Absolute Income Inequality and Rising House Prices

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Abstract

Income inequality and house prices have risen sharply in developed countries. We argue that this co-movement is no coincidence but that inequality has driven up house prices on the grounds that it raises the total demand for houses. Our results suggest that absolute inequality and house prices in most OECD countries indeed were positively correlated and cointegrated during 1975-2010, and for most countries absolute inequality Granger-caused house prices. Relative inequality, on the other hand, is not cointegrated with house prices, which is expected given that total house demand depends on the absolute amount of investible income. Moreover, our results confirm previous findings that in some countries low short-term real interest rates also contributed to the surge in house prices, whereas real GDP growth did not.

Key Words: Personal Income Inequality, Absolute Inequality, House Prices, Asset Price Inflation, Asset Bubbles, Panel Cointegration

JEL Classification: C22, C23, D31; G12; R21

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1. Introduction

Variations in house prices can have important macroeconomic effects. Rising house prices stimulate consumption expenditure and economic growth when they increase the security feeling of homeowners and ease access to credit—so called wealth and collateral effects (Case *et al.*, 2005, 2013; Campbell, and Cocco, 2007; Hryshko *et al.*, 2010). However, at the same time, easier access to credit can foster unsustainable debt-driven growth models and declining house prices can lead to large reductions in household consumption and prolonged recessions. Indeed, all of these effects have been observed prior to and after the Great Recession (Hryshko *et al.*, 2010; Mian *et al.*, 2013; Jordá *et al.*, 2014; Mian and Sufi, 2015; Goda *et al.*, 2017).

Moreover, starkly rising prices can make housing unaffordable. This especially concerns the most productive urban areas and low income households (Dewilde and Lancee, 2013; Gyourko *et al.*, 2013)[1]. Finally, house price inflation can translate into retail price inflation (Stroebel and Vavra, 2014), which can have important implications for monetary policy and is also seen to affect mainly low income households (see Easterly and Fischer (2001) on the latter).

Considering these potential socio-economic effects, it is not surprising that a vast literature on the dynamics of house prices exist (especially in the aftermath of the US Subprime Crisis). Typically, the level and growth of income is identified as an important long-run determinant of house prices (Case and Shiller, 2003; ECB, 2003; Sommer *et al.*, 2013). However, in developed countries since "the final decades of the twentieth century, house price growth outpaced income growth by a substantial margin" (Knoll *et al.*, 2017: 338). Recent literature

[1] In the UK, for example, "Homes in popular towns and London boroughs have risen to 10 and 20 times local incomes, while rents account for up to 78% of earnings" (Collinson, 2015). However, it is important to note that "while the increase in house prices has been most pronounced in cities, it is not exclusively an urban phenomenon" (Knoll *et al.*, 2017: 343).

suggests that this phenomenon is mainly explained by low real interest rates coupled with credit expansion (Taylor, 2007; Goodhart and Hofman, 2008; Gerdesmeier *et al.*, 2010; Agnello and Schuknecht, 2011; Bordo and London-Lane, 2013a). Other studies also consider financial innovation and deregulation (Dokko *et al.*, 2011; Bordo and London-Lane, 2013b), and global liquidity (Sá *et al.*, 2014; Cesa-Bianchi *et al.*, 2015) as explanatory factors.

All of these determinants have in common that they are seen to increase the total demand for houses, which leads to increasing prices taking into account that land and house supply is restricted. However, another common feature of all of these determinants is that their effects mostly took place in the first decade of the twenty-first century, while house prices are increasing strongly since the 1970s. The aim of the present paper is to assess rising income inequality as an alternative contributing factor for the strong increase in house prices during the period 1975-2010.

Theoretical models provide two potential mechanisms that link inequality to house prices: (i) with rising inequality the number of households that are willing to pay higher prices for their homes increases (Gyourko *et al.*, 2013; Määttänen and Terviö, 2014); (ii) houses are an investment good for the upper part of the income distribution and in more unequal countries the investment demand is higher (Nakajima, 2005; Zhang, 2016). In both cases, the change in demand is expected to drive up house prices when supply restrictions are considered.

It is well established that house ownership is very unevenly distributed. In OECD countries the top 10% of the income distribution typically owns between 40% (Italy) and 60% (US) of houses, while the Gini coefficient ranges between 0.6 and 0.7 (Cowell *et al.*, 2012), even rising to above 0.9 when only non-primary residences are considered (Bonesmo Fredriksen, 2012). It is also well established that income inequality increased starkly in most developed countries after 1980, especially due to income concentration at the top (OECD, 2015). Our hypothesis is that the co-movement of income inequality and house prices is no coincidence, but that the increase in inequality has driven up house demand and, in turn, their prices.

To our best knowledge, no previous study has empirically tested if the stark increase in real house prices in developed countries during the last decades was partly driven by rising income inequality. To close this gap in the literature and test our hypothesis, the present study conducts cointegration tests for a panel of 18 OECD countries for the period 1975-2010. Given the evident nonstationarity of house prices and inequality we use cointegration based methods to avoid problems of spurious regression.

A second novelty of our study is that we will use both absolute and relative inequality measures to test our hypothesis.[2] The difference between relative and absolute inequality measures is that the former report proportional income differences (e.g. the Gini coefficient), while the latter refer to income differences in absolute terms (e.g. the variance). Studies that investigate the impact of inequality on socio-economic variables like growth and crime typically only account for relative inequality measures. However, absolute and relative inequality trends can be quite different (see Ravallion, 2004; Atkinson and Brandolini, 2010, Goda and Torres, 2017), and because "it is the absolute level of resources, not their relative distribution, that affects access to housing" (Dewilde and Lancee, 2013: 1189), we expect that absolute inequality measures are more suitable for our purpose.

Indeed, we find that absolute income inequality and house prices in OECD countries are positively correlated and cointegrated (with the notable exception of Germany, Japan, and Korea), whereas the relative inequality measures are not cointegrated. Importantly, we find that for the vast majority of our sample countries absolute inequality Granger-causes house

[2] Our measures for overall inequality changes are the Gini coefficient and the variance, while our measures for changes in the concentration of income are the top 5% income share and the top 5% market income.

prices, whereas house prices do not Granger-cause income inequality. In other words, the increase in absolute income inequality has driven up house prices, whereas in most countries the increase in house price seemingly has not contributed to the observed inequality increase.

Moreover, our results confirm previous findings that falling short-term real interest rates have also contributed to the long-term increase in real house prices (at least in some countries). Real GDP, on the contrary, shows no signs of cointegration with OECD house prices, which is in line with the above mentioned observation that real house price growth in OECD countries has been much higher than income growth.

The layout of this paper is as follows. Section 2 details the theoretical link between inequality and house prices. Section 3 gives an overview of the research design. Section 4 presents and discusses the results. Section 5 concludes the paper.

2. The theoretical link between (absolute) inequality and house prices

The models that examine whether inequality affects house prices are typically general equilibrium models that have three main conditions in common: First, the existence of heterogeneous agents, so that inequalities can be analysed. Second, house supply is assumed to be at least very inelastic, so that the house market adjusts to demand shocks by price changes. Third, the presence of frictions that limit access to the housing market.

According to these models, inequality can affect house prices via two demand mechanisms: (i) when houses are considered as consumption goods, an increase in income inequality raises the amount of people that are willing to pay high prices for their residence; and (ii) when houses are considered as rent generating assets, inequality is expected to increase the absolute amount of savings, which in turn raises the total demand for houses as investment good.

Regarding the first mechanism, Gyourko *et al.* (2013) presents a model in which two types of houses exist. The first type has an elastic supply, whereas the second type has an inelastic supply and is preferred by households. The model also differentiates between low and high wage earners. When the wage distribution changes in favour of high wage earners, more people desire to live in (and can pay for) the preferred houses. As a result, the price of preferred houses and the average house price increase, given that the other type of houses experiences a quantitative adjustment.

Määttänen and Terviö (2014) present a related model but differentiate houses according to their quality. The quality is defined as a continuous spectrum, which implies that for each house-quality type the supply is perfectly inelastic. Agents are assumed to maximize their utility choosing between goods consumption and the quality of their residence.[3] With increasing inequality low income households' willingness to pay for quality houses decreases, whereas the willingness of high income households to pay for quality houses increases. The outcome is that rising inequality leads to lower prices for low quality houses and to higher prices for high quality houses. The overall effect on house prices depends on which of these two opposing effects dominates.

Finally, Matlack and Vigdor (2008) present a model that considers the importance of land as a production factor (that can be transformed into houses without any cost) and of houses as consumption goods. The model assumes that the quantity of land is constant, that workers are divided according to their skills (high- and low-skilled), that wages equal marginal productivity, and production has a neoclassical production function.[4] Considering this

[3] The model assumes that each household only owns one house and that it chooses the quality level according to its income.

[4] More specifically, the authors assume the following production function: $Y = H^{\alpha}L^{\beta}A^{\varphi}K^{1-\alpha-\beta-\varphi}$, where H are high-skilled workers, L are low-skilled workers and A is land. Changes in α, β or φ, not only change the marginal productivity of each factor, but also the participation in total income. Hence, a variation in the values of these parameters changes the distribution of income.

setting, rising wage inequality leads to an increase in house prices when the share of land in the production is constant. This is the case because house demand of high-skilled workers increases by more than the demand from low-skilled workers decreases.

The second line of research considers houses as assets. Nakajima (2005) uses a life-cycle general equilibrium model economy composed of workers and retirees, who decide how to allocate their savings between housing and a financial asset. The return of each asset is determined by the ratio of the total return in terms of the available quantity. Houses are assumed to be inelastic, whereas the financial asset is assumed to be elastic with a decreasing marginal productivity. Rising income inequality, or the increase of the volatility of income, increases the precautionary savings of workers and their demand for housing (whereas the housing demand of the retired does not change significantly). The increasing demand for houses, in turn, increases their price.

Zhang (2016), on the other hand, proposes an incomplete market model with heterogeneous households and an exogenously given house supply. In the same vein as Nakajima (2005), Zhang treats houses as an asset that competes against an alternative asset (i.e. bonds) but in his model houses have a higher rate of return than the investment alternative. The reason why the return is higher is that houses are assumed to be a risk-free investment and that entry barriers to the market exist. Given its relatively high return, households always want to invest in the house market. However, the poor have insufficient income to enter the market and middle-income households can only hold a minimum amount of houses. Top income households, on the other hand, are not constrained and increase investment income in the house market when their income goes up. Rising inequality thus leads to increasing house demand and, in turn, to an increase in their prices.

All of these potential mechanisms have in common that an increase of the *absolute* level of income at the top leads to an increase in land and house demand and thus might partly explain why "the late twentieth century surge in house prices was due to sharply rising land prices" and not due to rising construction costs (Knoll *et al.*, 2017: 349).

It therefore seems important to distinguish between relative and absolute income inequality when empirically studying the impact of income inequality on house prices. The most widely used relative inequality index is the Gini coefficient (1), whereas the variance (2) is typically used when measuring absolute income differences (see Chakravarty, 2001; Goda and Torres García, 2017). The main difference between these two indices is that the Gini coefficient normalizes the sum of income differences with the mean income (μ), whereas the variance subtracts μ from individual incomes.

$$\text{Gini} = \frac{1}{\mu}\frac{1}{N^2}\frac{1}{2}\sum_{i=1}^{N}\sum_{j=1}^{N}\left|\gamma_j - \gamma_i\right| \tag{1}$$

$$\text{Variance} = \frac{1}{N}\sum_{i=1}^{N}(\gamma_i - \mu)^2 \tag{2}$$

where N is population size, μ is mean income, γ_i is income of the *i-th* individual, and γ_j is the income of the *j-th* individual.

An important property of the Gini coefficient is that its value is independent of overall income (i.e. it is scale invariant), whereas the opposite is true for the variance. To make this more palpable, Figure 1 shows the income distribution of two countries that are both assumed to have a population size of five. Although the income per capita of Country B is much higher than that of Country A, the Gini coefficient of both countries is identical (0.38). On the contrary, the value of the variance of both countries is quite distinct: the variance of Country B is 4,616, while that of Country A is only 46.

< Figure 1 >

As well as overall inequality the absolute amount of spendable/investible income of the upper part of the distribution in Country B is much higher than that of Country A. According to the theories discussed above, one would therefore expect that households in Country B would pay higher prices for houses as consumption goods and would also have a higher demand for houses as an investment good. As a result house prices in Country B should be higher than in Country A (*ceteris paribus*).[5]

Indeed, Figure 2 and 3 suggest that in most OECD countries real house prices are positively correlated with inequality, and that the correlation between absolute inequality and house prices is stronger than that of relative inequality. The remainder of this paper tests whether either absolute or relative income inequality and house prices in OECD countries are statistically positively correlated in the long-run and with the direction of causation running from inequality to house prices. To this end we use cointegration based methods that deal with nonstationarity and avoid the associated problem of spurious regression.

< Figure 2 >

< Figure 3 >

3. Research design

3.1 General specification and data

Consider the following potential cointegrating equation of interest:

$$ln(HP_t) = \beta_1 + \beta_2 INEQ_t + u_t \tag{3}$$

[5] Similarly, when the Gini coefficient in both countries increases by the same amount, the absolute amount of income in the upper part of the distribution would increase by more in Country B than in Country A, and consequently house prices in Country B are expected to increase by more than in Country A.

where $ln(HP_t)$ is the natural logarithm of real house prices and $INEQ_t$ denotes different income inequality measures. The house price data are yearly averages of the OECD real house price index. The two relative inequality measures considered are the Gini coefficient ($Gini_t$) and the top income share ($Top5\%_t$), while the two absolute inequality measures are the income variance in constant PPP prices ($ln(Var_t)$), and the market income of the top 5% income earners in constant US$ ($ln(Top5\$_t)$).

Data on real house prices and Gini coefficients are available on a yearly basis for 18 countries: Australia (AUS), Belgium (BEL), Canada (CAN), Denmark (DEN), Finland (FIN), France (FRA), Germany (GER), Ireland (IRE), Italy (ITA), Japan (JAP), the Netherlands (NET), New Zealand (NEW), Norway (NOR), South Korea (KOR), Spain (SPA), Sweden (SWE), the UK (UKD) and the USA (USA).

The market Gini coefficient is retrieved from Solt's Standardized World Income Inequality Database (SWIID, V5.0). The SWIID combines and adjusts Gini coefficients from different sources and currently is the most extensive publicly available database of income Gini coefficients that are comparable across countries and time. SWIID data have been widely used in previous studies concerned with income inequality.[6]

The income variance for each country is calculated as follows:

$$Var = \frac{1}{20}\sum_{p=1}^{20}\big((x_p * GDP) - GDPpc\big)^2 \qquad (4)$$

where GDP_{pc} is the mean per capita income of the country, x_p is the income share of the *p-th* population ventile of the country, and GDP is the total income of the country.[7]

[6] See, for example, Bergh and Nilsson (2010), Fox and Hoelscher (2012), Agnello and Soussa (2014), Herzer *et al. (*2014), Chon (2015) and Goda and Torres García (2017).
[7] Please note that x_p is not readily available on a frequent basis, so it was calculated using the same method as a recent study that estimates global changes in absolute inequality (Goda and Torres, 2017).

3.2 Determining stationary, trend stationary and nonstationary series

To establish if the necessary condition for cointegration between real house prices ($ln(HP_t)$) and the inequality measures ($INEQ_t$) is satisfied, first Pesaran's (2007) panel unit root test (based upon truncated CADF statistics) is applied to the natural log of real house prices and the four inequality variables. In a second step Cerrato *et al.*'s (2009, 2011) heterogeneous nonlinear panel unit root test is used. Both tests account for cross-sectional dependence.

For both panel unit root tests the sequential panel selection method (SPSM), proposed by Chortareas and Kapetanios (2009), is applied to identify which cross-sections (countries) are stationary and which are nonstationary[8]. The null hypothesis is that all countries' series are I(1) and the alternative is that at least one country's series is I(0).

The SPSM essentially involves applying the panel unit root test to all N countries, the test statistic is denoted $\bar{t}(N,T)$, and if the null cannot be rejected the procedure stops and all countries' series are I(1). However, if the null hypothesis is rejected at least one country's series is I(0) and we exclude the country that rejects the I(1) null the most, which is the one that has the smallest (most negative) individual country test statistic, $min\{t_i(N,T)\}$. The panel unit root test statistic, $\bar{t}(N-1,T)$, is calculated for the remaining $N-1$ countries. The test is repeated for the remaining countries and the process continues until the panel unit root test cannot reject the null. All countries' series included in the last test are I(1) and all countries' series excluded from the last test are I(0).

To finally determine which series is stationary, trend-stationary or nonstationary we use the following procedure: if the unit root null is rejected using the test including only an

[8] Chortareas and Kapetanios (2009) apply the SPSM procedure to the Im *et al.* (2003) panel unit root test that does not account for cross-sectional dependence.

intercept as a deterministic term the series is stationary. However, if the null is not rejected, the unit root test including both an intercept and trend is considered. If the null of this test is rejected the series is trend stationary, whereas if the null is not rejected the series has a unit root.

The Cerrato *et al.* (2009, 2011) test assumes nonlinear adjustment (possibly approximating structural breaks) whereas the Pesaran (2007) test assumes linear adjustment. Since each test is most powerful for the adjustment it is designed for we infer stationarity if either test indicates stationary. Further, if either test suggests trend stationarity and neither indicates stationarity we will infer trend stationarity. Otherwise, we infer a unit root.

3.3 Determining cointegration and causality

We proceed to test for cointegration between $ln(HP_t)$ and $INEQ_t$ by applying Westerlund's (2007) panel cointegration test. We use the `xtwest` command, provided by Persyn and Westerlund (2008), with the Stata 14 IC program, to produce all of the reported results associated with Westerlund's (2007) method. Westerlund's (2007) tests use the following model assuming a single cointegrating vector:

$$\Delta y_{i,t} = \delta_{1,i} + \delta_{2,i}t + \alpha_i y_{i,t-1} + \boldsymbol{\lambda}_i' \boldsymbol{x}_{i,t-1} + \sum_{j=1}^{p_i} \alpha_{ij} \Delta y_{i,t-j} + \sum_{j=-q_i}^{p_i} \boldsymbol{\gamma}_{ij}' \Delta \boldsymbol{x}_{i,t-j} + e_{i,t} \quad (5)$$

where, $\boldsymbol{x}_{i,t}' = (x_{1,i,t} \quad x_{2,i,t} \quad \cdots \quad x_{K,i,t})$ is a vector containing K explanatory variables that are assumed to be weakly exogenous while the inclusion of q_i lead values prevents the violation of strict exogeneity. The number of leads and lags is chosen to minimise Akaike's information criterion (AIC) as implemented with Persyn and Westerlund's (2008) Stata program.

The null of no cointegration for any cross-sectional unit, H_0: $\alpha_i = 0 \ \forall \ i$, is tested against two different alternative hypotheses. The two group mean statistics, denoted G_τ and G_α, specify the alternative as cointegration *for at least one cross-sectional unit*: $H_1^G: \alpha_i < 0$ *for at least one* i. G_α utilises a heteroscedasticity and autocorrelation consistent (HAC) adjustment where we set the bandwidth parameter using: $M_i = 4\left(\frac{T}{100}\right)^{2/9}$, giving $M_i = 3$.[9] The two panel statistics, denoted P_τ and P_α, specify the alternative hypothesis that there is cointegration *for all cross-sectional units*, that is, $H_1^P: \alpha_i < 0\ \forall\ i$.[10]

The four panel cointegration statistics are normalised using the asymptotic moments reported in Table 1 of Westerlund (2007) and have an asymptotic standard normal distribution. Any normalised statistic that is less negative (greater) than the left-tail critical value implies that the no cointegration null should not be rejected. We report bootstrapped probability values (using 800 replications), that are robust to very general forms of cross-sectional dependence, as produced by Persyn and Westerlund's (2008) program.

The Westerlund (2007) test assumes weak exogeneity and we asses this assumption by, firstly, applying the Westerlund (2007) test to the reverse regression of inequality on house prices. However, this method of assessing weak exogeneity is only suggestive. The reasons for this include the following. First, the cointegrating equations in the autoregressive distributed lag (ADL) models are different when the difference of $ln(HP_t)$ is the dependent variable and when the difference of inequality is the dependent variable. Second, only leads and lags of the differenced regressors (and not the dependent variable) are included in the ADL model. A more typical test for weak exogeneity is based on the error-correction form of a vector autoregression (VAR), typically referred to as the *restricted* vector error correction model (VECM) or VEC.[11] The VEC, assuming one cointegrating equation with

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[9] We set the maximum number of lead and lags in (5) to 3.

[10] We find that when cointegration is supported it is based on at least one of the panel statistics suggesting cointegration for the whole panel of countries.

[11] The restricted VECM, or VEC, imposes the number and form of cointegrating equations on the unrestricted VECM.

(unrestricted) intercept and no trend, in this two variable system would be specified as follows for country i:

$$\begin{aligned}\Delta ln(HP_t) &= \gamma_{11} + \textstyle\sum_{j=1}^{p} \gamma_{12j} \Delta\, INEQ_{t-j} + \sum_{j=1}^{p} \gamma_{13j} \Delta ln\big(HP_{t-j}\big) + \alpha_1 [ln(HP_{t-1}) - \beta_2\, INEQ_{t-1}] \\ \Delta\, INEQ_t &= \gamma_{21} + \textstyle\sum_{j=1}^{p} \gamma_{22j} \Delta\, INEQ_{t-j} + \sum_{j=1}^{p} \gamma_{23j} \Delta ln\big(HP_{t-j}\big) + \alpha_2 [ln(HP_{t-1}) - \beta_2\, INEQ_{t-1}]\end{aligned} \qquad (6)$$

The null of weak exogeneity, and no long-run Granger non-causality (LRGNC), uses t-tests on α_1 and α_2 with the following alternative hypotheses:

- $H_A^1\colon \alpha_1 \neq 0$ implies that $ln(HP_t)$ is not weakly exogenous with respect to the parameters in the equation for $\Delta\, INEQ_t$ and so $INEQ_t$ Granger-causes $ln(HP_t)$ in the long-run.
- $H_A^2\colon \alpha_2 \neq 0$ implies that $INEQ$ is not weakly exogenous with respect to the equation for $\Delta ln(HP_t)$ and that $ln(HP_t)$ Granger-causes $INEQ_t$ in the long-run.

In applying the LRGNC tests we estimate system (6) for each country with time-series regressions using previously specified cointegrating equations to define the error-correction terms.[12] We subtract the mean of these error-correction terms to produce new zero mean error-correction terms to be used in a slightly modified version of (6) when applying the LRGNC tests. The lag lengths for each country are determined using the AIC with a maximum lag of $p = 3$.

[12] LRGNC tests are only applied to models with evident cointegration because the error-correction term will only be stationary if there is cointegration. Without cointegration, application of LRGNC tests involve the regression of a stationary dependent variable on a non-stationary error-correction term. This will make standard critical values used in the LRGNC tests inappropriate due to a spurious significance problem (Stewart, 2011) and hence bias inference towards rejecting weak exogeneity.

3.4 *Robustness checks*

Finally, we conduct two robustness checks. The first is a bivariate cointegration analysis that tests if income (real GDP, retrieved from AMECO) is a determinant of $ln(HP_t)$:

$$ln(HP_t) = \beta_1 + \beta_2 \ln(GDP)_t + u_t \tag{7}$$

As mentioned in the introduction, traditionally the level and growth of income have been suggested to be an important house price determinant (ECB, 2003; Sommer *et al.*, 2013). Moreover, Figure 1 illustrates that with a given income distribution absolute inequality increases when overall income increases. Hence, it might be that OECD house prices were co-moving with income rather than with absolute inequality.

The second robustness check applies cointegration tests to the following two trivariate models that account for monetary policy (r) as a proxy for credit market conditions:

$$ln(HP_t) = \beta_1 + \beta_2 INEQ_t + \beta_3 \mathrm{r}_t + u_t \tag{8}$$

$$ln(HP_t) = \beta_1 + \beta_2 \ln(GDP)_t + \beta_3 \mathrm{r}_t + u_t \tag{9}$$

It is expected that low interest rates increase the access to and lower the financing costs of mortgages, leading to an increase in the demand for houses. Hence expansionary monetary policy is the most studied potential driver behind the upsurge in OECD house prices (see e.g. Taylor, 2007; Goodhart and Hofman, 2008; Gerdesmeier *et al.*, 2010; Agnello and Schuknecht, 2011; Dokko *et al.*, 2011, Bordo and Landon-Lane, 2013a). The proxy used is the nominal 3-month nominal interbank interest rate (adjusted with consumer price inflation), which is readily available from OECD.stat and the St. Louis Fed (Germany and Ireland).

4. The impact of (absolute) income inequality on house prices

4.1 Are Real House Prices and Income Inequality Cointegrated?

The panel unit root tests suggest that $ln(HP_t)$ and our four inequality measures are at least I(1) for the vast majority of the 18 countries (see Appendix). To be more precise, the number of countries that are found to be I(1) according to at least one of the two test are: 13 for the logged real house price index ($ln(HP_t)$), 16 for the Gini coefficient ($Gini_t$), 15 for the top 5% income share ($Top5\%_t$), 17 for the logged income variance ($ln(Var_t)$), and 16 for the logged market income of the top 5% ($ln(Top5\$_t)$).

That not all countries' variables are I(1) may be due to factors such as Type I errors. Hence, we treat all series as if they are I(1), satisfying the necessary condition for cointegration, and proceed to conduct tests of cointegration.[13] If the assumption that the necessary condition for cointegration being satisfied is incorrect this should manifest itself in the rejection of cointegration.

We therefore proceed to test for cointegration between $ln(HP_t)$ and $INEQ_t$ by applying Westerlund's (2007) panel cointegration test (Table 1). For the two relative inequality measures, $Gini_t$ and $Top5\%_t$, all four tests for both sets of deterministic terms cannot reject the null hypothesis. Hence, it is unambiguous that there is no evidence of cointegration between $ln(HP_t)$ and $Gini_t$ and $ln(HP_t)$ and $Top5\%_t$. For $ln(Var_t)$ and $ln(Top5\$_t)$ all four tests indicate cointegration at the 5% level when the intercept is the only deterministic term included in the model. When both an intercept and trend are included in the model two tests, G_τ and P_τ, indicate cointegration at the 5% level for both absolute inequality variables.

< Table 1 >

[13] If some of the series are I(0) this should not be an issue because the ADL method can identify error-correction relationships when some series are I(1) and others are I(0), see Pesaran *et al.* (2001).

4.2 The Long-Run Relationship between House Prices and Absolute Inequality

Given the general evidence in favour of cointegration with homogeneous long-run coefficients across all 18 countries for both absolute measures of inequality we report their implied estimated homogeneous long-run relationships in Table 2. When both an intercept and trend are included in the model the trend term is not significant. This suggests that the trend term can be excluded from the long-run equation and that cointegrating equations including a trend should not be favoured. This is consistent with the model including both intercept and trend providing less support for cointegration than the model where the intercept is the only deterministic term (see Table 1). Hence, we favour inference from the models where the intercept is the only deterministic term. This finding also suggests that there are no omitted variables from the long-run equations that approximately follow a linear trend.

< Table 2 >

In the two long-run models where the only deterministic term included is an intercept the inequality measures are both significant at the 1% level and exhibit the expected positive coefficient sign. Given the double log specifications the coefficients can be interpreted as elasticities. According to the overall absolute inequality measure (Var_t) a 1% rise in absolute inequality leads to around a 0.39% increase in real house prices, while a 1% rise in the top 5% market income ($Top5\$_t$) leads to an approximate 0.78% increase in real house prices.

Table 3 (second row) reports panel dynamic OLS (DOLS) estimates of the long-run relationships assuming homogeneous coefficients across countries and with only an intercept included as a deterministic term in the model for both inequality measures where cointegration was found. Both inequality measures are significant at the 1% level and the estimated elasticities are around 0.30 for Var_t and 0.61 for $Top5\$_t$. Whilst slightly lower than the estimates implied by the Westerlund (2007) model they are not too dissimilar. This suggests that the results are broadly robust in the sense of positive and significant coefficients on the inequality measures as well as the coefficient on $Top5\$_t$ being around twice as large as that on Var_t.

< Table 3 >

Whilst our tests suggest that the cointegrating equations are homogeneous across all 18 countries we also report DOLS estimates of the long-run equations for each of the individual countries in Table 3 (rows 3-20). The general results are robust across both absolute inequality measures in the following ways. First, for 14 countries the coefficient on absolute inequality (however measured) is positive and significant at the 1% level. Second, for one country (SWE) this coefficient is positive and only significant at the 10% level. Third, for one country (JAP) this coefficient is positive if insignificant. Fourth, for two countries (GER and KOR) the coefficient on inequality is negative and significant. Hence, while these results may arguably be interpreted as supporting the homogeneity of the coefficient on inequality for 15 of the countries (in the sense that it is positive and significant) there are doubts that this homogeneity extends to GER, JAP, KOR.

4.3 Direction of Causation

The Westerlund (2007) panel cointegration tests on the reverse regression with inequality as the dependent variable regressed on $ln(HP_t)$ reject cointegration for all four inequality measures regardless of the deterministic specification (Table 4). Hence, for all measures of inequality this suggests that inequality is weakly exogenous with respect to $ln(HP_t)$ and that the cointegration results reported in Table 1 are not subject to low power due to the violation of weak exogeneity. A further implication of the suggestion of the two absolute measures of inequality ($ln(Var_t)$ and $ln(Top5\$_t)$) being weakly exogenous with respect to $ln(HP_t)$ is that there is uni-directional long-run Granger-causality from absolute inequality to $ln(HP_t)$ and no reverse causality in the opposite direction.

< Table 4>

The individual country probability values of t-tests for LRGNC based on time-series regressions (Table 5) confirm the above finding that for the overwhelming majority of countries there is no evident violation of the weak exogeneity assumption, which implies that the cointegration results from the Westerlund (2007) tests reported above are valid. To interpret our results we use a 5% level of significance. For 13 countries there is evidence that, in the long-run, $ln(Var_t)$ Granger-causes $ln(HP_t)$ and that $ln(HP_t)$ does not Granger-cause $ln(Var_t)$. For two countries (CAN and NEW) there is bi-directional long-run Granger-causality, for three countries (IRE, JAP and SPA) there is evidence of no long-run Granger-causality in either direction, and for no country is there evidence of uni-directional long-run Granger-causality from $ln(HP_t)$ to $ln(Var_t)$.

< Table 5 >

For the $ln(Top5\$_t)$ measure of inequality there is evidence of uni-directional long-run Granger causality to $ln(HP_t)$ for 11 countries. For one country (NEW) there is evidence of bi-directional long-run Granger-causality, for five countries (AUS, GER, IRE, JAP and SWE) there is no long-run Granger causality in either direction and for one country (SPA) there is evident uni-directional long-run Granger causality from $ln(HP_t)$ to $ln(Top5\$_t)$.

Overall, the time-series Granger-causality test results from Table 5 show that for the vast majority of countries in our sample the direction of Granger-causality is from absolute inequality to real house prices. The anomalies found may be due to small (time-series) sample effects, Type I errors and questionable cointegrating equations in the case of GER and JAP (as reported in Table 3).

4.4 Is it Inequality or Income that Drives House Prices?

In Section 4.1 it was established that the two absolute inequality measures cointegrated with $ln(HP_t)$ on their own. We next consider whether the natural logarithm of real GDP (denoted $ln(GDP_t)$) also cointegrates with $ln(HP_t)$.

It is first important to note that $ln(GDP_t)$ is I(1) for 16 countries and at least I(2) for the other 2 countries (FIN and UKD) according to at least one of the two panel unit root tests (see Appendix). Therefore, it is possible that $ln(GDP_t)$ cointegrates on its own with $ln(HP_t)$ given that they generally have the same orders of integration.

Table 6 reports the bivariate Westerlund (2007) statistics for the null of no cointegration between $ln(HP_t)$ and $ln(GDP_t)$, and shows that there is no evidence of cointegration. The lack of evident cointegration between $ln(HP_t)$ and $ln(GDP_t)$ is in line with Knoll *et al.*'s (2017) observation that real house price growth has significantly outpaced income growth during the period under study. This, as well as our finding that the absolute inequality measures cointegrate with house prices, implies that house price growth is not due to overall income growth but instead due to an increasingly unequal distribution of income.

< Table 6 >

4.5 Does the Inclusion of Monetary Policy Change the Results?

We next consider whether these results stay robust when the real short-term interest rate (r_t) is considered as a covariate. First, we find that the interest rate series is unlikely to be cointegrated on its own with $ln(HP_t)$ because in many cases they have a different order of integration. While $ln(HP_t)$ is at least I(1) for the majority of countries, the Cerrato *et al.* (2009, 2011) and Pesaran (2007) based tests suggests that (r_t) is I(0) for 11 countries (BEL, CAN, FIN, GER, ITA, KOR, NEW, NOR, NET, SPA and SWE) and I(1) for the remaining seven countries (see Appendix).

That the real interest is I(0) for many countries is consistent with the Fisher hypothesis, (Malliaropulos (2000), Costantini and Lupi (2007), Omay and Yuksel (2015), Panopoulou and Pantelidis (2016)). However, (r_t) can still potentially form part of the cointegrating relationship with $ln(HP_t)$ when it is considered a covariate with another I(1) explanatory variable – see, for example, Pesaran *et al.* (2001) for a discussion of including I(0) and I(1) variables in an ADL model's equilibrium.

Table 7 reports Westerlund's (2007) cointegration tests for trivariate regressions of $ln(HP_t)$ on r_t and $INEQ_t$, and $ln(HP_t)$ on r_t and $ln(GDP_t)$. The results unambiguously indicate no evident cointegration for the models involving the two relative inequality variables. The same is true when the trivariate regressions include $ln(Top\$_t)$ and r_t as covariates. This is a surprising result given that $ln(HP_t)$ cointegrates with $ln(Top\$_t)$ in bivariate regressions. A potential explanation could be reduced efficiency due to increased covariates that raise (lower) the coefficient standard error (t-ratio) of the adjustment coefficient upon which the cointegration tests are based. Another potential explanation could be that a violation of weak exogeneity may reduce the power Westerlund's cointegration test. However, unreported results (available upon request) do not suggest cointegration in the regression of $ln(Top\$_t)$ on $ln(HP_t)$ and r_t based on Westurlund's tests. This is indicative of

$ln(Top\$_t)$ being weakly exogenous with respect to the parameters in the regression of $ln(HP_t)$ on r_t and $ln(Top\$_t)$.[14]

The trivariate model containing the explanatory variables $ln(GDP_t)$ and r_t also does not suggest cointegration at the 5% level. Hence, the only trivariate regression that suggests evidence of cointegration at the 5% level contains $ln(Var_t)$ and r_t. While there is some ambiguity over the support for cointegration (four out of eight tests indicate cointegration at the 5% level)[15], these specifications exhibit the most convincing evidence favouring cointegration of the trivaraite models. Hence, the results of Table 7 broadly confirm our main finding of Section 4.1 that absolute inequality is seemingly cointegrated with real house prices, whereas relative inequality and income are not.

< *Table 7* >

We therefore proceed in estimating the long-run relationship for the regression of $ln(HP_t)$ on $ln(Var_t)$ and r_t. Given our relatively small time-series dimension panel DOLS equilibrium estimates are arguably more efficient than those obtained from Westerlund's ADL model. The panel DOLS results are reported in Table 8 (row 2). The coefficients on both $ln(Var_t)$ and r_t have the expected sign although only $ln(Var_t)$ is significant at the 5% level. The regression also has coefficients of plausible magnitudes that suggest a 1% rise in inequality induces an increase in house prices of around 0.4%, and a 1% rise in real interest rates causes house prices to fall by about 1.2% (if this interest rate effect is strictly insignificant).

< *Table 8* >

[14] Similarly, unreported results indicate that r_t is weakly exogenous with respect to the parameters in the regression of $ln(HP_t)$ on r_t and $ln(Top\$_t)$.
[15] This ambiguity may be due to some loss of efficiency because of the number of variables included in the estimated models.

The coefficient of $ln(Var_t)$ from this trivariate regression is lower than that obtained with the bivariate cointegration results (reported in Table 3). This suggests that the addition of interest rates has impacted this estimated coefficient. Nevertheless, this trivariate regression supports cointegration between house prices and absolute inequality and is consistent and confirming of our bivariate cointegration analysis.

Finally, Table 8 (row 3-20) also presents individual country long-run relationships for $ln(HP_t)$ on $ln(Var_t)$ and r_t. With the exception of GER, JAP and KOR all countries exhibit a significant and positive relationship between absolute inequality and real house prices. Again, these findings are in line with the bivariate results (presented in Table 3). The real interest rate, on the other hand, is only significant and has the expected negative sign in six out of the 18 countries. Hence, there is some ambiguity as to whether the interest rate is part of the equilibrium relationship because it is insignificant for most countries.

Investigation of this potential heterogeneity would be an avenue where future research could be directed. However, regardless of whether we prefer the trivariate regression of $ln(HP_t)$ on $ln(Var_t)$ and r_t or the bivariate regression of $ln(HP_t)$ on $ln(Var_t)$ it is clear that $ln(Var_t)$ is significant and has a positive coefficient in the vast majority of our sample of countries.

5. Conclusions

Our results provide two novel insights. First, increasing income inequality contributed to the rise in real house prices in 15 out of 18 OECD countries during the period 1975-2010. Second, the results are sensitive to the use of relative and absolute inequality measures.

To be more precise, the bivariate cointegration analysis suggests that the natural logarithm of the variance ($ln(Var_t)$) and the natural logarithm of the market income of the top 5%

($ln(Top5\$_t)$) individually form irreducible cointegrating equations with $ln(HP_t)$ with theoretically plausible coefficients in 15 out of the 18 countries. There is little ambiguity in these cointegration results and the causation is from inequality to house prices.

Together with absolute inequality, the short-term real interest rate (r_t) also shows some evidence of cointegration with $ln(HP_t)$ if the interest rate only has a significant and negative coefficient in six countries. The two relative inequality measures used, on the other hand, do not show any signs of cointegration. The same is true for real GDP, which suggests that the significance of the absolute inequality measures cannot be attributed to an overall growth in income but to its increasingly unequal distribution.

The finding that the recent surge in house prices was partly driven by rising absolute income inequality contributes to a growing literature that finds that the recent inequality increase in developed countries has important socio-economic effects (see e.g. OECD, 2015). Moreover, it suggests that the current focus on relative inequality measures is unduly restrictive and that more attention should be given to alternative inequality measures like the ones presented in this article.

An area for future research may be in the application of time-series cointegration testing by country to further investigate the degree of homogeneity of the relationship between house prices and inequality across countries. This will become possible in the future as time-series with a sufficient sample size to obtain reliable results become available. However, the time-series dimension available to us here made the application of panel cointegration methods more appropriate.

APPENDIX

To establish the (non-)stationarity of the data the unit root tests of Pesaran (2007) and Cerrato *et al.* (2009, 2011) are used. Pesaran's (2007) test assumes linear adjustment, can deal with cross-sectional dependence and is based upon the following time-series regression estimated for each i:

$$\Delta y_{i,t} = a_i^P + \alpha_i^P t + b_i^P y_{i,t-1} + c_{i,0}^P \bar{y}_{t-1} + \sum_{j=1}^{p_i} c_{i,j}^P \Delta y_{i,t} + \sum_{j=0}^{p_i} d_{i,j}^P \Delta \bar{y}_{t-j} + u_{i,t}^P \qquad \text{(1A)}$$

where, $i = 1, 2, \dots, N$; $t = 1, 2, \dots, T$, $\Delta \bar{y}_t = \frac{1}{N}\sum_{i=1}^{N} y_{i,t}$ and $\bar{y}_{t-1} = \frac{1}{N}\sum_{i=1}^{N} y_{i,t-1}$.

The null hypothesis is that there is a unit root for all cross-sectional units, $b_i^P = 0 \; \forall \; i$ while the alternative is that $y_{i,t}$ is stationary for at least one cross-section, $b_i^P < 0$ for *at least* one i. The CADF statistic for each cross-section is the ordinary least squares (OLS) t-ratio corresponding to b_i^P, denoted $t_i^P(N,T) = \frac{\hat{b}_i^P}{s_{\hat{b}_i^P}}$. The panel test statistic, $CIPS$, is:

$$CIPS = \frac{1}{N}\sum_{i=1}^{N} t_i^P(N,T) \qquad \text{(2A)}$$

The version of the test that we use is, following the scheme given in Pesaran (2007) and denoted $t_i^{P^*}(N,T)$, thus:

$$CIPS^* = \frac{1}{N}\sum_{i=1}^{N} t_i^{P^*}(N,T) \qquad \text{(3A)}$$

Cerrato *et al.*'s (2009, 2011) heterogeneous nonlinear panel unit root tests involves estimating the following nonlinear auxiliary regression:

$$\Delta y_{i,t} = a_i^C + \alpha_i^C t + b_i^C y_{i,t-1}^3 + c_{i,0}^C \overline{y_{t-1}^3} + \sum_{j=1}^{p_i} c_{i,j}^C \Delta y_{i,t} + \sum_{j=0}^{p_i} d_{i,j}^C \Delta \bar{y}_{t-j} + u_{i,t}^C \qquad \text{(4A)}$$

where, $\overline{y_{t-1}^3} = \frac{1}{N}\sum_{i=1}^{N} y_{i,t-1}^3$. A time trend, t, is included following Cerrato *et al.* (2013) and the lag length, p_i, can be determined using information criteria.

The null hypothesis is $b_i^C = 0 \; \forall \; i$, while the alternative is $b_i^C < 0$ for *at least* one i. The t-ratios for each cross-section, denoted $t_i^C(N,T) = \frac{\hat{b}_i^C}{s_{\hat{b}_i^C}}$, where $\hat{b}_i^C$ is the OLS estimate of b_i^C and $s_{\hat{b}_i^C}$ is the corresponding OLS coefficient standard error, are used to calculate the panel test statistic, thus:

$$\bar{t}(N,T) = \frac{1}{N}\sum_{i=1}^{N} t_i^C(N,T) \qquad (5A)$$

If the test statistic is not more negative than the critical value, reported in Cerrato *et al.* (2009 and 2011), the null hypothesis cannot be rejected. Simulations indicate that this test has superior size and power than Pesaran's (2007) test when the data generating process is nonlinear.

Table A.1 reports Pesaran's (2007) panel unit root test using the SPSM procedure applied to $ln(HP_t)$, $GINI_t$, $Top5\%_t$, $ln(VAR_t)$, $ln(Top5\$_t)$, $ln(GDP_t)$ and r_t. The null hypothesis that all 18 countries' series in the panel are $I(1)$ cannot be rejected for all variables, except r_t, regardless of the deterministic specification of the test equations. In the case of r_t, 7 countries' series are I(0) around a constant mean (BEL, CAN, ITA, KOR, NET, SPA and SWE), no countries' series are trend stationary and the 11 remaining countries' series are at least I(1).

< *Table A.1* >

Table A.2 reports Pesaran's (2007) panel unit root test using the SPSM procedure applied to the first difference of the variables tested in Table A.1. The unit root test results based upon the Pesaran (2007) method suggest that $ln(HP_t)$ is at least $I(2)$ for virtually all countries whereas the inequality measures are $I(1)$ for most nations. For $ln(GDP_t)$ 8 countries' series are I(1) (AUS, BEL, CAN, DEN, IRE, ITA, KOR and the USA) and the 10 remaining countries' series are at least I(2). For r_t 7 countries' series are I(0) around a constant mean (BEL, CAN, ITA, KOR, NET, SPA and SWE), no countries' series are trend stationary and the 11 remaining countries' series are $I(1)$.

< *Table A.2* >

The results based on the Pesaran (2007) test suggest that the necessary condition for cointegration between $ln(HP_t)$ and all of the inequality measures as well as r_t is violated for most countries. However, given that we expect $ln(HP_t)$ to be $I(1)$ we consider the possibility that this result is due to low power (possibly due to structural breaks) and, in our cointegration analysis, we treat $ln(HP_t)$, and all measures of inequality, as if they are $I(1)$ for all countries.[16] If $ln(HP_t)$ is $I(2)$ for most countries it will not cointegrate with the generally $I(1)$ inequality variables for those countries and our cointegration test results will reveal this.

Similarly, the finding that many countries' $ln(GDP_t)$ series are at least $I(2)$ may also be due to structural breaks and we consider the possibility that this is $I(1)$ in our cointegration analysis. However, r_t being $I(0)$ for many countries and no more than $I(1)$ for any country suggests that it is unlikely to cointegrate with $ln(HP_t)$, on its own, for many nations. Further, when we apply unit root tests that allow for nonlinear adjustment we find far more widespread evidence that most countries' series are $I(1)$ according to at least one test. These results are presented below.

Table A.3 reports Cerrato *et al.*'s (2011) heterogeneous nonlinear panel unit root tests (using the SPSM procedure) for the levels of the variables tested above. This test accommodates cross-sectional dependence and extends the Pesaran *et al.* (2007) method that assumes a linear adjustment process by allowing nonlinear adjustment. Such nonlinear adjustment could look like structural breaks without being confined to a single once-and-for-

[16] It may be that for some countries both $ln(HP)$ and the inequality variable are found to be $I(2)$ according to our test results and in fact are both $I(1)$ around structural breaks and they are found to cointegrate because they cointegrate and co-break.

all jump at one particular point in time. Hence, when there is such nonlinear adjustment Cerrato *et al.*'s (2011) test should be more powerful than that of Pesaran (2007).

The results reported in Table A.3 indicate that all series are at least $I(1)$ for all countries except for $ln(HP_t)$, where there is evidence of stationarity for 3 countries (Finland, Japan and New Zealand), $GINI_t$ where there is evidence of stationarity for 1 country (the Netherlands) and r_t where there is evidence of stationarity for 8 countries (Belgium, Finland, Germany, Italy, New Zealand, Norway, Spain and Sweden) and trend stationarity for 1 country (Canada). Excepting these minor anomalies (that may be due to, for example, Type I errors) these results broadly confirm those from Pesaran's (2007) test that all 5 (non interest rate) series are at least $I(1)$ for all countries.

< *Table A.3* >

Table A.4 reports Cerrato *et al.*'s (2011) heterogeneous nonlinear panel unit root test using the SPSM procedure applied to the first difference of the variables tested in Table A.3. Considering both panel unit root tests (allowing for both linear and nonlinear adjustment) we find evidence that all series are $I(1)$ for the vast majority of countries (with the exception of interest rates where 11 countries series are $I(0)$ and 7 are $I(1)$). Further, any anomalies may be due to factors such as Type I errors. Hence, we treat all series as if they are $I(1)$, satisfying the necessary condition for cointegration, and proceed to conduct tests of cointegration.[17] If the assumption that the necessary condition for cointegration being satisfied is incorrect this will cause our tests to reject cointegration.

< *Table A.4* >

[17] If some of the series are $I(0)$ this should not be an issue because the ADL method can identify error-correction relationships when some series are $I(1)$ and others are $I(0)$ – although the critical values in Westerlund (2007) assume $I(1)$ variables. Further, if some series are trend stationary this can be accounted for in our application of the Westerlund (2007) procedure because we apply the method incorporating just an intercept and both an intercept and linear trend.

REFERENCES

Agnello, L. and Schuknecht, L. (2011). Booms and busts in housing markets: Determinants and implications. *Journal of Housing Economics*, 20(3): 171-90.

Atkinson, A. B. and Brandolini, A. (2010). On analyzing the world distribution of income. *World Bank Economic Review*, 24(1): 1–37.

Bonesmo Fredriksen, K. (2012). “Less Income Inequality and More Growth – Are they Compatible? Part 6. The Distribution of Wealth”. OECD Economics Department Working Papers, No. 929.

Bordo, M.D. and Landon-Lane, J. (2013a). Does Expansionary Monetary Policy Cause Asset Price Booms; some historical and empirical evidence. NBER Working Paper, No. 19585.

Bordo, M.D. and Landon-Lane, J. (2013b). What Explains House Price Booms?: History and Empirical Evidence. NBER Working Paper, No. 19584.

Campbell, J.Y. and Cocco, J.F. (2007). How do house prices affect consumption? Evidence from micro data. *Journal of Monetary Economics*, 54(3): 591-621.

Case, K.E. and Shiller, R.J. (2003). Is There a Bubble in the Housing Market? *Brookings Papers on Economic Activity*, 34(2): 299-362.

Case, K.E., Quigley, J.M. and Shiller, R.J. (2005). Comparing Wealth Effects: The Stock Market versus the Housing Market. *B.E. Journal of Macroeconomics*, 5(1): 1-34.

Case, K.E., Quigley, J.M. and Shiller, R.J. (2013). Wealth Effects Revisited 1975-2012. *Critical Finance Review*, 2(1): 101-128.

Cerrato, M., Peretti C.D., Larsson R. and Sarantis, N. (2009). “A nonlinear panel unit root test under cross section dependence”, Discussion Paper 2009-28, Department of Economics, University of Glasgow.

Cerrato, M., Peretti C.D., Larsson R. and Sarantis, N., (2011). “A Nonlinear Panel Unit Root Test under Cross Section Dependence”, Scottish Institute for Research in Economics Discussion Papers, No. 2011-08.

Cesa-Bianchi, A., Cespedes, L.F. and Rebucci, A. (2015). Global Liquidity, House Prices, and the Macroeconomy: Evidence from Advanced and Emerging Economies. *Journal of Money, Credit and Banking*, 47(1): 301-335.

Chakravarty, S.R. (2001). The variance as a subgroup decomposable measure of inequality. *Social Indicators Research*, 53(1): 79-95.

Chortareas, G. and Kapetanios G. (2009). Getting PPP right: Identifying mean-reverting real exchange rates in panels. *Journal of Banking and Finance*, 33: 390-404.

Collinson, P. (2015). “Average house price rises to 8.8 times local salary in England and Wales”. The Guardian online, 6 August.

Costantini, M. and Lupi, C. (2007). An analysis of inflation and interest rates. New panel unit root results in the presence of structural breaks. *Economics Letters*, 95(3): 408–414.

Cowell, F., Karagiannaki, E. and McKnight, A. (2012). ‘Accounting for Cross-Country Differences in Wealth Inequality’. LWS Working Paper Series, No. 13.

Dewilde, C. and Lancee, B. (2013). Income Inequality and Access to Housing in Europe. *European Sociological Review*, 29(6): 1189–1200.

Dokko, J., Doyle, B.M., Kiley, M.T., Kim, J., Sherlund, S., Sim, J., and van Den Heuvel, S. (2011). Monetary policy and the global housing bubble. *Economic Policy*, 26(4): 233-283.

Easterly, W. and Fischer, S. (2001). Inflation and the Poor. *Journal of Money, Credit and Banking*, 33(2): 160-178.

ECB (2003). *Structural Factors in the EU Housing Market*. Frankfurt: European Central Bank.

Gerdesmeier, D., Reimersz, H.-E. and Roffia, B. (2010). Asset Price Misalignments and the Role of Money and Credit. *International Finance*, 13(3): 377–407.

Goda, T. and Torres García, A. (2017). The Rising Tide of Absolute Global Income Inequality During 1850–2010: Is It Driven by Inequality Within or Between Countries?. *Social Indicators Research*, 130(3): 1051-1072.

Goda, T., Onaran, O. and Stockhammer, E. (2017). Income Inequality and Wealth Concentration in the Recent Crisis. *Development and Change*, 48(1): 3-27.

Goodhart, C. and Hofmann, B. (2008). House prices, money, credit, and the macroeconomy. *Oxford Review of Economic Policy*, 24(1): 180–205.

Gyourko, J., Mayer, C. and Sinai, T (2013). Superstar Cities. *American Economic Journal: Economic Policy*, 5(4): 167-99.

Hryshko, D., Luengo-Prado, M.J. and Sørensen, B.E. (2010). House prices and risk sharing. *Journal of Monetary Economics*, 57(8): 975-987.

Im, K. S., Pesaran, M. and Shin, Y. (2003). ‘Testing for unit roots in heterogeneous panels’, *Journal of Econometrics*, 115: 53–74.

Jordà, O., Schularick, M. and Taylor, A.M. (2014). The Great Mortgaging: Housing Finance, Crises, and Business Cycles. NBER Working Paper, No. 20501.

Knoll, K., Schularick, M. and Steger, T.M. (2017). No Price Like Home: Global House Prices, 1870-2012. *American Economic Review*, 107(2): 331-351.

Määttänen, N. and Terviö, M. (2014). Income distribution and housing prices: An assignment model approach. *Journal of Economic Theory*, 151: 381-410.

Malliaropulos, D. (2000). A note on nonstationarity, structural breaks, and the Fisher effect. *Journal of Banking and Finance*, 24(5): 695–707.

Matlack, J.L. and Vigdor, J.L. (2008). Do rising tides lift all prices? Income inequality and housing affordability. *Journal of Housing Economics*, 17(3): 212-224.

Mian, A. and Sufi, A. (2015). *House of Debt: How They (and You) Caused the Great Recession, and How We Can Prevent It from Happening Again*. Chicago: University of Chicago Press.

Mian, A., Rao, K. and Sufi, A. (2013). Household Balance Sheets, Consumption, and the Economic Slump. *Quarterly Journal of Economics*, 128(4), 1687–1726.

Nakajima, M. (2005). ‘Rising Earnings Instability, Portfolio Choice and Housing Prices’. Mimeo, University of Illinois, Urbagna Champaign.

OECD (2015). In It Together: Why Less Inequality Benefits All. Paris: Organisation for Economic Co-operation and Development.

Omay, T. and Yuksel, A. and Yuksel, A. (2015). An empirical examination of the generalized Fisher effect using cross-sectional correlation robust tests for panel cointegration. *Journal of International Financial Markets, Institutions and Money*, 35: 18–29.

Panopoulou, E. and Pantelidis, T. (2016). The Fisher effect in the presence of time-varying coefficients. *Computational Statistics and Data Analysis*, 100: 495–511.

Persyn, D. and Westerlund, J. (2008). Error Correction Based cointegration Tests for Panel Data. *Stata Journal*, 8(2): 232-241.

Pesaran, M. H. (2007). A simple panel unit root test in the presence of cross section dependence. *Journal of Applied Econometrics*, 22(2): 265-312.

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): 289–326.

Ravallion, M. (2004). Competing concepts of inequality in the globalization debate. In Collins, S.M. and Graham, C. (Eds.), *Brookings trade forum 2004. Globalization, poverty, and inequality*. Washington: Brookings Institution Press, pp. 1–38.

Sá, F., Towbin, P. and Wieladek, T. (2014). Capital inflows, financial structure and housing booms. *Journal of the European Economic Association*, 12(2): 522–546.

Sommer, K., Sullivan, P. and Verbrugge, R. (2013). The equilibrium effect of fundamentals on house prices and rents. *Journal of Monetary Economics*, 60(7): 854-870.

Stewart, C. (2011) ‘A note on spurious significance in regressions involving $I(0)$ and $I(1)$ variables’, *Empirical Economics*, 41(3): 565–571.

Stroebel, J. and Vavra, J. (2014). ‘House Prices, Local Demand and Retail Prices’. NBER Working Paper, No. 20710.

Taylor, J.B. (2007). ‘Housing and Monetary Policy’. NBER Working Paper, No. 13682.

Westerlund, J. (2007). Testing for Error Correction in Panel Data. *Oxford Bulletin of Economics and Statistics*, 69(6): 709-748.

Zhang, F. (2016). Inequality and House Prices. University of Michigan Department of Economics, Job Market Paper.

Table 1. Robust p-values for Westerlund's (2007) panel cointegration tests of $ln(HP_t)$ on $INEQ_t$

	$Gini_t$		$Top5\%_t$		$ln(Var_t)$		$ln(Top5\$_t)$	
	Int	*Trend*	*Int*	*Trend*	*Int*	*Trend*	*Int*	*Trend*
G_τ	0.888	0.310	0.910	0.326	**0.000***	**0.013**	**0.003***	**0.029**
G_α	0.764	0.279	0.814	0.191	**0.023**	0.088*	**0.048**	0.270
P_τ	0.359	0.194	0.319	0.143	**0.000***	**0.010**	**0.001***	**0.064**
P_α	0.328	0.341	0.299	0.268	**0.005***	0.080*	**0.003***	0.115
Leads	1.28	1.28	1.28	1.28	1.22	1.50	1.22	1.61
Lags	1.00	1.00	1.06	1.00	1.11	1.22	1.11	1.28

Table 1 notes. The first row denotes the inequality measure involved in the potential cointegrating equation with $ln(HP_t)$ as the dependent variable. The second row specifies the deterministic terms included in the cointegration equation as Int when only an intercept is included and Trend when both an intercept and trend are included. G_τ and G_α denote the tests when the alternative hypothesis is that there is cointegration for at least one country in the panel. P_τ and P_α denote the tests when the alternative hypothesis is that there is cointegration for all 18 countries in the panel. All four tests are based on either OLS or heteroscedasticity and autocorrelation consistent (HAC) coefficient standard errors, respectively. The average number of leads and lags (selected with the AIC) used in the 18 countries' error-correction models are specified in the rows labelled Leads and Lags, respectively. A maximum of 3 leads and lags are allowed. *, ** and *** denote rejection of the non-cointegration null at the 10%, 5% and 1% levels, respectively.

Table 2. Estimated panel long-run relationship and short-run adjustment for $ln(HP_t)$

	$ln(Var_t)$		$ln(Top5\$_t)$	
	Int	*Trend*	*Int*	*Trend*
$INEQ_t$	**0.387***** (4.41)	0.209 (0.70)	**0.783***** (4.57)	0.131 (0.21)
Intercept	-3.380* (-1.93)	-10.992 (-0.28)	-5.078** (-2.47)	-29.656 (-0.82)
Trend		0.006 (0.25)		0.016 (0.75)
Adjustment	-0.164*** (-6.61)	-0.251*** (-6.32)	-0.158*** (-6.44)	-0.246*** (-8.22)

Table 2 notes. See notes to Table 1. The estimated long-run coefficients, with t-ratios given in parentheses, are reported for each measure of inequality specified in the top row, where $ln(HP_t)$ is the dependent variable.

Table 3. Bivariate DOLS long-run relationship of $ln(HP_t)$ on $INEQ_t$

	$ln(Var_t)$	$ln(Top5\$_t)$
PANEL	0.302*** (10.61)	0.612*** (10.99)
AUS	0.531*** (7.58)	1.103*** (7.73)
BEL	0.624*** (5.64)	1.089*** (4.40)
CAN	0.348*** (12.05)	0.720*** (10.48)
DEN	0.427*** (3.97)	0.861*** (4.17)
FIN	0.263*** (5.69)	0.522*** (5.64)
FRA	0.486*** (3.0)	0.874*** (2.91)
GER	-0.117*** (-3.48)	-0.193*** (-3.17)
IRE	0.508*** (8.30)	0.989*** (9.92)
ITA	0.227*** (3.53)	0.446*** (3.77)
JAP	0.011 (0.23)	0.047 (0.48)
KOR	-0.136*** (-4.72)	-0.204*** (-3.48)
NET	0.924*** (5.21)	1.881*** (6.33)
NEW	0.428*** (12.81)	0.874*** (11.18)
NOR	0.409*** (7.27)	0.822*** (7.19)
SPA	0.554*** (7.15)	1.188*** (7.20)
SWE	0.242* (1.91)	0.473* (1.84)
UKD	0.531*** (8.43)	1.035*** (7.63)
USA	0.147*** (4.91)	0.315*** (5.24)

Table 3 notes. The DOLS estimated long-run coefficients, with t-ratios based on HAC standard errors given in parentheses, are reported for each measure of inequality specified in the top row, where $ln(HP_t)$ is the dependent variable. Leads and lags are chosen using the AIC with a maximum of 3 leads and 3 lags with only an intercept included as a deterministic term. The second row shows the estimated long-run relationship of the whole panel and rows 3 – 20 show the individual country effects.

Table 4. Robust p-values for Westerlund's (2007) panel cointegration tests of $\boldsymbol{INEQ_t}$ on $\boldsymbol{ln(HP_t)}$

	$\boldsymbol{Gini_t}$		$\boldsymbol{Top5\%_t}$		$\boldsymbol{ln(Var_t)}$		$\boldsymbol{ln(Top5\$_t)}$	
	Int	*Trend*	*Int*	*Trend*	*Int*	*Trend*	*Int*	*Trend*
G_τ	0.760	0.239	0.793	0.495	1.000	0.408	1.000	0.278
G_α	0.936	0.565	0.946	0.809	1.000	0.410	1.000	0.199
P_τ	0.611	0.503	0.620	0.478	0.989	0.924	0.973	0.811
P_α	0.564	0.555	0.600	0.546	0.990	0.873	0.975	0.718
Leads	1.06	1.00	1.06	1.00	1.39	1.50	1.17	1.33
Lags	0.56	0.50	0.39	0.33	0.89	0.72	0.44	0.61

Table 4 notes. See notes to Table 1, except the first row denotes the inequality measure that is the dependent variable in the potential cointegrating equation with $ln(HP_t)$ as the regressor.

Table 5. Time-series long-run GNC tests

	$ln(Var_t)$			$ln(Top5\$_t)$		
	Lag	$INEQ_t$ to $ln(HP_t)$	$ln(HP_t)$ to $INEQ_t$	Lag	$INEQ_t$ to $ln(HP_t)$	$ln(HP_t)$ to $INEQ_t$
AUS	2	0.033**	0.142	3	0.072*	0.438
BEL	1	0.007***	0.425	1	0.007***	0.270
CAN	3	0.002***	0.033**	3	0.001***	0.066*
DEN	1	0.014**	0.987	1	0.013**	0.850
FIN	1	0.003***	0.805	1	0.003***	0.628
FRA	1	0.008***	0.445	1	0.011**	0.479
GER	3	0.043**	0.296	3	0.092*	0.898
IRE	2	0.058*	0.075*	2	0.059*	0.062*
ITA	2	0.000***	0.644	2	0.000***	0.632
JAP	2	0.146	0.766	2	0.167	0.827
KOR	2	0.000***	0.412	2	0.000***	0.397
NET	1	0.006***	0.082*	1	0.003***	0.105
NEW	1	0.013**	0.009***	1	0.023**	0.003***
NOR	1	0.015**	0.083*	1	0.017**	0.050*
SPA	1	0.112	0.050*	1	0.079*	0.046**
SWE	1	0.003***	0.470	3	0.139	0.840
UKD	1	0.021**	0.438	1	0.020**	0.539
USA	3	0.014**	0.982	3	0.008***	0.804

Table 5 notes. The probability value of a t-test on the error-correction term are reported. Lag denotes the VAR lag length chosen according to AIC criterion. $INEQ_t$ to $ln(HP_t)$ refers to tests of the measure of inequality Granger-causing $ln(HP_t)$ while $ln(HP_t)$ to $INEQ_t$ refers to tests of $ln(HP_t)$ Granger-causing inequality.

Table 6. Robust p-values for Westerlund's (2007) panel cointegration tests of $\boldsymbol{ln}(\boldsymbol{HP_t})$ on $\boldsymbol{ln}(\boldsymbol{GDP_t})$

	$\boldsymbol{ln}(\boldsymbol{GDP_t})$	
	Int	*Trend*
G_τ	0.625	0.985
G_α	0.629	0.910
P_τ	0.235	0.960
P_α	0.075*	0.888
Leads	1.44	1.33
Lags	1.13	1.11

Table 6 notes. See notes to Table 1, except the column labelled $ln(GDP_t)$ denotes the logarithm of real income as covariate involved in the potential cointegrating equation with $ln(HP_t)$ as the dependent variable.

Table 7. Robust p-values for Westerlund's (2007) test of $ln(HP_t)$ on r_t and $INEQ_t$ or $ln(GDP_t)$

		$Gini_t$		$Top5\%_t$		$ln(Var_t)$		$ln(Top5\$_t)$		$ln(GDP_t)$	
		Int	*Trend*	*Int*	*Trend*	*Int*	*Trend*	*Int*	*Trend*	*Int*	*Trend*
r_t	G_τ	0.964	0.990	0.996	0.995	0.542	0.924	0.325	0.513	0.840	0.888
	G_α	0.856	0.714	0.939	0.701	0.749	0.989	0.050*	**0.037****	0.858	0.783
	P_τ	0.109	0.555	0.301	0.626	0.213	0.739	**0.001*****	**0.004*****	0.119	0.229
	P_α	0.169	0.539	0.443	0.664	0.238	0.491	0.060*	**0.034****	0.269	0.420
	Leads	1.89	2.06	1.83	2.17	1.50	1.89	1.72	2.00	2.06	2.17
	Lags	2.06	2.17	1.83	2.33	1.89	2.22	2.06	2.33	2.17	2.17

Table 7 notes. See notes to Table 1 and Table 6, except the columns labelled r_t denotes the short-term real interest rate as covariate involved in the potential cointegrating equations.

Table 8. Trivariate DOLS long-run relationship of $ln(HP_t)$ on r_t and $ln(Var_t)$

	$ln(Var_t)$	r_t
PANEL	0.382*** (10.04)	-1.176 (-1.58)
AUS	0.559*** (13.22)	-3.351*** (-4.76)
BEL	0.642*** (5.34)	-4.170* (-1.75)
CAN	0.436*** (15.63)	1.897** (2.70)
DEN	0.134** (2.76)	-11.041*** (-10.30)
FIN	0.597*** (12.46)	9.739*** (14.59)
FRA	1.018*** (5.98)	-1.648 (-1.16)
GER	-0.104*** (-4.09)	1.666*** (3.01)
IRE	0.545*** (36.88)	-4.495*** (-16.22)
ITA	0.549*** (7.33)	4.492*** (3.07)
JAP	-0.114* (-1.91)	-3.537 (-1.68)
KOR	-0.067*** (-4.18)	0.865 (1.73)
NET	0.811*** (14.20)	-10.259*** (-13.08)
NEW	0.515*** (10.51)	-0.739 (-1.04)
NOR	0.531*** (6.22)	-1.636 (-0.46)
SPA	1.030*** (6.74)	5.098 (1.64)
SWE	0.361*** (5.85)	-4.697*** (-5.90)
UKD	0.587*** (9.05)	-1.451** (-2.43)
USA	0.249*** (3.30)	5.620 (1.70)

Table 8 notes. See notes to Table 3.

Table A.1. Pesaran's (2007) panel unit root test applied to levels data with the SPSM procedure

		Intercept only		Intercept and trend	
Variable	**N**	**Statistic**	**5% critical**	**Statistic**	**5% critical**
$ln(HP_t)$	18	-2.563	-3.336	-2.780	-3.857
$GINI_t$	18	-2.530	-3.335	-3.064	-3.855
$Top5\%_t$	18	-2.491	-3.335	-2.987	-3.855
$ln(VAR_t)$	18	-2.720	-3.334	-2.790	-3.855
$ln(Top5\$_t)$	18	-2.520	-3.334	-2.634	-3.855
$ln(GDP_t)$	18	-2.169	-3.335	-2.355	-3.855
Country excluded	**Variable:** r_t				
SPA/SPA	18	-3.979**	-3.335	-4.298**	-3.857
SWE/CAN	17	-3.895**	-3.339	-4.201**	-3.858
CAN/BEL	16	-3.785**	-3.344	-4.060**	-3.860
KOR/SWE	15	-3.597**	-3.347	-3.878**	-3.861
ITA/ITA	14	-3.579**	-3.347	-3.860**	-3.859
BEL/None	13	-3.502**	-3.346	-3.842	-3.858
NET/None	12	-3.362**	-3.345		
None/None	11	-3.227	-3.344		

Table A.1 notes. The column headed Variable indicates the variable that the tests are applied to and the column headed N denotes the number of countries included in the panel unit root tests. The tests are applied with two sets of deterministic terms being only an intercept (reported in the column headed Intercept only) and an intercept and trend (reported in the column headed Intercept and trend). The truncated panel unit root test statistics (CIPS) are reported in the columns headed Statistic while the corresponding 5% critical value (interpolated from those reported in Pesaran, 2007, and those for a standard ADF test when N=1) are given in the columns headed 5% critical. ** indicates the rejection of the null hypothesis that all the countries' series in the panel are $I(1)$ at the 5% level. When the null hypothesis is rejected the first column, headed Country excluded, gives the three letter country identifier for the country to be excluded from the next test in the SPSM sequence. The first country specified is for the Intercept only case and the second country identifier is given for the Intercept and trend case.

Table A.2. Pesaran's (2007) panel unit root test applied to differenced data with the SPSM procedure

Variable: $\Delta ln(HP_t)$						
	Intercept only			**Intercept and trend**		
N	**Country excluded**	**Statistic**	**5% critical**	**Country excluded**	**Statistic**	**5% critical**
18	UKD	-3.369**	-3.337	None	-3.728	-3.859
17	None	-3.230	-3.341			
Variable: $\Delta GINI_t$						
18	GER	-4.891**	-3.337	GER	-5.215**	-3.858
17	KOR	-4.775**	-3.341	FIN	-4.974**	-3.860
16	FRA	-4.679**	-3.345	KOR	-4.923**	-3.860
15	FIN	-4.466**	-3.348	FRA	-4.821**	-3.862
14	JAP	-4.293**	-3.348	NEW	-4.697**	-3.861
13	SWE	-4.162**	-3.347	JAP	-4.551**	-3.859
12	CAN	-4.285**	-3.347	SWE	-4.424**	-3.857
11	NEW	-4.165**	-3.346	CAN	-4.231**	-3.854
10	NOR	-3.986**	-3.346	NOR	-4.107**	-3.852
9	USA	-3.822**	-3.303	USA	-3.975**	-3.819
8	ITA	-3.707**	-3.259	UKD	-3.855**	-3.786
7	NET	-3.572**	-3.216	None	-3.686	-3.753
6	DEN	-3.512**	-3.172			
5	None	-3.124	-3.129			
Variable: $\Delta Top5\%_t$						
18	GER	-5.072**	-3.336	GER	-5.058**	-3.858
17	KOR	-4.734**	-3.340	FRA	-4.871**	-3.860
16	FRA	-4.633**	-3.344	FIN	-4.770**	-3.862
15	FIN	-4.493**	-3.348	NEW	-4.635**	-3.863
14	CAN	-4.320**	-3.348	CAN	-4.553**	-3.862
13	JAP	-4.206**	-3.347	AUS	-4.460**	-3.860
12	SWE	-4.069**	-3.347	ITA	-4.388**	-3.857
11	USA	-4.156**	-3.346	DEN	-4.268**	-3.855
10	NOR	-4.073**	-3.346	USA	-4.198**	-3.853
9	NEW	-3.906**	-3.303	UKD	-4.034**	-3.820
8	ITA	-3.741**	-3.259	NOR	-3.937**	-3.788
7	NET	-3.576**	-3.216	KOR	-3.825**	-3.755
6	DEN	-3.515**	-3.172	None	-3.687	-3.720
5	None	-3.099	-3.129			

Table A.2 (continued)

Variable: $\Delta ln(VAR_t)$						
	Intercept only			**Intercept and trend**		
N	**Country excluded**	**Statistic**	**5% critical**	**Country excluded**	**Statistic**	**5% critical**
18	FIN	-5.180**	-3.336	FIN	-5.258**	-3.857
17	GER	-5.075**	-3.340	GER	-5.191**	-3.859
16	FRA	-4.972**	-3.344	FRA	-4.963**	-3.861
15	NEW	-4.786**	-3.349	NEW	-4.876**	-3.863
14	JAP	-4.748**	-3.349	DEN	-4.842**	-3.862
13	SWE	-4.648**	-3.348	USA	-4.683**	-3.860
12	NOR	-4.584**	-3.347	KOR	-4.631**	-3.858
11	KOR	-4.575**	-3.347	IRE	-4.534**	-3.856
10	ITA	-4.413**	-3.347	JAP	-4.445**	-3.854
9	IRE	-4.230**	-3.304	SWE	-4.238**	-3.820
8	NET	-4.104**	-3.260	NET	-4.158**	-3.786
7	CAN	-4.044**	-3.217	NOR	-4.043**	-3.753
6	USA	-4.006**	-3.174	None	-3.699	-3.721
5	BEL	-4.015**	-3.131			
4	DEN	-3.880**	-3.087			
3	None	-2.949	-3.042			
Variable: $\Delta ln(Top5\$_t)$						
18	FIN	-5.154**	-3.336	FIN	-5.230**	-3.857
17	GER	-5.046**	-3.340	GER	-5.159**	-3.859
16	FRA	-4.944**	-3.344	FRA	-5.050**	-3.860
15	NEW	-4.764**	-3.348	DEN	-4.867**	-3.863
14	DEN	-4.631**	-3.348	NEW	-4.657**	-3.861
13	KOR	-4.557**	-3.348	KOR	-4.597**	-3.859
12	ITA	-4.508**	-3.348	USA	-4.548**	-3.857
11	NET	-4.479**	-3.346	ITA	-4.478**	-3.855
10	IRE	-4.429**	-3.346	NET	-4.467**	-3.852
9	JAP	-4.349**	-3.303	IRE	-4.394**	-3.819
8	SWE	-4.160**	-3.259	BEL	-4.324**	-3.786
7	CAN	-3.845**	-3.216	SWE	-4.282**	-3.753
6	NOR	-3.691**	-3.172	JAP	-4.107**	-3.720
5	BEL	-3.442**	-3.129	None	-3.669	-3.687
4	USA	-3.122**	-3.086			
3	None	-2.753	-3.042			
Variable: $\Delta \boldsymbol{lgdp_t}$						
18	BEL	-4.071**	-3.336	BEL	-4.239**	-3.858
17	AUS	-3.954**	-3.340	AUS	-4.165**	-3.860
16	DEN	-3.834**	-3.344	KOR	-4.004**	-3.861
15	ITA	-3.806**	-3.348	None	-3.814	-3.863
14	USA	-3.762**	-3.348			
13	CAN	-3.664**	-3.347			
12	KOR	-3.563**	-3.347			
11	IRE	-3.377**	-3.347			
10	None	-3.301	-3.346			

Table A.2 (continued)

Variable: Δr_t						
	Intercept only			Intercept and trend		
N	Country excluded	Statistic	5% critical	Country excluded	Statistic	5% critical
18	SWE	-5.813**	-3.338	SWE	-5.815**	-3.860
17	NEW	-5.762**	-3.342	NEW	-5.802**	-3.862
16	SPA	-5.749**	-3.346	SPA	-5.792**	-3.863
15	DEN	-5.672**	-3.350	UKD	-5.699**	-3.865
14	UKD	-5.725**	-3.349	DEN	-5.752**	-3.863
13	NOR	-5.694**	-3.349	NOR	-5.716**	-3.861
12	CAN	-5.597**	-3.349	CAN	-5.573**	-3.859
11	AUS	-5.540**	-3.348	AUS	-5.509**	-3.856
10	IRE	-5.510**	-3.347	IRE	-5.461**	-3.854
9	JAP	-5.416**	-3.305	JAP	-5.293**	-3.822
8	KOR	-5.317**	-3.262	KOR	-5.189**	-3.789
7	NET	-5.140**	-3.217	NET	-5.023**	-3.754
6	FRA	-4.894**	-3.174	FRA	-4.820**	-3.723
5	USA	-4.678**	-3.130	FIN	-4.541**	-3.689
4	GER	-4.901**	-3.086	GER	-4.537**	-3.655
3	FIN	-4.873**	-3.043	USA	-4.584**	-3.621
2	ITA	-4.309**	-2.999	ITA	-4.234**	-3.588
1	BEL	-7.014**	-2.951	BEL	-7.290**	-3.548

Table A.2 notes. See Table A.1 notes except the column headed Country excluded indicates a country's series identified as not rejecting the unit root null. An entry of "None" in this column means that the unit root null cannot be rejected for all N remaining countries included in the panel unit root test.

Table A.3. Cerrato *et al.*'s (2011) nonlinear panel unit root test applied to levels data with SPSM

Variable: $ln(HP_t)$						
	Intercept only			**Intercept and trend**		
N	**Country excluded**	**Statistic**	**5% critical**	**Country excluded**	**Statistic**	**5% critical**
18	NEW	-2.385**	-2.021	None	-2.263	-2.379
17	JAP	-2.151**	-2.027			
16	FIN	-2.040**	-2.033			
15	None	-1.864	-2.039			
Variable: $GINI_t$						
18	NET	-2.113**	-2.023	None	-2.102	-2.380
17	None	-1.810	-2.029			
Variable: $Top5\%_t$						
18	None	-2.014	-2.023	None	-2.119	-2.381
Variable: $ln(VAR_t)$						
18	None	-1.922	-2.024	None	-1.888	-2.381
Variable: $ln(Top5\$_t)$						
18	None	-1.940	-2.024	None	-2.036	-2.381
Variable: $ln(GDP_t)$						
18	None	-1.375	-2.023	None	-1.780	-2.381
Variable: r_t						
18	NEW	-2.822**	-2.023	NEW	-2.981**	-2.381
17	FIN	-2.772**	-2.030	FIN	-2.841**	-2.377
16	SPA	-2.426**	-2.036	CAN	-2.650**	-2.373
15	GER	-2.709**	-2.042	ITA	-2.463**	-2.370
14	BEL	-2.587**	-2.048	None	-2.337	-2.366
13	SWE	-2.456**	-2.055			
12	NOR	-2.450**	-2.060			
11	ITA	-2.144**	-2.066			
10	None	-2.002	-2.072			

Table A.3 notes. See Table A.2 notes except critical values are interpolated from those reported in Cerrato *et al.* (2011), Table 13 and Table 14, as well as Cerrato *et al.* (2013).

Table A.4. Cerrato *et al.*'s (2011) nonlinear panel unit root test applied to difference data with SPSM

Variable: $\Delta ln(HP_t)$						
	Intercept only			Intercept and trend		
N	Country excluded	Statistic	5% critical	Country excluded	Statistic	5% critical
18	NET	-3.212**	-2.021	JAP	-3.524**	-2.377
17	JAP	-3.001**	-2.027	NET	-3.345**	-2.373
16	SWE	-2.882**	-2.033	DEN	-3.176**	-2.369
15	KOR	-2.712**	-2.039	SWE	-2.963**	-2.366
14	NOR	-2.692**	-2.045	NOR	-2.757**	-2.362
13	DEN	-2.700**	-2.050	AUS	-2.843**	-2.358
12	AUS	-2.462**	-2.057	KOR	-2.707**	-2.354
11	ITA	-2.449**	-2.063	ITA	-2.614**	-2.351
10	UKD	-2.332**	-2.070	USA	-2.614**	-2.351
9	GER	-2.343**	-2.076	UKD	-2.408**	-2.343
8	USA	-2.278**	-2.082	GER	-2.504**	-2.340
7	CAN	-2.226**	-2.088	CAN	-2.393**	-2.336
6	SPA	-2.151**	-2.094	None	-2.326	-2.332
5	None	-1.890	-2.100			
Variable: $\Delta GINI_t$						
18	NOR	-3.727**	-2.021	NOR	-3.904**	-2.378
17	NEW	-3.659**	-2.027	NEW	-3.709**	-2.374
16	UKD	-3.614**	-2.034	UKD	-3.632**	-2.371
15	FRA	-3.424**	-2.040	FRA	-3.409**	-2.367
14	FIN	-3.285**	-2.045	FIN	-3.276**	-2.363
13	CAN	-3.073**	-2.051	CAN	-3.071**	-2.359
12	JAP	-2.904**	-2.057	IRE	-2.912**	-2.355
11	NET	-2.848**	-2.063	JAP	-2.804**	-2.351
10	SWE	-2.715**	-2.069	NET	-2.881**	-2.347
9	IRE	-2.732**	-2.075	SWE	-2.728**	-2.343
8	DEN	-2.618**	-2.081	DEN	-2.575**	-2.339
7	USA	-2.319**	-2.087	None	-2.281	-2.335
6	AUS	-2.241**	-2.093			
5	None	-2.058	-2.099			
Variable: $\Delta Top5\%_t$						
18	UKD	-3.411**	-2.021	UKD	-3.473**	-2.378
17	NOR	-3.550**	-2.027	NOR	-3.595**	-2.374
16	CAN	-3.417**	-2.033	NEW	-3.468**	-2.370
15	FRA	-3.299**	-2.039	CAN	-3.351**	-2.367
14	FIN	-3.181**	-2.045	FIN	-3.214**	-2.363
13	NEW	-2.935**	-2.051	FRA	-2.973**	-2.359
12	NET	-2.769**	-2.057	NET	-2.785**	-2.355
11	JAP	-2.655**	-2.063	JAP	-2.663**	-2.351
10	SWE	-2.748**	-2.070	SWE	-2.627**	-2.347
9	DEN	-2.645**	-2.075	DEN	-2.666**	-2.343
8	USA	-2.404**	-2.081	IRE	-2.440**	-2.339
7	IRE	-2.297**	-2.087	None	-2.316	-2.335
6	ITA	-2.251**	-2.093			
5	GER	-2.759**	-2.100			
4	None	-2.044	-2.105			

Table A.4. (continued)

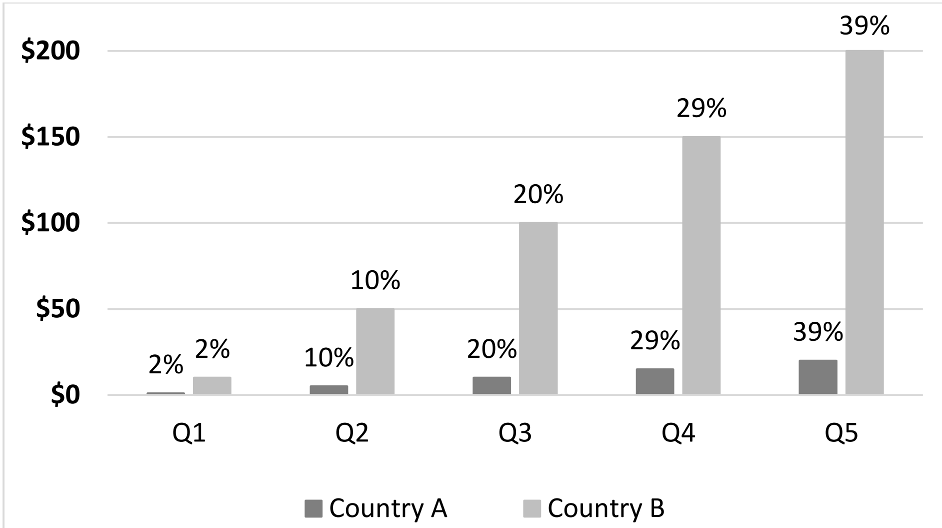
Variable: $\Delta ln(VAR_t)$						
	Intercept only			**Intercept and trend**		
N	**Country excluded**	**Statistic**	**5% critical**	**Country excluded**	**Statistic**	**5% critical**
18	NOR	-3.763^{**}	-2.022	NOR	-3.780^{**}	-2.378
17	FRA	-3.569^{**}	-2.028	FRA	-3.370^{**}	-2.374
16	FIN	-3.472^{**}	-2.033	FIN	-3.492^{**}	-2.370
15	SPA	-3.259^{**}	-2.039	GER	-3.290^{**}	-2.366
14	DEN	-3.287^{**}	-2.046	DEN	-3.374^{**}	-2.363
13	IRE	-2.607^{**}	-2.051	UKD	-2.908^{**}	-2.359
12	UKD	-2.499^{**}	-2.057	NEW	-2.726^{**}	-2.355
11	NEW	-2.381^{**}	-2.063	IRE	-2.726^{**}	-2.351
10	CAN	-2.182^{**}	-2.069	KOR	-2.857^{**}	-2.348
9	None	-1.964	-2.074	JAP	-2.631^{**}	-2.343
8				SWE	-2.494^{**}	-2.339
7				CAN	-2.437^{**}	-2.336
6				None	-2.094	-2.331
Variable: $\Delta ln(Top5\$_t)$						
18	NOR	-3.912^{**}	-2.022	NOR	-3.938^{**}	-2.379
17	FRA	-3.727^{**}	-2.028	FRA	-3.758^{**}	-2.375
16	FIN	-3.620^{**}	-2.033	GER	-3.654^{**}	-2.370
15	GER	-3.416^{**}	-2.039	FIN	-3.404^{**}	-2.367
14	DEN	-3.444^{**}	-2.046	DEN	-3.470^{**}	-2.363
13	IRE	-3.067^{**}	-2.052	UKD	-3.097^{**}	-2.359
12	CAN	-2.975^{**}	-2.058	IRE	-2.942^{**}	-2.355
11	UKD	-2.892^{**}	-2.064	SWE	-2.830^{**}	-2.351
10	SWE	-2.758^{**}	-2.070	JAP	-3.031^{**}	-2.348
9	JAP	-2.940^{**}	-2.076	CAN	-2.590^{**}	-2.344
8	NEW	-2.369^{**}	-2.082	NEW	-2.364^{**}	-2.340
7	KOR	-2.140^{**}	-2.088	None	-2.134^{**}	-2.336
6	USA	-2.213^{**}	-2.094			
5	ITA	-2.131^{**}	-2.100			
4	BEL	-2.109^{**}	-2.106			
3	None	-1.866	-2.112			
Variable: $\Delta lgdp_t$						
18	GER	-3.147^{**}	-2.021	GER	-3.107^{**}	-2.378
17	KOR	-3.294^{**}	-2.027	KOR	-3.263^{**}	-2.374
16	SPA	-3.590^{**}	-2.034	SPA	-3.601^{**}	-2.371
15	BEL	-3.510^{**}	-2.040	USA	-3.535^{**}	-2.367
14	USA	-3.415^{**}	-2.045	ITA	-3.430^{**}	-2.363
13	ITA	-3.360^{**}	-2.052	SWE	-3.350^{**}	-2.359
12	SWE	-3.239^{**}	-2.058	NET	-3.324^{**}	-2.355
11	NET	-3.167^{**}	-2.064	BEL	-3.219^{**}	-2.351
10	AUS	-3.047^{**}	-2.070	AUS	-3.099^{**}	-2.348
9	JAP	-2.800^{**}	-2.075	JAP	-2.813^{**}	-2.343
8	DEN	-2.653^{**}	-2.081	DEN	-2.631^{**}	-2.339
7	FRA	-2.623^{**}	-2.088	FRA	-2.634^{**}	-2.336
6	NEW	-2.485^{**}	-2.094	NEW	-2.512^{**}	-2.332
5	CAN	-2.552^{**}	-2.100	CAN	-2.554^{**}	-2.328
4	NOR	-2.148^{**}	-2.104	None	-2.085	-2.322
3	None	-1.823	-2.109			

Table A.4. (continued)

Variable: Δr_t						
	Intercept only			**Intercept and trend**		
N	**Country excluded**	**Statistic**	**5% critical**	**Country excluded**	**Statistic**	**5% critical**
18	SWE	-2.913^{**}	-2.020	SWE	-2.868^{**}	-2.376
17	NEW	-2.718^{**}	-2.026	NEW	-2.642^{**}	-2.372
16	JAP	-3.040^{**}	-2.032	JAP	-3.143^{**}	-2.369
15	IRE	-2.535^{**}	-2.535	IRE	-2.657^{**}	-2.365
14	AUS	-2.346^{**}	-2.043	None	-2.302	-2.360
13	SPA	-2.149^{**}	-2.049			
12	CAN	-2.330^{**}	-2.056			
11	BEL	-2.353^{**}	-2.062			
10	FRA	-2.306^{**}	-2.068			
9	ITA	-2.096^{**}	-2.072			
8	None	-1.887	-2.080			

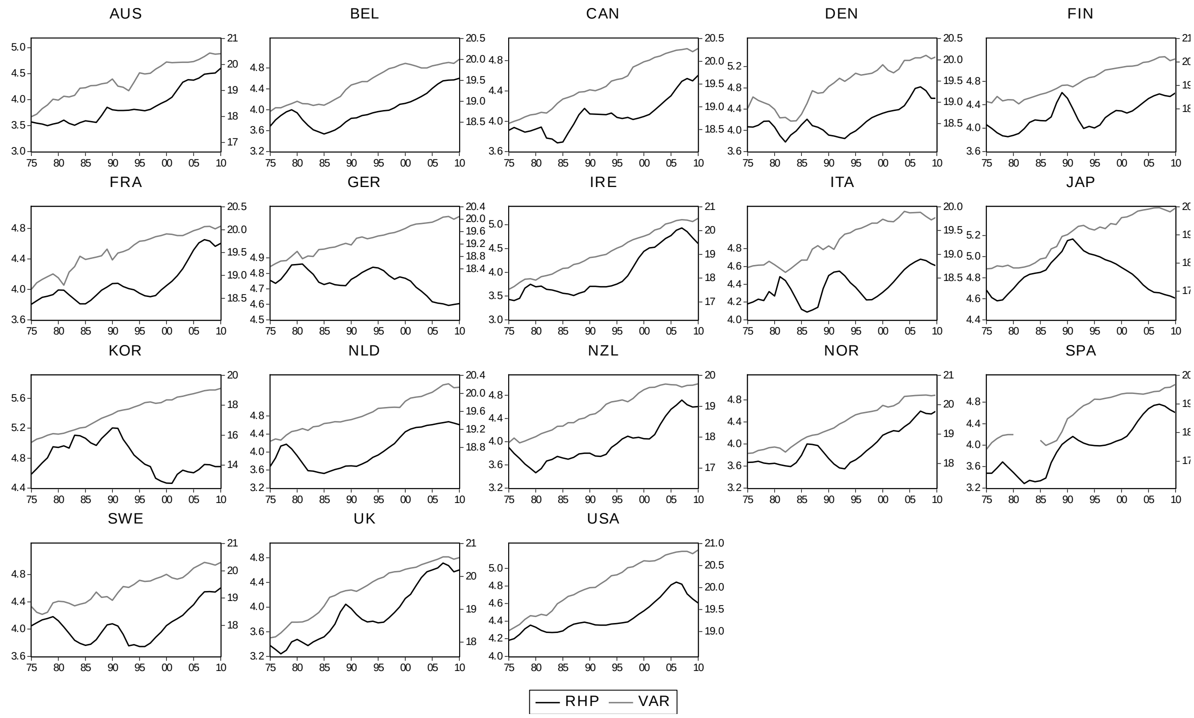
Table A.4 notes. See Table A.2 and Table A.3 notes.

Figure 1: Relative vs. Absolute Inequality



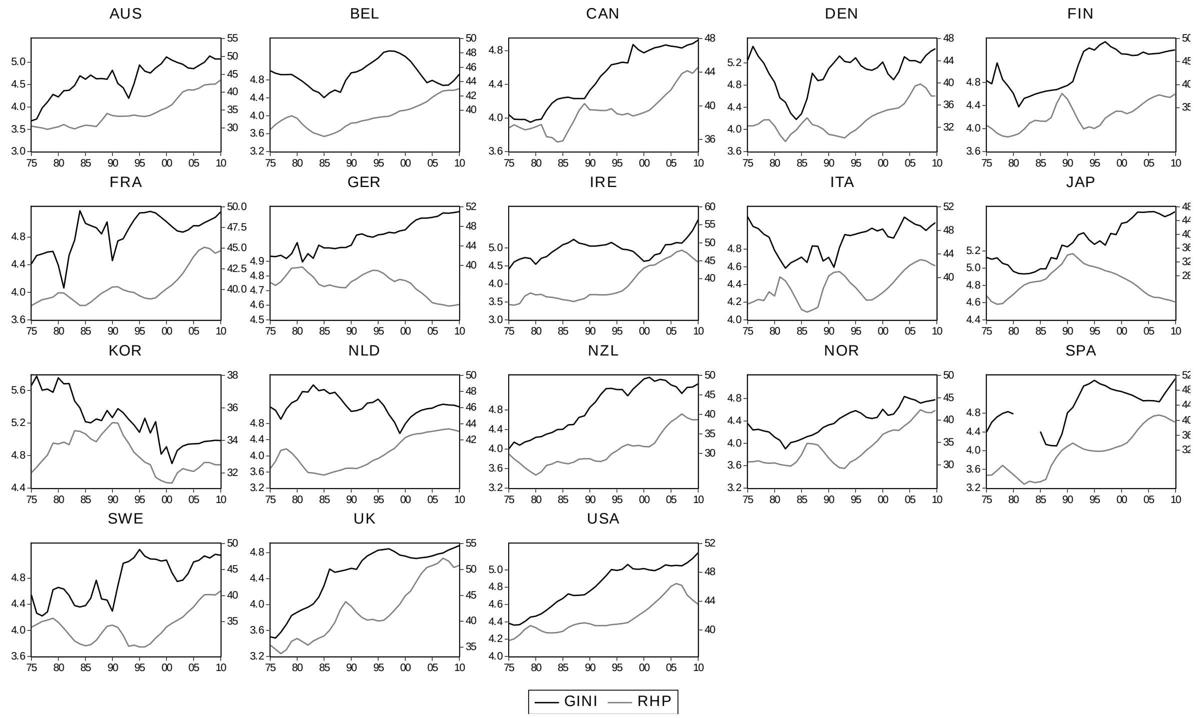
Note: This graph illustrates that two countries with different per capita incomes and equal income shares have distinct absolute income differences.

Figure 2: Real house prices and absolute inequality in OECD countries, 1975-2010



Note: This graphs shows the evolution of the logarithm of real house prices (RHP, left axis) and the logarithm of the income variance (VAR, right axis) in 18 selected OECD countries during the period 1975-2010.

Figure 3: Real house prices and relative inequality in OECD countries, 1975-2010



Note: This graphs shows the evolution of the logarithm of real house prices (RHP, left axis) and the Gini coefficient (Gini, right axis) in 18 selected OECD countries during the period 1975-2010.