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Testing for persistence in housing price-to-income and price-to-rent ratios in 16 OECD countries[$#]

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Abstract

Housing price-to-income and price-to-rent ratios are among the most widely monitored indicators of housing market conditions. While these ratios tend to fluctuate around a constant level or a mild trend over the long term, they also tend to deviate from these benchmarks for protracted periods. Traditional unit root tests often indicate the presence of a unit root. This paper uses the framework of fractional integration to test the persistence of price-to-income and price-to-rent ratios in a sample of 16 OECD countries spanning four decades. The results indicate that the ratios are highly persistent. The possibility that persistence estimates may be affected by structural breaks in the series is also considered, but evidence of such breaks is found only in a very limited number of cases. Policy action may be required if high price-to-income and price-to-rent ratios have adverse social and economic consequences. Policies should be guided by a careful analysis of the factors behind high ratios.

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Keywords: Housing; Price-to-income ratio; Price-to-rent ratio; Fractional integration; Persistence; Long memory.

1. Introduction

Ratios of housing prices to income and rents are often used as indicators of overvaluation of housing prices. For example, such ratios are regularly published in *The Economist* and the *OECD Economic Outlook*. Economic commentators often express concern over high price-to-income or price-to-rent ratios. The economic interpretation of these ratios is fairly straightforward. Price-to-income ratios are a measure of the affordability of housing. Increases in housing prices cannot deviate indefinitely from growth in the income of potential buyers. If housing prices outpace income growth, at some point households will no longer be able to afford buying and demand will dry up, bringing prices down. Similarly, households can choose to own or to rent their homes. This choice is affected by the relative levels of prices and rents. High prices relative to rents should push more people into renting, alleviating pressures on housing prices and pushing rents up.

Over the past four decades in OECD countries, the ratios have tended to fluctuate around a constant level or a mild trend (Figure 1). Nevertheless, deviations of price-to-income or price-to-rent ratios from their long-term benchmarks tend to be protracted. Standard unit root tests generally find that these ratios are non-stationary. For example, Girouard *et al.* (2006) find that in a set of 18 OECD countries spanning more than thirty years, the presence of a unit root in the ratios could in most cases not be rejected. Malpezzi (1999), performing a panel unit root test for the price-to-income ratio in 133 US major

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metropolitan areas over the period 1979-1996, could not reject the presence of a unit root. Similarly, Gregoriou *et al.* (2013), using linear and non-linear unit root tests and allowing for structural change, find that UK aggregate and regional house price to earnings ratios are non-stationary over the period 1983Q2-2009Q1. Black *et al.* (2006) find stationarity at the 10% significance level for the aggregate UK price-to-income ratio over the period 1973Q4-2004Q3. Real house prices and real disposable income are cointegrated at the 5% significance level. However, deviations from equilibrium are protracted, reflecting both intrinsic bubbles and house price dynamics driven by momentum trading. Gallin (2006) finds no evidence of cointegration between house prices and *per capita* income in the United States, whether using national level data from 1975 to 2002 or a panel of 95 metropolitan areas from 1978 to 2000. On the contrary, Gallin (2008) finds cointegration between US house prices and rents over the period 1970-2005. But Mikhed and Zemčik (2009a), performing panel unit root tests robust to cross-section correlation, find no cointegration between house price and rents in 23 US metropolitan areas over the period 1978-2006. The result of no cointegration still holds when additional fundamental determinants of house prices are included (Mikhed and Zemčik, 2009b). Taipalus (2006) finds unit roots in price-to-rent ratios in Finland, Germany, Spain, the United Kingdom and the United States over a period starting from 1968 to 1987 depending on the countries examined and ending in 2004.

[Figure 1. Price-to-income and price-to-rent ratios]

Non-stationarity in price-to-income and price-to-rent ratios can be rationalised in several ways. While income is a major determinant of housing demand, other factors also play an important role. In particular, as most households need to borrow to buy houses, mortgage rates and credit conditions have a strong impact on housing demand. The responsiveness of housing supply, often restricted by land-use planning constraints, also affects prices. Hence, there is no simple relation between housing prices and income (Meen, 2008). A simple asset pricing model clearly shows the dependence of the price-to-rent ratio on interest rates and expectations of capital gains on residential assets (Poterba, 1984). Therefore, a permanent shift in interest rates could lead to a level shift in the price-to-rent ratio. A permanent loosening of credit conditions would have the same effect. Furthermore, arbitrage between renting and owning tends to be imperfect, especially because housing units in the two markets differ (Glaeser and Gyourko, 2007).

The fact that the economic literature points to a lack of stationarity in price-to-income and price-to-rent ratios calls for caution in interpreting these ratios in terms of overvaluation or undervaluation of housing prices. Econometric models have the advantage over simple ratios of taking into account all determinants of prices (provided the model is well specified). However, they have their own weaknesses. Notably, fundamentals explaining housing prices may be unsustainable, leading to the deceptive impression that prices in line with fundamentals are not vulnerable to sharp falls. Such unsustainable fundamentals may include levels of income, interest rates and the architecture of credit (Muellbauer, 2012). Shiller (2006) notes that fundamentals often cited in support of confident assessments of the housing market are surprisingly weak at explaining historical prices.

While price-to-income and price-to-rent ratios may not be very good indicators of the state of the housing market in the short term, they nevertheless seem useful as long-term benchmarks. As shown in Figure 1, over the period from the 1970 to 2012, ratios usually ended up reverting to their long-term average. Friggit (2008) shows that secular home price indices for Paris, Norway's main cities and the United States move broadly in line with income per household over the long term. Shiller (2006), however, shows that the price-to-rent ratio in the United States has trended up since 1913. Nevertheless, the trend is mild over recent decades and the recent crisis has brought back US price-to-income and price-to-rent ratios close or even below their 40-year average.

Overall, price-to-income and price-to-rent ratios provide a useful long-term perspective on housing market conditions. Nevertheless, they deviate for protracted periods from their long-term level, calling for a careful investigation of their persistence. As mentioned above, a number of studies have tested the presence of a unit root in these ratios and found that in most cases this hypothesis could not be rejected. The approach can be extended to gain further information about persistence, or "long memory" in time series, by using the more general statistical framework of fractional integration (Caporale and Gil-Alana, 2007 and 2008; Gil-Alana *et al*., forthcoming). Traditional integration analysis tests for stationarity in series differentiated d times, with d an integer value. Fractional integration allows estimating the value of d necessary to get a stationary I(0) series, which may be fractional. A value below one indicates mean-reversion, implying that exogenous shocks have temporary effects, while a value equal to or above one implies that exogenous shocks have permanent effects.

A difficulty with the analysis of "long memory" processes is that empirically they can easily be confused with regime switching processes (Diebold and Inoue, 2001; Granger and Hyung, 2004). This is particularly relevant in housing markets, as structural changes have taken place over the past 30 years in the countries included in the sample. These changes include deregulation of mortgage markets, changes in rental market regulations, property taxation and land-use planning rules, which may generate breaks in the price-to-income and price-to-rent series. For example, if mortgage market deregulation leads to permanently lower mortgage rates, it would be expected to shift the ratios to higher levels permanently. The literature provides evidence that structural changes can affect house price dynamics. Cook and Vougas (2009) find structural change in UK house prices and show that contrary to traditional unit root tests smooth transition-momentum threshold autoregressive (ST-MTAR) tests reject the presence of a unit root in UK house prices. Ganoulis and Giuliodori (2011) find that financial liberalisation in Europe has resulted in a "de-linking" of short-term house price dynamics from income and argue that "the housing market may have become more similar to a financial asset market, with interest rates and expectations of capital gains playing a more prominent role". Canarella *et al*. (2012) find structural breaks in the rate of capital gain from the sales of houses in the United States. Hence, the analysis of fractional integration needs to be complemented by searching for breaks in the series.

A number of papers have applied the fractional integration framework to housing prices. Gil-Alana *et al.* (forthcoming) test for persistence, breaks and outliers in South African house prices, finding persistence, though not in all market segments. Barros *et al.* (2012) investigate the relationship between US house prices at state and national level, finding a wide range of values for the order of integration d across states and no co-integration between states or between states and national indices. Osterrieder and Schotman (2011) test for fractional integration in rent-to-price ratios in a neighbourhood of Amsterdam from 1650 to 2005. They find a fractional integration order $d = 0.75$, indicating that the series, while non-stationary is mean reverting.

The present paper adds to the literature by testing for fractional integration in housing price-to-income and price-to rent ratios in 16 OECD countries over the past four decades. It is organised as follows: Section 2 and 3 describe, respectively, the dataset and the methodology; Section 4 presents the results and Section 5 concludes.

2. Data

This paper uses a set of quarterly data compiled by the Economics Department of the OECD, covering the period from 1970 to 2011 for most countries. For some countries, data are available over a slightly shorter period. However, the shortest period covers 26 years, which is sufficient to span several housing cycles. As there is no international harmonised dataset of housing prices, series have been selected among various available national data sources, in most cases government bodies (Table 1). The methodologies and the coverage of these series vary widely. Series differ in terms of transaction mix and quality adjustment. An average or median price index is affected by the share of various types of homes in transactions. To overcome this problem, mix-adjusted, repeat-sales or hedonic indices are produced in some countries. Coverage varies from most transactions in the country to selected transactions (*e.g.* certain types of dwelling, homes financed through conventional mortgages) or metropolitan areas. It is also worth noting that, as long-term country-wide housing price series are not available for Japan, an urban land price index is used as a proxy. Despite differences in methodology and coverage, the series are thought to be the most representative of each national market for existing homes and are among the most closely monitored by policymakers. Price-to-income ratios are calculated by dividing housing prices by the average household disposable income *per capita* from the national accounts data compiled in the OECD Economic Outlook database. Price-to-rent ratios are computed by dividing housing prices by the rent component of the consumer price index extracted from the OECD Main Economic Indicators dataset.

[Table 1. Housing price series entering price-to-income and price-to-rent ratios]

3. Methodology

Denoting any of the series by y_t we describe its behaviour through the following model:

$$y_t = \beta^T z_t + x_t, \quad t = 1, 2, \ldots, \tag{1}$$

where z_t is a (kx1) vector of deterministic terms that may be an intercept (z_t = 1) or an intercept with a linear trend (i.e., $z_t = (1,t)^T$), and x_t are the regression errors, which follow an I(*d*) model of the form:

$$(1 - L)^d x_t = u_t, \quad t = 1, 2, \ldots \tag{2}$$

where L is the lag operator, *d* can be any real number, and u_t is supposed to be I(0), defined for the purpose of the present work as a covariance stationary process with spectral density function that is positive and bounded at the zero frequency, including thus potentially ARMA structures. Note that this model is very general, in the sense, that if we impose, for example, $d = 0$ and $z_t = (1,t)^T$, we obtain the classical "trend stationary" I(0) representation suggested by many authors to describe the behavior of economic time series (e.g. DeJong *et al.,* 1992), while if $d = 1$, we obtain the "unit root" or I(1) model advocated by many other authors (Nelson and Plosser, 1982). However, in this work, we also allow *d* to be a fractional number. Thus, the parameter *d* might be 0 or 1, but it may also take values between these two numbers or even above 1. Note that the polynomial $(1–L)^d$ in (2) can be expressed in terms of its binomial expansion, such that, for all real *d*:

$$(1 - L)^d = \sum_{j=0}^{\infty} \psi_j L^j = \sum_{j=0}^{\infty} \binom{d}{j} (-1)^j L^j = 1 - d\, L + \frac{d(d-1)}{2} L^2 - \ldots,$$

and thus:

$$(1 - L)^d x_t = x_t - d x_{t-1} + \frac{d(d-1)}{2} x_{t-2} - \ldots .$$

In this context, *d* plays a crucial role, since it will be an indicator of the degree of dependence of the series. Thus, the higher the value of *d* is, the higher the level of association will be between the observations. Processes with $d > 0$ in (2) display the property of "long memory", so-named because of the strong degree of association between observations far distant in time. They are also characterised by autocorrelations which decay hyperbolically slowly and a spectral density function unbounded at the origin. The origin of these processes is found in the 1960s, when Granger (1966) and Adelman (1965) pointed out that most aggregate economic time series have a typical shape where the spectral density increases dramatically as the frequency approaches zero. However, differencing the data frequently leads to over-differencing at the zero frequency. Fifteen years later, Robinson (1978) and Granger (1980) showed that aggregation could be a source of fractional integration through the aggregation of heterogeneous autoregressive (AR) processes. Since then, fractional processes have been widely employed to describe the dynamics of time series (Diebold and Rudebusch, 1989; Sowell, 1992a; Baillie, 1996; Gil-Alana and Robinson, 1997). Gil-Alana and Hualde (2009) present an updated review of fractional integration and its applications in economic time series.

The methodology employed in the paper to estimate the fractional differencing parameter is based on the Whittle function (an approximation to the likelihood function) in the frequency domain (Dahlhaus, 1989). We also employ a testing procedure developed by Robinson (1994) that allows for testing any real value of d in I(d) models. The latter is a Lagrange Multiplier (LM) procedure which is considered to be the most efficient procedure in the context of fractional integration. It tests the null hypothesis H_o: $d = d_o$ for any real value d_o in (1) and (2) and different types of I(0) disturbances, and given the fact that the test statistic follows a standard (normal) limit distribution, it is possible to easily construct confidence bands for the non-rejection values. The functional form of this method is specified in various empirical applications (Gil-Alana and Robinson, 1997; Gil-Alana, 2000; Gil-Alana and Henry, 2003). Other parametric methods, such as Sowell's (1992b) maximum likelihood estimation in the time domain, along with a semi-parametric Whittle method in the frequency domain (Robinson, 1995; Abadir *et al.,* 2007) were also employed in the paper leading to essentially the same results.

4. Results

The order of integration d is estimated using the methodology described in section 3. From Figure 1, in the long term price-to-rent and price-to-income ratios generally hover around a constant. However, a trend for some series cannot be ruled out through visual inspection. Therefore, both the model with only an intercept ($z_t = 1$) and the one with an intercept and a linear time trend (($z_t = (1, t)^T$) are estimated. The model is first estimated assuming that u_t is a white noise disturbance. For the model with only an intercept, the order of integration of the price-to-income ratio ranges from slightly above one in Canada and Germany to 1.72 in New Zealand and 1.85 in Italy (Table 2). In all cases, the value of one is outside the 95% confidence interval. Hence the price-to-income appears not only non-stationary, but also non mean-reverting. Results are similar when a linear time trend is included. Assuming that u_t is a white noise disturbance may, however, be unrealistic, as residuals may be autocorrelated. Therefore, an alternative model with a Bloomfield disturbance is estimated (Table 3). Under that specification, the point estimate of the order of integration of the price-to-income ratio is lower than one in Australia (0.89), Canada (0.95) and Italy (0.73). However, for Australia and Canada, the 95% confidence interval includes the value of one, implying that the hypothesis of no mean-reversion cannot be rejected. For Italy, the order of integration is below one at the 95% level of confidence. Hence, the price-to-income ratio, though not stationary, is mean-reverting in Italy. For countries other than Australia and Canada and Italy, the assumption of mean-reversion can clearly be rejected. The highest values of integration are for Belgium (1.62), the Netherlands (1.68) and France (1.90). Results are almost unaltered by the inclusion of a linear time trend.

[Table 2. Estimates of d and 95% confidence intervals for the price-to-income ratios (White noise disturbances)]

[Table 3. Estimates of d and 95% confidence intervals for the price-to-income ratios (Bloomfield disturbances)]

The order of integration of the price-to-rent ratio in the model with only an intercept and u_t assumed to be a white noise disturbance ranges from 1.21 in Canada and 1.37 in Germany to 1.87 in France and 2.02 in Japan (Table 4). The results are similar when a linear time trend is included. In all cases the value of one is outside the 95% confidence interval. As for the price-to-income ratio, a model with a Bloomfield disturbance is estimated to account for autocorrelation in residuals (Table 5). The point estimate of the order of integration is below one only in Canada (0.97). But even in that case, the hypothesis of no mean-reversion cannot be rejected, as the value of one lies within the 95% confidence interval. In Australia and Italy, while the point estimate (respectively 1.16 and 1.05) is above one, the hypothesis that the order of integration is equal to one cannot be rejected at the 95% level of confidence. For all other countries, the order of integration is clearly higher than one, with the highest values in France (1.78), Switzerland (1.90) and the Netherlands (1.95). Again, including a linear time trend has little influence on the results.

[Table 4. Estimates of d and 95% confidence intervals for the price-to-rent ratios (White noise disturbances)]

[Table 5. Estimates of d and 95% confidence intervals for the price-to-rent ratios (Bloomfield disturbances)]

Overall, no price-to-income or price-to-rent series is stationary and very few of them are mean-reverting over a period spanning several decades. As noted in Section 1, "long memory" processes can easily be confused with regime switching processes. Therefore, it is necessary to test for the presence of structural breaks in the series. This is done by producing recursive estimates of *d* (Figure 2 and 3). Starting with a sample containing the first 80 observations, four observations (one year) are added at each iteration till the end of the sample. The order of integration is remarkably stable for most countries. However, in Italy, the order of integration increases over time, although it remains below one. The increase is already visible in the 1990s (the first ten observations on the chart), but accelerates in the 2000s. A similar increase in the integration order is also visible in Ireland, where *d* was close to one on the sample ranging to the mid-1990s and has increased to 1.37 on the whole sample. This suggests the creation of the euro area, which was followed by a sharp fall in interest rates, may have increased persistence in the price-to-income ratio. Housing price increases triggered by easier lending conditions may have fed into price expectations, leading to further price increases. It is too early to say how the current euro area crisis will affect persistence in price-to-income ratios going forward.

There are abrupt changes in the estimates of *d* for Australia in the early 2000s and in Denmark in the middle of that decade. In both cases, the price-to-income ratio jumped up. In Australia, underlying factors include falling interest rates, financial deregulation and innovation improving access to finance, increases in household wealth, changes in property taxation, high population growth and supply constraints (Yates, 2011). Low interest rates and innovations in housing finance have also played an important role in the Danish housing boom. Interest-only loans introduced in 2003 proved very popular, boosting household debt and housing demand (Lunde, 2012). A freeze in the calculation basis for the property value tax may also have contributed to encourage investment in housing. Interestingly, evolutions in the two countries have been very different after the boom. In Australia, the price-to-income ratio more or less stabilised at a level about 30% above its long-term average. In Denmark, it fell back sharply and is now close to its long-term average. This illustrates the difficulty to assess the sustainability of housing prices from price-to-income ratios alone. The recursive estimates of the order of integration for the price-to-rent ratio are also very stable, except for the countries where there is also instability in the price-to-income ratio and for New Zealand. Overall, the order of integration of the price-to-income and price-to-rent series is extremely stable over time and one can safely conclude that the absence of mean-reversion found for most countries is not the result of structural breaks.

[Figure 2. Recursive estimates of *d* in the price-to-income ratio]
[Figure 3. Recursive estimates of *d* in the price-to-rent ratio]

Our results are in line with the literature using non-fractional unit root tests reviewed in the introduction of this paper, which most often rejects the hypothesis of stationarity of price-to-income and price-to-rent ratios. Using the fractional integration framework allows going a step further than rejecting stationarity, as more protracted mean-reversion, consistent with “long memory”, can also be ruled out in most cases. This result confirms on a wider set of countries the high persistence in house prices found in studies testing for fractional integration in South Africa (Gil-Alana *et al*., forthcoming) and the United States (Barros *et al*., 2012). It is also consistent with other strands of the literature pointing to persistence in house prices, price-to-income and price-to-rent ratios. Hui and Zheng (2012) use multivariate stochastic volatility (MSV) models to show that the volatility of prices is significantly higher than that of rents across different segments of the Hong Kong property market. Furthermore, they find that correlations between prices and rents are time-varying. Cunningham and Kolet (2011) study duration dependence of house prices in a panel of 137 US and Canadian cities. As the authors note, “duration dependence is linked to mean reversion in that significant duration dependence would allow one to predict the timing of prices reverting to their mean”. Controlling for house price determinants, including income, they find duration dependence in expansions in the United States, but not in Canada and not in recessions in either countries. They suggest that speculative activity may prolong expansions, ultimately leading to overshooting of house prices. This interpretation is consistent with the fact that expectations of house prices tend to be extrapolative (Cho, 1996; Muellbauer, 2012) and with some evidence of bubbles and momentum trading in housing markets (Black *et al*., 2006). Sommer *et al.* (2013) shed light on the increase in the US price-to-rent ratio between 1995 and 2006 by developing a dynamic equilibrium model with endogenous house prices and rents. In their model, the combination of low interest rates and smaller down-payment requirements pushes up house prices much more than rents, and hence leads to higher price-to-rent ratios. Changes in fundamentals, however, explain only about half of the increase in the price-to-rent ratio over the period, suggesting a further role for over-optimistic expectations about house price increases. In sum, the high persistence in price-to-income and price-to-rent ratios found using fractional integration tests may be explained by a combination of shifts in fundamentals and expectation-driven house price dynamics.

5. Concluding remarks

This paper has investigated the persistence of housing price-to-income and price-to-rent ratios in 16 OECD countries over a 40-year period, using a fractional integration framework. Even though these ratios tend to fluctuate around a stable level over the very long term, they are generally not found to be mean-reverting over the sample. This result is not caused by structural breaks, which are found only in a very limited number of countries. Some consequences may be inferred for housing market analysis and economic policy. As the order of integration of price-to-income and price-to-rent ratios is above one for most countries, exogenous shocks to these ratios will be permanent. Moreover, the integration order is in most cases significantly higher than one, suggesting that shocks are in fact amplified.

If the housing system has weak stabilising properties, policy intervention may be warranted. High housing prices may result from different causes, like low interest rates, lax lending conditions, demographic pressures or tightness of supply. There is no reason *a priori* to consider that high price-to-income or price-to-rent ratios should be brought back to their historical averages. Nevertheless, high ratios may be a problem from different perspectives. For example, excessively high housing prices may cause social problems if access to decent housing becomes unaffordable for many households, leading in particular to overcrowding and homelessness. Social and economic inequalities may be exacerbated by unequal access to homeownership. In a tight market, the wealth gap between those able to put a foot on the housing ladder and those who rent may keep on widening. High housing costs may prevent a smooth functioning of the labour market and erode the competitiveness of the economy. Housing price increases which look unsustainable may raise concerns about financial stability. Altogether, the need for policy action should be assessed on the basis of the consequences of high price-to-income or price-to-rent ratios.

The appropriate policy instruments to use will depend on the nature of the problems associated with high price-to-income or price-to-rent ratios. High housing prices may result from rigid housing supply linked to tight supply of land for development, as in the United Kingdom and Australia. In that case, supply-side measures, such as reforming land-use planning or developing infrastructure, may be warranted. A number of other structural factors, such as taxation and regulations may raise the volatility of housing markets (Andrews *et al*., 2011). Buoyant prices can also result from unsustainable demand, which could be reined in by monetary policy tightening or macro-prudential policies, and may indicate risks of financial crisis. Borio (2012) finds that the financial cycle is most parsimoniously described in terms of credit and property prices. Overall, housing price-to-income and price-to-rent ratios are useful indicators to monitor, as deviations from their long-term average may reflect unsustainable developments in housing or mortgage markets. Nevertheless, the ratios are very persistent. If high ratios have adverse social and economic consequences, policy action guided by a careful analysis of underlying factors, may be warranted.

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Table 1: Housing price series entering price-to-income and price-to-rent ratios

Country	Source	Series	Original frequency	Seasonally adjustemnt
Australia	Australian Bureau of Statistics	House Price Indexes: Eight Capital Cities	Quarterly	OECD
Belgium	Banque National de Belgique	Residential property prices, existing dwellings, whole country	Quarterly	OECD
Canada	Teranet, National Bank, National Composite House. Price index Since 1999Q2. Department of Finance for earlier data	Average existing home prices	Quarterly	Original
Denmark	StatBank	Price index for sales of property	Quarterly	OECD
Finland	Statistics Finland	Prices for dwellings	Quarterly	OECD
France	Institute National de la Statistique et des Études Économiques INSEE	Indice trimestriel des prixdes logements Anciens _ France métropolitaine _ Ensemble _ Indice brut	Quarterly	Original
Germany	Deutsche Bundesbank	Residential property prices in Germany	Annual[1]	Original
Ireland	Central Statistics Office	Residential property price index	Monthly	OECD
Italy	Nomisma	13 main metropolitan areas. Average current prices of used housing	Semi-Annual[2]	OECD
Japan	Japan Real State Institute	Urban Land Price Index _ Nationwide	Semi-Annual[2]	OECD
Netherlands	Kadaster	House Price Index for existing own	Monthly	OECD
New Zealand	Reserve Bank of New Zealand	House Price Index _ All residential	Quarterly	OECD
Sweden	Statistics Sweden	Real estate price index for one and two Dwelling buildings for permanent living	Quarterly	OECD
Switzerland	Banque Nationale Suisse	Real state price indices	Quarterly	OECD

U.K.	Department for communities and local goverment	Mix adjusted house price index	Quarterly	OECD
U.S.A.	Federal Housing Finance Agency FHFA	Purchase and all transactions indices[3]	Quarterly	Original

1 Quarterly series for prices for owner occupier apartments in seven cities are used to determine the quarterly profile from 2008 on; linear interpolation is used for earlier data.
2 Quarterly series are derived through linear interpolation.
3 FHFA. Purchase index from 1991 and all transaction index before (seasonally adjusted by the OECD)

Table 2: Estimates of d and 95% confidence intervals for the price to income ratios (White noise disturbances)

PRICE TO INCOME	An intercept	A linear time trend
Australia	1.171 (1.049, 1.332)	1.172 (1.040, 1.335)
Belgium	1.327 (1.252, 1.425)	1.325 (1.255, 1.417)
Canada	1.096 (1.009, 1.232)	1.097 (1.005, 1.235)
Denmark	1.535 (1.411, 1.687)	1.531 (1.401, 1.697)
Finland	1.621 (1.498, 1.793)	1.619 (1.480, 1.790)
France	1.531 (1.444, 1.642)	1.532 (1.446, 1.648)
Germany	1.087 (1.014, 1.156)	1.089 (1.011, 1.188)
Ireland	1.253 (1.165, 1.360)	1.253 (1.160, 1.362)
Italy	1.848 (1.506, 2.267)	1.845 (1.505, 2.262)
Japan	1.336 (1.244, 1.447)	1.337 (1.255, 1.442)
Netherlands	1.504 (1.414, 1.615)	1.504 (1.411, 1.619)
New Zealand	1.721 (1.554, 1.970)	1.721 (1.554, 1.973)
Sweden	1.319 (1.244, 1.411)	1.318 (1.233, 1.415)
Switzerland	1.236 (1.154, 1.333)	1.236 (1.154, 1.338)
U.K.	1.398 (1.293, 1.533)	1.398 (1.290, 1.536)
U.S.A.	1.248 (1.165, 1.354)	1.249 (1.160, 1.356)

Most of the series start at 1970/1 and end at 2011/4. The exceptions are Denmark, starting at 1981/1; Germany, at 1980/1; Finland, at 1975/1; France, at 1978/1; UK, at 1975/1; Ireland, at 1977/1 and New Zealand, at 1986/1.

Table 3: Estimates of d and 95% confidence intervals for the price to income ratios (Bloomfield disturbances)

PRICE TO INCOME	An intercept	A linear time trend
Australia	0.892 (0.744, 1.110)	0.892 (0.744, 1.110)
Belgium	1.619 (1.402, 1.911)	1.597 (1.401, 1.888)
Canada	0.947 (0.808, 1.132)	0.947 (0.809, 1.137)
Denmark	1.501 (1.255, 1.854)	1.506 (1.243, 1.910)
Finland	1.388 (1.109, 1.786)	1.384 (1.103, 1.866)
France	1.901 (1.602, 2.388)	1.902 (1.609, 2.411)
Germany	1.423 (1.231, 1.654)	1.417 (1.239, 1.661)
Ireland	1.368 (1.172, 1.611)	1.367 (1.172, 1.622)
Italy	**0.734** **(0.589, 0.967)**	**0.741** **(0.588, 0.960)**
Japan	1.598 (1.323, 1.968)	1.614 (1.334, 1.972)
Netherlands	1.681 (1.424, 2.089)	1.681 (1.422, 2.090)
New Zealand	1.349 (1.121, 1.677)	1.349 (1.122, 1.679)
Sweden	1.570 (1.343, 1.831)	1.569 (1.344, 1.845)
Switzerland	1.579 (1.294, 2.011)	1.580 (1.294, 2.011)
U.K.	1.439 (1.213, 1.798)	1.438 (1.213, 1.788)
U.S.A.	1.379 (1.191, 1.621)	1.381 (1.201, 1.666)

In bold, statistical evidence of mean-reversion ($d < 1$) at the 5% level.

Table 4: Estimates of d and 95% confidence intervals for the price to rent ratios (White noise disturbances)

PRICE TO RENT	An intercept	A linear time trend
Australia	1.768 (1.566, 1.990)	1.772 (1.590, 1.988)
Belgium	1.376 (1.300, 1.475)	1.385 (1.312, 1.487)
Canada	1.208 (1.093, 1.366)	1.210 (1.095, 1.366)
Denmark	1.539 (1.421, 1.682)	1.539 (1.422, 1.688)
Finland	1.648 (1.504, 1.833)	1.649 (1.506, 1.832)
France	1.869 (1.740, 2.020)	1.861 (1.733, 2.011)
Germany	1.370 (1.270, 1.485)	1.369 (1.283, 1.484)
Ireland	1.539 (1.403, 1.710)	1.539 (1.400, 1.712)
Italy	1.648 (1.455, 1.911)	1.648 (1.455, 1.910)
Japan	2.017 (1.884, 2.187)	2.015 (1.875, 2.182)
Netherlands	1.415 (1.344, 1.500)	1.416 (1.344, 1.509)
New Zealand	1.657 (1.511, 1.844)	1.658 (1.514, 1.843)
Sweden	1.714 (1.601, 1.861)	1.667 (1.563, 1.818)
Switzerland	1.391 (1.300, 1.493)	1.389 (1.305, 1.488)
U.K.	1.663 (1.535, 1.823)	1.663 (1.534, 1.826)
U.S.A.	1.430 (1.344, 1.534)	1.431 (1.354, 1.532)

Most of the series start at 1970/1 and end at 2011/4. The exceptions here are Australia, starting at 1973/1; Belgium, at 1977/1; and Sweden, at 1980/1.

Table 5: Estimates of d and 95% confidence intervals for the price to rent ratios (Bloomfield disturbances)

PRICE TO RENT	An intercept	A linear time trend
Australia	1.159 (0.970, 1.616)	1.189 (0.960, 1.688)
Belgium	1.585 (1.395, 1.858)	1.618 (1.424, 1.874)
Canada	0.970 (0.821, 1.196)	0.969 (0.817, 1.180)
Denmark	1.449 (1.214, 1.756)	1.449 (1.212, 1.758)
Finland	1.259 (1.034, 1.623)	1.260 (1.036, 1.625)
France	1.782 (1.510, 2.221)	1.759 (1.513, 2.166)
Germany	1.461 (1.254, 1.709)	1.451 (1.263, 1.704)
Ireland	1.219 (1.013, 1.485)	1.219 (1.021, 1.486)
Italy	1.048 (0.888, 1.291)	1.047 (0.887, 1.302)
Japan	1.729 (1.411, 2.159)	1.739 (1.437, 2.177)
Netherlands	1.957 (1.684, 2.322)	1.977 (1.711, 2.366)
New Zealand	1.362 (1.188, 1.674)	1.372 (1.188, 1.670)
Sweden	1.699 (1.451, 2.060)	1.617 (1.400, 1.910)
Switzerland	1.900 (1.554, 2.303)	1.871 (1.544, 2.290)
U.K.	1.483 (1.215, 1.880)	1.482 (1.214, 1.886)
U.S.A.	1.635 (1.414, 1.890)	1.642 (1.433, 1.981)

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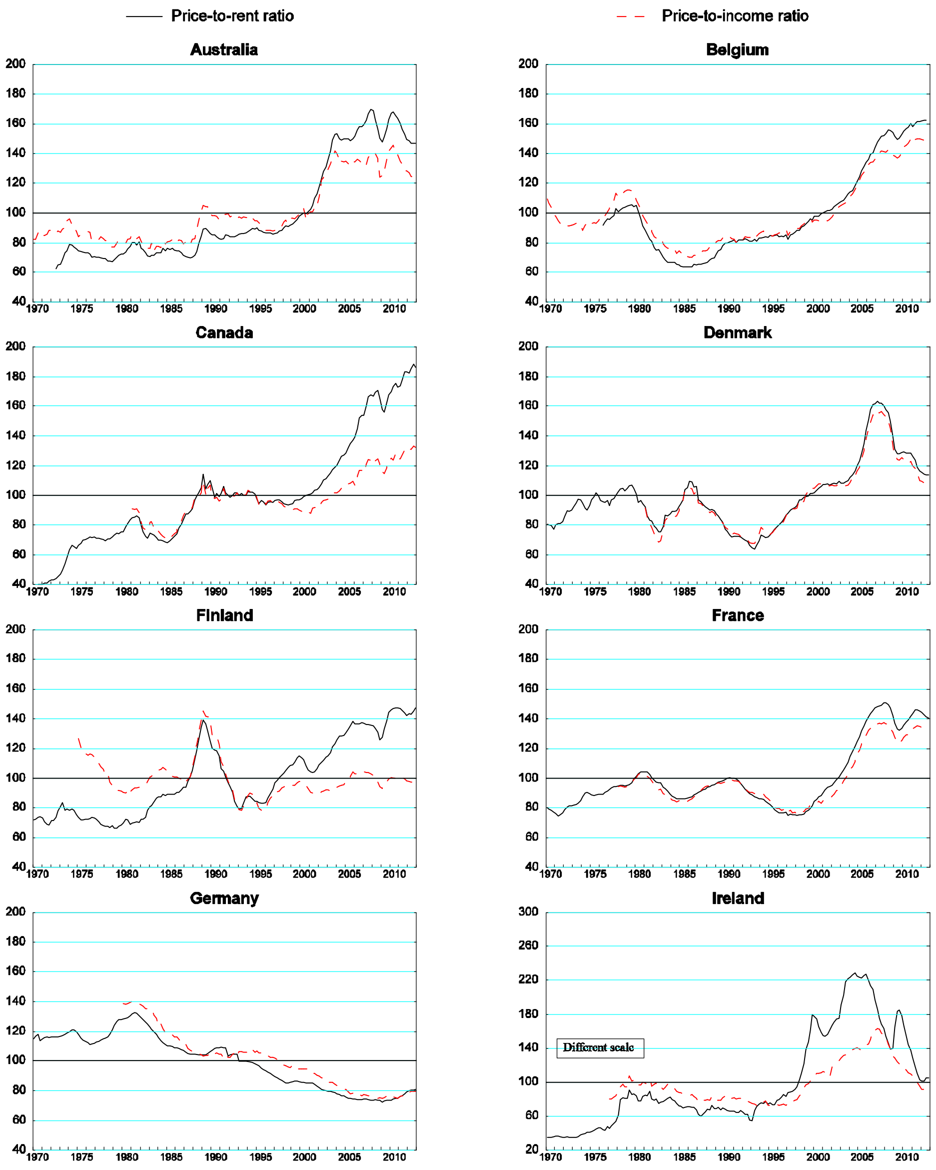
Figure 1. Price-to-income and price-to-rent ratios

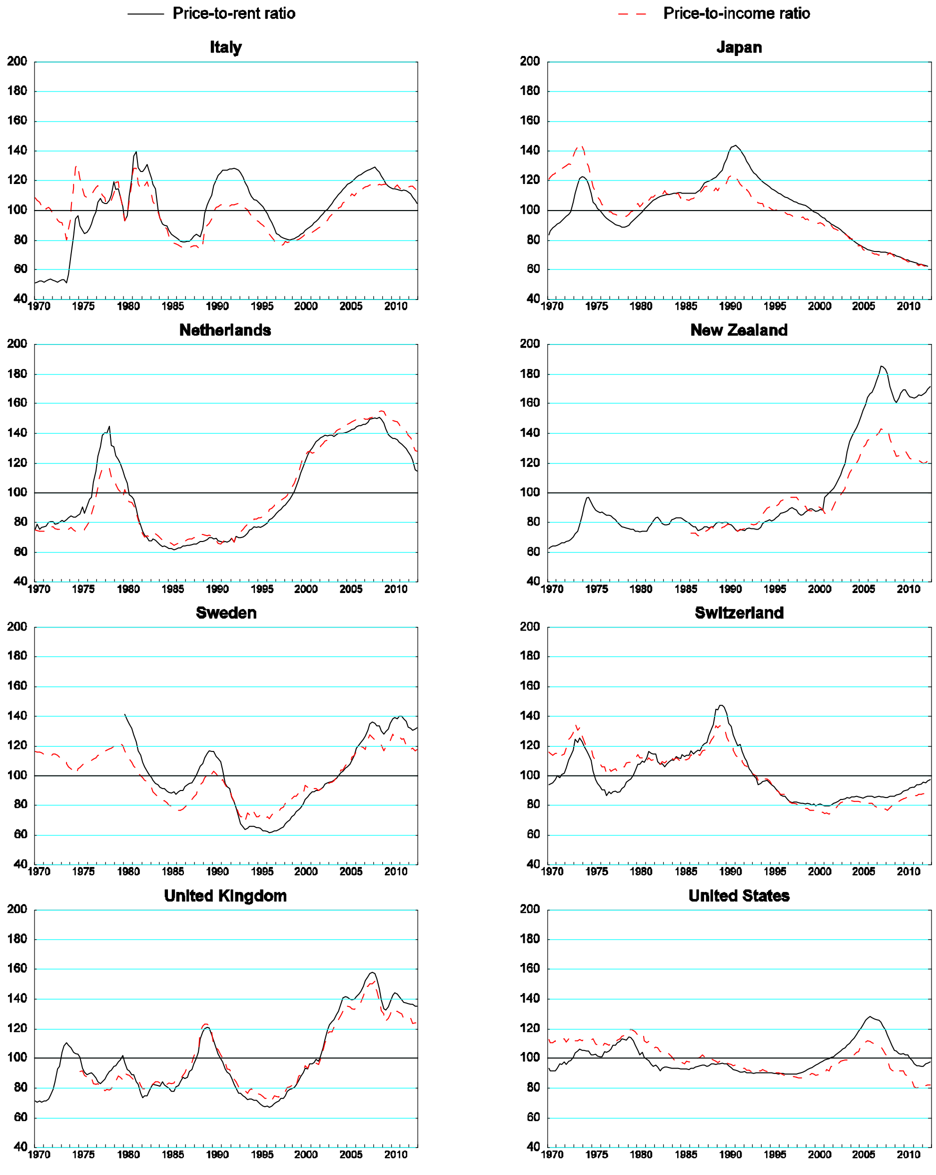
Sample average = 100



Source: OECD calculations.

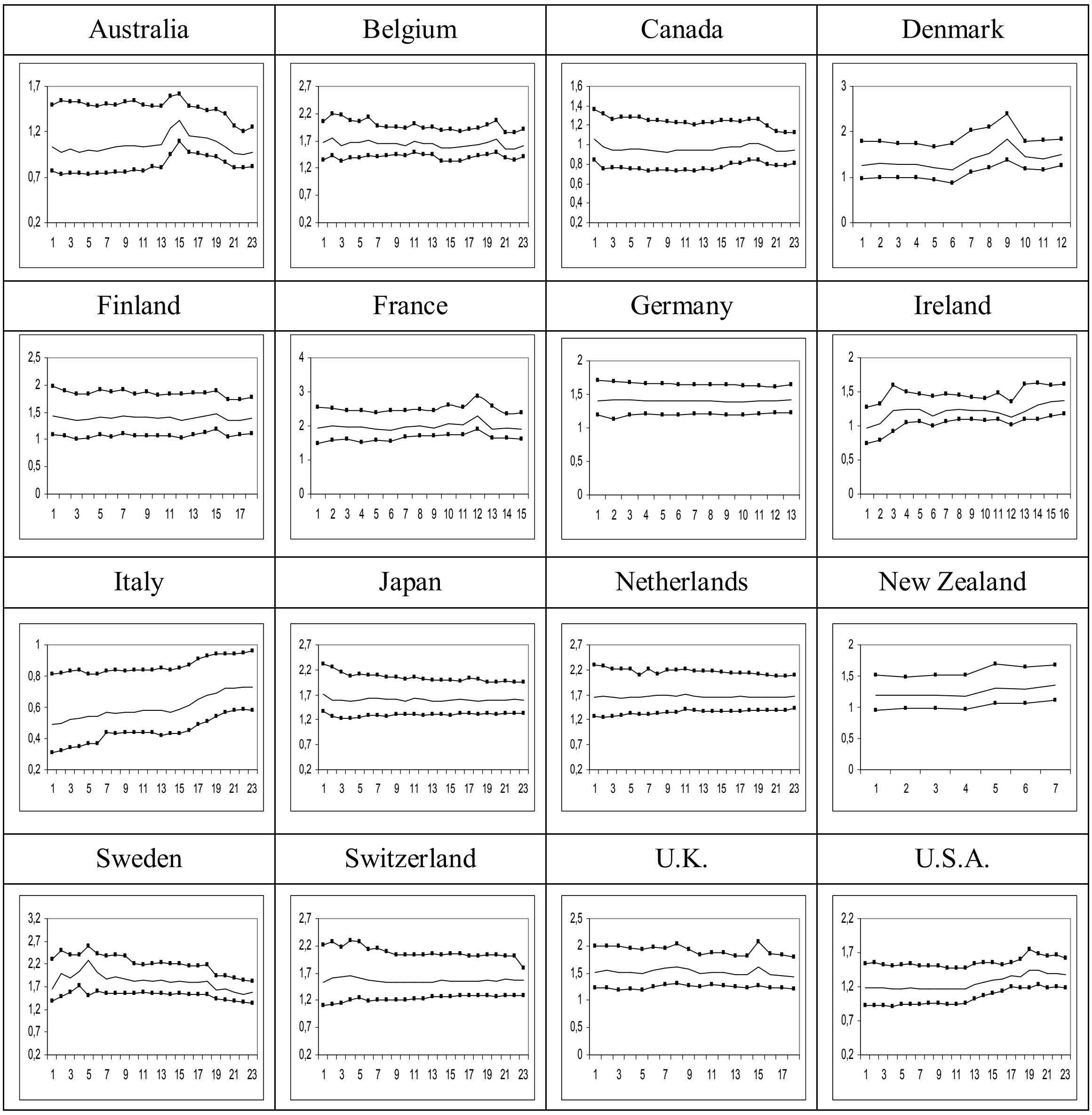
Figure 1. Price-to-income and price-to-rent ratios (cont.)

Sample average = 100



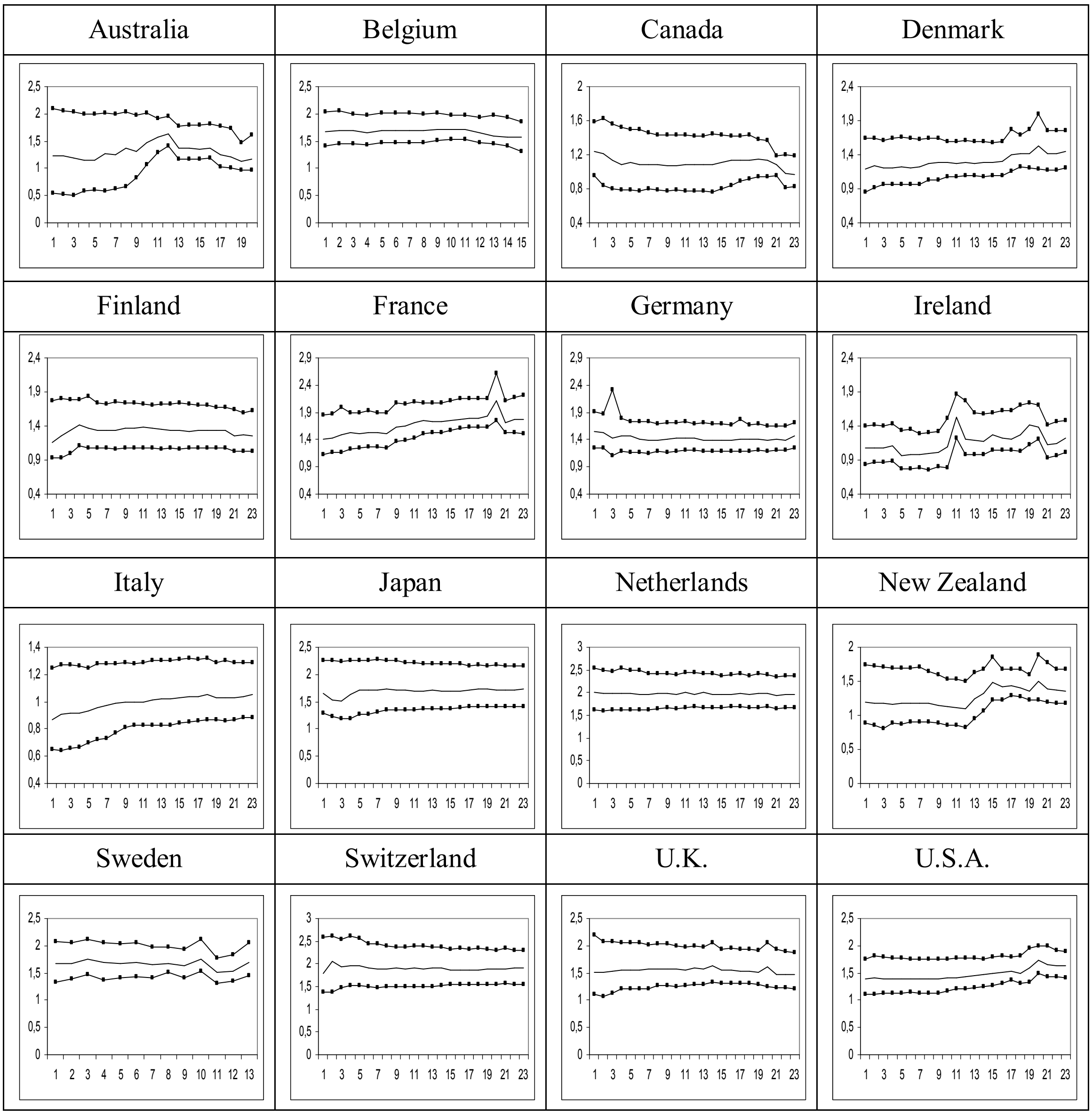
Source: OECD calculations.

FIGURE 2: RECURSIVE ESTIMATES OF D IN THE INCOME RATIOS



The dotted lines refer to the 95% confidence intervals.

FIGURE 3: RECURSIVE ESTIMATES OF D IN THE RENT RATIOS



The dotted lines refer to the 95% confidence intervals.