Candidate Number: 3163G

An analysis into the price adjustment process in housing markets

Abstract: This paper conducts single-country ARDL and Pooled Mean Group panel estimations on a sample of 17 OECD countries with the aim of identifying equilibrium relationships between housing prices and its "fundamental" determinants, such as disposable income, interest rates, unemployment, credit, and the property tax rate. After establishing a cointegrating relationship, an analysis of the short-run dynamics follows, focusing on cross-country differences, the role of expectations and the possible sources of booms and busts. Our results have the potential to provide an internationally robust model of housing markets.

Word Count: 6532 words including equations, excluding $3\frac{3}{4}$ pages of tables and figures ~ 960 words.

1. Introduction

The value of property owned by households is a core component of their financial wealth and therefore has a significant effect on household consumption patterns and financial stability, which in turn impacts the credit conditions of the banking sector. Since the 2008-financial crisis, which was largely attributed to the over-leveraging of the housing market,[1] there has been an increased interest in trying to understand housing dynamics and its relationship with the overall macroeconomy.[2] More generally, housing is an area of interest from a social policy perspective, as an investment vehicle, and as a macroeconomic variable.

Several papers[3] have looked at the overvaluation of the housing market using indicators such as price-to-rent or price-to-income ratios. However, such indicators are often too reductive, meaning a more comprehensive model is necessary to determine the true fundamentals and their effect on house prices, so that we can disentangle long-run and short-run dynamics. Understanding these dynamics has implications for the estimated deepness of affordability issues, the extent of financial risk stemming from the housing market and the choice and effectiveness of policy tools trying to remedy them.

This paper uses an Error Correction Model approach to model the dynamics. Our approach is novel in that it conducts single country cointegration testing and estimation of several OECD countries, using the ARDL bounds testing method developed by Pesaran, Shin and Smith (2001), to establish reliable fundamental relationships, then carries out Pooled Mean Group estimations by Pesaran et al. (1999) to make the coefficient estimates more robust.

We find that the core long-run forcing variable of house prices is disposable income, while unemployment also plays a significant role as an indicator of financial stability. In the short-run, changes in income, size of credit stock and expectations are the main drivers, with the latter two having the most sizeable effects leading up to housing busts. Relative to these, the strength of the adjustment to equilibrium is of comparable size, with an estimated 8 years to correct a deviation. We found no strong evidence of short-run effects of unemployment, interest rates and inflation.

[1] Demyanyk (2011) provides an analysis emphasising the unsustainable subprime lending growth pre-2008 as an important reason for the breakdown.
[2] Brunnermeier et al. (2016) describe the "diabolical loop" leading to financial crises following a severe deterioration of banks' balance sheet, which was increasingly exposed to risky collateral pre-2008
[3] Girouard (2006) for an overview

The rest of this paper is organised as follows. Section 2 provides an overview of the literature on the equilibrium modelling of housing markets, Section 3 and 4 establish the theoretical model that this paper builds on, Section 5 discusses the data sources and issues, Section 6 describes the estimation methodology, Section 7 presents the results and Section 8 evaluates. Finally, Section 9 offers concluding remarks.

2. Literature Review

Several of the early core papers of the housing literature such as Abraham and Hendershott (1996), Capozza et al. (2002) or Malpezzi (1999) simply assumed the presence of cointegration and proceeded to run two-step estimation methods, based on the error correction term they derived from the assumed long-run relationship, and used interaction terms to test for the asymmetry of the error correction process and the presence of bubbles. Later, Meen (2002) carried out single-step ARDL error correction model estimations but still relied on inconclusive Johansen (1988) tests to try to establish a cointegration relationship. Girouard (2006) reports a review of the cointegration literature from OECD countries but there is an overarching trend of a lack of prudent cointegration testing, likely due to the short time span of available data for most countries and the low power of cointegration tests – given small sample sizes – available at the time, such as Engle Granger's (1987). We ran rudimentary testing of both Johansen and Engle Granger methods on our panel of 17 OECD countries and were unable to get conclusive results.[4]

More conclusive evidence of cointegration started to emerge with the proliferation of international panel methods. Adams and Füss (2010) conducted tests formulated by Pedroni (2004) – which allows the long-run coefficients of the cointegrating vector to be heterogeneous – and were able to reject the null hypothesis of no cointegration. Ott (2014) conducted Westerlund (2007) cointegration tests to reject no cointegration and estimated their models using the panel ECM method developed by Pesaran et al. (1999) – that this paper also builds on.

[4] Arestis (2014) finds cointegration for 18 of the 18 OECD countries they test for using the Johansen method. However, they appear to choose their cointegrating variables sporadically, often omitting factors otherwise believed to be fundamental, such as income.

Regarding estimation results, the core heterogeneity comes from the supply side variables included in the estimation. Adams and Füss (2010) use construction costs and find a coefficient of 1.3 but observe significant variation in the coefficient estimates on their "Economic Activity" variable across countries, due to possible over-aggregation. Meen (2002) and Ott (2014) use housing stock and find coefficients between 2 and 3. Meen (2002) observes that the omission of the supply variables has significant effect on his income elasticity estimates. Finally, Philiponnet and Turrini (2017) use housing investment in a DOLS estimation but find positive coefficient estimates instead of their expected negative ones. A discussion around such results will follow in this paper.

3. Housing market theory

The modelling approach used in this paper builds on a theoretical model of real estate markets developed by DiPasquale and Wheaton (1996), and also used by Adams and Füss (2010). It encapsulates the heterogeneous nature of housing by connecting the market for the use of housing services with the asset market of housing as an investment and, lastly, the property development market such that a long-run general equilibrium can be illustrated.

The first diagram (*quadrant I*) is of the equilibrium state in the market for the use of built housing stock, where rent is the market clearing price such that the quantity of housing demanded at those rental prices is equal to the amount of housing stock available at that given time (S). The quantity of housing services demanded at a given price level, among other things, depends on the population's purchasing power, which is a function of their disposable income (DPI), unemployment (Un), degree of inequality and wealth (W)) and demographic factors, such as household formation (HH). A strong positive relationship between income and rents is expected due to the possibility of extracting "economic rent" when it comes to housing. Being a basic good and in limited quasi-fixed supply, landlords can charge in significant excess of marginal costs. Lower unemployment is expected to have an inflationary effect on house prices as it is associated with an expectation of permanent real income gains and lower inequality.

$$R = f(DPI, un, S, HH, W)$$

The second diagram (*quadrant II*) is of the arbitrage condition in the asset market between the present value of the user cost associated with owning a property and the stream of earnings derived from the particular asset, i.e., the rental price. The user cost is a function of the price of the property and the capitalization rate, which encapsulates the foregone real interest of investing in alternative assets (IR), taxes on property (tax), the depreciation rate of the real value of the housing unit (d), perceived risk of owning real estate, and the expected capital gains on housing investment (g). Mortgage rates also play a demand role by influencing the financing costs of owning a house. A “steeper” line in this quadrant is a graphical expression of higher capitalization rate.

$$\mathrm{R} = \mathrm{HP} * (\mathrm{IR} + \mathrm{tax} + \mathrm{d} - \mathrm{g})$$

Combining these two quadrants, by substituting in for the rental price, gives us the demand side of the housing market:

$$HP = \alpha_0 + \alpha_1 \mathrm{DPI} + \alpha_2 \mathrm{S} + \alpha_4 \mathrm{Un} - \alpha_5 \log(\mathrm{IR} + \mathrm{tax} + \mathrm{d} - \mathrm{g})$$

The third and fourth diagram (*quadrants III and IV*) represent the property development industry, which serves to reconcile the long-run imbalances between demand for housing and the existing supply of housing stock. The line in quadrant III is the short-run supply curve of the construction industry, representing the willingness of developers to build housing units given the final selling price and the cost of construction (cc). Quadrant IV relates the flow of new housing units, represented by investment (I), with the existing stock, in a capital accumulation equation of motion, where the long-run steady-state has been imposed. Hence, the line in this quadrant has a slope equal to the depreciation rate.

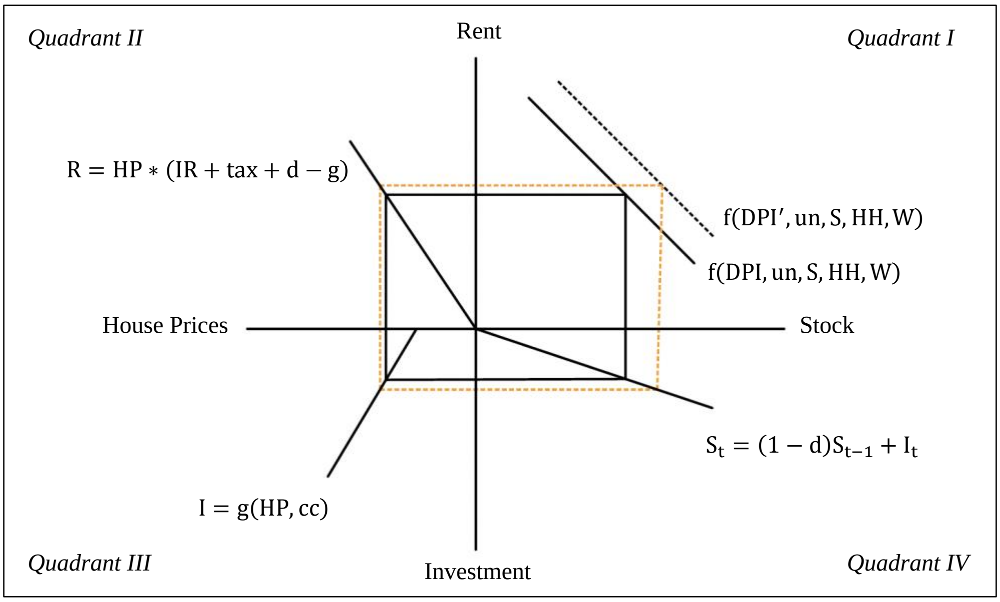
$$\mathrm{S_t} = (1 - \mathrm{d})\mathrm{S_{t-1}} + \mathrm{I_t}$$

According to DiPasquale and Wheaton, the supply side can be combined to yield the following equilibrium equation:

$$\mathrm{S_t} = \beta_1 + \beta_2 \mathrm{HP_t} - \beta_3 cc_\mathrm{t} + \epsilon_\mathrm{t}$$

A graphical representation of a general housing market equilibrium is presented in Figure 1 along with an alternative state of general equilibrium following a ceteris paribus increase in the average disposable income. The dynamic adjustment of the values of the endogenous variables can be tracked in a counterclockwise direction.

Figure 1: DiPasquale and Wheaton housing market diagram



4. Modelling approach and discussion of the theoretical model

The DiPasquale and Wheaton (DW) model combines several partial equilibria into one general equilibrium. If the variables in the diagram are integrated of order one (which will be tested) then it is posited that each quadrant represents a cointegrating relationship and that "combining" quadrants retains the cointegration structure. Estimation of partial equilibria, then, would theoretically require systems estimation to account for the simultaneity between the endogenous variables.

Given the limited data availability of supply side variables, this paper argues that there's an often-overlooked dichotomy in the DW model that causes a break in the error correction process. While the time span of the demand side and the property development quadrant is short-to-medium term, the stock-flow equation is much slower. Considering that the housing stock is relatively stable around its growth path, while investment displays significant swings influenced by the state of the developer market, – i.e., house and construction prices – even in the decades long time-period that we investigate, the cointegrating relationship between the two is unlikely to manifest in the data. Therefore, while the effect of housing stock on house prices will take effect in the short run, the change in house prices could take several years to influence the amount of housing stock. Hence, solving for the full equilibrium and using construction costs, as Adams and Füss (2010) do, or combining quadrants I, II, IV and using investment, as Philiponnet and Turrini (2017) do, is expected to lead to more bias – due to mixing short-run and long-run effects – than if we solve for the demand side and treat housing stock as weakly exogenous, meaning that it doesn't contemporaneously error correct. Thereby, with a further assumption of weak-exogeneity of the non-axis variables, the demand side partial equilibrium could be estimated as a single equation – as is attempted in this paper – instead of a system.

5. Data Sources and Issues

Our dataset covers 17 OECD countries on an annual basis between 1970 and 2018 with varying time span and coverage across the following countries: Belgium, Finland, France, Germany, Ireland, Italy, Netherlands, Spain, Australia, Canada, Denmark, Japan, Norway, New Zealand, Sweden, Switzerland, United Kingdom. Summary of the data is found in Table 1. Although the use of annual data is at the cost of lower degrees of freedom, due to the slow-moving dynamics in the housing market, the additional information content of quarterly data is limited, therefore results using annual data are argued to be consistent. Campbell and Perron (1991) demonstrate that the time span of the data is significantly more important in the power of unit root tests – the basis of cointegration testing – than the frequency of the data. Results of previous studies using annual data also accord with the rest of the literature that uses quarterly frequency.[5]

Due to simultaneous movements in prices and the characteristics of properties, the house price index has been quality adjusted – with hedonic regression or repeat sales methods, depending on the nature of the data–, so that only non-quality price movements are reflected in the index. Disposable income (DPI) was chosen over GDP since the latter includes the house price index, which would introduce endogeneity. DPI, however, doesn't consider in-kind public provisions covered from taxes that were deducted from the indicator. Mortgage rate availability is weak across countries, therefore long-term interest rates are used as a composite measure. While mortgage rates tend to react quickly to changes in the bank rate, the full adjustment can take much longer, so interest rates are expected to underestimate the true effect of mortgage rates.

[5] See Girouard (2006), Andrews (2010), OECD Economic Survey: Netherlands (2005)

Some of the most important variables in the pricing equation, the supply side variables, are also the most difficult to source in a standardised manner. Housing stock is unavailable for most countries before 2000 (EMF[6]) and are often reported with long-periods of imputations (ECB[7]). Therefore, they are not directly available for our estimation. Construction costs were also unavailable in a long enough panel fashion.[8] Gross fixed capital formation is used as an indicator of housing investment and is reported in constant prices.

For the purposes of estimation, all nominal variables have been expressed in real terms. Nominal house price and personal disposable income have been deflated by the Private Consumer Expenditure Deflator. Real long-term interest rates were derived by taking the difference between the 10-year nominal bond interest rates and the CPI based inflation rate. All variables are used in logarithmic form, except interest rates and inflation.

Table 1: Data Description

Variable	**Indicator**	**Source:**
House price index	HP	OECD
Disposable Personal Income	DPI	AMECO: Code: HVGTP
Real long term (10-year) bond yields	IR	AMECO and OECD
Unemployment rate	Un	OECD: EO
Ratio of property tax revenues to house prices	Tax ratio	OECD
Inflation	Inflation	OECD
Domestic Credit to Household Sector (% of GDP)	Credit	BIS
Gross Fixed Capital Formation (Dwellings)	GFCF	AMECO
Stock of Dwellings per capita	Stock	EMF

6. Methodology

The following long-run relationship was derived from the theoretical discussion of the DW model

$$p = \alpha_0 + \beta_1 \mathrm{DPI} - \beta_2 S + \beta_4 \mathrm{Un} - \beta_5 \log(\mathrm{IR} + \mathrm{tax} + \mathrm{d} - \mathrm{g})$$

However, housing stock isn’t reliably available pre-2000, therefore it cannot be included in the equation. By reviewing single-country studies where housing stock was more readily available, we see coefficients between -1.5 and -3.[9] In contrast, the elasticity of income is usually between 1 and 2.5. Using the stock data for the countries and years available to us, we derive that the

[6] European Mortgage Federation
[7] European Central Bank
[8] Eurostat publishes data from 2005 for a number of EU countries
[9] Meen (2002) and Jacobsen (2005) find stock elasticities of 2.9 and 1.4, respectively.

relative variation of disposable income to housing stock is a factor of 14.3,[10] on average. Putting these together, while the housing stock may theoretically be non-stationary, its role in the error correction equation – relative to the other explanatory variables – could be small enough such that its exclusion from the long-run equation is expected not to erase the stationarity of the error correction term empirically.

Meen (2002) finds that the exclusion of housing stock has a significant effect on the coefficient of income in the UK. Considering that Ott (2014) finds construction costs to be stationary, we could make the argument that house prices and housing stock are together cointegrated, and thus driven by the same exogenous long run forcing demand variables. Consequently, it is argued that the stock variable in Meen's (2002) estimation picked up the drift of income, potentially causing his income elasticity estimate to be biased. Considering these points, out of necessity, we proceed with the omission of the housing stock, acknowledging that this is expected to cause moderate bias in the estimated error correction term but may provide better elasticity estimates for income.

A long-run relationship between the variables of interest may be then represented by an ARDL equation of the following form:

$$\mathrm{y_t} = c_0 + c_1 t + \sum_{i=1}^{p} \lambda_j y_{t-i} + \sum_{j=1}^{k} \sum_{l_j=1}^{q_j} \boldsymbol{\beta}'_{j l_j} \boldsymbol{x}_{j,t-l_j} + \varepsilon_t \,, \qquad (1)$$

where y_t stands for the house price index, and x_t is the vector of k potential regressors: income (DPI), interest rates (IR), unemployment (un), credit (credit) and ratio of property taxes to the value of dwellings (tax ratio), while q_k is the regressor specific lag order, and ε_t satisfies the standard assumptions.

However, a traditional OLS estimation of this equation suffers from the issue of spurious regressions in the presence of unit root variables, if in truth there is no cointegration.[11] Relying on the use of the Beveridge-Nelson Decomposition, ARDL specification (1) can be reparametrized into an equation of stationary variables and an error correction term as follows:

[10] Derived by taking the ratio of the standard deviations of first differences of the log of both series indexed to 2005.
[11] Granger and Newbold (1974)

$$\Delta y_t = c_0 + c_1 t + \alpha(y_{t-1} - \boldsymbol{\theta} x_{t-1}) + \sum_{i=1}^{p-1} \psi_j \Delta y_{t-i} + \sum_{j=1}^{k} \sum_{l_j=1}^{q_j-1} \boldsymbol{\omega}'_{jl_j} \Delta \boldsymbol{x}_{j,t-l_j} + \varepsilon_t \quad (2)$$

It is apparent from this formulation that the estimation of this equation is only appropriate if the error correction term $y_{t-1} - \boldsymbol{\theta} x_{t-1}$ (ECT) is also stationary, therefore a pre-estimation test of this cointegration condition is necessary. In the presence of cointegration, the ECT represents the long-run equilibrium relationship, while α is the adjustment parameter responsible for restoring the long-run equilibrium and is thus expected to be negative and significant in non-degenerate cases.

As shown by Stock and Watson (1993), the Engle Granger 2-step residual-based test has low power to reject the null hypothesis of no cointegration, – especially in the context of the housing market where slow error correction[12] means that the null and the alternative hypotheses are too close to distinguish – due to small sample issues. Further, housing prices are subject to "periodically collapsing bubbles" in the spirit of Blanchard and Watson (1982) and Evans (1991). Therefore, the EG method will essentially have to decide between the presence of bubbles against the existence of a slow equilibrium process without bubbles, even though the truth is expected to be a combination of the two. The ARDL bounds test method presented in Pesaran et al. (2001) – that this paper builds on – is able to account for both these issues by taking account of the short-run dynamics and incorporating the serial correlation of house prices, in the form of persistence variables.

The ARDL bounds test estimates equation (2) and runs a preliminary F-test of the joint significance of the coefficients in $\boldsymbol{\theta}$ along with the adjustment parameter, α. The tests reported in this paper are based on Narayan (2005)'s critical values adjusted to small samples. The null hypothesis is rejected if the F-statistic is above the upper bound, meaning that there is a single cointegrating vector and the exogenous variables are purely I(1), while a result between the bounds means that a VECM approach is necessary due to potential cointegration among the explanatory variables. A second test of significance of the adjustment parameter alone is also run to exclude the degenerate case of an equilibrium relationship without a reversion to this equilibrium.

[12] Annett (2005) and Adams and Füss (2010) estimate correction to take 7 to 13 years.

This method relies on a restriction assumption – of weak exogeneity of the explanatory variables – that is key to the validity of the methodology of this paper. The weak exogeneity assumption states that the coefficient vector on the level of house prices in the marginal ECM equations of the explanatory variables is 0, thereby constraining the error correction to be done solely by the house prices. Therefore, relative to the Johansen (1988) test which estimates a general equilibrium, this method trades off robustness against simultaneity issues for gains in power in small samples. Interest rates, property tax ratios and unemployment are reasonably exogenous in this regard based on theoretical considerations. However, property being one of the largest assets of households, changes in house prices are likely to have wealth effects on consumption, introducing some degree of endogeneity, that is expected to cause an upwards bias on the long-run income coefficient. Meanwhile, as shown later on, including investment is likely to significantly violate the exogeneity assumption and is clearly inappropriate as a proxy for the supply side.

Based on our ARDL results, we carry out Pooled Mean Group (PMG) estimations developed by Pesaran et al. (1999). Pooling the data is desirable considering the slow equilibrium dynamics of the housing market and the limited data of 48 annual observations. Philiponnet and Turrini (2017) further demonstrate that pooled estimates can more robustly estimate the long-run equilibrium of individual countries, considering, as Pesaran et al. (1999) explain, the bias caused by group-specific omitted variables and measurement biases. The less synchronised the international housing markets and thereby the bias-inducing correlations are, the more robust pooling is, ceteris paribus. While the PMG estimation constrains the long-run coefficients to be equal across countries, it allows the intercept, short-run coefficients, and error variances to differ. The cost of restriction of homogeneity is also diminished by only running the regressions on a group of countries that had similar long-run coefficient estimates in the ARDL stage. The allowed heterogeneity in the short-run coefficients mutes the issue of overestimation and reflects the evident differences in the housing market dynamics across countries. Therefore, the results are both more robust and realistic. To confirm the validity of the homogeneity assumption, adjustment parameter-based Westerlund (2007) cointegration tests – based on a similar panel ECM estimation as the PMG method – were also conducted. Of the four test statistics, the results for Gt and Pt (Gt tests the existence of cointegration for at least one country, while Pt tests for cointegration across the entire panel) are reported, as they are more appropriate for finite samples.[13]

7. Estimation

7.1 Unit root testing

Although the often highlighted benefit of the ARDL bounds testing is that it doesn't require all the variables to be I(1), we conduct unit root tests as a reference point for the findings of the bounds test and to gain a better understanding of the data generating process. We present, in Table 2, the results for the single-country Dickey-Fuller GLS tests proposed by Elliott, Rothenberg, and Stock (1992), who showed that the DF-GLS method has significantly more power in small samples than the Augmented Dickey-Fuller test. We also ran panel unit root tests developed by Im et al. (2003) for the group of countries included in the panel estimations, which was chosen because it allows heterogeneity in the autoregressive slope parameter as well as in the lag structure of individual countries. Its null hypothesis is that all the panels have unit roots while the alternative is that some are stationary. The results of these tests can be found at the bottom of Table 2. Overall, the results point to unit roots in disposable income, credit, and tax ratio. There is weak evidence of stationarity in some countries for inflation and interest rates, while GFCF are effectively found stationary by both methods. There is no evidence of stationarity in unemployment based on single-country tests, but the panel test rejects the unit root hypothesis. A graphical inspection of the data reveals that there were frequent structural breaks in unemployment over the period, weakening the findings of unit root tests. The null of the Im et al. (2003) test is rejected at the 10% for house prices, but not at the 5%, which aligns with previous studies'[14] findings that the error correcting nature of house prices makes it difficult to distinguish it from a trend-stationary process.

[13] The Westerlund (2007) test's hypothesis is of H_0: $\alpha_i = \alpha > 0\ for\ all\ i,$ which is a stronger assumption than simply H_0: $\alpha_i > 0\ for\ all\ i.$

[14] Philiponnet and Turrini (2017) get a p-value of 0.06 for house prices using Levin-Lin-Chu test.

Table 2: Single-country DF-GLS and panel Im et al. (2003) unit root tests

	HP	DPI	IR	Un	Credit	Tax ratio	GFCF	Inflation
Australia	-1,738	-2,405	-2,847***	-0,813	-1,579	-1,919	-2,114	-1,689
Belgium	-1,740	-0,707	-1,614	-0,770	-2,998*	-2,659	-3,712**	-2,151*
Canada	-2,259	-2,240	-1,872	-1,410	-3,616**	-1,521	-2,134	-1,740
Denmark	-1,398	-2,611	-1,002	-0,554	-1,288	-3,581**	-2,419	-1,715
Finland	-2,729	-2,160	-2,654***	-1,177	-1,524	-2,044	-1,813	-2,184*
France	-2,245	-1,009	-1,911	-0,550	-2,002	-2,885*	-2,502	-2,263*
Germany	-0,958	-4,817***	-0,438	-0,292	-1,041	-0,533	-1,544	-0,395
Ireland	-2,523	-2,219	-2,291**	-2,345*	-1,424	-1,850	-2,452	-1,102
Japan	-1,994	-0,998	-1,852	-0,889	-0,720	-0,887	-2,060	-2,086*
Netherlands	-2,021	-1,891	-0,975	-0,517	-0,638	-1,475	-3,454**	-2,515**
New_Zealand	-2,536	-1,330	-1,115	-1,491	-1,230	-3,380**	-4,052***	-0,720
Norway	-2,244	-1,679	0,259	-1,568	-2,295	-2,181	-2,123	-0,929
Spain	-3,496**	-2,183	-0,866	-1,306	-2,143	-3,735**	-2,409	-0,004
Sweden	-1,060	-2,362	-1,453	-1,034	-1,301	-1,450	-2,182	-1,076
Switzerland	-1,648	-1,138	-3,917***	-0,785	-2,505	-1,945	-2,359	-2,463**
United_Kingdom	-2,661	-1,667	-2,179*	-1,286	-2,796	-3,550**	-3,036*	-1,496
DF-GLS trend spec.	trend	trend	constant	constant	trend	trend	trend	constant
Im et al. (2003) test								
With constant	1,620	0,451	**-1,308***	**-5,112****	0,682	-0,103	-2,534***	**-1,448***
	(0,947)	(0,674)	**(0,095)**	**(0,000)**	(0,753)	(0,459)	(0,006)	**(0,074)**
with trend	**-1,822***	**0,473**	0,664	-1,339*	**0,617**	**-0,847**	**-3,790*****	-4,712***
	(0,064)	**(0,682)**	(0,747)	(0,090)	**(0,731)**	**(0,199)**	**(0,000)**	(0,000)

Note: p-values are shown in brackets, ***, **, * indicate that the statistic is significant at 1%, 5%, and 10%, respectively
In the DF-GLS and Im et al. (2003) regression's the lags are chosen by AIC (Akaike information criterion) with a maximum lag of 4.
All variables are expressed in logarithmic form except interest rates and inflation

7.2 Results

In an attempt to identify a reliable long-run relationship among fundamental variables, a number of different model variations of the ARDL approach were estimated for the sample of 17 countries separately. The results of these estimations are found in Table A1 in the appendix.

Testing for cointegration based on the estimates of Model 1 (*dpi, ir*) and Model 2 (*dpi, ir, credit*), we find an equilibrium relationship for 5 and 8 countries respectively, out of the 17 in our sample. While the long-run coefficients for countries with the finding of cointegration are along the right lines in these two models, the size of the coefficients – especially in the case of disposable income – are below the levels found in existing literature, and the significance of the long-run coefficients are weak and sporadic.

For Model 3 (*dpi, ir, gfcf*) the PSS (2001) test finds cointegration in 11 countries, however, the formulation of this model is arguably inappropriate as it violates the PSS (2001) weak exogeneity assumption of no reverse causality from house prices in the direction of investment. This is evidenced by the positive coefficients estimated for investment. Although, in the short-run, investment is expected to have a positive impact on house prices due to a perceived increase in the demand for new housing, in the long-run, higher rates of investment lead to more housing stock, thereby lowering house prices. Therefore, the model seems to identify the reverse causality of higher house prices improving the profitability of property development, leading to more investment.

Models 4 (*dpi, ir, tax ratio*) and 5 (*dpi, ir, unemployment*) find cointegration for 9 and 13 countries, respectively. For the common variables of disposable income and interest rates, when their coefficient is significant at the 10% level in both models for a given country, the size of the coefficient is broadly similar. The coefficient on income is highly significant in both models with elasticities typically between 1 and 2 for countries with cointegration. In Model 5, unemployment and interest rate are found to be significant in 5 and 4 countries, respectively. Meanwhile, in Model 4, the coefficients on tax ratio were very robustly significant with the expected negative sign in all countries where cointegration was found and the coefficients on interest rate also became significant in six of the nine cointegrating countries. These findings lend support for the theoretical model of the asset pricing approach extended by the user demand quadrant.

However, though not presented in this paper for reasons of brevity, the short-run coefficients of these ARDL models are inconclusive, with sporadic signs and significance and weak support for economic intuition. Overall, we conclude that while the ARDL method is able to adequately estimate the long-run equilibrium relationship of a single-country, it cannot account for the short-term dynamics. Given the simultaneous estimation of the two, the long-run coefficients are expected to be biased in small samples, despite being around their true locus. Without a specific economic understanding of the country-specific housing market – and thereby careful selection of variables and lag orders –, simply increasing the frequency of the data is not expected to remedy this issue, nor is extending the time span considering the Lucas critique (1976).

In order to gain robustness, the two most promising models (4 and 5) and their combination were also estimated using the Pooled Mean Group by Pesaran et al. (1999). Considering that the PMG method restricts the long-run coefficients to be homogeneous across countries, we carefully consider which countries to include in the panel stage. We only include countries where cointegration was found by the corresponding ARDL approach, and make further selection based on the observed ARDL results. In the unemployment model, Japan stands out for having a positive coefficient on interest rates in both models and an elasticity of 0.58 for disposable income. This could be explained by the events following the 1989 Japanese stock market crash. Specifically, banks were nationalised and one of the first international instances of Quantitative Easing, an unconventional expansionary monetary policy tool, were introduced in the 1990s, thus making Japan an outlier. Ireland and Norway have an estimated income elasticity of around 0.8, which are also considerably below the group average.[15]

For the four core panel models we eventually estimate, the results of the Westerlund panel cointegration tests are reported at the bottom of Table 3. There's strong evidence of cointegration at the 5% significance level for both the unemployment and the tax ratio formulation. Combining the two models into one also gives some evidence of cointegration, but not as reliable as the other two models. Nonetheless, we ran regressions using all three models. For the lag selection of the panel estimation, we adopt the methodology used by Ludwig and Slok (2004), by taking averages of the country specific ARDL lags for each variable.

The first differences of credit and inflation were also included as exogenous variables to reflect a more comprehensive model. Given that the panel estimation restricts the lag structure to be the same across countries, we set the lag of the exogenous variables at 1. An alternative approach is to run the long-run estimation in a first step and run single-country stationary ARDL estimations using the resulting error correction term in a second step. However, considering that the essence of the PSS method is the simultaneous determination of the long-run and short-run coefficients, this approach is abandoned.

The results of the PECM estimations are reported in Table 3. We observe that following the inclusion of credit, the long-run coefficient on the tax ratio becomes insignificant. Considering the instability of the tax coefficient in response to changes in the model specification, while the unemployment model retains its stability with estimated coefficients strongly in accordance

[15] New Zealand was also excluded based on inspection of the PMG results pointing to outlier behaviour.

with economic theory, we proceed with the results of the extended unemployment estimation (PMG Model 2) and their evaluation in the following section.

Table 3. Long-run coefficient results of the Pooled Mean Group estimation by Pesaran, Shin and Smith (1999) and Westerlund (2007) cointegration test results

	PMG Model 1	PMG Model 2	PMG Model 3	PMG Model 4	PMG Model 5
Specification	LR: DPI, IR, un	LR: DPI, IR, un SR: credit, inflation	LR: DPI, IR, tax	LR: DPI, IR, tax SR: credit, inflation	LR: DPI, IR, un, tax
DPI	1,782*** (0,000)	2,018*** (0,000)	1,778*** (0,000)	2,014*** (0,000)	1,935*** (0,000)
IR	-0,037*** (0,000)	-0,024*** (0,000)	-0,017*** (0,000)	-0,027*** (0,003)	-0,037*** (0,000)
un	-0,153*** (0,002)	-0,157*** (0,002)			-0,136*** (0,000)
tax			-0,260*** (0,000)	0,014 (0,681)	-0,052*** (0,000)
Westerlund (2007)					
Gt	-2,085[2]*** (0.019)	-2,085*** (0.019)	-4,467*** (0,000)	-4,467*** (0,000)	-2,705*** (0,003)
Pt	-1,855*** (0.032)	-1,855*** (0.032)	-2,260*** (0,012)	-2,260*** (0,012)	-1,292* (0,098)
Countries	AU, BE, CA, FR, FI, IT, NL, ES, SE, UK	AU, BE, CA, FR, FI, IT, NL, ES, SE, UK	BE, FR, FI, NL, ES, SE	BE, FR, FI, NL, ES, SE	BE, FR, FI, NL, ES, SE
	Note: p-values are shown in brackets, ***, **, * indicate that the statistic is significant at 1%, 5%, and 10%, respectively [2] z-value statistic from Westerlund (2007)				

8. Evaluation

8.1 Long-run equilibrium

The Pooled Mean Group panel estimated coefficients of the long-run forcing variables – disposable income, interest rates, unemployment – are found in column 2 of Table 3. The country-specific short-run coefficients are presented in Table 4, along with the average short-run coefficients in the final column. Essentially, all coefficients follow economic intuition – in terms of size, sign, and significance level – and fall within the range of findings from prior literature. To aid our interpretation, Figures 2, 3, and 4 are also reported. Figure 2 shows the estimated and actual first-difference of house prices, while Figure 3 breaks the estimated first-difference of house prices into its components. Figure 4 illustrates the simulated equilibrium and actual price level of the aggregated panel.

In the selected group of developed countries, we find a pooled long-run coefficient of 2.01 on disposable income, meaning that a 1% increase in disposable income is expected to raise the equilibrium price by around 2%. As expected, as households become wealthier, their share of income available for non-basic spending (on top of food, utilities, transport, minimum housing etc.) increases, which is then largely captured by the housing market due to the significant market power of owners of housing. Adams and Füss (2010) estimate an average elasticity of 1.53% for income, while Meen (2002) estimates 2.5 after including housing stock.

Further, our estimated coefficients on interest rates and unemployment are -0.023 and -0.157, respectively. Given that interest rates are nominated in percentages, not logs, this means that a 1 percentage point increase in interest rates results in a 0.02% decrease in the equilibrium long-run house prices. As interest rates fall, the financing of a housing purchase becomes cheaper, while the return on alternative investments falls, leading to an increase in housing demand. This result agrees with the findings of 0.02% from Annett (2005) and 0.035% from Adams and Füss (2010). The finding that a 1% increase in unemployment leads to a 12% decrease in house prices indicates that not only the mean level of income, but the distribution of financial stability in the form of permanent income expectation is also an important determinant of housing demand. Schnure (2004) also includes unemployment in their estimation of the US market and finds an elasticity of 1% using levels, which corresponds to 10-20% elasticity in logs, aligning with our results.

8.2 Short-run dynamics

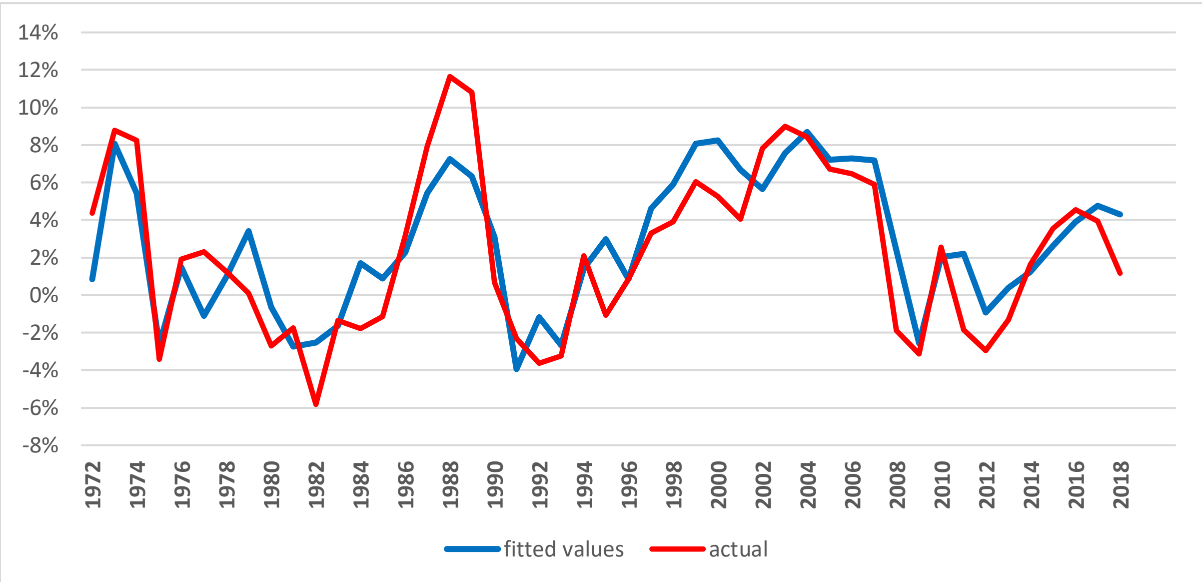
Table 4. Single-country and aggregate short-run coefficient results of the Pooled Mean Group estimation by Pesaran, Shin and Smith (1999)

	Australia	Belgium	Canada	Finland	France	Italy	Netherlands	Spain	Sweden	UK	Group
ECT	-0,118***	-0,206***	-0,063**	-0,057*	-0,143***	-0,073	-0,180***	-0,161***	-0,163***	-0,123***	-0,129***
	(0,001)	(0,000)	(0,032)	(0,092)	(0,000)	(0,210)	(0,001)	(0,000)	(0,000)	(0,006)	(0,000)
ΔHP_{-1}	0,175	0,207	0,355***	0,484***	0,722***	0,089	0,630***	0,458***	0,264**	0,478***	0,386***
	(0,262)	(0,101)	(0,009)	(0,003)	(0,000)	(0,528)	(0,000)	(0,000)	(0,050)	(0,000)	(0,000)
ΔHP_{-2}	-0,138	-0,103	0,093	-0,230	-0,230**	0,185*	-0,046	0,212**	0,158	-0,064	-0,016
	(0,329)	(0,219)	(0,417)	(0,131)	(0,023)	(0,088)	(0,760)	(0,048)	(0,243)	(0,483)	(0,757)
ΔDPI_{-1}	-0,360	0,393**	0,122	1,074***	0,441	0,578	0,062	1,137**	0,802***	1,582***	0,583***
	(0,337)	(0,022)	(0,741)	(0,003)	(0,110)	(0,208)	(0,853)	(0,010)	(0,009)	(0,000)	(0,001)
ΔIR	0,004	0,010***	-0,004	-0,005	0,001	0,021***	0,005	0,012**	0,001	0,007	0,005**
	(0,557)	(0,000)	(0,526)	(0,430)	(0,739)	(0,003)	(0,178)	(0,010)	(0,721)	(0,223)	(0,034)
Δun	-0,151*	0,026	-0,228**	0,049	0,024	-0,178	-0,066	-0,164*	-0,057	-0,048	-0,079***
	(0,059)	(0,403)	(0,014)	(0,587)	(0,700)	(0,244)	(0,180)	(0,051)	(0,388)	(0,564)	(0,010)
Δun_{-1}	0,135*	0,056*	0,145**	-0,048	0,135**	-0,036	0,056	0,236***	0,030	0,077	0,078***
	(0,054)	(0,062)	(0,045)	(0,448)	(0,014)	(0,792)	(0,187)	(0,001)	(0,532)	(0,314)	(0,004)
Δinf	-0,002	0,004	0,002	-0,006	0,004	0,024***	-0,004	-0,015**	-0,002	0,004	0,001
	(0,785)	(0,118)	(0,786)	(0,354)	(0,300)	(0,000)	(0,578)	(0,011)	(0,528)	(0,296)	(0,781)
$\Delta credit$	0,646***	0,525***	0,330*	0,038	0,625***	-0,076	0,551***	0,101	0,253	0,362*	0,336***
	(0,006)	(0,000)	(0,060)	(0,705)	(0,000)	(0,668)	(0,005)	(0,456)	(0,104)	(0,087)	(0,000)
cons	-0,491***	-0,877***	-0,250**	-0,243*	-0,637***	-0,298	-0,830***	-0,720***	-0,700***	-0,555***	-0,560***
	(0,001)	(0,000)	(0,034)	(0,085)	(0,000)	(0,227)	(0,002)	(0,000)	(0,000)	(0,002)	(0,000)

Note: p-values are shown in brackets, ***, **, * indicate that the statistic is significant at 1%, 5%, and 10%, respectively
Results that were insignificant are shown in lighter colour for clearer overview.

The error correction term stands for the difference between the estimated equilibrium house price and the actual house price we observe. Therefore, in the presence of an equilibrium relationship, the coefficient on the ECT, also called the adjustment parameter, is expected to be negative and significant as an indicator of mean reversion. For every country, except Italy, we find a negative adjustment parameter that is highly significant at the 1% level. It is argued that what other studies, such as Ott (2014), pick up as a short-run effect of housing stock, our model picks up in the ECT term. Fastest error correction is found in Belgium, Sweden, and Netherlands, such that the correction following a deviation from the equilibrium is estimated to take 5-6 years, which indicates a more dynamic housing market. Slowest error correction is found in Finland and Canada, where approximate full correction is expected to take 14-16 years. The average adjustment period is around 8 years for the entire panel, which falls between the findings of 7 years by Annett (2005) and 14 years of Adams and Füss (2010) and is responsible for around a fourth of the total short-run dynamics. Finally, from Figure 3, we observe an apparent structural break around 1992, at which point the ECT reversed from being generally negative to being generally positive. This finding could be explained by the Maastricht treaty of 1992, i.e., expectations of more favourable credit conditions since the eurozone's inception.

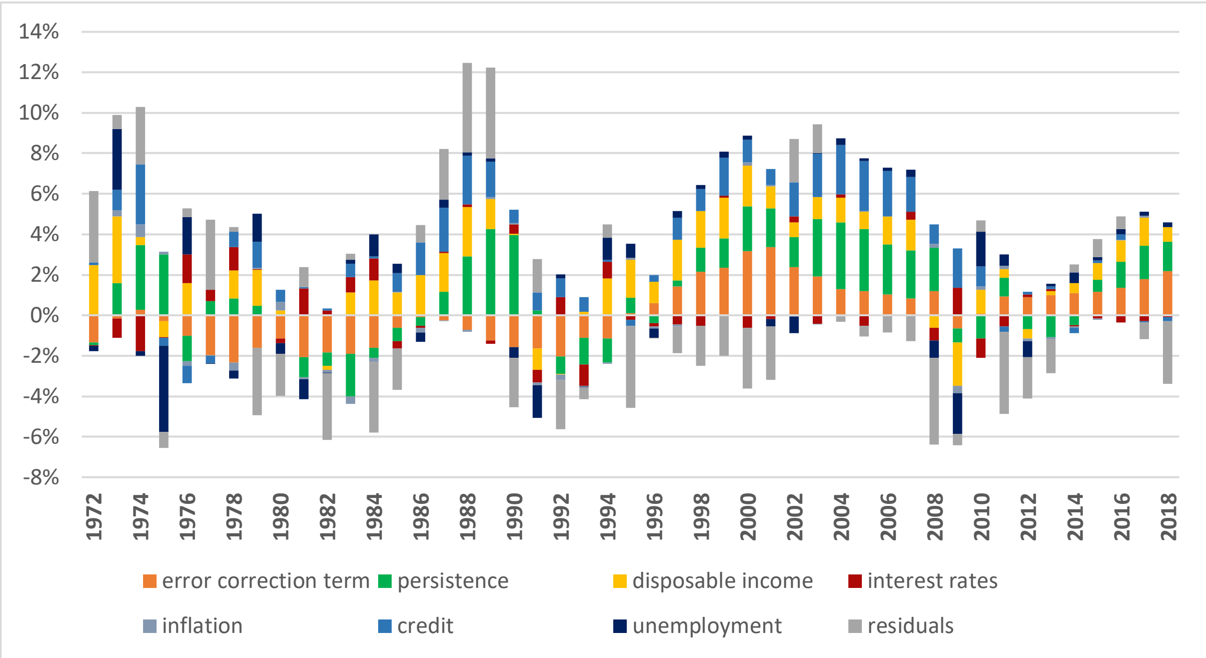
Figure 2: Estimated and actual mean changes in house prices



The short-run coefficients with the greatest impact in addition to the adjustment parameters, are on the disposable income, the lagged difference in house prices, and the ratio of credit to GDP. The coefficient on DPI is positive in all countries where it is significant at the 10% level.

As an average consumer's disposable income increases, their financing capacity improves, and so they can meet banks' terms for higher mortgage loans in a continuous effort to climb the "property ladder".[16] In the opposite direction, a fall in income leads to an inability to finance mortgages especially in the case of liquidity constrained households.

Figure 3. Component break-down of average estimated short-run change in house prices



Note: Components were calculated by factoring the mean change in the given variable with its mean short-run coefficient.

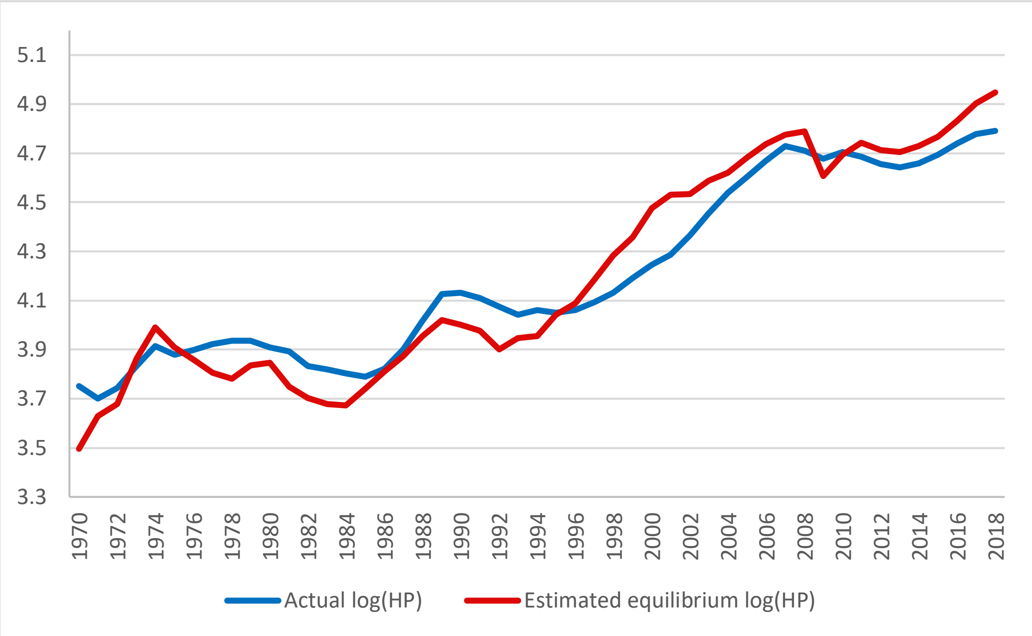
The coefficient on credit is significant in seven of the ten countries in the sample, with uniformly positive coefficients. Figure 3 demonstrates that leading up to all three housing busts in our sample (1973 first oil crisis, 1990 recession, 2008 subprime crisis) the role of credit was increasing and had significant impact relative to the total short-run effect. This result corroborates the findings of Schularick and Taylor (2012) who find, based on 140 years of data, that lagged credit growth is a highly significant predictor of financial crises. Built around this, the new macroprudential view, proposed by Borio (2003), is that credit bubbles need to be pricked before they get too big, to short-circuit a vicious cycle. While the amount of credit has been stable following the great financial crisis, it is argued to be due to banks' balance sheets being constrained by bad credit from pre-2008; banks are therefore still subject to instability[17]

[16] Simon and Stone (2017)

[17] Gertler and Karadi (2011) present the macroeconomic conditions in the presence of balance sheet constrained private banks and the role of unconventional monetary policy in dealing with it

From Figure 4, we see that actual house prices are estimated to be undervalued throughout the period of the 2008 global financial crisis. While this could be a result of misspecification, we still find that there was a strong catch-up of prices to the estimated equilibrium, in large part driven by persistence, as can been seen from Figure 3. Further, it is argued that the 2008 housing crisis could come about as a result of an unsustainable equilibrium price propping up real house prices, thereby resulting in a rational bubble. In general, we observe from Figure 3 that leading up to all three housing busts in our sample, the role of persistence was increasing, which provides some evidence that speculative behaviour had a role in propping up prices. The persistence coefficient is positive and significant in seven of the ten countries.

Figure 4. Actual house price level and estimated equilibrium price level



Inflation is found to be insignificant in most of the countries in the short-run and has negligible effect on average. This could be due to a combination of factors: housing could be a considered a hedge against high inflation pushing prices up, while mortgage rates, especially for adjustable mortgages, go up in response to higher inflation, the two forces cancelling out. Inflation has also been low and stable in our sample period. However, given the current conditions of an emerging energy crisis and Russia's invasion of Ukraine, this result may reverse as inflation is reaching new heights. Unemployment was also found to have limited effect in the short run, which could be due to the slow build-up of financial security after finding a job.

Lastly, by looking at the individual equilibrium price levels of the countries, our model appears to overestimate the equilibrium price level following the 2008 Great Financial Crisis. This is arguably due to the introduction of Quantitative Easing measures that were instituted in the presence of the zero lower bound post-2010. However, QE was unable to fully substitute for the effects of lower interest rates, resulting in diminished house prices.

9. Conclusion

In summary, this paper aimed at identifying a robust equilibrium relationship between house prices and macroeconomic fundamentals by testing for the presence of cointegration that is stable across a substantial number of developed countries. Using the ARDL bounds testing approach developed by Pesaran, Shin and Smith (2001), we found disposable income, interest rates and unemployment to be long-run forcing variables of housing prices across 13 of the 17 OECD countries in our sample, with income being the main determinant. Subsequently applying the Pooled Mean Group estimation on a subset of countries with evidence of cointegration, we derive more robust estimates for the long-run coefficients and get country-specific short-run results that are – unlike those from the ARDL step – predominantly in line with economic intuition and previous work on the topic. In addition to disposable income, credit availability is found to be an important determinant of short-run dynamics, especially preceding housing busts. There is some further evidence that the proliferation of unconventional monetary policy after 2008 may have caused a structural shift that needs to be factored into new models. Since the role of unemployment has been somewhat under-researched in the literature, this paper contributes to filling that hole.

There are, however, several shortcomings to our approach that limit a more nuanced analysis of the housing market dynamics. First, we were evidently limited by the lack of housing stock data available, eliminating the possibility of an important benchmark model. Further, the ARDL technique makes significant identifying exogeneity assumptions especially on the relationship of house prices with income and credit that could introduce bias in the estimates. Therefore, sophisticated systems approaches with better data availability could be valuable tools for a more comprehensive understanding of the causal relationships. Lastly, it is expected that there are significant nonlinearities in housing markets and business cycles in general, such as the assumed downward stickiness of house price, or the asymmetry between the generation and deterioration of financial stability. Methods, such as the non-linear ARDL cointegration technique by Shin, Yu and Greenwood-Nimmo (2014) could potentially account for some of these issues.

Appendix:

Table A1: Long-run coefficient estimates and F-stats and t-stats of ARDL bounds test, with outcome of cointegration decision

	Australia	Belgium	Canada	Denmark	Finland	France	Germany	Ireland	Italy	Japan	Netherlands	New Zealand	Norway	Spain	Sweden	Switzerland	UK
Model 1: Disposable income, Interest rates																	
adj	-0,304***	-0,260***	-0,130**	-0,092*	-0,272***	-0,085**	0,134**	-0,128***	-0,300***	-0,097***	-0,104**	-0,396***	-0,271***	-0,117	-0,150***	-0,092**	-0,033
DPI	1,626***	1,316***	1,548***	0,690	0,874***	1,681***	-0,168	0,024	0,788***	0,317*	1,100***	2,051***	0,929***	1,920***	1,457***	0,342	2,623
IR	-0,014**	-0,062***	-0,051**	-0,119*	-0,007	-0,059***	0,075	-0,228**	-0,019	0,068***	-0,099**	-0,008*	-0,092***	-0,028	-0,083***	0,082	-0,038
F-stat	0,014	0,000	0,272	0,181	0,007	0,284	0,001	0,011	0,006	0,000	0,145	0,000	0,000	0,586	0,183	0,156	0,939
p-stat	0,010	0,000	0,351	0,578	0,004	0,286	1,000	0,190	0,013	0,012	0,303	0,000	0,000	0,714	0,180	0,458	0,933
Coint.	1	1	0	0	1	0	0	0	1	1	0	1	1	0	0	0	0
Model 2: Disposable income, Interest rates, credit																	
adj	-0,161*	-0,230***	0,016		-0,301***	-0,272***			-0,255***	-0,155***	-0,137*		-0,405***	-0,349***	-0,277***		-0,376***
DPI	5,621	2,197***	-11,858		0,702***	1,001**			0,656**	-0,261	0,502		0,856***	1,024***	0,799***		0,261
IR	-0,016	-0,022**	0,394		-0,011	-0,042***			-0,009	0,011	-0,012		-0,061***	0,021	-0,027		-0,076***
credit	-2,052	-0,146	7,247		0,125	0,361*			-0,023	1,088***	0,771		0,388**	0,495***	0,873***		1,158***
F-stat	0,031	0,000	0,156		0,026	0,011			0,007	0,000	0,724		0,000	0,029	0,001		0,004
p-stat	0,659	0,000	0,989		0,031	0,027			0,029	0,034	0,574		0,001	0,010	0,064		0,011
Coint.	0	1	0		1	1			1	1	0		1	1	1		1
Model 3: Disposable income, Interest rates, GFCF																	
adj	-0,240***	-0,089***	-0,022	-0,265***	-0,199***	-0,175***		-0,275***	-0,308***	-0,171***	-0,252***	-0,363***	-0,333***	-0,249***	-0,349***		-0,035
DPI	1,740***	4,231**	-6,523	1,018***	0,761***	1,518***		1,223***	0,732***	0,313***	1,293***	1,893***	0,562***	1,332***	1,330***		2,373
IR	-0,012	-0,034*	0,067	-0,011	0,009	-0,015		-0,006	-0,017	0,042***	-0,055***	-0,004	-0,027	0,037*	-0,081***		-0,033
GFCF	-0,117	-2,313*	13,388	0,738***	0,275	1,998***		0,362***	0,398	0,539***	0,992***	0,335**	0,813***	0,501**	0,055		0,588
F-stat	0,067	0,000	0,001	0,013	0,120	0,075		0,005	0,010	0,000	0,000	0,000	0,053	0,007	0,050		0,979
p-stat	0,040	0,255	0,932	0,045	0,129	0,028		0,031	0,017	0,006	0,000	0,000	0,039	0,003	0,027		0,948
Coint.	1	0	0	1	0	1		1	1	1	1	1	1	1	1		0
Model 4: Disposable income, Interest rates, Tax ratio																	
adj	-0,131***	-0,273***	-0,110**	-0,399***	-0,353***	-0,118***		-0,177		-0,200***	-0,282***	-0,363***	-0,325***	-0,356***	-0,297***		-0,059
DPI	1,318***	1,139***	1,310***	0,633***	1,528***	2,059***		0,421		1,312***	1,689***	1,548***	0,736***	1,819***	1,850***		1,540*
IR	0,004	-0,059***	0,020	-0,054***	-0,037***	-0,035*		-0,112		0,034***	-0,003	-0,010**	-0,097***	-0,002	-0,033**		-0,026
tax	-0,659*	0,088*	-1,115***	-0,896***	-0,281***	-0,350*		-0,678***		-0,839***	-0,394***	-0,391***	-0,186*	-0,418***	-0,070***		-0,661
F-stat	0,415	0,000	0,569	0,002	0,001	0,168		0,034		0,000	0,001	0,000	0,000	0,029	0,007		0,963
p-stat	0,224	0,000	0,494	0,001	0,010	0,127		0,648		0,006	0,000	0,000	0,000	0,059	0,003		0,915
Coint.	0	1	0	1	1	0		0		1	1	1	1	1	1		0
Model 5: Disposable income, Interest rates, Unemployment																	
adj	-0,280***	-0,246***	-0,175***	-0,061	-0,287***	-0,131***	-0,045	-0,254***	-0,257***	-0,119***	-0,187***	-0,310***	-0,279***	-0,263***	-0,186***	-0,060	-0,031
DPI	1,637***	1,553***	1,034***	-0,294	1,370***	2,120***	-0,085	0,791***	1,516***	0,581***	1,584***	2,067***	0,911***	2,064***	1,443***	7,614	2,134*
IR	0,003	-0,055***	-0,044*	-0,249	0,017*	-0,015	-0,029	0,015	-0,006	0,069***	-0,041	-0,010	-0,092***	0,034*	-0,055***	0,031	0,125
un	-0,175	-0,084	-0,411	0,664	-0,349***	-0,365*	-2,061	-0,637***	-0,717***	-0,365**	-0,015	-0,021	-0,051	-0,328**	-0,082	-1,018	-1,888
F-stat	0,065	0,000	0,030	0,208	0,005	0,061	0,000	0,006	0,006	0,000	0,022	0,001	0,000	0,082	0,039	0,002	0,791
p-stat	0,045	0,000	0,084	0,851	0,005	0,029	0,870	0,033	0,058	0,004	0,028	0,000	0,000	0,031	0,049	0,758	0,952
Coint.	1	1	1	0	1	1	0	1	1	1	1	1	1	1	1	0	0

Note: ***, **, * indicate that the statistic is significant at 1%, 5%, and 10%, respectively. In the Cointegration line, 1 indicates that the null of no cointegration was rejected, while 0 is that it couldn't be rejected.
Empty columns mean that data availability was too weak to run estimation

References

Abraham, J.M. and Hendershott, P.H., 1994. Bubbles in metropolitan housing markets.

Adams, Z. and Füss, R., 2010. Macroeconomic determinants of international housing markets. *Journal of Housing Economics*, *19*(1), pp.38-50.

Andrews, D., 2010. Real house prices in OECD countries: The role of demand shocks and structural and policy factors.

AMECO available at:
http://ec.europa.eu/economy_finance/db_indicators/ameco/index_en.htm

Arestis, P. and González, A.R., 2014. Modelling the housing market in OECD countries. *International Review of Applied Economics*, *28*(2), pp.131-153.

Blanchard, O.J. and Watson, M.W., 1982. Bubbles, rational expectations and financial markets.

Borio, C., 2003. Towards a macroprudential framework for financial supervision and regulation?. *CESifo Economic Studies*, *49*(2), pp.181-215.

Brunner, K. and Meltzer, A., 1983. Econometric policy evaluation: A critique. In *Theory, Policy, Institutions: Papers from the Carnegie-Rochester Conferences on Public Policy* (Vol. 1, p. 257). North Holland.

Brunnermeier, M.K., Garicano, L., Lane, P.R., Pagano, M., Reis, R., Santos, T., Thesmar, D., Van Nieuwerburgh, S. and Vayanos, D., 2016. The sovereign-bank diabolic loop and ESBies. *American Economic Review*, *106*(5), pp.508-12.

Campbell, J.Y. and Perron, P., 1991. Pitfalls and opportunities: what macroeconomists should know about unit roots. *NBER macroeconomics annual*, *6*, pp.141-201.

Capozza, D.R., Hendershott, P.H., Mack, C. and Mayer, C.J., 2002. Determinants of real house price dynamics.

Colwell, P.F., 2002. Tweaking the DiPasquale-wheaton model. *Journal of Housing Economics*, *11*(1), pp.24-39.

Demyanyk, Y. and Van Hemert, O., 2011. Understanding the subprime mortgage crisis. *The review of financial studies*, *24*(6), pp.1848-1880.

DiPasquale, D. and Wheaton, W.C., 1996. *Urban economics and real estate markets* (Vol. 23, No. 7). Englewood Cliffs, NJ: Prentice Hall.

Égert, B. and Mihaljek, D., 2007. Determinants of house prices in central and eastern Europe. *Comparative economic studies*, *49*(3), pp.367-388.

Elliott, G., Rothenberg, T.J. and Stock, J.H., 1992. Efficient tests for an autoregressive unit root.

Engle, R.F. and Granger, C.W., 1987. Co-integration and error correction: representation, estimation, and testing. *Econometrica: journal of the Econometric Society*, pp.251-276.

Evans, G.W., 1991. Pitfalls in testing for explosive bubbles in asset prices. *The American Economic Review*, *81*(4), pp.922-930.

Girouard, N., Kennedy, M., Van Den Noord, P. and André, C., 2006. Recent house price developments: the role of fundamentals.

Im, K.S., Pesaran, M.H. and Shin, Y., 2003. Testing for unit roots in heterogeneous panels. *Journal of econometrics*, *115*(1), pp.53-74.

Euro Area Policies: Selected Issues. *IMF Staff Country Reports* 2005, 266, A003

Gallin, J., 2006. The long-run relationship between house prices and income: evidence from local housing markets. *Real Estate Economics*, *34*(3), pp.417-438.

Gertler, M. and Karadi, P., 2011. A model of unconventional monetary policy. *Journal of monetary Economics*, *58*(1), pp.17-34.

Jacobsen, D.H. and Naug, B.E., 2005. What drives house prices?.

Johansen, S., 1988. Statistical analysis of cointegration vectors. *Journal of economic dynamics and control*, *12*(2-3), pp.231-254.

Lambiri, D. and Rovolis, A., 2014. Real estate, and housing markets. *Handbook of regional science*, pp.125-145.

Ludwig, A. and Sløk, T., 2004. The relationship between stock prices, house prices and consumption in OECD countries. *Topics in Macroeconomics*, *4*(1).

Malpezzi, S., 1999. A simple error correction model of house prices. *Journal of housing economics*, *8*(1), pp.27-62.

Meen, G.P., 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), pp.1-23.

Meen, G., 2002. The time-series behavior of house prices: a transatlantic divide?. *Journal of housing economics*, *11*(1), pp.1-23.

Narayan, P.K., 2005. The saving and investment nexus for China: evidence from cointegration tests. *Applied economics*, *37*(17), pp.1979-1990.

OECD database available at: https://stats.oecd.org/

Ott, H., 2014. Will euro area house prices sharply decrease?. *Economic Modelling*, *42*, pp.116-127.

Pedroni, P., 2004. Panel cointegration: asymptotic and finite sample properties of pooled time series tests with an application to the PPP hypothesis. *Econometric theory*, *20*(3), pp.597-625.

Pesaran, M.H., Shin, Y. and Smith, R.P., 1999. Pooled mean group estimation of dynamic heterogeneous panels. *Journal of the American statistical Association*, *94*(446), pp.621-634.

Pesaran, M.H., Shin, Y. and Smith, R.J., 2001. Bounds testing approaches to the analysis of level relationships. *Journal of applied econometrics*, *16*(3), pp.289-326.

Philiponnet, N. and Turrini, A., 2017. *Assessing house price developments in the EU* (No. 048). Directorate General Economic and Financial Affairs (DG ECFIN), European Commission.

Sarantis, N. and Stewart, C., 2003. Liquidity constraints, precautionary saving and aggregate consumption: an international comparison. *Economic Modelling*, *20*(6), pp.1151-1173.

Schnure, C., 2005. Boom-bust cycles in housing: The changing role of financial structure.

Schularick, M. and Taylor, A.M., 2012. Credit booms gone bust: Monetary policy, leverage cycles, and financial crises, 1870-2008. *American Economic Review*, *102*(2), pp.1029-61.

Shin, Y., Yu, B. and Greenwood-Nimmo, M., 2014. Modelling asymmetric cointegration and dynamic multipliers in a nonlinear ARDL framework. In *Festschrift in honor of Peter Schmidt* (pp. 281-314). Springer, New York, NY.

Simon, J. and Stone, T., 2017. *The property ladder after the financial crisis: The first step is a stretch but those who make it are doing oK*. Economic Research Department, Reserve Bank of Australia.

Stock, J.H. and Watson, M.W., 1993. A simple estimator of cointegrating vectors in higher order integrated systems. *Econometrica: journal of the Econometric Society*, pp.783-820.

Tsatsaronis, K. and Zhu, H., 2004. What drives housing price dynamics: cross-country evidence. *BIS Quarterly Review, March*.

Westerlund, J., 2007. Testing for error correction in panel data. *Oxford Bulletin of Economics and statistics*, *69*(6), pp.709-748.