Lancaster University Management School

Economics Working Paper Series

2017/012

Exuberance in the U.K. Regional Housing Markets

Alisa Yusupova, Efthymios Pavlidis, Ivan Paya, David Peel

The Department of Economics
Lancaster University Management School
Lancaster LA1 4YX
UK

Exuberance in the U.K. Regional Housing Markets

Alisa Yusupova, Efthymios Pavlidis, Ivan Paya, David Peel*

March 2017

Abstract

We combine the estimation of a structural model with inference based on recently developed recursive unit root tests to analyse the behaviour of regional real estate markets in the U.K. over the last four decades. We find two episodes, the late 1980s and the early and mid-2000s, when all regional house prices experienced explosive dynamics above and beyond factors such as housing supply relative to demographics, income, regional spillovers and credit availability. This is the first econometric analysis to provide evidence that would endorse the view that 'bubbles', with a particular spatial pattern, are a feature of UK regional housing markets.

Keywords: Regional house prices, Structural housing model, Cointegration, Generalised supremum ADF, Speculative bubbles

JEL Classification: C22, G12, R30, R31

*Alisa Yusupova (corresponding author): Department of Economics, Lancaster University, Lancaster, LA1 4YX, UK. a.yusupova@lancaster.ac.uk. Efthymios G. Pavlidis: Department of Economics, Lancaster University, Lancaster, LA1 4YX, UK. e.pavlidis@lancaster.ac.uk. Ivan Paya: Department of Economics, Lancaster University, Lancaster, LA1 4YX, UK. i.paya@lancaster.ac.uk. David A. Peel: Department of Economics, Lancaster University, Lancaster, LA1 4YX, UK. d.peel@lancaster.ac.uk.

1. Introduction

House prices in the UK have recently climbed to unprecedentedly high levels, surging ahead of their 2007 peak values. The price growth in London is even more dramatic. According to Nationwide figures, real estate prices in the metropolis have nearly doubled since the trough of 2009, being now 50% above their pre-crisis levels.[1] A concern that property prices in the UK and, in particular, in UK metropolitan areas, might be growing too quickly and can soon rise to unsustainable levels has been expressed by international organisations, central banks and housing market observers (see, e.g., the IMF 2014, 2016 Article IV Consultation reports, and the 2016 U.K. stress testing exercise of the Bank of England).[2] In this context, understanding the dynamics of real estate prices, what are the factors that are driving house price movements, and the nature of historical episodes of property price exuberance becomes particularly important.

A popular approach to analysing property price movements is to use a dynamic error-correction model to estimate fundamental real estate prices and, in turn, compare the estimates to the actual house price series (for UK housing market applications see, e.g., IMF (2003,2005), Barrell et al. (2004), Meen (2002), Cameron et al. (2006)). The rational of this approach is that, if house price movements do not reflect movements in economic fundamentals then the other factors above and beyond fundamental determinants are driving the dynamics of real estate markets. A complication with all house-pricing models is that there is no general consensus about the set of house price fundamentals. As a consequence, these models are subject to the problems related to omitted variables. Simple models, which consider a limited range of fundamentals (households' income, interest rates and lagged values of property prices), fail to accommodate important factors (such as supply-side effects) and spatial effects, which may well lead to erroneous inference about the presence of non-fundamental dynamics in real estate markets (e.g., IMF (2003,2005), Barrell et al. (2004)). To mitigate this problem, Cameron et al. (2006) propose a comprehensive model of regional property prices. In addition to the conventional set of demand-side variables, the model incorporates credit availability, demographics, regional spillover effects as well as supply-side factors.

[1]Nationwide is the UK based world's largest building society and one of UK's largest mortgage providers. The Nationwide database, which stretches back as far as 1952, contains data on UK national and regional house prices and housing affordability estimates. Other sources of UK regional and national house price data, such as Halifax and Land Registry, provide similar result.

[2]In the 2014 annual consultation report, IMF economists have articulated the potential adverse effects of rapid house price growth on the UK economy, stating that "there are few of the typical signs of a credit-led bubble in the housing market" (IMF, 2014). The report warns that raising residential and commercial property prices in London can potentially spread out to the rest of the country and threaten financial and macroeconomic stability.

In this paper, we employ the model of Cameron et al. (2006) in order to examine to what extent a rich set of fundamental factors has been driving the behaviour of UK regional property prices. We find that, like simpler models, this comprehensive model fails to adequately explain house price movements, when the sample is extended to include the latest boom-bust episode. Our results suggest that house prices and their fundamental determinants do not form a cointegrating system. The lack of evidence in favour of a stable long-run equilibrium is consistent with the rational bubble hypothesis.

Rational house price bubbles emerge when property prices are determined not only by the economic fundamentals but are also driven by the expectation of gains from future price increases, which introduces explosiveness in the house price series. The explosive nature of bubble processes has a direct implication for empirical tests: if economic fundamentals are at most integrated of order one, then the presence of explosive dynamics in the series of property prices constitutes evidence in favour of speculative bubbles. On the basis of this rational, in their seminal paper, Diba and Grossman (1988) suggest testing for the presence of asset price bubbles simply by applying right-tailed unit root tests to the asset price series. Unfortunately, as demonstrated by a number of authors, standard unit root tests suffer from low power in detecting periodically collapsing bubbles (see, e.g., Evans (1991), Gurkaynak (2008), Phillips and Yu (2011), Phillips et al. (2015), Pavlidis et al. (2016), Engsted et al. (2016)). That is, they fail to distinguish such periodically collapsing behaviour from non-explosive, unit root processes and hence, may often erroneously indicate the absence of a bubble when the data actually contains one. Recently, Phillips and Yu (2011) and Phillips et al. (2015) developed recursive unit root tests (the supremum *ADF*, *SADF*, and the Generalized supremum *ADF*, *GSADF*), which mitigate these problems. These recursive testing strategies are based on a repeated application of the right-tailed unit root test on a forward-expanding sample sequence, and have substantially higher power than the conventional procedures. Another appealing feature is that they enable to not only test for exuberance in the underlying series but also to shed light on the chronology of its origination and collapse.

We employ the tests of Phillips et al. (2011,2015) to formally examine whether UK real estate markets in the past were subject to explosive behaviour. In summary, our results indicate the presence of explosiveness in all regional real estate markets under consideration, while a panel version of the *GSADF* procedure, developed by Pavlidis et al. (2016), uncovers the overall, nationwide exuberant behaviour of UK property prices. The associated date-stamping strategies reveal two explosive episodes in the history of nearly four decades of the UK house price dynamics, namely, in the late 1980s and in the early and mid-2000s. These two episodes coincide with the periods of the largest deviations from the long-run equilibrium property prices in the model of Cameron et al. (2006). A conclusion that emerges from our analysis is that the fundamental model of housing does not explain the property price movements during the exuberant phases and hence, suggests that the house price exuberance was driven by non-fundamental factors in those time periods. The empirical analysis also shows that the error-correction terms from the estimated fundamental value model of housing are explosive, thereby providing further evidence that factors above and beyond fundamental determinants have induced exuberance in the dynamics of the real estate markets. These findings highlight the critical importance of monitoring housing market developments and are of particular relevance to policymakers and market participants.

The remainder of the paper is structured as follows. A description of the housing data is presented in Section 2. The structural model of regional real estate prices of Cameron et al. (2006) is outlined in Section 3. Section 4 discusses model estimation results and the implications of no cointegration between property prices and their fundamental determinants for the analysis of rational bubbles. Section 5 presents the univariate and the panel recursive right-tailed unit root tests' results and discusses the chronology of exuberance identified with the associated date-stamping mechanisms. This section also describes the application of the recursive unit root tests to the error-correction terms from the structural model of real estate prices. Finally, Section 6 provides concluding remarks.

2. Stylised Facts. 1975-2012

The house price data used in this paper is from the Nationwide House Price Database.[3] The Nationwide Database reports quarterly mix-adjusted regional house price indices for thirteen regional real estate markets: the North (NT), Yorkshire and Humberside (YH), North West (NW), East Midlands (EM), West Midlands (WM), East Anglia (EA), Outer South East (OSE), Outer Metropolitan (OM), Greater London (GL), South West (SW), Wales (WW), Scotland (SC) and Northern Ireland (NI). We adopt the Nationwide's regional

[3]Details of the methodology used to construct regional house price indices is available from the Nationwide web page: http://www.nationwide.co.uk/ /media/MainSite/documents/about/house-price-index/nationwide-hpi-methodology.pdf

classification, and refer the reader to the Nationwide web page for details on the regional composition. Nominal house price indices are deflated by the Consumer Price Index (all items) obtained from the OECD Database of Main Economic Indicators. In our application, we use the log of the regional real house price series. Figure 1 illustrates the evolution of regional real house price indices over the entire sample period, from the first quarter of 1975 until the fourth quarter of 2012, together with their linear time trends. We observe similar patterns of house price behaviour across regions, with a number of price upswings and downturns. For the vast majority of regions, the sample can be split into two distinct sub-periods: late 1980s - early 1990s, and early and mid-2000s.

In the first sub-period, house prices increased dramatically, reaching a maximum of 224% of the corresponding trend value in the late 1980s. This increase was accompanied by low interest rates, removal of credit and exchange controls and easing of prudential regulation.[4] At the same time, income of households failed to keep pace with growing residential prices, and examination of Figure 2 reveals a dramatic increase in the ratio of real estate prices to personal disposable income during the first sub-period. The average value of the reported price-to-income statistic across all regional markets of the country rose from about 68% in 1987 to nearly 98% by the middle of 1989, while in some regional markets, in particular Greater London and East Anglia, the peak value of housing affordability measure stood at nearly 130% in 1989:Q1.[5] Interestingly, the diagrams of housing prices and the price-to-income ratios indicate that Northern Ireland was the only regional market with no signs of a housing boom during the period under consideration (property prices in the region were, in fact, below the estimated linear trend at the end of 80s).

[INSERT FIGURES 1 & 2]

From the house price diagrams, we observe a prolonged downturn in regional real estate prices following the period of housing expansion in the end of 1980s. On average, a 60% fall in house prices occurred across UK regions from the peak of 1989 to the trough of 1993. As evident from Figure 2, the ratios of house prices to personal disposable income display a similar pattern, with an average fall of 34% fall over the bust years.

Regional real estate markets started to recover from the recession in the mid-1990s. According to re-

[4]The 1988 Basel I Capital Accord documented a requirement for banks to maintain capital of at least 8% of their risk-weighted assets. The regulatory framework imposed a 100% risk weight on unsecured loans, while mortgage lending received a preferred status with 50% risk weight assigned to loans secured on residential property.

[5]Regional income data is obtained from the Family Expenditure Survey (FES). Please refer to Table A1 of the Data Appendix for details.

gional property price diagrams, housing prices grew gradually after the first half of 1995, which marked the beginning of a prolonged period of house price growth that prevailed until 2007:Q3. According to Kuenzel and Bjørnbak (2008), population growth, higher income of households, low mortgage rate, financial deregulation and increased credit availability were among the key factors that fuelled another round of property price expansion. During the upswing of the early and mid-2000s, the average real house prices across all regional markets of the country doubled relative to the previous peak value of the statistic in 1989. Northern Ireland, in particular, recorded the biggest increase in residential and commercial property prices over the period: housing prices in the area in 2007:Q3 were nearly six times higher than in 1989:Q1. At the national level, house prices stood at about 109% above the estimated linear trend values in 2007:Q3. The fact that real personal disposable income was not growing at a comparable rate led to rapid deterioration in housing affordability. As evident from the price-to-income diagrams, in all regions of the country, with the exception of East Anglia, the ratio of real house prices to real personal disposable income reached unprecedentedly high levels. Notably in 2007:Q3, the mean value of the price-to-income statistic across 13 regions of the UK was nearly 70% above the historical average.

Following the start of the sub-prime mortgage crisis in the US, all housing markets of the UK experienced a sharp downturn in residential and commercial property prices. By the first quarter of 2009 regional real house price indices dropped, on average, by nearly 20% from their 2007:Q3 peak values. The housing market of Northern Ireland again stands out as the most volatile: the region recorded the biggest fall in real estate prices (around 30%) across all property markets in our sample.

The overall conclusion that emerges from the examination of the regional diagrams is that UK property markets were subject to substantial instability over the last four decades. To examine to what extent fundamental factors were driving the behaviour of UK real estate markets, we next employ the structural model of Cameron et al. (2006).

3. The Model of Real Estate Prices: Formulation and Estimation Results

The model of Cameron et al. (2006) is constructed as the system of inverted housing demand equations, one for each region of the country, where each regional house price equation is modelled as a dynamic error-correction relationship. Let $i = 1, \ldots, 13$, denote the regional index, and $\Delta lrhp_{i,t}$ stand for the growth in log real house prices, then the basic specification of the housing regression equation is given by

$$\begin{aligned}\Delta lrhp_{i,t} = \alpha \times (&\beta_{0,i} + \beta_1 \times lrynhs_{i,t} + \beta_2 \times MACCI_t + (1 - \varphi \times MACCI_t) \times (\beta_3 \times \Delta^2 labmr_t + \\ &+ \beta_4 \times (labmr_t - mean.labmr)) + \beta_5 \times MACCI_t \times (rabmr_t - mean.rabmr) + \\ &+ \beta_6 \times rabmr_t - lrhp_{i,t-1}) + \beta_7 \times \Delta clrhp_{i,t-1} + \beta_8 \times \Delta lrpdin_t + \beta_9 \times \Delta lrpdin_{t-1} + \\ &+ \beta_{10} \times MACCI_t \times \Delta lrpdin_t + \beta_{11} \times \Delta^2 lpc_t + \beta_{12} \times \Delta lrftse_t + \beta_{13} \times \Delta lrftseneg_t + \\ &+ \beta_{14} \times ror.neg_{i,t} + \beta_{15} \times \Delta pop2039_{i,t-1} + \beta_{16} \times \Delta(lwpop_{i,t} - lhs_{i,t-1}) + \beta_{17} \times D88 + \\ &+ \beta_{18} \times D08 + \epsilon_{i,t},\end{aligned} \tag{1}$$

where the right-hand side variables include regional real households' income ($lrynhs_{i,t}$), nominal ($labmr_t$) and real ($rabmr_t$) mortgage rates, credit availability indicator ($MACCI_t$), last period's house price growth in the neighbouring regions ($clrhp_{i,t-1}$), national-level personal disposable income ($lrpdin_t$), negative returns on housing ($ror.neg_{i,t}$), the supply of new constructions relative to the growth in working age population ($lwpop_{i,t} - lhs_{i,t-1}$), inflation acceleration (lpc_t), change in the real FTSE index ($\Delta lrftse_t$), negative changes in the real value of the FTSE ($\Delta lrftseneg_t$) and demographic effects ($pop2039_{i,t-1}$). The basic specification includes two dummy variables for 1988 and 2008 ($D88$, $D08$).[6] The former corresponds to the introduction of the Poll Tax system and the latter to the collapse of Lehman Brothers. There are some cases when the specification of regional equations differs from the one outlined in Eq.(1). These cases will be discussed in detail below. To assist the reader, the annotated model of regional house prices is included also in the Appendix with detailed description of the fundamental variables, their expected effects and the data sources.

We follow a two-stage estimation strategy suggested by Cameron et al. (2006). In the first step, the system is estimated using the Seemingly Unrelated Regressions (SUR) method and the estimated error covariance matrix is stored.[7] The second stage involves a Generalised Least Squares (GLS) approach, where

[6] We examine the unit root properties of the data and conclude that all regional real house price series are non-stationary in levels: we are not able to reject the unit root hypothesis of the *ADF* test at all conventional significance levels. Tax adjusted mortgage rates, indicator of credit availability and all regional income series entering the long-run equilibrium are I(1). Furthermore, all variables of the short-run dynamics: national income, demographics, the number of housing starts etc. are also I(1). These variables enter the house price model in the form of the first differences. The unit toot test results are available from the authors on request.

[7] Following Cameron et al (2006), in the first stage, we assign the value of 1.6 to the long-run income elasticity of housing β_1.

the unknown covariance matrix is replaced by the estimate from the first step. The chosen methodology accounts for contemporaneously correlated disturbances, which is particularly important, since the assumption of uncorrelated shocks in regional real estate markets appears unrealistic.

While Cameron et al. (2006) employ annual data and their estimation period ends in 2003, our data is sampled quarterly and covers the period from the first quarter of 1975 to the fourth quarter of 2012 to incorporate the latest boom and bust in the housing market and the Great Recession. Table 1 reports the estimation results. We first focus on the long-run equilibrium determinants and, in particular, credit availability, nominal and real mortgage rates and their interactions with the credit conditions index. We then discuss the estimation results for the variables that affect house prices with a time lag, which implies the dynamic effects.

[INSERT TABLE 1]

3.1. Long-run Equilibrium Determinants

Credit Availability Indicator One of the key elements of the long-run equilibrium is the index of credit conditions. This indicator, designed as a linear spline function, was proposed by Fernandez-Corrugedo and Muellbauer (2006) to capture the shifts in the supply of credit, changes in lending policy and prudential regulation.[8] The estimated effect of credit availability is positive and statistically significant and the magnitude of the coefficient is close to that reported by Cameron et al. (2006). The easing of prudential regulation and liberalisation of lending policy, therefore, encourage mortgage borrowing and lead to an increase in real estate prices.

Nominal and Real Interest Rates It is important to control, not only for the direct impact of credit policy changes on housing prices, but also for interaction effects of mortgage rates with the index of credit conditions. Failure to take these facts into account results in model misspecification and incorrect inference about the magnitude and direction of the interest rate effects. The long-run solution includes two mortgage rate

By doing so we, on the one hand, save degrees of freedom, and, on the other hand, retrieve the long-run coefficients when checking for the interaction effects. In the estimation we allow each regional equation to have a region-specific intercept $\beta_{i,0}$.

[8]The index of credit conditions can be estimated from the two-equation system of secured and unsecured lending. A detailed description of the methodology and the results of the credit availability estimation as well as the sources of the data used in the exercise can be found in a supplementary appendix on the authors' webpage.

measures: nominal and real interest rates of building societies adjusted for the cost of tax relief. The interest rates enter the house price model on their own and in interaction with the indicator of credit availability. The estimated positive interaction effect with real interest rate, reported in Table 1, suggests that while an increase in the cost of credit *per se* discourages mortgage borrowing and reduces house prices, this negative effect weakens with the removal of lending constraints and easing of prudential regulation. In other words, when credit policy is relaxed it becomes easier for households to find an opportunity to refinance the debt and deal with the burden of interest payments in the near-term.

In the model outlined in Eq.(1), the nominal interest rate effect is conditional on credit availability. Intuitively, inflation growth raises nominal rates and the burden of mortgage loan in the first few years of the contract. The risk of not being able to service the debt deters potential house buyers from mortgage borrowing and results in a fall in real estate prices in the long-run. However, easing of prudential regulation and liberalisation of credit conditions allow households to gain access to numerous refinancing opportunities, thus reducing the negative effect of an increase in nominal interest rates on real estate prices. According to the results presented in Table 1, a rise in nominal interest rates reduces commercial and residential property prices more when credit is constrained than when credit conditions are relaxed. In other words, the negative nominal interest rate effect weakens with credit liberalisation and access to more dynamic and competitive markets of mortgage lending.

3.2. Dynamic Effects

The group of variables that affect housing prices with a time lag includes last period's property price growth in the neighbouring regions, national-level personal disposable income effects, negative returns on housing, the supply of new houses relative to the growth in working age population, inflation acceleration, stock market effects and demo graphic factors. By analogy with the elements of the long-run equilibrium, the dynamic effects will be discussed in succession.

Spatial Dynamics One of the key effects of the short-run dynamics is a composite variable ($\Delta clrhp_{i,t-1}$) that is computed as a weighted sum of the last period's price growth in a region, regions contiguous to it (average house price growth across neighbouring areas) and Greater London. The weights on lagged growth rates are allowed to take any value on the unit interval and sum up to one. To take into account regional spillover effects, the weights vary by area and are assigned based on proximity to London. The southern regions (Outer Metropolitan, Outer South East, East Anglia and South West) attach 100% weight on last period's growth in London house prices, while as we move farther away from the metropolitan area the weight on London price growth becomes smaller, reflecting the fact that it takes more than a quarter for London house price impulses to reach distant areas.[9] This weighting scheme allows to capture the so-called ripple effect, widely documented in the empirical literature on UK housing markets, which implies that house price shocks emanating from Greater London have a tendency to spread out and affect neighbouring regions with a time lag (see, e.g., MacDonald and Taylor (1993), Alexander and Barrow (1994), Drake (1995), Meen (1999), Cook and Thomas (2003), Holly et al. (2010), *inter alia*). In line with previous studies, the estimated dynamic effect of the composite variable is positive, therefore last period's house price growth in a region (neighbouring areas and Greater London) leads to further price appreciation in this real estate market.

National-level Personal Disposable Income With regard to the other dynamic effects, we control for both direct effects of current and previous quarter's growth in personal disposable income (measured at the national-level) and for the interaction effect of the former variable with the indicator of credit availability. The inclusion of the interaction term is motivated by the argument that income changes matter less with removal of lending constraints, easing of prudential regulation and access to various financing opportunities. We show that both direct income effects are positive and statistically significant, however the indicator of credit availability does not provide additional explanatory power to the model when interacted with the growth in personal disposable income.

Downside Risk Each regional equation incorporates a dynamic measure of downside risk ($ror.neg_{i,t}$) defined as a four-quarter moving average of past negative returns on housing in the corresponding real estate market. Table 1 shows a small but significant positive effect of the dynamic downside risk measure, which suggests that the four consecutive quarters of negative housing returns depress current real estate prices

[9] Table A1 of Appendix A contains information about the composition of $\Delta clrhp_{i,t-1}$. Please refer to the notes to Table 1 for the regional weights.

above and beyond the own lag effect.

Inflation Acceleration To capture the dynamic effect of an increase in the general price level, we use the two-period change in the log of consumer expenditure deflator ($\Delta^2 lpc_t$). Intuitively, inflation acceleration leads to mortgage rate uncertainty, discourages mortgage borrowing and eventually results in lower residential prices. We report a significant negative effect of inflation acceleration, with the estimated coefficient implying that a one percentage point increase in the general price level results in a 0.2% fall in the price of housing.

Supply-side Effects The literature which deals with modelling real estate prices in the UK often ignores supply-side effects (e.g. models of Barrell et al. (2004) and IMF (2003, 2005)). However, a number of studies, including Glaeser et al. (2008), Hilbert and Vermeulen (2016), *inter alia*, demonstrate that it is important to incorporate the supply-side factors when analysing property price dynamics. Hilbert and Vermeulen (2016), who examine the impact of planning policies and local regulatory and geographical constraints on house prices in England, show that the rigidity of housing supply and the existing physical constraints on new developments are crucial factors behind the latest boom in the real estate markets. The authors demonstrate that residential prices in England would have been nearly 35% lower in 2008 had the regulatory constraints on local development been removed. Moreover, international organisations argue that limits on the supply of houses in the UK are among the key aspects of concern, responsible for the recent house price volatility (IMF Article IV Consultation report, 2014, 2016). Cameron et al. (2006) introduce the effect of changes in the supply of new houses relative to the growth in working age population $\Delta(lwpop_{i,t} - lhs_{i,t-1})$. Intuitively, if the supply of new homes fails to keep pace with the demographic growth it leads to an increase in the price of houses. Perhaps surprisingly, we find no significant effects of this measure in our application.[10]

Stock Market Effects We also examine whether the returns on financial investments are important determinants of the short-run house-price dynamics. Two indicators of the stock market behaviour are considered.

[10] Another effect of the short-run dynamics that we find not statistically significant is the growth in the share of people aged between 20 and 39 in the total working age population ($\Delta pop2039_{i,t-1}$). The 20-39 age group represents potential first-time home buyers and hence, growth in the proportion of households in this age segment can, potentially, have a positive effect on the demand for housing and on the real estate prices. Our results do not support this hypothesis.

First is the change in the real FTSE index ($\Delta lrFTSE_t$) which is included to test the assumption that higher returns on equity raise the wealth of financial investors (potential house-buyers) and eventually lead to an increase in the real estate prices. The second indicator, $\Delta lrFTSEneg_t$, is equal to $\Delta lrFTSE_t$ when the latter is negative and is zero otherwise and examines the effect of portfolio re-balancing in the face of stock market downturns. Cameron et al. (2006) show that the stock market effects are important only in London and in the South - centres of investment, equity ownership and well-paid employees - and have little impact on the rest of the country. Perhaps surprisingly, we fail to find a significant effect of stock market dynamics in Greater London. However, we do find this effect to be important in Outer Metropolitan region.[11] Our results suggest an asymmetric response of real estate prices to positive and negative shocks in the equity market.

Time Dummies Finally, each regional equation includes dummy variables to control for the shocks to the demand for and the supply of housing. The 1988 year dummy captures the introduction of the Poll Tax system in replacement of the local domestic rates taxation.[12] Furthermore, this variable picks up the effect of budget announcement in March of 1988, limiting the number of mortgage interest relief claims to one per property. We report a positive and significant effect of the 1988 dummy, which is consistent with the estimate of Cameron et al. (2006). We introduce a dummy variable for 2008 to pick up the effect of the Lehman Brothers collapse in September 2008 followed by a turmoil in the financial markets. According to our estimates, the 2008 dummy has a significant negative effect on the dynamics of the UK regional house prices.

4. Episodes of Exuberance in UK Housing Markets

With regard to the model specification, Cameron et al. (2006) design each regional house price equation as an error-correction relationship, and implicit in this formulation is that prices and fundamentals converge to a stable long-run equilibrium. In the context of our paper, there exists a stable long-run equilibrium relationship between house prices and economic fundamentals if the error-correction terms from the estimated

[11] In the final model specification, the two effects of the stock market dynamics enter the Outer Metropolitan equation only and are assumed zero in all remaining regions.

[12] Since the Poll Tax reform concerned only England and Wales, the 1988 dummy is set to zero in the equations of Scotland and Northern Ireland.

regional house price models are stationary. Otherwise, prices and their fundamental determinants are not cointegrated and, hence, the error-correction models should not be used to model the behaviour of the real estate prices.

We test the regional error-correction terms for stationarity (see Table 2) and our results indicate that, except for the North and Wales, where the deviations from the long-run equilibrium proved I(0) at 5% level of significance, for all remaining regions the unit root hypothesis cannot be rejected.[13] Therefore, we conclude that even for the comprehensive model of Cameron et al. (2006) house prices and economic fundamentals are not found to be cointegrated.

[INSERT TABLE 2]

The finding of non-stationary deviations from the long-run equilibrium prices is consistent with the existence of rational bubbles. The theory of asset price bubbles postulates that any cointegrating relationship between asset prices and fundamentals breaks down when the price series under investigation contains an explosive non-fundamental component. It can be shown that in the presence of bubbles, property prices are not solely determined by economic fundamentals but are also driven by the expectation of a gain from future price increases (see, e.g., Diba and Grossman (1988), Case and Shiller (2003), LeRoy (2004), Pavlidis et.al. (2016))

$$P_t = F_t + B_t, \tag{2}$$

where F_t is the fundamental-based property price, that is driven only by housing fundamentals, and B_t is a bubble process that satisfies

$$E_t(B_{t+1}) = (1+\rho)B_t. \tag{3}$$

According to Eq. (3), the bubble component is explosive on expectation, since the constant discount factor is positive ($\rho > 0$). The explosive nature of the bubble process introduces explosiveness in the house price series, and in the deviation of prices from their fundamental component ($P_t - F_t$).

The fact that our results indicate the absence of a cointegrating relationship between the property prices and the fundamental determinants is indicative of the existence of nonstationary dynamics in the house price series that is not explained by the economic fundamentals.

[13] For these two regions (North and Wales) the null of a unit root cannot be rejected at 1% significance level.

4.1. Recursive Unit Root Tests

We apply the test of Phillips et al. (2011, 2015) to the series of regional real house prices and the ratios of real prices to real personal disposable income in order to examine whether the UK regional real estate markets were explosive during the period under consideration. The reader is referred to Appendix B for the details of the *SADF* and the *GSADF* test procedures.

The upper panel of Table 3 reports the test statistics of the univariate *SADF* and *GSADF* for both variables under consideration together with the finite sample critical values, obtained by Monte Carlo experiments with 2000 replications. The reported *GSADF* results provide strong evidence of exuberance in the regional real house prices: the null of a unit root is confidently rejected at all conventional significance levels in all regions but two - Outer Metropolitan and Greater London, where we can only reject the null at the 5% level of significance. When we turn to the ratio of prices to income, the indication of explosiveness remains strong in most of the regional markets with the exception of East Anglia, for which the unit root hypothesis cannot be rejected. Comparing the results of the *SADF* and the *GSADF* test procedures, we notice that for the former the evidence of exuberance is weaker and even more so when we look at the statistics of the price-to-income ratio (we fail to reject the null in 7 regions out of 13). This finding is expected given the higher power of the *GSADF* test, as documented by Phillips et al. (2015).

[INSERT TABLE 3]

In order to identify the origination and termination dates of exuberance, we adopt the date-stamping strategy suggested by Phillips et al. (2015). Figures 3 and 4 plot the series of the Backward sup *ADF* statistics (*BSADF*) for the real house prices and the price-to-income ratios, respectively, together with the sequence of 95% critical values. For the convenience of the reader, we shade the periods when the estimated *BSADF* lies above the sequence of critical values, which implies that a series displays explosive dynamics. We also present the timeline of regional exuberance for both variables under consideration in Figure 5.

[INSERT FIGURES 3 to 5]

Overall, we observe a remarkably similar pattern across regions. The date-stamping mechanism reveals two explosive episodes: one in the late 1980s and another in the first half of 2000s. With regard to the first episode, Greater London and East Anglia were the first regions to enter the exuberant phase in the second quarter of 1987, followed by Outer Metropolitan (1987:Q3) and the contiguous areas of Outer South East and South West (1988:Q1). Within a year the southern regions were joined by the Midland areas (1988:Q2), Wales and Yorkshire & Humberside (1988:Q4). Exuberance reached the North West and the North regions by the first and second quarter of 1989, respectively, but did not reach Scotland and Northern Ireland.[14] The identified timeline of exuberance is consistent with the literature that documents the existence of a strong regional interconnectedness between the real estate markets in the UK. MacDonald and Taylor (1993) and Alexander and Barrow (1994), *inter alia*, demonstrate the tendency of house price shocks emanating from the southern regions, in particular London and South West, to spread out northward and affect the rest of the country (the so-called ripple-effect). What is particularly interesting is the striking synchronisation in the termination of the first explosive episode. The signal of property price collapse spread out and affected all regional housing markets almost simultaneously, within the first two quarters of 1989.

Turning to the second episode, we notice that all regional house prices became explosive in the first half of 2000s. The propagation of exuberance closely resembles the pattern observed in the late 80s, with exuberance originating from the southern regions (Greater London and Outer Metropolitan (2000:Q1), South West (2000:Q3) and Outer South East (2001:Q1)) and transmitting through the midland areas (East Anglia, East and West Midlands (2001:Q2), Wales (2001:Q4)) to the northern parts of the country (the North, North West (2002:Q2) and Scotland (2002:Q3)).[15] Contrary to the late 1980s, however, the termination of the second episode was less synchronised, as indicated by the estimated end dates. We observe a gradual collapse of the regional statistics over a five-year period. Our results indicate that the *BSADF* statistic of the Outer Metropolitan region was the first to fall below the sequence of the corresponding critical values in the third quarter of 2003. However, it was not until the end of 2004 - beginning of 2005 that the second identified episode of exuberance collapsed in Greater London and Outer South East (in 2004:Q4 and 2005:Q1, respectively). As can be seen from Figure 5, the *BSADF* statistics of Northern Ireland and Scotland, where exuberance prevailed until the third quarter of 2008, were the last to fall below the exuberance threshold.

Turning to the results for the price-to-income ratios, we observe that the *GSADF* statistic is significant

[14] Comparing the results for different autoregressive lag lengths we note that the *GSADF* date-stamping estimation with no lags detects a short period of exuberance in real house prices of Scotland and Northern Ireland in the end of 1980s and locates the dates of its origination as 1989:Q3 and 1990:Q1 respectively. In general, the duration of house price explosiveness is longer in the no lag case.

[15] We note that Northern Ireland, where the origination date of exuberance is located at the second quarter of 1997, is the first region to enter the exuberant phase.

at the 5% level for all regions but one, East Anglia. Thus, even after controlling for fundamentals there is evidence of exuberance in housing markets. However, the duration of exuberance is, in general, shorter. Most notably, for Scotland, Northern Ireland, North West, Yorkshire & Humberside and Wales the date-stamping strategy reveals no sign of explosiveness in the early 1990s; and for Outer Metropolitan it indicates a very short period of explosive dynamics in the 2000s. With regard to the synchronisation of regional markets and the pattern of northward propagation of housing dynamics, we observe a similar behaviour of house prices and price-to-income ratios.

We, finally, examine the overall, nationwide exuberance in the UK regional housing markets by using the panel version of the *GSADF* methodology proposed by Pavlidis et al. (2016) (see Appendix B). The bottom section of Table 3 reports the panel *GSADF* statistics together with the corresponding finite sample critical values computed for both the real house price series and the ratio of real prices to real disposable income. The null hypothesis of a unit root is rejected in favour of the explosive alternative for both variables under consideration providing strong evidence of nationwide explosiveness in the housing markets of the UK. As can be seen from Figure 6, the evolution of the panel *BSADF* statisticsis is in line with the pattern displayed by the individual regional *BSADF* series. Irrespective of the variable under examination (real house prices or price-to-income ratios), we observe two episodes of overall exuberance during the sample period: one in the late 80s and another in the early and mid-00s. The phases of the overall exuberance in the price-to-income ratios are somewhat shorter than those detected in the house price series, which is again consistent with the univariate date-stamping results.

[INSERT FIGURE 6]

Exuberance in Deviations From the Long-Run Equilibrium As a last exercise, we apply the recursive unit root procedure of Phillips et al. (2011, 2015) to the error-correction terms from the structural model of Cameron et al. (2006), estimated in Section 3. Because we examine estimated residual series, standard finite sample *SADF* and the *GSADF* critical values are no longer valid. To draw statistical inference, we adopt a

Monte Carlo simulation approach.[16]

[INSERT TABLE 4]

Table 4 reports the regional *SADF* and the *GSADF* statistics together with their respective finite sample critical values. The *GSADF* test results indicate that the null of a unit root is confidently rejected in favour of the explosive alternative at all conventional significance levels for all regional error-correction series. The results of the *SADF* test procedure are somewhat less unanimous. We notice very few rejections of the null, which, as discussed above, is consistent with the lower power of the *SADF* test. Figure 7 plots the *BSADF* series against the sequence of 95% critical values obtained by repeated application of the test procedure to the series of simulated cointegrating residuals. We note that all regional *BSADF* sequences lie above the series of critical values during the latest boom in the housing market. Generally, the regional *BSADF* series cross the explosive threshold around 2000-2001 and fall below the respective critical value sequence just before the downturn in the housing market, around 2005-2006. This chronology corresponds to the timeline of the second period of explosiveness in the series of property prices and the price-to-income ratios, uncovered by the univariate *GSADF* procedures (see Figures 3 and 4). Overall, the econometric results suggest that the error-correction terms from the structural model of regional real estate prices, are, in fact, explosive.

[INSERT FIGURES 7 & 8]

We complete our analysis by combining the evidence of the structural model with the results of Phillips et al. (2011) and Phillips et al. (2015) test procedures. Figure 8 displays the error-correction terms from the structural model (blue solid line) together with the periods during which the corresponding price-to-income ratio exhibited explosive dynamics (shaded areas). Visual examination of the regional diagrams suggests that the identified episodes of explosive dynamics generally correspond to the periods of the largest deviations from the long-run equilibrium prices. The fact that the economic fundamentals do not explain the

[16] Specifically, we simulate from a bivariate cointegrated system for prices and fundamentals, $Y_t = (y_{1t}, y_{2t})\prime$, with cointegrating vector $\beta = (1, -\beta_2)\prime$ using Phillips (1991) triangular representation of the form

$$y_{1t} = \beta_2 y_{2t} + u_t,$$
$$y_{2t} = y_{2t-1} + v_t,$$

where $\beta_2 = 1$, $u_t = 0.75u_{t-1} + \epsilon_t, \epsilon_t \sim iidN(0, 0.5^2), v_t \sim iidN(0, 0.5^2)$. We then estimate the system, and apply the *SADF* and *GSADF* to the error-correction term. We set the number of Monte Carlo simulations equal to 5000.

exuberant behaviour of the regional real estate prices in the end of 80s and, in particular, in the first half of 00s is consistent with the conjecture of the presence of bubbles in property prices.

In summary, the unit root test results for prices, price-to-income ratios, and error-correction terms provide strong and consistent evidence in favour of explosive dynamics in UK housing markets.

5. Conclusion

In this paper, we analysed the behaviour of UK national and regional house prices, by employing both a structural model of property prices and formal econometric tests for explosive dynamics. We began by estimating the fundamental value model of housing suggested by Cameron et al. (2006), that incorporates a wide range of the national-level and the regional-level house price determinants, the impact of credit liberalisation, as well as regional spillover effects. We estimated the model over a period which covers the recent boom and bust in the housing market, and found that, although the direction and the magnitude of the estimated effects are, generally, compatible with those reported by Cameron et al. (2006), the model fails to explain a large part of the variation in house prices. Visual examination of the deviations of the regional property prices from their respective long-run equilibria reveals that the model was not able to capture the regional house price dynamics in the late 1980s and the early and mid-2000s. By formally testing for cointegration between house prices and fundamentals, we found that there does not exist a stable long-run equilibrium relationship between the regional property prices and their fundamental determinants. The evidence of non-stationary deviations from the long-run equilibrium is consistent with the presence of rational bubbles.

Rational bubbles, if they exist, create exuberance in housing markets. To this end, we examined the time-series properties of regional house price series. The tests of Phillips et al. (2011, 2015) strongly supported the hypothesis of exuberance in all regional real estate prices, while the panel modification of the test procedure, suggested by Pavlidis et al. (2016), indicated the presence of nationwide exuberance in the UK housing market. With regard to the timeline of exuberance, we found two episodes of explosive dynamics (in the late 1980s and in the early and mid-2000s) and a spatial pattern of northward propagation. Finally, by applying the recursive unit root procedure to the error-correction terms from the structural model of Cameron et al. (2006), we found explosiveness in the deviations of all regional property prices from their respective long-run equilibria. This finding provides further support to the hypothesis that exuberance in the house price series was driven not by the economic fundamentals but by non-fundamental explosive elements of real estate prices.

References

Alexander C. and Barrow M. (1994). Seasonality and Cointegration of Regional House Prices in the UK. *Urban Studies*, vol. 31(10), pp. 1667-1689.

Bank of England (2016). *Stress Testing the UK Banking System: Key Elements of the 2016 Stress Test*, London: Bank of England.

Barrell et al. (2004). The Current Position of UK House Prices. *National Institute Economic Review*, vol. 189, pp. 57-65.

Blanchard O. and Watson M.W. (1982). Bubbles, Rational Expectations and Financial Markets. *National Bureau of Economic Research Working Papers no. 945*.

Cameron G., Muellbauer J. and Murphy A. (2006). Was There a British House Price Bubble? Evidence from a Regional Panel. *Centre for Economic Policy Research, Discussion Paper 5619*.

Case K.E. and Shiller R.J. (2003). Is There a Bubble in the Housing Market? *Brookings Papers on Economic Activity*, vol.34(2), pp. 209-362.

Clark S.P. and Coggin T.D. (2011). Was there a U.S. house price bubble? An econometric analysis using national and regional panel data. *The Quarterly Review of Economics and Finance*, vol.51(2), pp. 189-200.

Cook S. and Thomas C. (2003). An Alternative Approach to Examining the Ripple Effect in UK House Prices. *Applied Economics Letters*, vol. 10(13), pp.849-851.

Diba B.T. and Grossman H.I. (1988). Explosive Rational Bubbles in Stock Prices? *The American Economic Review*, vol.78(3), pp. 520-530.

Drake L. (1995). Testing for Convergence between UK Regional House Prices. *Regional Studies*, vol. 29(4), pp.357-366.

Engsted T., Hviid S.J. and Pedersen T.Q. (2016). Explosive Bubbles in House Prices? Evidence from the OECD Countries. *Journal of International Financial Markets, Institutions & Money*, vol. 40, pp.14-25.

Evans G.W. (1991). Pitfalls in Testing for Explosive Bubbles in Asset Prices. *The American Economic*

Review, vol.81(4), pp.922-930.

Fernandez-Corugedo E. and Muellbauer J. (2006). Consumer Credit Conditions in the United Kingdom. *Bank of England Working Paper No. 314*.

Gallin J. (2006). The Long-Run Relationship between House Prices and Income: Evidence from Local Housing Markets. *Real Estate Economics*, vol. 34(3), pp. 417-438.

Glaeser E.L., Gyourko J. and Saiz A. (2008). Housing Supply and Housing Bubbles. *Journal of Urban Economics*, vol.64, pp.198-217.

Gurkaynak R.S. (2008). Econometric Tests of Asset Price Bubbles: Taking Stock. *Journal of Economic Surveys*, vol. 22(1), pp. 166-186.

Hilber C.A.L. and Vermeulen W. (2016). The Impact of Supply Constraints on House Prices in England. *Economic Journal*, vol. 126(591), pp. 358-405.

Holly S., Pesaran M.H. and Yamagata T. (2010). Spatial and Temporal Diffusion of House Prices in the UK. *Institute of the Study of Labour (IZA) Discussion Paper 4694.*

Homm U. and Breitung J. (2012). Testing for Speculative Bubbles in Stock Markets: A Comparison of Alternative Methods. *Journal of Financial Econometrics*, vol. 10(1), pp. 198-231.

International Monetary Fund (2003). United Kingdom: Selected Issues. *IMF Country Report no. 03/47*.

International Monetary Fund (2005). United Kingdom: Selected Issues. *IMF Country Report no. 05/81*.

International Monetary Fund (2014). United Kingdom: 2014 Article IV Consultation - Press Release, Staff Report; and Statement by the Executive Director for the United Kingdom. *IMF Country Report no. 14/233*.

International Monetary Fund (2016). United Kingdom: 2015 Article IV Consultation - Press Release, Staff Report; and Statement by the Executive Director for the United Kingdom. *IMF Country Report no. 16/57*.

Kuenzel R. and Bjørnbak B. (2008). The UK Housing Market: Anatomy of a House Price Boom. *ECFIN Country Focus: Economic Analysis from the European Commission's Directorate-General for Economic and Financial Affairs*, vol.5(11), pp.1-10.

LeRoy S.F. (2004). Rational Exuberance. *Journal of Economic Literature*, vol. 42(3), pp. 783-804.

MacDonald R. and Taylor M.P. (1993). Regional House Prices in Britain: Long-Run Relationships and Short-Run Dynamics. *Scottish Journal of Political Economy*, vol. 40(1), pp. 43-55.

Meen G. (1999). Regional House Prices and the Ripple Effect: A New Interpretation. *Housing Studies*, vol.

14(6), pp.733-753.

Meen G. (2002). The Time-Series Behaviour of House Prices: A Transatlantic Divide? *Journal of Housing Economics*, vol. 11(1), pp. 1-23.

Moon H.R. (1999). A Note on Fully-Modified Estimation of Seemingly Unrelated Regressions Models with Integrated Regressors. *Economics Letters*, vol. 65, pp. 25-31.

Pavlidis E.G. et al. (2016). Episodes of Exuberance in Housing Markets: In Search of the Smoking Gun. *The Journal of Real Estate Finance and Economics*, vol. 53(4), pp. 419-449.

Phillips P.C.B. (1991). Optimal Inference in Cointegrated Systems. *Econometrica*, vol.59(2), pp. 283-306.

Phillips P.C.B. and Yu J. (2011). Dating the Timeline of Financial Bubbles During the Subprime Crisis. *Quantitative Economics*, vol. 2(3), pp. 455-491.

Phillips P.C.B., Shi S. and Yu J. (2015). Testing for Multiple Bubbles: Historical Episodes of Exuberance and Collapse in the S&P 500. *International Economic Review*, vol.56(4), pp. 1043-1078.

Appendix A: Data Description and Sources

Table A1: The Model of Regional House Prices.

Variable	Description	Data Sources
$\Delta lrhp_{i,t} =$	Dependent variable is the log growth in the regional real house price index.	The nominal house price data is from the Nationwide database. To transform the data into real values regional indices are deflated using the CPI (all items) from the OECD Main Economic Indicators.
$+\alpha \times \beta_{0i}$	α is the speed of adjustment. β_{0i} is the region-specific intercept term.	
$+\alpha \times [\beta_1 \times lrynhs_{i,t} - lrhp_{i,t-1}]$	Error-Correction Term. Following Cameron et al. (2006), the long-run income elasticity β_1 is set to 1.6.	Annual regional income data is extracted from the Family Expenditure Survey (FES) (FES runs from 1961 to 2001, from 2001 it was replaced by the Expenditure and Food Survey (EFS), which became the Living Costs and Food Survey (LCF) from 2008). For each year in our sample we split the dataset by region and extract the data on average total household's weekly expenditure (Outer Metropolitan and Outer South East are assumed to correspond to the South East region in the regional classification adopted by the FES). The annual data is then interpolated to obtain quarterly series.
$+\alpha \times \beta_2 \times MACCI_t$	Effect of credit liberalisation. The moving average of the index is computed as $MACCI_t = \frac{CCI_t + CCI_{t-1}}{2}$.	The sources of data used for the *CCI* estimation are available on request from the authors.
$+\alpha \times (1 - \varphi \times MACCI_t) \times (\beta_3 \times (labmr_t - mean.labmr) + \beta_4 \times \Delta_2 labmr_t)$	The nominal tax-adjusted mortgage rate is defined as $abmr_t = bmr_t - \frac{CostofMortgageTaxRelief_t}{SD_t}$, where SD_t is the stock of mortgage debt. $labmr_t$ is the natural log of the $abmr_t$ series. The nominal interest rate enters the equation in interaction with the index of credit availability. We expect a positive coefficient in front of the interaction term since liberalisation of credit markets weakens the negative effect of an increase in the nominal mortgage rate.	The data on total amount of lending secured on dwellings is available via ONS.The cost of interest relief can be accessed via HM Revenue & Customs. The data is interpolated to obtain quarterly series. The cost of mortgage relief is zero from 2000Q1 onwards. The source of the building societies' mortgage rate data (bmr_t) is OECD: Main Economic Indicators.
$+\alpha \times [\beta_5 \times MACCI_t \times (rabmr_t - mean.rabmr)]$	The real tax-adjusted mortgage rate is defined as $rabmr_t = abmr_t - \Delta \log Deflator_t$. We interact the real rate with the index of credit availability. The negative effect of real interest rates weakens with credit liberalisation.	
$+\alpha \times \beta_6 \times rabmr_t$	We expect a negative effect of real interest rates on house prices.	

Table A1: The Model of Regional House Prices.(Continued)

Variable	Description	Data Sources
$+\beta_7 \times \Delta clrhp_{i,t-1}$	Positive effect of lagged house price growth in the neighbouring regions. $\Delta clrhp_{i,t-1} = (1 - w_{1,i} - w_{2,i}) \times \Delta lrhp_{i,t-1} + w_{1,i} \times \Delta lrhp_{CR,t-1} + w_{2,i} \times \Delta lrhp_{GL,t-1}$, where $w_{1,i}$, $w_{2,i}$ are region-specific weights that are assigned based on proximity to London. Please refer to the notes to Table 1 for $w_{1,r}$, $w_{2,r}$ values.	The data on regional house prices is available from the Nationwide database. $\Delta lrhp_{CR,t-1}$ is the last period's average growth rate in real house prices in contagious regions and $\Delta lrhp_{GL,t-1}$ is the lagged growth rate in London real estate prices.
$+\beta_8 \times \Delta lrpdin_t$ $+\beta_9 \times \Delta lrpdin_{t-1}$ $+\beta_{10} \times MACCI_t \times \Delta lrpdin_t$	We expect a positive effect of current and last period's national income growth on housing inflation. We expect a negative coefficient in front of the interaction term, since households' non-property income matters less when credit becomes freely available.	The Index of Personal Disposable Income adjusted for inflation is from Federal Reserve Bank of Dallas' International House Price Database. We take the log of the reported series.
$+\beta_{11} \times \Delta^2 lpc_t$	The expenditure deflator is computed as consumption expenditure at current prices divided by consumption expenditure at constant prices. Acceleration in inflation rate discourages mortgage borrowing and hence has a negative effect on house prices.	The data on consumer spending is available via ONS.
$+\beta_{12} \times \Delta lrftse_t$ (Outer Met) $+\beta_{13} \times \Delta lrftseneg_t$ (Outer Met)	The stock market effect. Enters the OM equation only. $lrftseneg_t = lrftse_t$ if $lrftse_t < 0$ and zero otherwise. Fall in the stock market results in investors reallocating their wealth and choosing real estate as a safe alternative. We expect a negative coefficient in front of the $\Delta lrftseneg_t$ term.	FTSE data is accessed via Datastream. We deflate the nominal index by the CPI (all items) and take the log of the series.
$+\beta_{14} \times ror.neg_{r,t}$	Downside risk in the real estate market is defined as a four-quarter moving average of past negative returns on housing in the region. The rate of return on housing is defined as: $ror_{i,t} = \Delta_4 lhp_{i,t-1} + 0.02 - abmr$, where $\Delta_4 lhp_{i,t-1}$ is a four-quarter change in the log regional house price index lagged one period. Negative rate of return $ror.neg_{i,t} = ror_{i,t}$ if $ror_{i,t} < 0$ and zero otherwise.	
$+\beta_{15} \times \Delta pop2039_{i,t-1}$	Demographic effect is measured by the last period's change in the share of people aged 20-39 in the total working age population. Population increase has a positive effect on demand for housing and hence on real estate prices.	The data on population estimates by region, age and sex can be accessed via ONS webpage (EA - East, NT - North East, OM and OSE - split the South East values in the ONS classification.). The data is available annually and was interpolated to obtain quarterly series.
$+\beta_{16} \times \Delta(lwpop_{i,t} - lhs_{i,t-1})$	The ratio of working age population to housing stock in the previous period. Increase in the population relative to the existing stock of dwellings has a positive effect on real estate prices.	Population estimates by region, age and sex are available from the ONS database. Live tables on housing stock by tenure and region are available from the GOV.UK database.
$+\beta_{17} \times D88$ (ex.SC and NI) $+\beta_{18} \times D08$	The D88 dummy variable captures the Poll Tax reform and limits on mortgage interest relief claims introduced in 1988. The D08 variable picks up a turmoil in the financial markets following the collapse of Lehman Brothers and seizure of Fannie Mae and Freddie Mac by the US government in September 2008.	Time Dummies are constructed as time trends going from 0.25 in Q1 of the corresponding year to 1 in Q4 of that year. The variable is zero otherwise.

Appendix B: The SADF and the GSADF Test Procedures

The Univariate *SADF* and *GSADF*

Consider the time series y_t with $[r_1 T]$ and $[r_2 T]$ specifying the first and the last observation respectively, where T is the total sample size and r_1, r_2 are the fractions of the total sample. The conventional right-tailed *ADF* test, suggested by Diba and Grossman (1988), estimates the following regression equation:

$$\Delta y_t = \mu_{r_1,r_2} + \phi_{r_1,r_2} y_{t-1} + \gamma^1_{r_1,r_2} \Delta y_{t-1} + \cdots + \gamma^k_{r_1,r_2} \Delta y_{t-k} + \epsilon_t, \tag{4}$$

where k denotes the chosen lag length, $\epsilon_t \sim iidN\left(0, \sigma^2_{r_1,r_2}\right)$ and $\mu_{r_1,r_2}, \phi_{r_1,r_2}$ and $\gamma^j_{r_1,r_2}$, where $j = 1...k$ are the regression coefficients. The null hypothesis of the right-tailed *ADF* procedure is that the series y_t contains a unit root, $\underline{H_0} : \phi_{r_1,r_2} = 0$, which is tested against the explosive alternative, $\underline{H_1} : \phi_{r_1,r_2} > 0$.

The conventional test statistic that corresponds to the case when both starting and ending points of the sample are fixed at $r_1 = 0$ and $r_2 = 1$ is labelled as $ADF^{r_2}_{r_1} = ADF^1_0$. The test statistic is compared to the right-tailed critical value from the limit distribution of ADF^1_0 and rejection of the null hypothesis in favour of the alternative signals the presence of explosiveness in the series y_t.

The test has low power in detecting periodically collapsing bubbles - a special class of explosive processes simulated by Evans (1991) that never collapse to zero and restart after the crash. Conventional right-tailed unit root tests fail to distinguish periodically collapsing behaviour from unit root, non-explosive processes and hence, may often erroneously indicate absence of a bubble when the data actually contains one.

Phillips et al. (2011) proposed recursive supremum *ADF* (*SADF*) test that proved robust to detection of periodically collapsing behaviour. The new approach suggests repeated estimation of the regression equation (4) on a forward expanding sample. The first estimated subsample comprises $[r_0 T]$ observations, where r_0 is the predetermined minimum window size as fraction of the total sample.[17] The starting point of the forward expanding sample is fixed at the first observation in our sample $r_1 = 0$, as in the conventional *ADF*, while the ending point is allowed to change $r_2 \in [r_0, 1]$ being incremented by one observation at a pass. Recursive

[17] When the total number of observations is relatively small, the size of the smallest moving window should be large enough to ensure effective estimation. Following the paper by Phillips et al. (2015), in our application r_0 comprises 36 observations (24% of the 152 observations).

application of the right-tailed *ADF* yields a sequence of test statistics denoted by $ADF_0^{r_2}$.

Statistical inference is based on the value of the largest test statistic in a sequence of $ADF_0^{r_2}$, called supremum *ADF*(*SADF*):

$$SADF(r_0) = \underset{r_2 \in [r_0,1]}{sup} \left\{ ADF_0^{r_2} \right\}. \tag{5}$$

If the statistic exceeds the right-tailed critical value from the limit distribution of the *SADF*, we reject the null of a unit root in favour of the explosive alternative.

Phillips et al. (2011) demonstrate that the test has more power in distinguishing periodically collapsing behaviour from stationary, mean-reverting processes than the conventional *ADF*. The suggested methodology gives rise to the date-stamping mechanism (discussed below) that allows to identify the origination and termination dates of exuberance and is shown to produce consistent results when applied to the data series with a single explosive episode in the sample (Phillips et al., 2011, 2015).

However Phillips et al. (2015) argue that the *SADF* test is inconsistent and produces conflicting results when applied to long economic series with multiple periods of exuberance within the sample. The authors propose a new test procedure, called Generalised *SADF* (*GSADF*), that covers more subsamples than the earlier approach as both starting and ending points of the forward-expanding sample are allowed to change. The estimation begins with the subsample, the first and the last observation of which are set to $r_1 = 0$ and $r_2 = r_0$ respectively. Holding the beginning point fixed, the subsample is incremented by one observation at a time until $r_2 = 1$. Then we shift the starting point by one observation and repeat the estimation process on the new set of subsamples. The recursive estimation continues until $r_1 = r_2 - r_0$. The largest test statistic over the full range of estimated $ADF_{r_1}^{r_2}$ is labelled as $GSADF(r_0)$:

$$GSADF(r_0) = \underset{\substack{r_2 \in [r_0,1] \\ r_1 \in [0, r_2 - r_0]}}{sup} \left\{ ADF_{r_1}^{r_2} \right\}. \tag{6}$$

As in the test procedures discussed above, we reject the null hypothesis of a unit root if the *GSADF* statistic exceeds the right tailed critical value from its limit distribution.

The Date-Stamping Strategy The univariate *SADF* and *GSADF* procedures discussed above allow not only to test for explosiveness in the underlying series but also to locate the dates of its origination and collapse. The date-stamping strategy associated with the *SADF* methodology defines the starting point of exuberance $[\hat{r}_e T]$ as the first observation whose ADF_0^{r2} lies above the sequence of corresponding critical values (Phillips et al., 2011, 2015):

$$\hat{r}_e = \underset{r_2 \in [r_0,1]}{inf} \left\{ r_2 : ADF_{r_2} > cv_{r_2}^{\beta_T} \right\},$$

while the termination date of exuberance $[\hat{r}_f T]$ is defined as the first observation after $\hat{r}_e T + \log(T)$ whose ADF_0^{r2} falls below the sequence of critical values:

$$\hat{r}_f = \underset{r_2 \in [\hat{r}_e T + \log(T),1]}{inf} \left\{ r_2 : ADF_{r_2} < cv_{r_2}^{\beta_T} \right\},$$

where $cv_{r_2}^{\beta_T}$ denotes the $100(1 - \beta_T)\%$ critical value of the ADF_0^{r2} distribution and β_T is the chosen level of significance.

As noted above, Phillips et al. (2015) demonstrate that the *SADF* date-stamping strategy fails to consistently locate origination and collapse dates when the data contains multiple explosive episodes of a different duration. The authors propose date-stamping mechanism associated with the *GSADF* test that overcomes the problem of the earlier technique. The new strategy is based on the value of the largest test statistic from backward expanding sample, labelled *BSADF* and defined as:

$$BSADF_{r_2}(r_0) = \underset{r_1 \in [0, r_2 - r_0]}{sup} \left\{ ADF_{r_1}^{r_2} \right\}, \tag{7}$$

To identify the chronology of exuberance the authors propose comparing the series of *BSADF* statistics with the sequence of $100(1 - \beta_T)\%$ critical values of the *SADF* distribution. The origination date of exuberance is defined as the first observation whose *BSADF* exceeds the critical value (Phillips et al. 2015):

$$\hat{r}_e = \underset{r_2 \in [r_0,1]}{inf} \left\{ r_2 : BSADF_{r_2}(r_0) > scv_{r_2}^{\beta_T} \right\},$$

while the termination of exuberance is the first observation after $\hat{r}_e T + \delta \log(T)$ for which the *BSADF* falls below the sequence of critical values:

$$\hat{r}_f = \underset{r_2 \in [\hat{r}_e T + \delta \log(T), 1]}{inf} \left\{ r_2 : BSADF_{r_2}(r_0) < scv_{r_2}^{\beta_T} \right\},$$

where $scv_{r_2}^{\beta_T}$ denotes the $100(1 - \beta_T)\%$ critical value of the *SADF* distribution, β_T is the chosen level of significance and δ is the parameter that depends on the frequency of the data. The assumption that termination date of exuberance is at least $\delta \log(T)$ observations away from its date of origin $[\hat{r}_e T]$ imposes a restriction on the minimum duration of explosive episode.

The Panel *GSADF*

Pavlidis et al. (2016) propose the panel version of the *GSADF* test that provides a way of testing for the degree of global exuberance in the datasets with a large number of cross-sectional units. The new panel *GSADF* test and the associated date-stamping strategy are based on the regression equation (4) with notation adjusted for panel structure of the data as:

$$\Delta y_{i,t} = \mu_{i,r_1,r_2} + \phi_{i,r_1,r_2} y_{i,t-1} + \gamma_{i,r_1,r_2}^{1} \Delta y_{i,t-1} + \cdots + \gamma_{i,r_1,r_2}^{k} \Delta y_{i,t-k} + \epsilon_{i,t}, \tag{8}$$

where $i = 1 \ldots N$ denotes the number of cross-sections in the dataset.

The null hypothesis of the panel *GSADF* procedure is that all cross-sectional units contain a unit root, $\underline{H_0} : \phi_{i,r_1,r_2} = 0$, which is tested against the alternative of an explosive root, $\underline{H_1} : \phi_{i,r_1,r_2} > 0$.

Statistical inference is made on the basis of the panel *GSADF* statistic that is defined as:

$$\text{Panel } GSADF(r_0) = \underset{r_2 \in [r_0, 1]}{sup} \left\{ \text{Panel } BSADF_{r_0}(r_0) \right\},$$

where the panel *BSADF* is computed as the average of N individual supremum *ADF* statistics from the backward expanding sample sequence:

$$\text{Panel } BSADF_{r_2}(r_0) = \frac{1}{N} \sum_{i=1}^{N} BSADF_{i,r_2}(r_0),$$

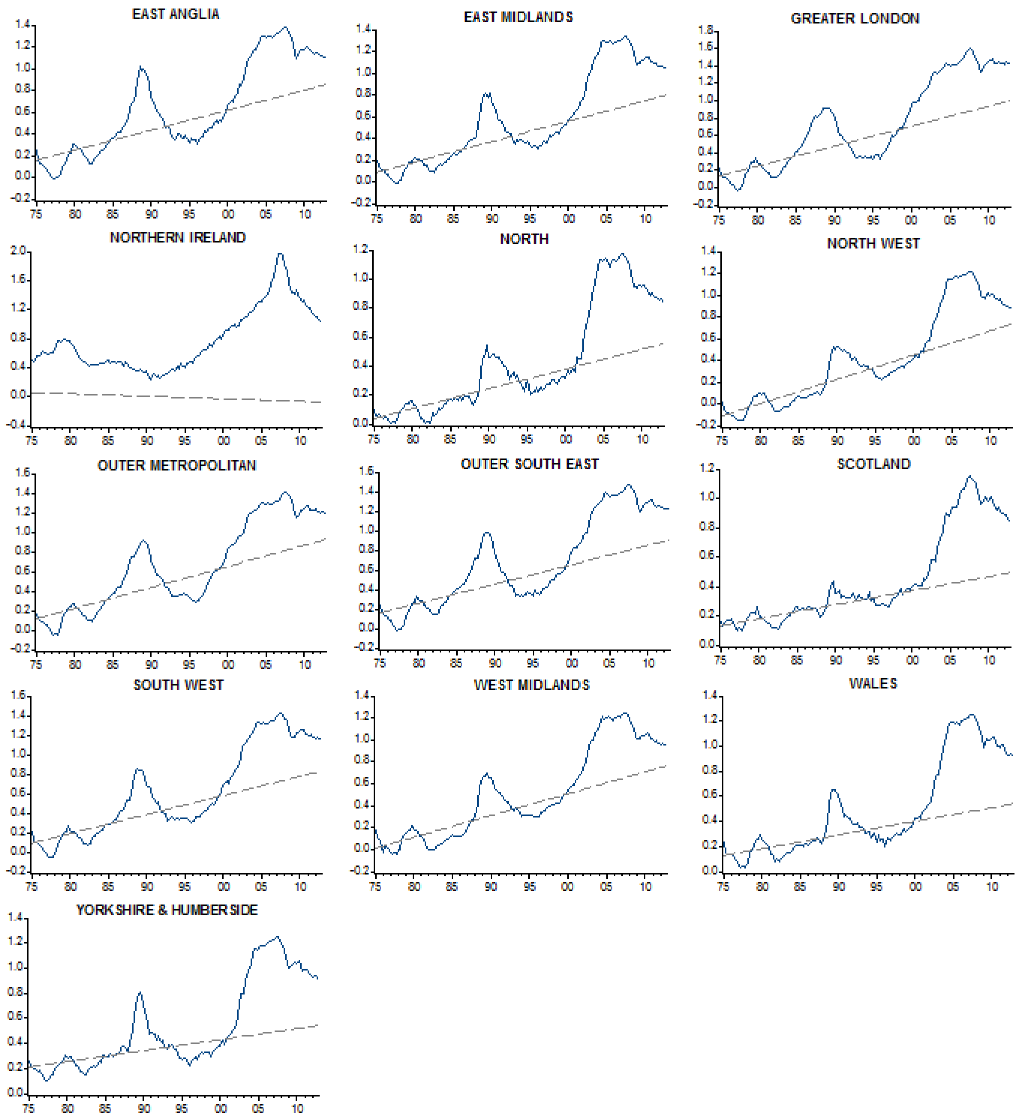
and individual $BSADF_{i,r_2}(r_0)$ is defined as in (7) with notation adjusted for the panel application as follows:

$$BSADF_{i,r_2}(r_0) = \underset{r_1 \in [0, r_2 - r_0]}{sup} \left\{ ADF_{i,r_1}^{r_2} \right\}.$$

The suggested date-stamping strategy compares the panel *BSADF* statistic with the sequence of $100(1-\beta_T)\%$ bootstrapped critical values.[18] By analogy with the univariate dating technique, the origination date of the overall exuberance is defined as the first observation that lies above the sequence of bootstrapped critical values, while its end date is located as the first observation that falls below the corresponding bootstrapped *BSADF* critical values.

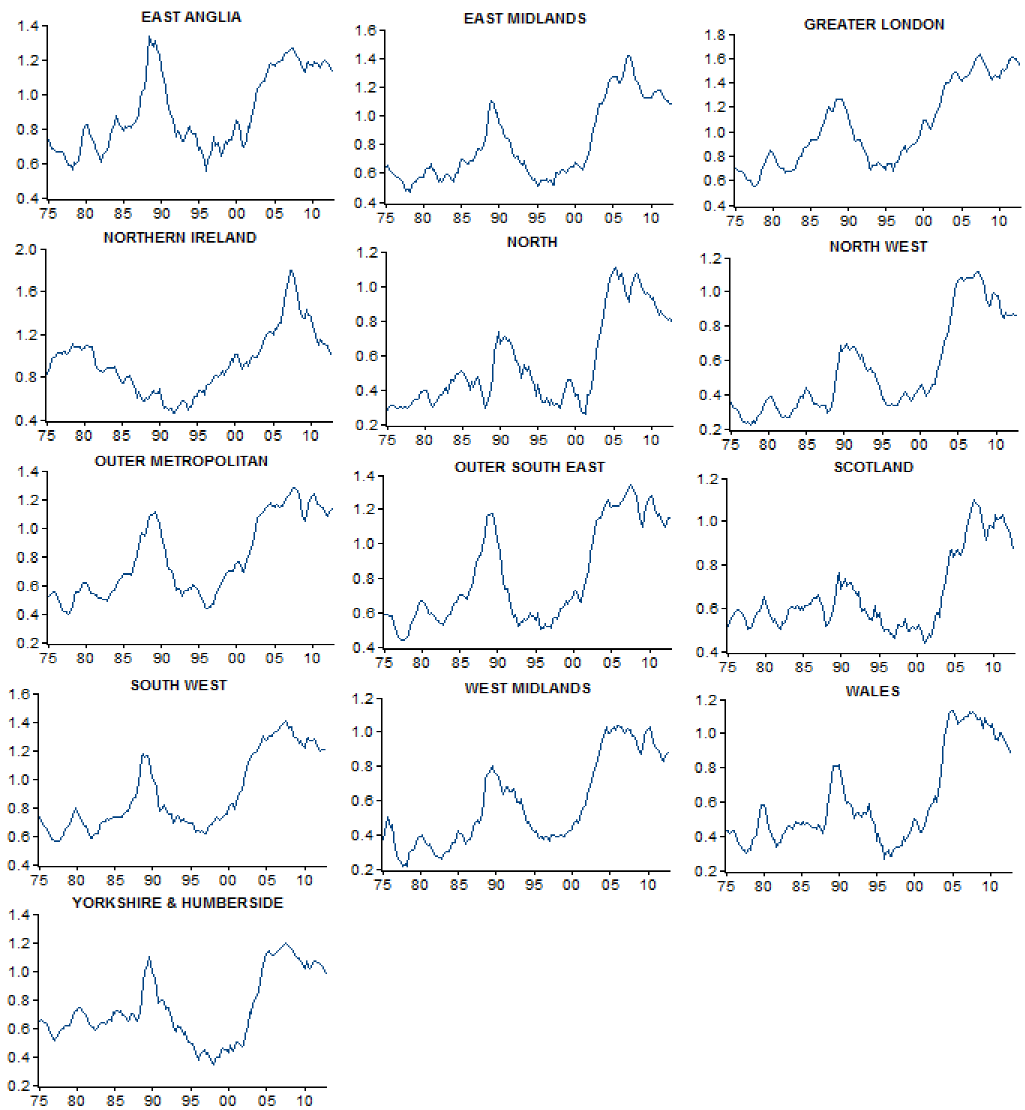
[18]See Appendix B of Pavlidis et al. (2016) for details of the bootstrap procedure

Figure 1: Real House Prices: Regional Series.



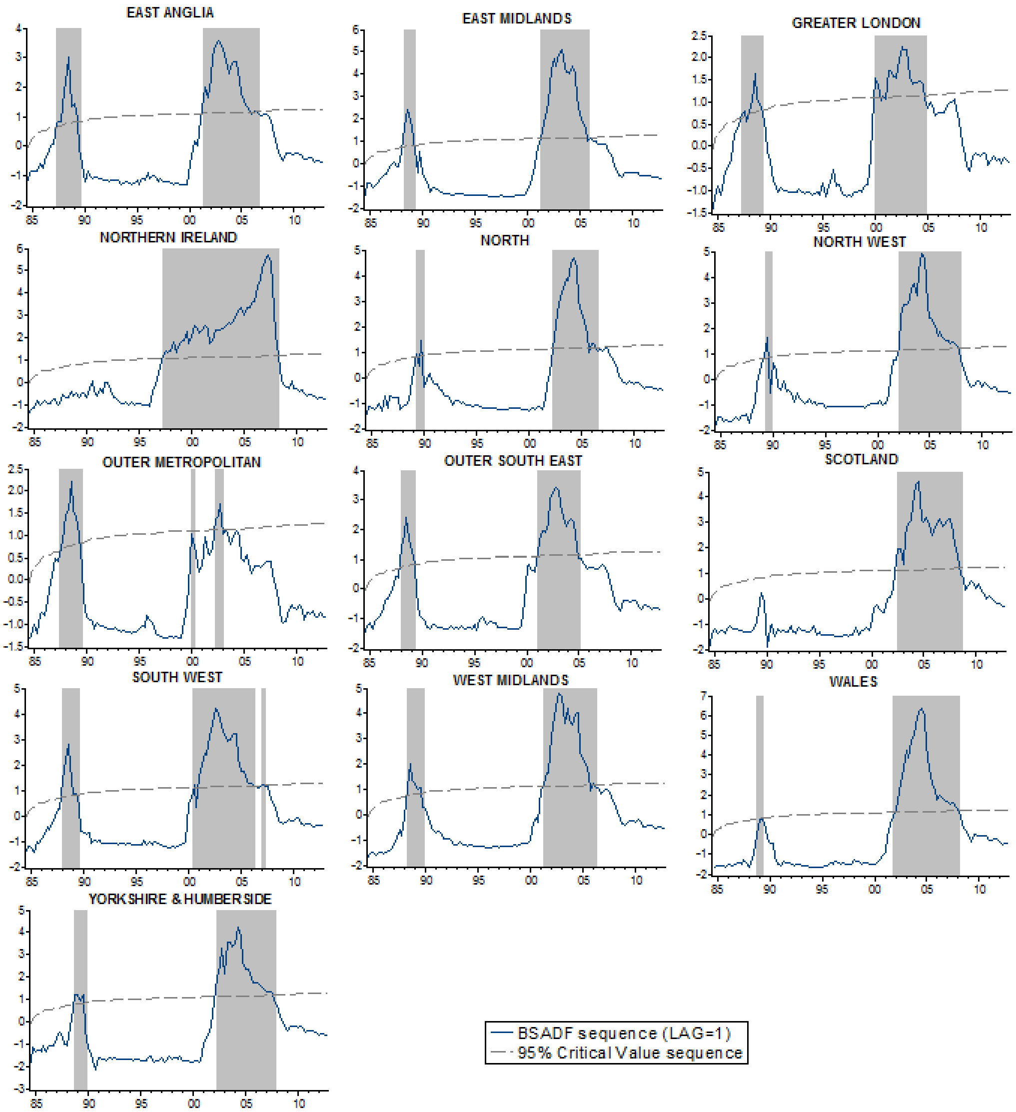
Note: The graph shows the evolution of the log real regional house price indices. The sample period: 1975:Q1-2012:Q4. Following IMF (2003, 2005) the linear time trend, estimated up to 1999:Q4 is added to each regional diagram (dashed line).

Figure 2: Price-to-Income Ratios: Regional Series.



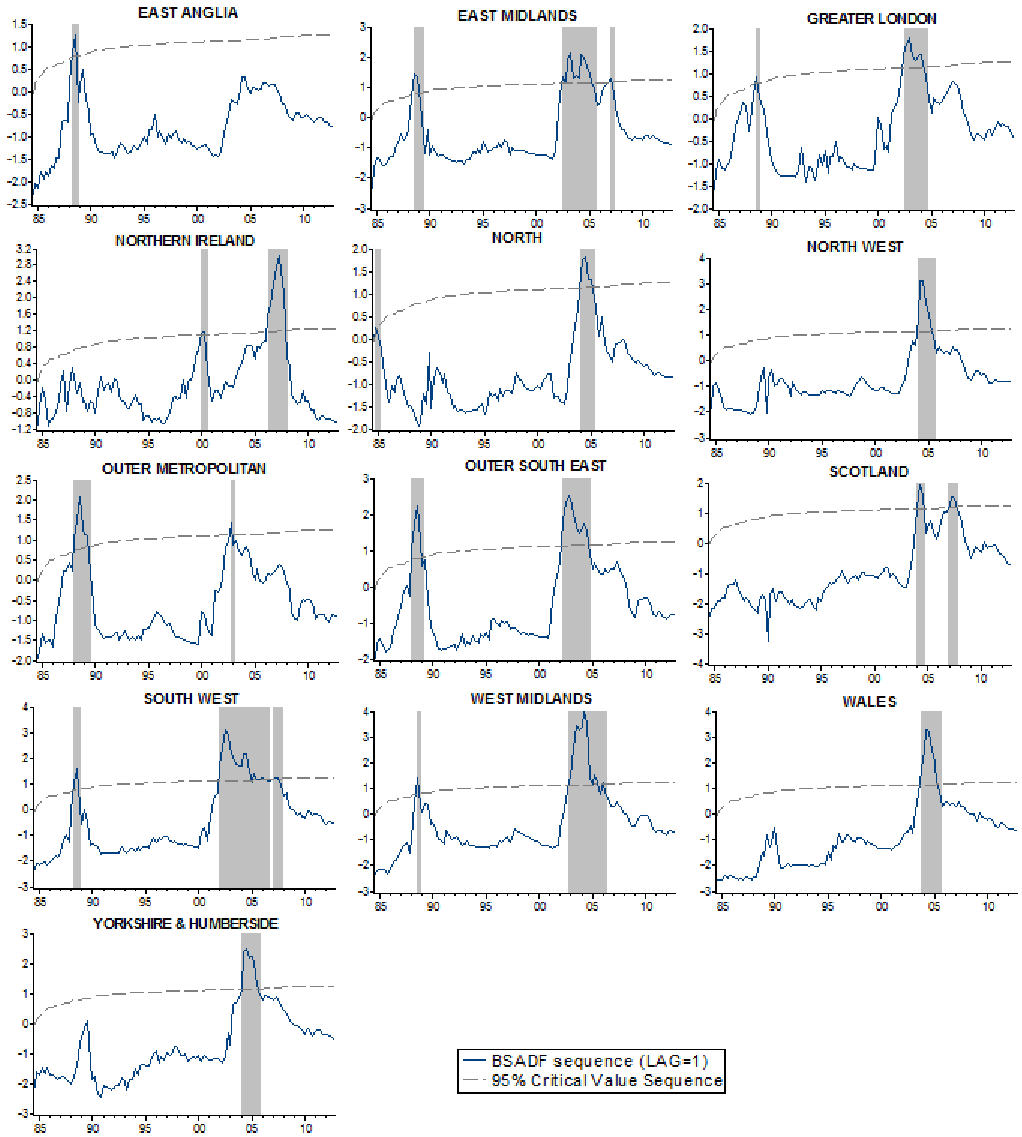
Note: Each regional diagram shows the evolution of the log of real house price to real personal disposable income ratio. The sample period: 1975:Q1-2012:Q4.

Figure 3: Regional Real House Prices: Date-Stamping of Explosive Episodes.



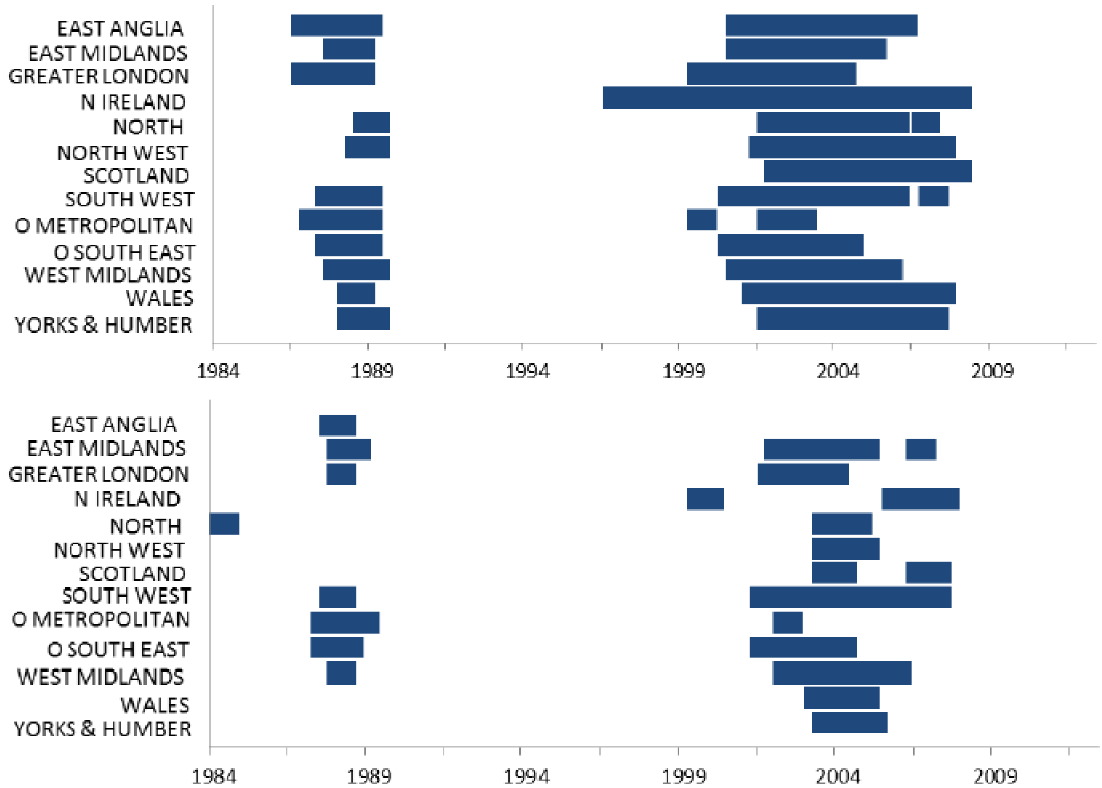
Note: Shaded areas indicate identified periods of exuberance (*BSADF* series is above the sequence critical value). The *BSADF* series are computed for the autoregressive lag length 1.

Figure 4: Ratio of Real House Prices to Real Personal Disposable Income: Date-Stamping of Explosive Episodes.



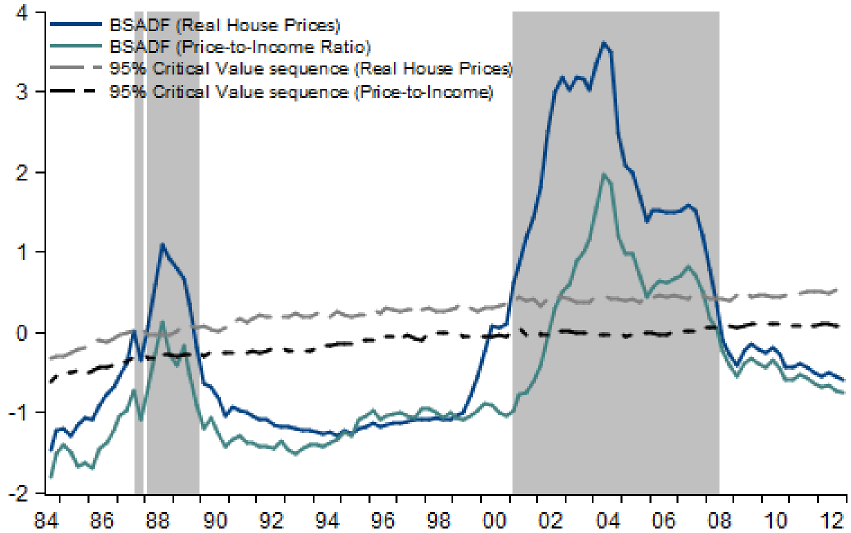
Note: Shaded areas indicate identified periods of exuberance (*BSADF* series is above the sequence critical value). The *BSADF* series are computed for the autoregressive lag length 1.

Figure 5: Date-Stamping of Explosive Episodes: Real House Prices and Price-to-Income Ratio



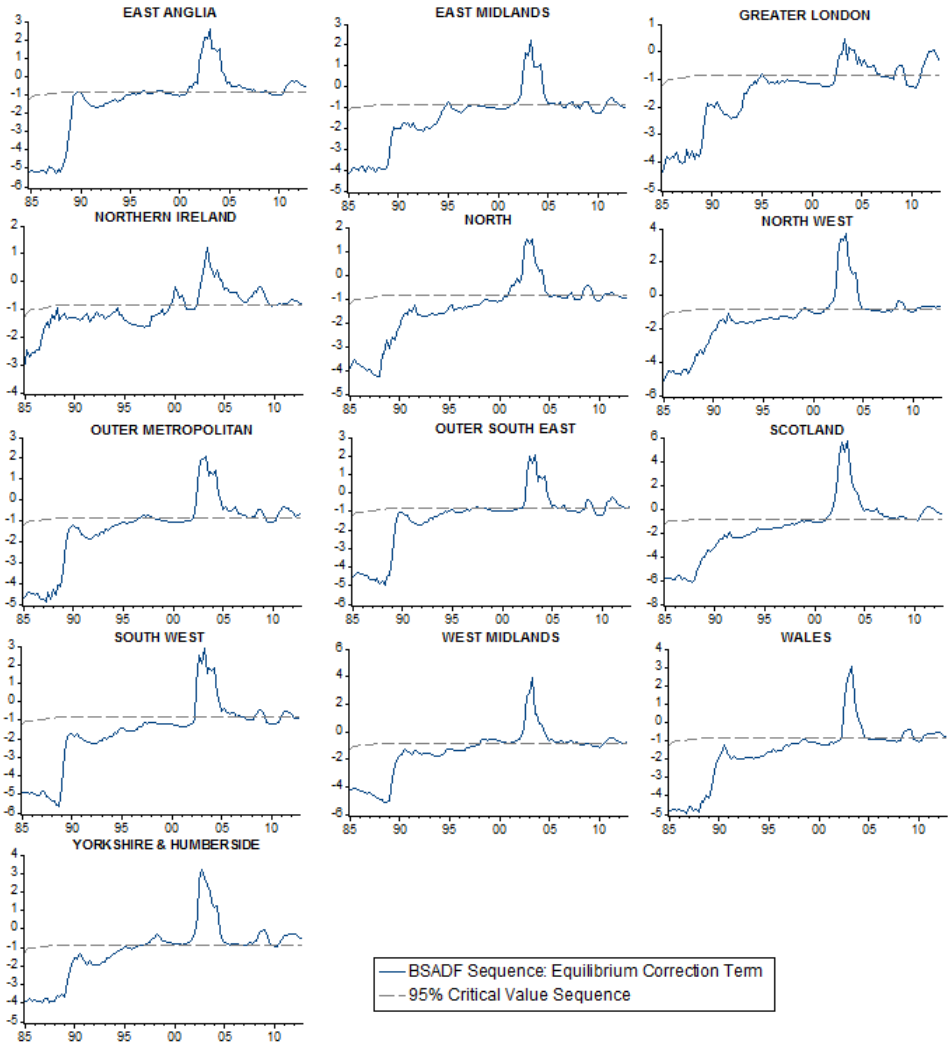
Note: Shaded areas indicate periods of exuberance identified by the *GSADF* test.

Figure 6: Date-Stamping Episodes of Nationwide Exuberance.



Note: Shaded areas indicate periods when the series of panel *BSADF* (real house prices) is above the sequence of critical values.

Figure 7: Model of Regional House Prices: Error-Correction Terms and Episodes of Exuberance.



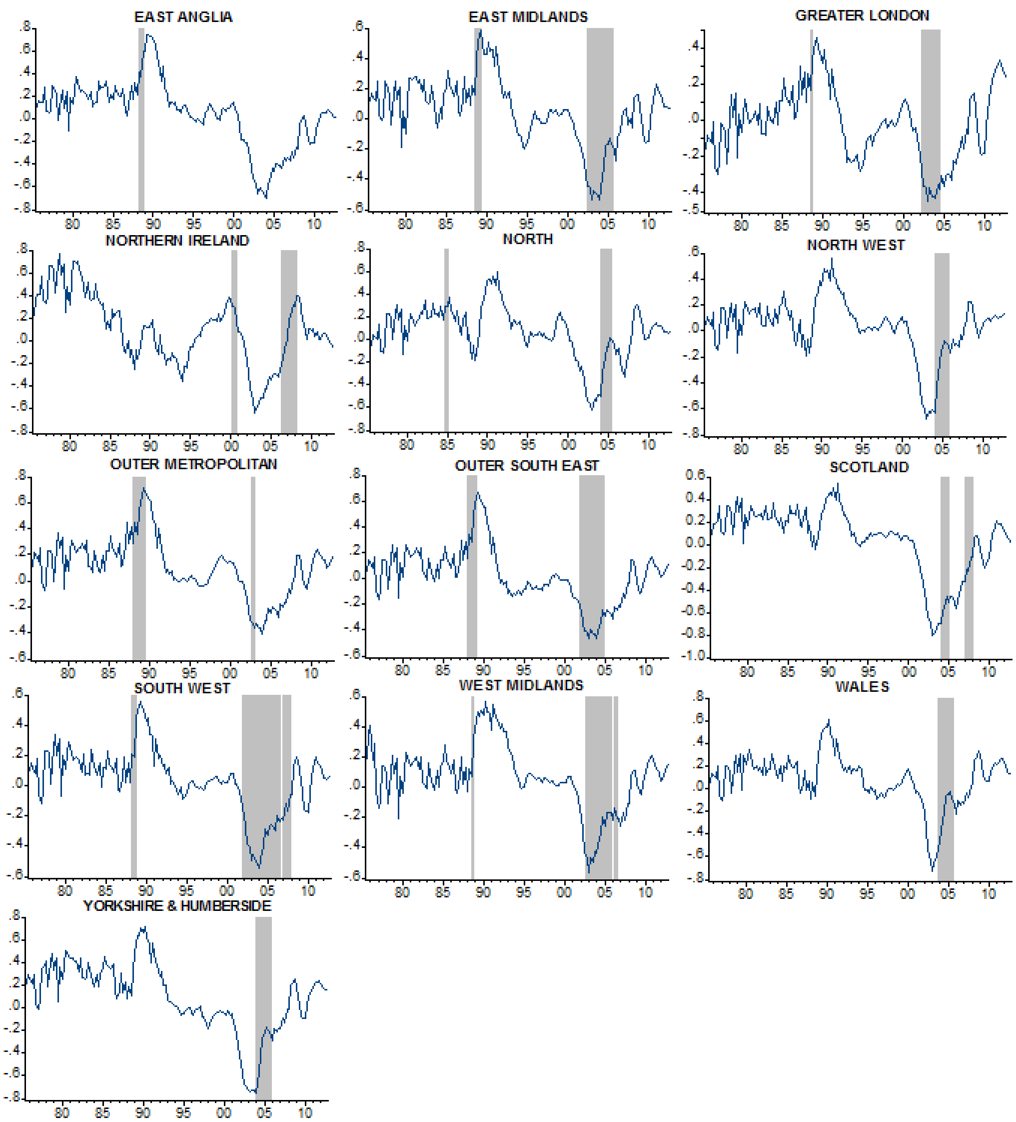
Note: The sequence of critical values is obtained by Monte Carlo experiments with 2000 replications. For each repetition, cointegrated system is simulated using Phillips' (1991) triangular representation as follows:

$$y_{1t} = \beta_2 y_{2t} + u_t, u_t = 0.75u_{t-1} + \epsilon_t, \epsilon_t \sim iidN(0, 0.5^2),$$
$$y_{2t} = y_{2t-1} + v_t, v_t \sim iidN(0, 0.5^2).$$

The *GSADF* test is then applied to the series of cointegrating residuals, computed as (Cameron et al. (2006)):

$$lrhp_{i,t-1} - \beta_{0,i} - \beta_1 \times lrynhs_{i,t} - \beta_2 \times MACCI_t - (1 - \varphi \times MACCI_t) \times (\beta_3 \times \Delta^2 labmr_t + \beta_4 \times (labmr_t - mean.labmr))$$
$$- \beta_5 \times MACCI_t \times (rabmr_t - mean.rabmr) - \beta_6 \times rabmr_t.$$

Figure 8: Model of Regional House Prices: Error-Correction Terms.



Note: Shaded areas indicate periods of exuberance identified by the *GSADF* date-stamping procedure applied to regional price-to-income ratios. In each regional diagram the error-correction term is computed as (Cameron et al. (2006)):

$$lrhp_{i,t-1} - \beta_{0,i} - \beta_1 \times lrynhs_{i,t} - \beta_2 \times MACCI_t - (1 - \varphi \times MACCI_t) \times (\beta_3 \times \Delta^2 labmr_t + \beta_4 \times (labmr_t - mean.labmr))$$
$$- \beta_5 \times MACCI_t \times (rabmr_t - mean.rabmr) - \beta_6 \times rabmr_t.$$

Table 1: Estimation Results for The Model of Regional House Prices.

	Variable	Estimate	t-Statistic
Speed of adjustment	α	0.03	7.78***
Index of credit conditions	$MACCI_t$	1.27	4.57***
Lagged house price growth	$\Delta clrhp_{i,t-1}$	0.24	10.15***
	$\Delta lrpdin_t$	0.81	3.34***
Income effects	$\Delta lrpdin_{t-1}$	0.33	2.26**
	$MACCI_t \times \Delta lrpdin_t$	-0.35	-0.83
	$(1 - \varphi \times MACCI_t) \times (labmr_t - mean.labmr)$	1.57	3.65***
	$(1 - \varphi \times MACCI_t) \times \Delta^2 labmr_t$	2.37	2.48**
Interest rate effects	φ	2.86	6.03***
	$MACCI \times (rabmr_t - mean.rabmr)$	23.29	1.71*
	$rabmr_t$	-8.42	-1.94*
Downside risk	$ror.neg_{i,t}$	0.09	4.78***
Inflation acceleration	$\Delta^2 lpc_t$	-0.22	-2.51**
Demographic effect	$\Delta pop2039_{i,t-1}$	-1.25	-0.94
Effect of new constructions	$\Delta(lwpop_{i,t} - lhs_{i,t-1})$	0.001	0.03
Stock market effect	$\Delta lrFTSE_t$	0.06	2.68***
(Outer Met)	$\Delta lrFTSEneg_t$	-0.06	-1.95*
Time	*D88*(ex.SC and NI)	0.06	6.15***
Dummies	*D08*	-0.05	-4.31***

Note: The dependent variable is the log regional real house price growth ($\Delta lrhp_i$). Each regional equation contains a region-specific intercept (estimates are not reported). The lagged house price growth effect $\Delta clrhp_{i,t-1}$ is computed as a weighted sum of last period's price growth in the region, regions contiguous to it and Greater London (see Table A1 in Appendix A). We follow Cameron et al. (2006) in their choice of regional weights:

Weights	Region												
	NT	YH	NW	EM	WM	EA	OSE	OM	GL	SW	WW	SC	NI
Own region	0.505	0	0.505	0.170	0.720	0	0	0	1	0	0	1	1
Greater London	0	0	0	0.112	0.280	1	1	1		1	0	0	0
Contig. regions	0.495	1	0.495	0.718	0	0	0	0	0	0	1	0	0

Table 2: Error-Correction Terms: Unit Root Test Results.

	Intercept	Trend and Intercept
Region	Test Statistic	Test Statistic
North	-2.99**	-3.28*
Yorks & Hside	-2.28	-2.61
North West	-2.71*	-2.77
East Midlands	-2.83*	-3.11
West Midlands	-2.78*	-2.68
East Anglia	-1.89	-2.34
Outer S East	-2.38	-2.49
Outer Met	-2.37	-2.49
Greater London	-2.42	-2.39
South West	-2.65*	-2.99
Wales	-2.99**	-3.14
Scotland	-1.92	-2.31
Northern Ireland	-2.36	-2.69
	Critical values	
90%	-2.58	-3.15
95%	-2.88	-3.44
99%	-3.48	-4.02

Note: The table reports Augmented Dickey-Fuller (ADF) test results. Superscripts *, ** and *** indicate rejection of the null hypothesis of non-stationarity at 10, 5 and 1 percent level of significance respectively.

Table 3: The *SADF* and the *GSADF* test results.

Panel A: Univariate SADF and GSADF statistics				
	Real House Prices		Price-to-Income Ratio	
Region	*SADF*	*GSADF*	*SADF*	*GSADF*
North	2.38^{***}	4.70^{***}	0.22	1.80^{**}
Yorks & Hside	1.02^{*}	4.21^{***}	-0.16	2.47^{***}
North West	1.33^{**}	4.89^{***}	0.22	3.11^{***}
East Midlands	1.70^{**}	5.10^{***}	1.28^{**}	2.11^{**}
West Midlands	1.46^{**}	4.78^{***}	-0.01	3.94^{***}
East Anglia	2.39^{***}	3.58^{***}	1.24^{*}	1.26
Outer S East	1.56^{**}	3.40^{***}	1.52^{**}	2.53^{***}
Outer Met	1.25^{*}	2.20^{**}	1.16^{*}	2.07^{**}
Greater London	0.86	2.23^{**}	0.55	1.78^{*}
South West	1.92^{***}	4.19^{***}	1.13^{*}	3.08^{***}
Wales	1.48^{**}	6.38^{***}	-0.26	3.26^{***}
Scotland	2.68^{***}	4.59^{***}	0.49	1.90^{**}
Northern Ireland	4.61^{***}	5.72^{***}	1.80^{***}	3.05^{***}
	Finite sample critical values			
90%	0.99	1.51	0.99	1.51
95%	1.27	1.78	1.27	1.78
99%	1.75	2.42	1.75	2.42
Panel B: Panel GSADF statistics				
	Real House Prices		Price-to-Income Ratio	
	3.62^{***}		1.96^{***}	
	Finite sample critical values			
90%	1.04		0.56	
95%	1.24		0.79	
99%	1.79		1.16	

Note: Superscripts *, ** and *** denote significance of the reported statistic at 10, 5 and 1 percent level of significance. Finite sample critical values for the sample of 150 observations are obtained from Monte Carlo simulations with 2000 replications. The smallest window r_0 corresponds to 24% of the data and comprises 36 observations. For both variables under consideration, reported univariate *SADF* and *GSADF* statistics as well as panel *GSADF* are computed for autoregressive lag length of one.

Table 4: Error-Correction Terms: The *SADF* and the *GSADF* test results.

Region	*SADF*	*GSADF*
North	-1.52*	1.52***
Yorks & Hside	-0.62***	3.24***
North West	-1.35*	3.71***
East Midlands	-1.58	2.23***
West Midlands	-1.87	3.96***
East Anglia	-0.75***	2.55***
Outer S East	-1.86	2.08***
Outer Met	-1.87	2.09***
Greater London	-2.44	0.45***
South West	-1.82	2.88***
Wales	-1.19**	3.07***
Scotland	-0.37***	5.65 ***
Northern Ireland	-1.07**	1.19***
Finite sample critical values		
90%	-1.54	-0.78
95%	-1.32	-0.56
99%	-0.79	-0.23

Note: Superscripts *, ** and *** denote significance of the reported statistic at 10, 5 and 1 percent level of significance. Finite sample critical values for the sample of 150 observations are obtained by Monte Carlo simulations with 2000 replications. For each repetition cointegrated system is simulated using Philips' (1991) triangular representation as follows:

$$y_{1t} = \beta_2 y_{2t} + u_t, u_t = 0.75u_{t-1} + \epsilon_t, \epsilon_t \sim iidN(0, 0.5^2),$$
$$y_{2t} = y_{2t-1} + v_t, v_t \sim iidN(0, 0.5^2).$$

The *SADF* and *GSADF* tests are then applied to the series of cointegrating residuals, computed as (Cameron et al. (2006)):

$$lrhp_{i,t-1} - \beta_{0,i} - \beta_1 \times lrynhs_{i,t} - \beta_2 \times MACCI_t - (1 - \varphi \times MACCI_t) \times (\beta_3 \times \Delta^2 labmr_t + \beta_4 \times (labmr_t - mean.labmr))$$
$$- \beta_5 \times MACCI_t \times (rabmr_t - mean.rabmr) - \beta_6 \times rabmr_t.$$