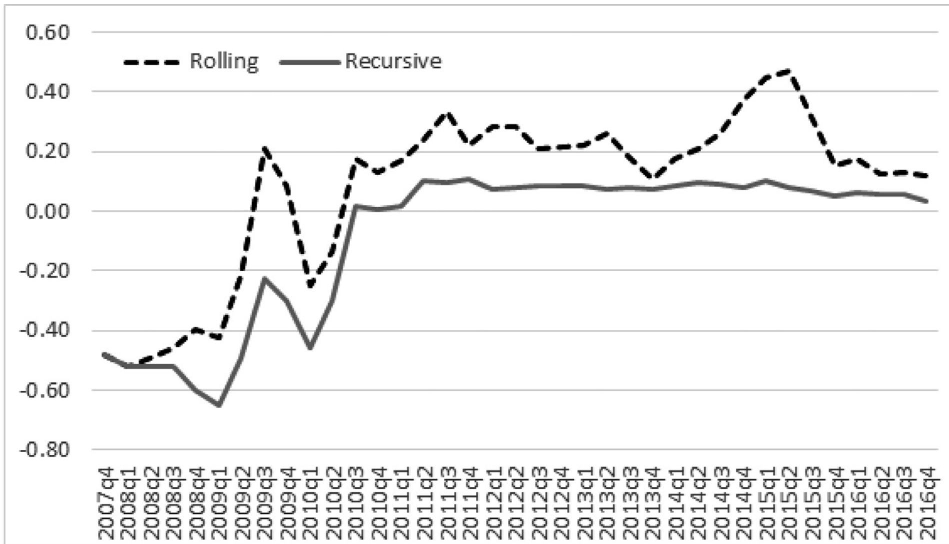

Fig. 12. Implied long-run coefficient on four-year change in housing prices, by window end-date.

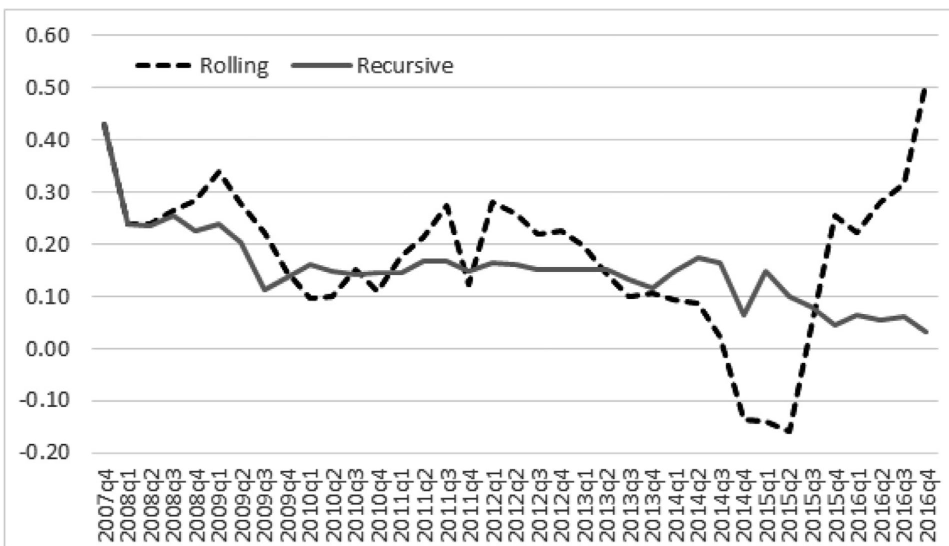

Fig. 13. Implied long-run coefficient on loan-to-value, by window end-date.

Fig. 14. Implied long-run coefficient on mortgage credit-deposit ratio, by window end-date.

analysis performed above – may affect their statistical significance somewhat, as they are correlated, particular in the first ten years of the sample. In the recursive regression, the LTV coefficient is relatively stable at about 0.15 when the sample extends beyond 2008 and up to 2015, but is rarely statistically significant. Furthermore, it is falling in the final two years, when Central Bank rules around minimum deposits were introduced. In the rolling regressions, the coefficient is mostly positive but never statistically significant. It falls below zero once the sample include 2015 but then rises sharply (but with larger CIs) at the end. In the regressive specification, the MCDR coefficient is consistent both in terms of size and sign (close to $+0.1$) and in statistical significance. Its counterpart in the rolling regressions is significantly more volatile, mostly notably in samples including 2014 and 2015. Overall, the coefficient displays very similar trend to LTV.

In all, the results point to an efficient housing market, in the sense that speed of adjustment remains relatively rapid (close to 20%) throughout sample analysed. Nonetheless, there are substantial uncertainties around the exact relation and its permanence. It is clear that there are at least two break-points in the relation. The first occurs around 2010 and is most noticeable in the lagged housing price appreciation terms but also in the speed of adjustment (recursive model). It is at this point that the one-year change in prices appears to become far less important, while the four-year change assumes the expected sign. As discussed above, this hints at different foundations of housing price expectations when prices are rising compared to falling.

The second break point appears to occur in late 2014/early 2015 and is apparent in all coefficients from rolling regression (which may hint at the role of omitting 2006 and early 2007) but also in the coefficients on credit conditions terms in the recursive model. In mid-2014, sale prices for housing were rising – in particular in the Dublin area – at double-digit rates and concerns about another housing bubble led the Central Bank to introduce macro-prudential rules, including per-borrower restrictions on loan-to-value and loan-to-income. These rules were introduced in early 2015, after which there was much more limited inflation (again especially in Dublin).

5. Conclusion

This paper has examined the housing price ratio in Ireland during its 2000–2012 housing boom and bust. Following a growing literature that focuses on the role of credit conditions, in particular Duca et al. (2011), the ratio of sale to renting prices for Irish housing was placed in an error-correction framework, where the key long-term determinants related to the user cost and credit conditions. This involved the use of new series of the typical loan-to-value paid by first-time buyers in Ireland, from 2000 on, using Central Bank of Ireland micro-data. Credit conditions were also measured at an aggregate level, using the system-wide ratio of mortgage credit to household deposits. The key time-varying terms in the estimated user cost are the net interest rate, minus a combination of 1-year and 4-year inflation, reflecting the stylised fact that expectations of future price changes appear to be to a large extent driven by recent changes.

A number of different potential series for the sale to rental ratio of housing prices in Ireland were considered. On the basis of consistent methodology over time and across segments, the daft.ie sale and rental series covering the entire Irish market were used, although the fit did not worsen dramatically if other datasets (including official series, with unclear weighting or methods) were used. The principal focus of the study was to examine the 2000–2012 period, one where both the housing price ratio and credit conditions varied dramatically. Nonetheless, the model was extended to include the post-2012 period, with both recursive and rolling window methods employed to examine regime and parameter stability.

In relation to the key boom/bust period in the Irish housing market, a relatively parsimonious long-run relationship emerges between sale and rental prices for Irish housing, where credit conditions matter, both for the long-run equilibrium relation and for short-run dynamics. In the preferred specification, the model explains over 86% of the variation observed in the housing price ratio, with a quarterly speed of adjustment to the long-run relation of 17% and a root mean square error of 0.012. The implied long-run relation implies that user cost includes a roughly equal measure of 4-year and 1-year lagged changes in housing prices, as well as a sizeable coefficient associated with interest rates, although this series is subject to minimal variation. The results also imply that a ten percentage point increase in the median first-time buyer loan-to-value was associated with a 9% rise in sale prices, holding other factors – including rental prices and the system wide ratio of credit to deposits – constant. Similarly, an increase in MCDR of 10 percentage points was associated with an increase in the housing price ratio of almost 5%.

Both these results are imprecisely estimated, given the relatively small sample size. Adding further quarters does not aid the precision of estimation, as the subsequent years (2012–2016) contain no substantial trend for the housing price ratio or the loan-to-value and only a mild downward drift in the ratio of credit to deposits. A thorough investigation of parameter stability suggests two potential break-points in the model, with coefficients in the long-run relation changing in 2010 and again in late 2014/2015.

In addition to contributing the first analysis of the housing price ratio of one of the world's most severe housing market cycles, this research suggests some broader findings for policymakers interested in housing market cycles and the role of macroprudential rules. Firstly, the results strongly support the recent move by policymakers, in particular Central Banks, to use macroprudential rules – both a borrower level, e.g. LTV, and at system-wide level, e.g. MCDR – to anchor sale prices of housing to the real economy, proxied here by the rental price of housing. Models of the housing price ratio where credit conditions are included as explanatory variables clearly outperform those where they are excluded, even if the exact coefficients are imprecisely estimated.

Secondly, the results provide further support for believing expected capital gains to be in part backward-looking. Both in rising and falling markets, recent housing price changes substantially improved the model, compared to a model where only interest rates were included. Further, the results suggest that the balance of immediate (1-year) and more prolonged (4-year) housing market history may play different roles in contributing to expected capital gains across boom and bust. The existence of survey data directly measuring expected housing price inflation in Ireland, from 2011 on, may also help address this issue in future.

The analysis presented here is limited to a focus entirely on the determinants of the housing price ratio, in particular credit conditions, and largely on those determinants during a particularly extreme housing market cycle. It has left for future work a separate but also very policy-relevant question on the determinants of credit conditions in Ireland during this period. One concern, explored in work such as Anundsen and Jansen (2013) and Cuestas (2017), is the potential for housing prices to be a key determinant of credit conditions. Error correction models of both the credit-deposit ratio and the median first-time buyer loan-to-value for Ireland 2000–2012 do not suggest any clear cut relation from housing prices to credit, but this does not diminish the importance of a more careful examination of these issues.[7] The use of a one-step error correction specification here is motivated largely by an attempt to focus on the determinants of the housing price ratio, allowing the data to reveal the true underlying relationship. Nonetheless,

[7] The results of these analyses are not shown here but area available from the author on request. Overall they suggest that the determinants of credit conditions in Ireland during this period do not overlap significantly with the determinants of the housing price ratio: compared to an adjusted R-squared of over 85% in the case of *dlhpr*, elements of the user cost and the housing price ratio explain at best one quarter of the change in median LTV. Nonetheless, it is likely that (net nominal) interest rates do have an effect on the housing price ratio. This is an important avenue for future research.

to the extent that further interactions may exist between the variables examined here – for example between interest rates and the median loan-to-value required by first-time buyers – this limits a causal interpretation on the coefficients presented in the results above.

The analysis here also avoided any discussion of the link between tenure choice and housing market outcomes. As suggested by Kim (2008), this is likely to be important. However, high-frequency data on tenure choice are not available for the Irish economy during this period. A suggestion for future research would be to explore spatial differences in tenure and housing market outcomes, in order to identify such effects.

Lastly, it is worth noting the speed with which housing prices in Ireland during this period adjusted to equilibrium. The preferred specification suggests that almost two thirds of the gap between the actual and equilibrium ratios was closed every year. The market was thus efficient in processing changes in credit conditions and expectations, even if those factors were subject to dramatic changes in a very short period of time. Understanding the factors affecting housing prices and other relevant outcomes in the more normal housing conditions is an obvious suggestion for future research.

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